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Women's wages and fertility revisited
Evidence from Norway

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Abstract:
Since the 1960s, Beckers' New Home Economics has provided a central theoretical framework for studies of fertility behaviour. New Home Economics predict a negative effect of female wages on fertility. This prediction has been tested in a number of studies over the past decades, but the results are far from unanimous. In this paper we review past evidence of the impact of female wages on their childbearing behaviour and supply new evidence from Norway. We estimate a simultaneous hazard rate model of transitions to first, second and third birth, including predicted wage as a time-dependent variable. Using a very large dataset covering all women born in Norway during the period 1955-74, we find that timing of births is associated with wage changes. The wage effect on the log hazard is U-shaped for all the four 5-year cohorts we are studying, but the effect varies across cohorts and parity. We also find that the relationship between timing of births and wages are not very sensitive to the omission of the women's non-labour income.

Keywords: Female fertility, wages, non-labor income, hazard model

JEL classification: J13, J30

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Sammendrag

Beckers “New Home Economics” har siden 1960-årene vært et viktig teoretisk rammeverk for empiriske studier av fertilitetsatferd. Ifølge teorien vil en isolert økning i timelønnen føre til redusert fertilitet. Prediksjonen har vært testet i en rekke empiriske studier over de siste tiårene, men resultatene gir ikke entydig støtte til teorien. I inneværende studie gir vi en oversikt over sentrale bidrag til litteraturen på feltet, samtidig som vi presenterer nye resultater basert på norske data som er spesielt velegnet for denne type studier. For å tallfeste sammenhengen estimerer vi en simultan hazardmodell for overganger til henholdsvis første, andre og tredje fødsel. Ved å bruke blant annet predikert timelønn og ikke lønnsinntekt som tidsavhengig forklaringsvariabel i modellen, sikrer vi oss at vi får best mulig konsistens med rammeverket til Becker. Basert på registerdata som dekker alle kvinner født i Norge i løpet av perioden 1955-74, finner vi at det er en sammenheng mellom endringer i timelønn og periodiseringen av fødsler. Effekten på logaritmen til hazarden er U-formet for alle de fire 5-års kohortene vi studerer, men effekten varierer over kohorter og paritet. Vi finner også at den tallfestede sammenhengen mellom timelønn og fødsler ikke er særlig sensitiv overfor utelatelse av kvinners arbeidsfrie inntekt.

Background

The most influential and formally coherent theory of fertility behaviour is no doubt Beckers' New Home Economics as set out in his original work from the early 1960s (Becker, 1960, 1965). In this framework children are recognized as providing utility to parents in much the same way as other consumer goods. Besides parents' preferences the crucial determinants of childbearing are the cost of a(nother) child (the "shadow price" of children) and the family's budget constraint. A wage increase has two offsetting effects, a positive income effect due to a larger family budget, and a negative substitution effect due to a higher "shadow price" of children. Since raising children is time-intensive, especially for mothers who are still the main care-givers in most societies, the substitution effect is expected to dominate the income effect for women. Thus, New Home Economics predict a negative effect of female wages on fertility.

However, as noted by several scholars, there are considerable challenges to testing the central propositions of Becker's fertility theory (see e.g. Hotz et al. 1997). A number of studies have provided empirical evidence of the effect of female wages on childbearing over the past decades, but the results are far from unanimous, and the wage-fertility nexus is not always estimated to be negative as predicted. Some of this divergence is clearly linked to the use of different indicators of women's opportunity costs, but even analyses using very similar models, methods and indicators sometimes arrive at quite different conclusions. The results seem e.g. quite sensitive to the wage series used to predict wages over the life-cycle, and this has been the subject of much debate (Tasiran 2002; Walker 2002).

In this paper we review past evidence of the impact of female wages on childbearing behaviour and supply new evidence from Norway based on a very large and high-quality dataset. Norway is an interesting case in this respect since Norwegian women differ from their "sisters" in many other countries by combining high fertility with high labour market participation. Today labour market participation among married and cohabiting women with children below the age of 16 is 84 percent (Statistics Norway 2013), and the total fertility rate is 1.78. Furthermore, cohort fertility remains high and stable at slightly above 2 children also for cohorts born in the 1960s (Statistics Norway 2014).

Our study is carried out by estimating a hazard rate model of birth transitions including the predicted wage and other variables that are known to affect fertility behaviour. We use longitudinal information on births, education and earnings extracted from various administrative registers combined with data on working hours from the Norwegian Labour Force Surveys. The analysis covers the period 1974 to

2009. Due to the comprehensive dataset, we are able to reduce biases from cohort effects by estimating the hazard functions separately for four different 5-years cohorts. By estimating the hazard functions for the three first births simultaneously, we also reduce biases due to the effects of unobserved heterogeneity across women. Heckman's selection model is used in the estimation of the wage equations.

Our main finding is that timing of births is associated with wage changes for all the first three births we are studying. Allowing a quadratic function in wages, we find that there is U-shaped relationship between wages and the log hazard function. The relationship varies across cohorts (and parity). Based on information about the average female wage during the calendar period studied, we conclude that many women most likely are on the downward sloping part of the wage curve when they are under risk of a first birth.

In a second part of the paper, we perform sensitivity analyses by 1) allowing interaction effects between wages and age, and by 2) including non-labor income of (married) women as covariate in the hazard rate model. The conclusion is that these modifications do not change the main picture from the simpler model specifications.

The paper proceeds as follows. In the next section we give a brief review of previous empirical findings on the association between women's wages and fertility. Then follows a discussion of challenges related to empirical analyses of fertility dynamics and a more detailed description of the data and methods used. Next we present the results, and the final section concludes with a brief summary and discussion.

Previous literature

Before the mid 1980s, micro evidence on the effect of women's wages on the timing and spacing of births was scarce because of the virtual absence of individual fertility histories linked with earnings data. Later several articles with access to such data have appeared. In this section we briefly review these articles and summarize the evidence so far.

One of the first articles based on individual data is Moffit (1984) who constructed a complete life-cycle model of fertility, labour supply and wages and estimated profiles over the life-cycle jointly for these outcomes. The data were from eight waves of the US National Longitudinal Survey of Young Women conducted annually from 1968. The women were aged 14-24 when first interviewed, implying they could only be followed till they were maximum 31 years. The joint estimation technique represented a novelty and no doubt an advantage as it took into account both heterogeneity of tastes, the selectivity problem of missing wage rates for non-workers, and endogeneity problems arising from the fact that fertility and labour supply decisions over the life-cycle are closely intertwined and fairly simultaneous processes. Moffit's results indicate that higher wages increase women's labour supply and decrease women's fertility, but the effect on employment is stronger than the effect on fertility.

Other early studies are Barmby and Cigno (1990) from the UK and Heckman and Walker (1990a and b) from Sweden. A common feature of these studies is that they combine individual fertility histories and other characteristics of the women with aggregate time series on income and earnings, and construct various proxies for women's potential market wage based on this information. Barmby and Cigno had access to a large random sample of British women aged between 16 and 59 years in 1980 and used macro data to define an indicator for women's net wage relative to men's over the study period, 1954 to 1980. The analysis were restricted to women who were married to the same man for at least ten years and only considered childbearing within marriage. Controlling also for child allowances, Barmby and Cigno found that a rise in the average female wage rate, while holding male wages constant, lengthened the expected time to first birth and reduced expected completed fertility.

Heckman and Walker (1990a and b) are based on the 1981 Swedish Fertility Survey which provided retrospective information on life-cycle fertility, employment, education and marital and cohabitational status from a representative sample of women. The cohorts included in the study were born from 1936 to 1955, and the period considered was 1948-1981. Heckman and Walker constructed a time series of wages using summary measures of personal tax returns by age and sex for selected years published by Statistics Sweden and generated a complete series for all the years studied by means of interpolation. They found that female wages play a strong role in Swedish fertility dynamics. The estimated effects were negative in all birth transitions (first to third), quantitatively large and almost always statistically significant at conventional test levels.

Taşiran (1995) used the same data as Heckman and Walker to study Swedish fertility dynamics, but constructed other wage and income series that combined individual wages with gender- and age-specific aggregate wage measures. Taşiran's estimated female wage effects were usually negative, but often not statistically significant and sometimes positive for transitions to second and third births. From these results, Taşiran inferred that Heckman and Walker's results were not robust to the introduction of micro-level data, and this has later been the subject of debate (Tasiran 2002; Walker 2002).

Replicating the study for Sweden by Heckman and Walker (1990a), Merrigan and St.-Pierre (1998) estimated Canadian fertility dynamics based on retrospective fertility and marital histories of women born between 1941 and 1960 from the Canadian 1990 General Social Survey. The survey did not contain life-cycle information on income or wages, so similar to the Swedish studies, Merrigan and St.-Pierre had to construct gender and age-specific real wages series from aggregate statistics. Their results for Canada were fairly similar to the results of Heckmand and Walker for Sweden. The female wage effect was negative and significant in all birth transitions (1st to 3rd birth), but the strongest effect was found for the transition to first birth. For later transitions, the wage effect was significant for the two oldest cohort groups only (women born 1941-45 and 1946-50). Somewhat weaker negative effects across cohorts were also observed in Heckman and Walker (1990a), and they attributed this cohort drift to the expansion of family- and work-related policies that made the female wage an increasingly less accurate proxy for the opportunity cost of childbearing over time.

In the last couple of decades, micro data on earnings and income have become easier available, and several studies have used such information to illuminate the relationship between female wages and fertility. These studies fall in two categories. In line with the above-mentioned studies, one category uses predicted wages as indicator of women's opportunity cost related to childbearing (Groot and Pott-Buter 1992; Francesconi 2002; Rønsen 2004; Rondinelli, Aassve and Billari 2010). The predicted wage is usually obtained by estimating a wage relation that is corrected for possible selection bias introduced by only having information on wages for women who are employed. Based on the estimated parameters of the wage relation, a potential wage for each woman in the sample is predicted and imputed. The second category of studies uses observed annual earnings or income as measures of the opportunity cost of childbearing (Andersson 2000; Vikat 2004; Andersson et al. 2009; Santarelli 2011). Not surprisingly, these studies arrive at quite different results, since predicted wages and observed annual earnings measure different things. Whereas predicted wage reflects the remuneration of market work per time input (usually hour), annual earnings is a product of wages and working hours. Women who have worked part time or part year may have low annual earnings even if they have relatively high wages, and annual earnings will therefore be an imprecise measure of opportunity costs. Moreover, the opportunity cost of non-working women will be zero, which is questionable since most people will have an alternative use of time, and this time has a positive value. In the following we therefore only consider the evidence from the relatively few studies based on individual-level data that use predicted wage as indicator.

In an early study, Groot and Pott-Buter (1992) analyse the transition to first birth in the Netherlands. The data is from a national longitudinal labour market survey from 1985, collecting work and maternity histories, earnings data and other background information for the period 1980-1985. The analysis includes women between 15 and 35 years in 1980 (cohorts 1945-65), who were married or cohabiting and who were still childless at that time. Groot and Pott-Buter include both the wage rate of the wife and the wage rate of the husband in the model, and find a negative effect on the transition to motherhood of both the female and the male wage rate. The latter is in apparent contrast to the New Home Economics theory which presupposes a positive effect on fertility of the husband's income. Instead Groot and Pott-Buter find that first birth rates in the Netherlands are positively associated with total household labour income.

Other studies are Rønsen (2004) and Rondinelli et al. (2010). Rønsen analyses the transition to first, second and third birth in Norway using retrospective survey information on fertility and employment merged with information on income from administrative registers. The analysis includes women born between 1945 and 1965, and the period considered is 1960-1988. The estimated wage effects are negative and significant for all transitions except for parity three, mainly because of a smaller sample size. Rondinelli et al. analyse the same birth transitions for Italy based on more recent data by combining information from two regular surveys, The Survey of Households' Income and Wealth and the Italian Labour Force Survey. The analysis covers the period 1983-2003 and includes women who were 15-40 years old in 2003 (cohorts born 1963-1988). Rondinelli et al. find a strong negative effect of the woman's wage on first birth, a less strong, but still significant negative effect on second birth, and no effect on third birth.

Using a similar approach as Moffit (1984), Francesconi (2002) constructs a joint dynamic model of fertility and work of married women. His data source is also the same, the National Survey of Young Women, but he has access to a longer panel, covering the years 1968-1991. The cohorts analysed are born 1944-1954, implying that they could be followed over most of their fecund period. Francesconi differentiates between full-time and part-time earnings profiles and finds that a downward shift of the full-time earnings profile increases fertility. Moreover, the steepness of the profile exhibit strong effects on the expected number of births, and a flatter wage profile is associated higher completed fertility.

The evidence so far thus generally corroborates the New Home Economics hypothesis of a negative effect on fertility of women's wages, but the effect seems to be larger for first birth than for transitions to higher parities and possibly larger for older cohorts than for younger cohorts. The evidence based on individual earnings data is still scant, however. More studies based on more recent data and from more countries are therefore needed. The present paper is a contribution in this respect.

Challenges in testing the central propositions of Becker's fertility theory empirically

As we have seen, the estimated effect of female wages on fertility is sensitive to the wage and income series used as shown by the analyses by Heckman and Walker (1990a and b) and Taşiran (1995) for Sweden. Besides, the results may of course differ because the analyses cover different time periods and different cohorts, and when analyzing different countries, several other factors come into play, e.g. different welfare state and cultural contexts, different tax systems and different labour market conditions.

In addition, all analysts of fertility processes are faced with the challenge of taking a theory that is basically static as their point of departure, as Becker originally modelled completed family size, and not the timing and spacing of births. Later several authors have proposed fertility models that blend features of the static model with other models of life-cycle behaviour (life-cycle consumption, life-cycle labour supply decisions and human capital investment and accumulation, e.g. Happel, Hill and Low 1984; Moffitt 1984; Hotz and Miller 1988; Cigno and Ermisch 1989; Cigno 1991; Walker 1995). However, these models often rely on strong assumptions, have no simple econometric specifications, and do not yield unambiguous predictions as to how life-cycle fertility is expected to vary with prices and income (Hotz et al. 1997).

An important distinction in life-cycle models of fertility is the assumption made about capital markets. It is either assumed that capital markets are perfect (the PCM assumption), i.e. parents are able to borrow and lend across time at a given real interest rate, or perfectly imperfect (the PICM assumption), in which case no borrowing or saving is possible. When capital markets are imperfect, *consumption smoothing* becomes a central motive, and this is achieved by synchronising the costs of a child with a time interval in which the income of the primary earner is relatively high. Consequently, the husband's earnings profile matters, not only the level of the present value of his lifetime earnings (Happel et al. 1984). Another consequence of this model is that if the husband's earnings increase continuously over time, births will be delayed to the biological limit.

When capital markets are perfect, the household relies on financial markets to 'smooth' or minimise the dispersion of its consumption profile. The optimal date of birth is then the one which maximises the wife's life-time earnings. The woman's *career planning* is a central motive in these models, and the optimal time of maternity is determined by several factors including accumulated human capital at the beginning of the planning period, the profile of human capital investments, the rate of return to these investments, the length of time spent out of the labour force, the rate at which job skills depreciate, and direct child expenditures (Gustafsson 2001). These determinants shape the level and the steepness of the woman's earnings profile, which is an important ingredient in models assuming perfect capital markets.

However, various authors arrive at somewhat different conclusions as to how the above determinants impact the timing of childbearing. Happel et al. (1984) argue e.g. that women with larger pre-parental work experience will tend to delay births because this will reduce the probability of losing all their job skill when staying home. Cigno and Ermisch (1989) reason on the other hand that women with more market-specific human capital will have fewer children, but sooner. The reason is that parents have a positive time preference for children, and the income effect makes an earlier childbirth economically viable. Depreciation of human capital plays an important role in Happel et al.'s model, while the Cigno and Ermisch model assumes that the depreciation rate is zero.

In Cigno and Ermisch (1989) the investment profile of human capital investment is central, and in their model a steeper earnings profile is assumed to make later births relatively less costly. Women with steeper earnings profiles have more to gain by accruing human capital early in the life-cycle and will therefore have a slower tempo of fertility. This hypothesis is corroborated by their empirical analysis which shows that women in occupation with steeper earnings profiles tend to have children later in life.

Walker (1995) on the other hand contends that a steeper earnings profile makes later births more costly. These two seemingly opposing results have to do with the emphasis put on different components of the shadow price of fertility. Walker constructs a profile of the shadow price which also takes into account taxes and family policies (parental leave benefits and child care subsidies), and the assumption made about the time absent from paid work following birth is vital in the construction of the price profiles. In Walker's model, a steeper profile implies that foregone earnings will be relatively larger later in the life-cycle, thus making earlier childbearing less costly. Changes which make the profile steeper will thus increase the tempo of fertility, while changes that flatten the profile will tend to delay fertility. Walker finds empirical support for this hypothesis in estimating the shadow price of fertility for Sweden, which shows that younger female cohorts have higher, but flatter earnings profiles than those faced by women in older cohorts. This is consistent with the delay in childbearing witnessed in Sweden during the period studied (1955-1989).

The above synopsis demonstrates some of the challenges related to the modelling of life-cycle fertility and exemplifies the sensitivity of the models to different assumptions and different emphasis given to the various components of the shadow price of fertility. In the present analysis, we make no attempt at building a comprehensive life-cycle model of fertility, but rest on the framework of Becker and the results from the previous life-cycle fertility literature to investigate fertility dynamics in Norway during the period 1974-2008. We shall assume that capital markets are perfect, and that the optimal timing of births is the one that maximises the woman's life-time earnings. We thus regard career planning to be a central motive. Unfortunately, many of the factors determining the optimal timing of births are not known (e.g. the profile of human capital investments and human capital appreciation and depreciation rates), but a large component of the opportunity cost of childbearing is the immediate foregone earnings due to a career break. Net foregone earnings again depend on taxes and family policies, but as illustrated in Walker (1995) it is no trivial task to construct a time series of the shadow price of fertility that also includes tax and policy elements. Still, as demonstrated by the analysis for Sweden (ibid.), the single most important determinant of the shadow price of fertility is the female's wage rate. The predicted wage over a woman's fertile period is thus our main fertility price indicator, but in the interpretation of the results we will also consider the changing family policy environment over our study period. Below follows a closer description of the data and the methods we have used.

Data and estimation procedure

Data

Estimating the impact of wages on fertility is very data demanding as it requires information about "lifetime" wage rates linked with women's fertility histories for a sufficient number of women over a sufficient number of years. Most researchers do not have access to this type of data, and estimation of childbearing decisions in line with the theoretical framework of Becker typically requires ad hoc assumptions that might bias estimation results. In contrast, the present study makes use of a rich and comprehensive dataset that is very suitable for this type of analysis.

The main data used in the analysis is an administrative dataset covering all women registered as living in Norway during the period 1974-2009. In this dataset we observe if and when the woman had her first, second and third birth, respectively. In the data we observe day of birth, but for the sake of simplicity we use year of birth in the model specifications. The dataset also includes information about the age of the woman and her immigration status (country of birth), as well as her birth region (county) if she is born in Norway. Educational information is linked to the dataset using register data from the Norwegian education data base (NUDB). NUDB is updated annually and includes information about the highest educational level of the woman and whether she is registered as being enrolled in education. Thus, we can treat education as a time varying covariate in the estimations.

A special feature of our methodological approach is that we have access to 36 waves of the Norwegian Labour Force Survey (NLFS) for every year 1974-2009, i.e., for the same period as for the register data on fertility. This survey includes detailed information about labour market participation and hours of work. Linking information about labour market participation and actual hours of work from this survey with information about actual wage incomes from the Norwegian Social Security Administration, we can calculate wages as the fraction of wage incomes and actual hours of work. This linkage allows us to estimate separate wage equations for each of the years 1974-2009. Based on the estimated equations we can impute year-specific wages for all women in the study. This approach solves two issues. The first is that we do not observe wages for those not participating in the labour market. The second is that observed wages might be correlated with the error term of the fertility model, which might bias estimation results. Using the predicted wage rate instead of actual wages, we also take into account that the opportunity cost of raising children is non-zero for women outside the labour market.

The specification of the wage equations is given in Equation 3 below. Potential work experience is measured as age minus years of education minus 6 while non-labour income is measured as the wage income of the husband. Both predicted wages as well as non-labour income are measured in real 2005-prices using the CPI from Statistics Norway.

Females that are self-employed according to the information in the labour market survey are omitted in the estimation of the wage equations. The sample size for the data used in these estimations varies across the period we are looking. For the period 1974-1988 the average sample size is 6400 women, while it is about 12000 during the period 1989-2009.

Both married and unmarried females are included in the estimation of the main hazard model. They are being followed from age 19 to 49. Females born outside Norway are excluded from the sample. A few females are excluded from the sample due to missing information about education.

In a second part of the paper, we check the robustness of the estimation result by including non-labour income of the female as a time varying regressor in the hazard function. In these estimations we only include females that are married at first birth.

All data sets are linked using a personal identification number.

Estimation procedure

To quantify the effect of wages on fertility, we estimate a proportional hazard model, where the effect of the covariates on the hazard of occurrence is multiplicative. Looking at a particular woman with given characteristics, the hazard for the woman's first birth at time t is the probability density that the woman gives birth to the first child in period t conditional on no prior (first) birth. We estimate hazard functions for first, second and third birth specified as

$$1) \quad \ln h(t) = \gamma T(t) + \beta' X(t),$$

where ln $h(t)$ is the logarithm of the hazard at time t, $X(t)$ is a vector of woman specific characteristics (wage, non-labour income, educational level, educational enrolment, region of birth) and β' is the corresponding parameter vector. The term $\gamma T(t)$ captures the baseline hazard duration dependence, where $T(t)$ is a piecewise linear spline and γ is a parameter including a constant term. For the first birth, $T(t)$ is the number of years since the woman become 19 years old while for the second and third birth it is the number of years since last birth (age of youngest child). For the second and third birth we also include age splines for the woman's age. Since no woman in the age cohort 1970-74 is older than 39 in 2009 (the last year covered by this data), the number of nodes in the splines are fewer for the youngest cohort than for the older cohorts.

In this paper we are primarily concerned about the effect of female wages on fertility. Assume that wages is the k'th element of $X(t)$ in Eq. 1. Then the corresponding parameter β_k measures the effect on the log hazard (ln $h(t)$) of a marginal increase in the wage rate. In the actual specification used in the analysis presented in the next section, it is assumed that the wage effect can be captured by a quadratic polynomial. The parameter corresponding to β_k is then measured by the marginal derivative of this polynomial. This derivative depends on the actual wage rate, and the slope of the curves in Figures 2–4 is a graphic measure of this derivative.

A specification corresponding to Eq. 1) is estimated separately for four cohorts of women, born 1955-59, 1960-64, 1965-69 and 1970-74. Since we have repeated observations (up to three births) of the women in the sample, we can consider unobserved heterogeneity by estimating the transitions to first, second and third birth simultaneously. This is done by assuming that unmeasured characteristics of the women can be divided into two components, one part which is woman-specific and constant across all her births, δ_i , and a second part, $\eta_{ij}(t)$, that is independent across all births (and observations). It is assumed that δ_i is normally-distributed. Formally, we estimate

$$2) \quad \ln h_{ij}(t) = \gamma_j T_{ij}(t) + \beta_j^{'} X_{ij}(t) + \delta_i + \eta_{ij}(t), \quad j = 1,2,3 \,,$$

where *j* is birth number and *i* is woman. The estimate of the standard deviation of the distribution of δ_i is denoted Sigdelta in the tables.

To predict wages we use a set of year specific wage equations

$$3) \quad \ln W_i(t) = \alpha_0(t) + \alpha^{'}(t) Z_i(t) + \mu_i(t),$$

where $\ln W_i(t)$ is the log wage rate, $Z_i(t)$ is a vector that includes length of schooling, potential work experience and experience squared, and $\mu_i(t)$ is an error term. Note that there is one equation with unique parameters for each of the years 1974—2009.

In line with the specification in Heckman (1979), we also consider selection effects in the estimation of the wage equations. The selection model includes the variables age, education (measured in years), number of children 0-2, 3-6 and 7-18 years, respectively, non-labour incomes, a dummy variable for immigrants from countries outside Western Europe and time of residence in Norway. Since estimation is performed separately for each of the years 1974-2009, we have chosen not to present the results. However, in Appendix A we provide the estimation results for a selected year as an example of the results obtained.

Main estimation results

First birth

Table 1 shows the estimation results when we estimate a parsimonious model with only splines in age and duration, and a quadratic function of wages. In addition we have controlled for region of birth, but we have omitted these estimates in the table in order to reduce the dimension of the table. By splitting the sample into four cohorts, we reduce estimation biases due to cohort effects. By estimating the hazard functions for the three first births simultaneously, we also reduce the effects of unobserved heterogeneity across women.

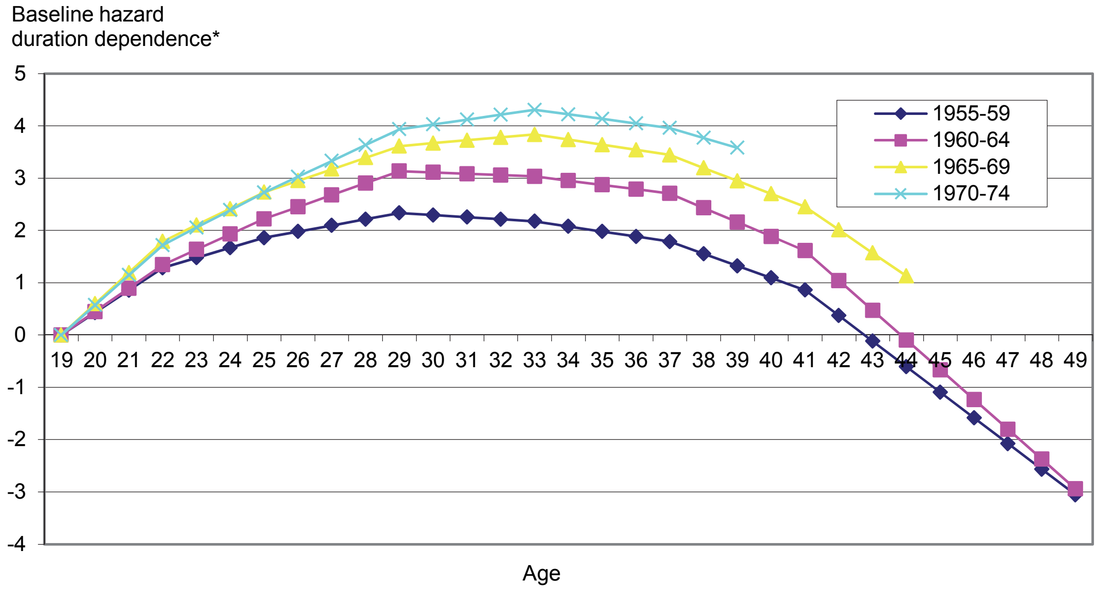
Table 1. Simultaneous estimation of transitions to first, second and third birth.[1] All females

	Cohort 1955-59		Cohort 1960-64		Cohort 1965-69		Cohort 1970-74	
Covariate	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic
First birth								
Constant	-0.28	-4.5	3.84	37.5	2.91	33.0	3.23	34.4
Spline age 19-22	0.43	55.7	0.45	62.8	0.60	77.8	0.57	69.6
Spline age 22-25	0.19	39.3	0.29	58.5	0.31	62.8	0.33	59.5
Spline age 25-29	0.12	28.8	0.23	56.5	0.22	58.5	0.30	76.5
Spline age 29-33	-0.04	-8.1	-0.03	-5.6	0.06	13.0	0.09	20.9
Spline age 33-37	-0.10	-14.8	-0.08	-13.4	-0.10	-17.9	-0.09	-14.9
Spline age 37-41	-0.23	-21.3	-0.27	-27.7	-0.25	-26.4	-0.19	-8.5
Spline age 41-49	-0.49	-28.2	-0.57	-32.5	-0.44	-12.9		
Wage (100 NOK)	-6.12	-50.4	-11.91	-75.8	-10.24	-81.0	-10.11	-81.5
Wage squared	2.35	44.8	4.00	74.6	3.04	81.1	2.72	81.1
Second birth								
Constant	-9.54	-53.9	-11.82	-63.1	-12.11	-68.9	-12.91	-66.2
Years since first birth (spline)								
0-2	3.80	54.5	4.13	58.6	4.38	62.0	4.63	59.3
2-4	0.46	58.3	0.39	50.9	0.35	48.1	0.37	49.1
4-6	-0.29	-35.7	-0.32	-39.4	-0.36	-45.0	-0.35	-40.9
6-8	-0.30	-29.9	-0.23	-23.4	-0.20	-20.4	-0.20	-17.3
8+	-0.17	-42.0	-0.17	-40.8	-0.17	-38.4	-0.18	-29.8
Spline age 19-24	0.11	9.8	0.15	13.0	0.13	11.0	0.17	12.2
Spline age 24-29	0.09	25.0	0.11	28.7	0.13	34.1	0.13	29.5
Spline age 29-34	0.05	10.1	0.05	12.6	0.06	14.8	0.08	19.1
Spline age 34-39	-0.10	-21.0	-0.09	-19.0	-0.08	-17.2	-0.03	-6.0
Spline age 39-49	-0.37	-44.1	-0.38	-45.4	-0.33	-29.3		
Wage (100 NOK)	-2.81	-12.8	-0.47	-2.5	-0.50	-3.3	-0.56	-3.6
Wage squared	1.60	19.3	0.58	9.4	0.44	10.2	0.41	9.8
Third birth								
Constant	-10.45	-25.8	-12.06	-27.9	-10.35	-26.4	-11.09	-24.0
Years since second birth (spline)								
0-2	3.36	32.3	3.58	35.3	3.76	36.8	3.66	33.3
2-4	0.09	7.1	0.06	5.2	0.05	4.1	0.08	6.1
4-6	-0.12	-10.1	-0.12	-10.2	-0.17	-15.2	-0.19	-14.4
6-8	-0.27	-22.0	-0.27	-22.2	-0.25	-19.5	-0.19	-11.9
8+	-0.16	-29.4	-0.15	-26.6	-0.13	-21.4	-0.13	-13.1
Spline age 19-24	-0.01	-0.1	0.05	0.7	-0.15	-2.3	-0.06	-0.7
Spline age 24-29	0.11	13.4	0.14	16.2	0.15	16.6	0.15	14.6
Spline age 29-34	0.03	6.7	0.01	2.7	0.01	2.9	0.00	-0.1
Spline age 34-39	-0.10	-18.7	-0.13	-24.4	-0.14	-27.1	-0.10	-16.3
Spline age 39-49	-0.45	-44.7	-0.42	-43.1	-0.34	-26.7		
Wage (100 NOK)	-1.91	-4.8	-0.20	-0.7	-1.21	-5.4	-0.91	-3.7
Wage squared	1.46	10.1	0.56	5.9	0.67	10.6	0.52	8.3
Sigdelta	0.82	62.5	0.89	79.3	0.89	83.1	0.95	87.0
Log Likelihood	-790649		-811377		-862193		-748027	

[1] Dummies for birth region are also included in the estimations.

From Table 1 we see that the spline in age (*$\gamma T(t)$*) for the first birth has an inverse U-shape. It rises rapidly when the women are in their early twenties, then flattens gradually and reaches a maximum around the age of 30, before turning down at higher ages. As can be seen from Figure 1, which is based on the estimation results in Table 1, the pattern is relatively similar for the four cohorts, but the downturn begins a bit later for the two youngest cohorts according to our results. This finding is consistent with a postponement of the first birth among younger cohorts of Norwegian women compared to older cohorts.

Figure 1. The shape of the first birth log baseline hazard for four different cohorts.



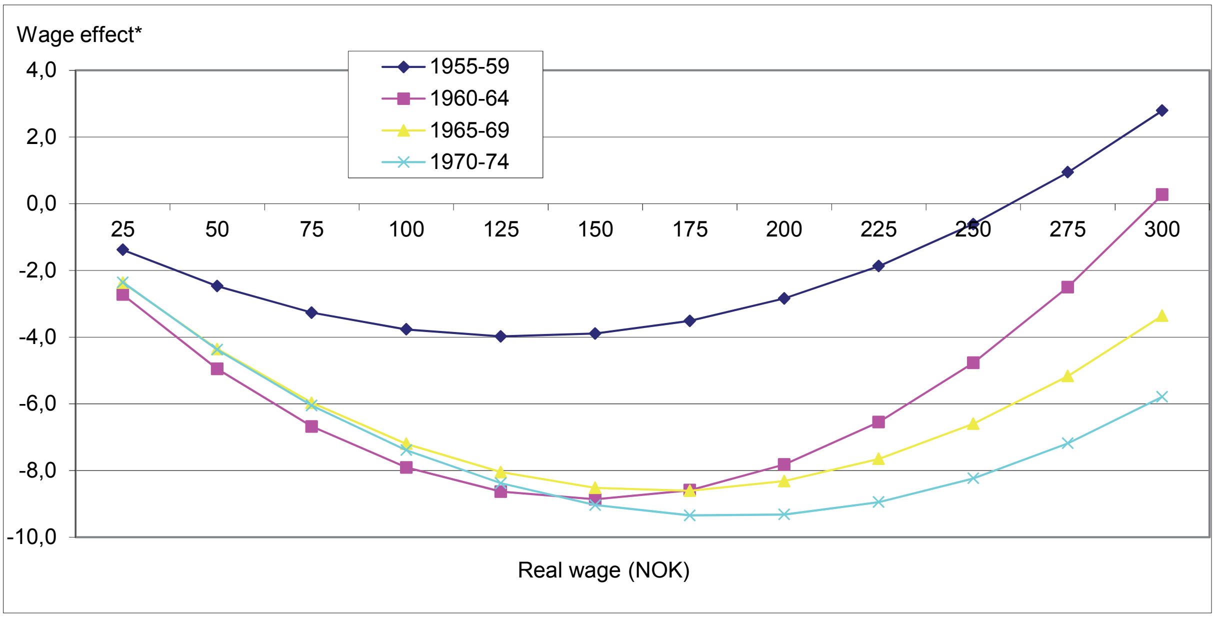
* Baseline hazard duration dependence is the estimated values of $\gamma_j T_{ij}$ in Eq. 2).

Our primary interest is, however, in the wage effect on the hazard of giving birth. From Table 1 we see that all the parameter estimates related to wages are highly statistically significant when we look at the results for the first birth.[1] The estimates suggest that there is a non-linear effect of wages on the hazard of becoming a mother for the first time, for all the four cohorts. The relationship is convex, see Figure 2. This means that for wages below a threshold, an increase in the wage rate reduces the log hazard rate (and hence the hazard rate) for the first birth, while for wage levels above the threshold, an increase in the wage rate increases the rate. In other words, among women with relatively low wage rates, an increase in the wage rate is associated with a longer duration of time until the first birth, while a wage increase is associated with a speeding up of first birth among high income women. According to our estimation results, the thresholds are about NOK 130, 149, 168 and 185 for the cohorts 1955–59, 1960–64, 1965–69 and 1970–74, respectively. To get an idea of the actual wage rate

[1] The standard errors have not been adjusted for the fact that we use predicted wages in the hazard model. The “true” T-statistics are therefore likely to be somewhat smaller than the ones reported in the tables. However, since the standard deviations in the estimation of the wage equations are relatively small, we would not expect this to change the findings.

among Norwegian women relative to these thresholds, note that a woman born in 1955 is 25 years in 1980. If she is born in 1960, she is 25 years in 1985 and so on. Using data from the Labour market section of National Accounts in Norway, we find that the average wage rate among all employed women in Norway, was NOK 118, 127, 138 and 151 in 1980, 1985, 1990 and 1995, respectively. Hence, the average wage rates are a bit to the left of the thresholds for the corresponding cohorts, indicating that many women are on the downward sloping part of the curves in Figure 2, in particular younger women with low education.

Figure 2. The shape of the wage effect on log hazard for the first birth, for four different cohorts



* Estimated as the contribution from wages through $\beta_j' X_{ij}(t)$ in Eq. 2).

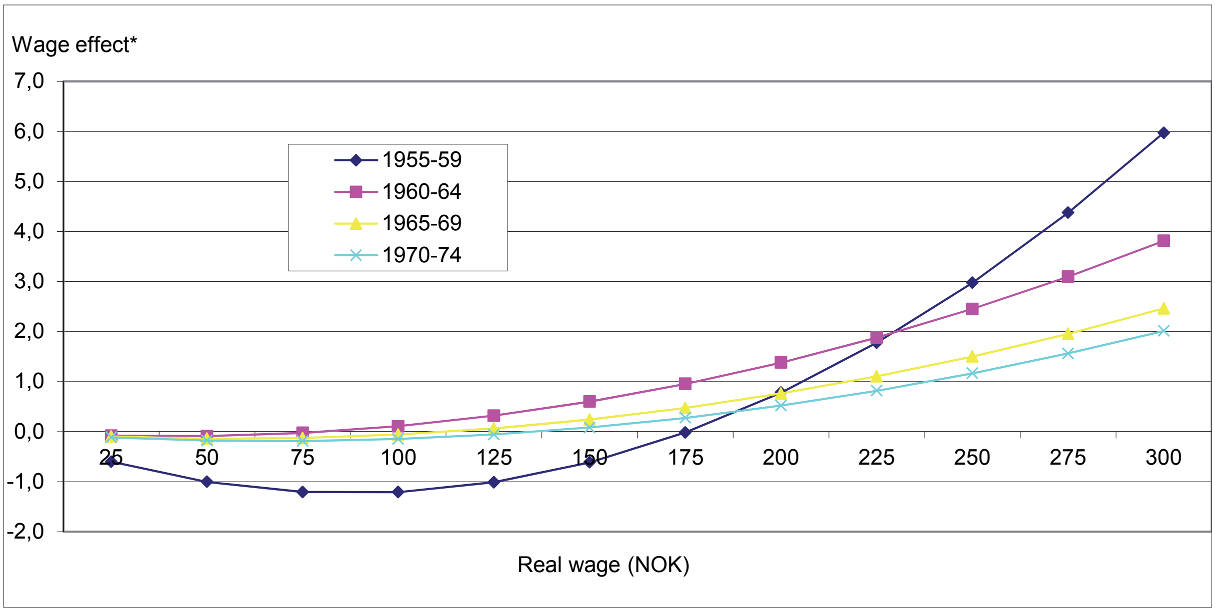
From the results in Table 1 we also notice that the parameter estimates related to wages are relatively stable across the three youngest cohorts, particularly for the cohorts 1960–64 and 1965–69. Note that for the youngest cohort, no woman is older than 35 years in the sample we are using in the estimation, and this may influence the results. We also see that the U-shape is more pronounced among the younger cohorts, in particularly compared to the 1955–59 cohort. This finding indicates that the timing of the first birth is more dependent of the wage rate now than it used to be. During the last decades Norwegian women have steadily increased their labour market participation, and an implication of this might be that our measure of the alternative cost of having children – the wage rate – is more appropriate for the younger cohorts than it is for the oldest cohort.

__Second and third birth__

Table 1 also shows the estimation results for second and third birth. Now duration is measured as years since the previous birth, and a spline with four knots (2,4,6 and 8 years) is used to capture this effect. In addition we control for the woman's age by using a spline with four knots, corresponding to the age 24, 29, 34 and 39.

First we look at the estimation results for the second birth. From the table we notice that all the parameters related to wages (and its squared) are significantly different from zero. Compared to the precision of the corresponding parameters for the first birth, the parameters are now less precisely determined. As for the first birth, there is a nonlinear and convex relationship between the wages and the log hazard for the second birth, see also Figure 3. Once more we find that the parameter estimates are relatively stable across the three youngest cohorts, while the estimates for the oldest 1955-59 cohort differ significantly from the estimates for the other cohorts.

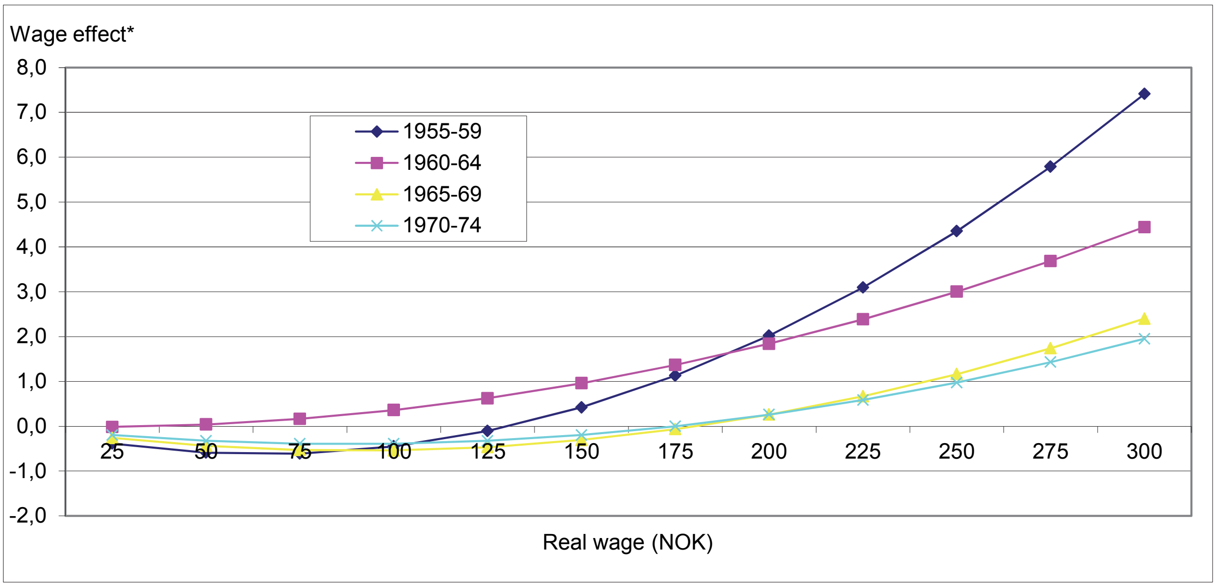
Figure 3. The shape of the wage effect on log hazard for second birth, for four different cohorts



* Estimated as the contribution from wages through $\beta_j^{'} X_{ij}(t)$ in Eq. 2).

By studying the graphs for the second birth in Figure 3, we see that the U-shape is most pronounced for the oldest cohort. However, at least for the women in the three younger cohorts, but also for many women in the 1955-59 cohort, it seems reasonable to assume that their wages lie in the upward sloping parts of the graphs. This means that a wage increase is associated with an increase in the hazard of giving birth to the second child. Compared to the results for the first birth, a shift in wages has a more unambiguous effect on the hazard for the second child since most women are on the upward sloping part of the wage effect curve by the second birth. The curvatures of the wage effect curves are, however, smaller for the second than for the first birth, which implies less sensitivity in the hazard for the second birth of a change in the wage rate.

Figure 4. The shape of the wage effect on log hazard for third birth, for four different cohorts



* Estimated as the contribution from wages through $\beta_j' X_{ij}(t)$ in Eq. 2).

It remains to give an account of the estimates for the third birth. If we look at the estimation results for this birth in Table 1, the overall impression is that the estimates of the wage effects are statistically significant. Figure 4 confirms our previous impression from Figure 3 and birth 2 that the U-shape is most pronounced for the oldest cohort. At least for the two youngest cohorts 1970-74 and 1965-69, but also for the cohort 1960-64 the moderate curvature of the graphs indicates that a wage change has a moderate effect on the hazard for the third birth.

By comparing the graphs in Figure 3 and 4, we see that the shapes of the wage effect curves are quite similar for second and third birth. Since the average wage rate is higher when women get their third child compared to the second one, most women are located more to the right with a higher wage rate in Figure 4 than in Figure 3. This fact tends to increase the effect of a wage increase by the third birth compared to the second one.

Sensitivity analyses

Specifications with age specific wage effects

One might hypothesize that the wage effect on the first birth varies with the age of the woman. In Table 2 we show the estimation results when the model from Table 1 is augmented with interaction terms between wages and dummy variables for different age groups. Women aged 19–24 are chosen as the reference group.

If we look at the estimation results for all the three births in Table 2, we notice that for the first and the second birth, the effects of the interaction terms are statistically significant for all age groups except ages 45-49. And once more we find that the parameter estimates are relatively stable across cohorts. There is a positive shift in the wage effect on the hazard for the first and the second birth, particularly among the age groups 30–44, compared to the reference group. For the oldest cohort, there is a clear positive trend across age in the magnitude of the estimates for both first and second birth. That is, the older the woman is, the less negative or the more positive is the effect of a wage change on the transition to first and second birth. This pattern is also evident for the cohorts 1960-64 and 1965-69 when we look at the second birth, while it is not that clear for the first birth.

For the third birth, we find that only some estimates are statistically significant, but the younger the cohort is, the less precise are the parameters determined. The overall impression is then that there does not seem to be any significant differences in the wage effects across age groups.

Specifications with non-labour income

Becker's analysis of demand for children predicts that non-labour income of the woman should be included in model specifications, and the theory predicts an increased demand for children when non-labour income increases. Since the effect of wages according to this theory can be decomposed into an income effect and a substitution effect, one might hypothesize that the income effect of a wage change might be confounded with the income effect of changes in non-labour income if we do not control for non-labour income. Specifications ignoring this variable might then suffer from a missing variable problem, which might bias the estimation results.

Table 2. Simultaneous estimation of transitions to first, second and third birth. Specifications with age specific wage effects.[1] All females

	Cohort 1955-59		Cohort 1960-64		Cohort 1965-69		Cohort 1970-74	
Covariate	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic
First birth								
Constant	-0.49	-7.9	3.87	34.0	2.83	31.3	3.16	33.3
Spline age 19-22	0.45	57.0	0.50	62.1	0.62	78.4	0.61	70.8
Spline age 22-25	0.15	25.8	0.23	34.5	0.27	44.8	0.28	41.7
Spline age 25-29	0.08	16.8	0.18	36.9	0.19	42.8	0.27	58.6
Spline age 29-33	-0.11	-12.1	-0.06	-6.4	0.00	0.5	0.05	6.9
Spline age 33-37	-0.13	-8.1	-0.09	-5.6	-0.08	-6.1	-0.09	-7.7
Spline age 37-41	-0.25	-14.6	-0.23	-13.9	-0.24	-17.1	-0.20	-8.7
Spline age 41-49	-0.44	-17.0	-0.51	-19.2	-0.42	-11.8		
Wage (100 NOK)	-5.63	-45.2	-11.90	-67.5	-10.08	-77.2	-9.98	-79.2
Wage squared	2.06	37.5	3.88	63.2	2.93	73.7	2.62	76.0
Interaction wage x age								
x age 25-29	0.18	13.6	0.26	18.9	0.15	13.2	0.17	14.8
x age 30-34	0.40	17.3	0.38	17.4	0.28	15.8	0.28	17.0
x age 35-39	0.54	14.8	0.40	12.1	0.26	10.3	0.31	14.5
x age 40-44	0.56	10.3	0.28	5.8	0.21	6.1		
x age 45-49	0.28	2.5	0.13	1.3				
Second birth								
Constant	-9.57	-53.6	-11.86	-58.0	-12.10	-68.3	-12.83	-65.4
Years since first birth (spline)								
0-2	3.79	54.6	4.13	54.4	4.37	62.1	4.62	59.3
2-4	0.46	58.2	0.38	46.6	0.35	47.9	0.37	49.0
4-6	-0.29	-35.9	-0.32	-36.7	-0.36	-45.2	-0.36	-41.2
6-8	-0.29	-29.7	-0.23	-21.4	-0.20	-20.3	-0.20	-17.1
8+	-0.17	-41.8	-0.17	-38.0	-0.17	-38.1	-0.18	-29.8
Spline age 19-24	0.11	9.5	0.14	11.0	0.12	10.1	0.16	11.3
Spline age 24-29	0.07	15.4	0.10	19.4	0.12	24.2	0.12	21.3
Spline age 29-34	0.02	3.9	0.05	7.4	0.05	9.0	0.09	14.7
Spline age 34-39	-0.11	-12.9	-0.11	-13.2	-0.11	-14.9	-0.08	-9.9
Spline age 39-49	-0.37	-20.0	-0.39	-21.2	-0.37	-17.0		
Wage (100 NOK)	-2.64	-11.7	-0.28	-1.3	-0.38	-2.5	-0.55	-3.4
Wage squared	1.48	17.2	0.47	6.8	0.37	8.2	0.37	8.8
Interaction wage x age								
x age 25-29	0.08	4.8	0.06	3.4	0.07	4.8	0.09	5.7
x age 30-34	0.17	7.6	0.10	4.3	0.11	5.7	0.08	4.3
x age 35-39	0.21	6.8	0.17	5.6	0.19	7.8	0.17	7.6
x age 40-44	0.28	5.8	0.26	5.9	0.27	7.7		
x age 45-49	0.05	0.5	0.03	0.4				
Third birth								
Constant	-10.57	-25.0	-12.37	-24.4	-10.58	-25.4	-11.26	-23.1
Years since second birth (spline)								
0-2	3.35	32.3	3.60	32.6	3.74	36.8	3.65	33.3
2-4	0.09	7.0	0.06	4.7	0.05	4.1	0.08	5.8
4-6	-0.12	-10.2	-0.12	-10.0	-0.18	-15.3	-0.19	-14.6
6-8	-0.27	-22.2	-0.27	-20.3	-0.25	-19.5	-0.19	-12.0
8+	-0.16	-29.0	-0.15	-23.7	-0.13	-21.2	-0.13	-12.9
Spline age 19-24	0.01	0.2	0.08	0.9	-0.10	-1.5	-0.04	-0.4
Spline age 24-29	0.11	11.4	0.15	13.5	0.15	14.5	0.15	13.0
Spline age 29-34	0.02	2.7	0.00	0.3	-0.01	-1.1	-0.01	-2.1
Spline age 34-39	-0.13	-14.9	-0.17	-18.2	-0.16	-19.5	-0.12	-12.3
Spline age 39-49	-0.45	-21.7	-0.43	-20.3	-0.33	-13.6		
Wage (100 NOK)	-1.81	-4.5	0.06	0.2	-1.11	-4.8	-0.81	-3.3
Wage squared	1.39	9.5	0.46	4.4	0.63	9.8	0.48	7.6
Interaction wage x age								
x age 25-29	-0.02	-0.5	-0.05	-0.9	-0.06	-1.6	-0.03	-0.6
x age 30-34	0.03	0.6	-0.02	-0.3	0.00	0.1	0.02	0.4
x age 35-39	0.11	2.0	0.08	1.4	0.07	1.4	0.07	1.5
x age 40-44	0.20	2.9	0.15	2.3	0.07	1.2		
x age 45-49	-0.17	-1.3	0.00	0.0				
Sigdelta	0.82	62.6	0.89	74.5	0.90	81.5	0.96	85.3
Log Likelihood:	-790393		-695545		-862003		-747807	

[1] Dummies for birth region are also included in the estimations.

In this section we look into this problem by estimating our reference model with a second order polynomial in wages, but now augmented with non-labour income of the female. For married or cohabiting women the wage income of their partner typically constitutes the main fraction of the woman's non-labour income. In Norway many women are cohabiting before they get married and give birth to the first child. Unfortunately we only observe the union status of married women during the period covered by our data. For practical purposes this means that we cannot identify the male partner for a large part of the women before first birth. Non-labour income is therefore only included in the hazard rate equations for second and third births, and these estimates are based on a subgroup of women who were married at first birth. If the women get divorced after first birth, they are censored at the time of the (first) divorce. The transitions to second and third birth are estimated simultaneously together with the transition to first birth, as before. The reason why we also include first birth in the estimation of the model even though we do not observe union status and non-labour income for these women, is that we want to take unobserved heterogeneity into consideration in the specification of the hazard rates.

Table 3 shows the estimation results. We have omitted the results for the first birth since they are almost identical to the ones in Table 1. The estimates of second and third birth are not directly comparable to those in Table 1, however, as they are based on a selection of women who were married at first birth. To facilitate comparison across model specifications we, therefore report two sets of estimations results for each cohort, one specification with and one specification without non-labour income.

If we look at the estimation results for the second birth, we notice that the parameter estimates related to non-labour income are statistically significant and positive as predicted by theory. This finding is true for all the four cohorts. Thus, a marginal increase in non-labour income increases the hazard rate for the second birth. For the third birth the results are more ambiguous. For the 1965-69 cohort the sign is not in accordance with economic theory, while the estimates are not statistically significant for the other cohorts.

In the introduction to this section we questioned whether the estimates related to the effect of the wage rate might be biased if non-labour income was omitted from model specifications. When considering the results for the second birth, we see that the parameter estimates for the oldest cohort 1955-59 and the two youngest cohorts 1965-69 and 1970-74 are not very much changed by the introduction of non-labour income. From an empirical point of view this means that non-labour incomes is more or less orthogonal to the wage rate. The parameter estimate related to wage squared for the 1965-69 cohort is more precisely determined when non-labour income is considered in model specifications, but for this cohort the estimates of the wage effects are not very convincing. Thus, we conclude that the estimated wage effects are not very sensitive to the omission on non-labour incomes.

Table 3. The effect of introducing non-labour income in the specification of the birth transitions

	Cohort 1955-59				Cohort 1960-64			
Covariate	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic
Second birth								
Constant	-9.15	-37.4	-8.97	-36.7	-10.91	-39.8	-10.40	-38.5
Years since first birth (spline)								
0-2	3.83	44.1	3.84	44.1	4.11	42.7	4.14	43.2
2-4	0.54	52.9	0.54	53.0	0.56	50.0	0.50	45.1
4-6	-0.25	-22.4	-0.26	-22.5	-0.22	-15.6	-0.28	-20.4
6-8	-0.30	-17.7	-0.30	-17.8	-0.28	-12.6	-0.30	-13.6
8+	-0.33	-29.7	-0.33	-29.9	-0.32	-20.4	-0.34	-21.6
Spline age 19-24	-0.03	-1.6	-0.03	-1.6	0.02	1.1	-0.03	-1.8
Spline age 24-29	0.05	7.5	0.05	8.0	0.08	12.0	0.07	9.2
Spline age 29-34	0.04	6.2	0.05	6.6	0.05	7.0	0.04	5.4
Spline age 34-39	-0.10	-16.1	-0.10	-15.8	-0.11	-15.8	-0.10	-13.6
Spline age 39-49	-0.28	-20.9	-0.28	-20.7	-0.37	-26.0	-0.30	-20.2
Wage (100 NOK)	-1.70	-5.4	-1.86	-5.9	-0.37	-1.2	-0.24	-0.8
Wage squared	1.05	8.8	1.13	9.6	0.52	5.2	0.43	4.3
Non-labor inc. (100 NOK)	0.05	11.6			0.02	5.7		
Third birth								
Constant	-11.57	-18.1	-11.57	-18.1	-13.20	-19.4	-12.66	-18.8
Years since second birth (spline)								
0-2	3.50	23.2	3.50	23.2	3.83	22.4	3.86	22.6
2-4	0.25	13.5	0.25	13.5	0.23	12.2	0.24	12.7
4-6	-0.09	-5.7	-0.09	-5.7	-0.10	-6.1	-0.10	-5.8
6-8	-0.27	-14.4	-0.27	-14.4	-0.32	-14.7	-0.31	-14.4
8+	-0.28	-23.4	-0.28	-23.4	-0.34	-22.2	-0.33	-21.3
Spline age 19-24	-0.15	-1.5	-0.15	-1.5	0.00	0.0	-0.07	-0.7
Spline age 24-29	0.00	-0.1	0.00	-0.1	0.09	6.9	0.03	2.0
Spline age 29-34	0.00	0.0	0.00	0.0	-0.01	-0.9	-0.01	-1.1
Spline age 34-39	-0.13	-14.3	-0.13	-14.3	-0.15	-16.3	-0.16	-16.2
Spline age 39-49	-0.47	-26.2	-0.47	-26.2	-0.47	-22.8	-0.47	-22.6
Wage (100 NOK)	1.03	1.8	1.03	1.8	0.91	1.9	1.37	2.9
Wage squared	0.50	2.4	0.50	2.4	0.31	2.0	0.13	0.8
Non-labor inc. (100 NOK)	0.00	-0.4			-0.01	-1.4		
Sigdelta	0.63	25.4	0.63	25.4	0.90	54.6	0.69	32.7
Log Likelihood	-379745		-379807		-555042		-299422	

Table 3. cont.

	Cohort 1965-69				Cohort 1970-74			
Covariate	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic
Second birth								
Constant	-11.24	-36.8	-10.91	-36.4	-12.74	-30.6	-12.74	-30.5
Years since first birth (spline)								
0-2	4.49	39.4	4.53	40.0	5.10	34.5	5.11	34.5
2-4	0.56	44.0	0.50	39.0	0.59	39.7	0.59	39.7
4-6	-0.27	-15.8	-0.34	-19.9	-0.26	-11.7	-0.26	-11.8
6-8	-0.28	-9.6	-0.29	-10.0	-0.26	-6.3	-0.27	-6.3
8+	-0.31	-13.5	-0.33	-14.3	-0.42	-8.4	-0.42	-8.4
Spline age 19-24	-0.05	-1.8	-0.13	-5.2	0.02	0.6	0.03	0.6
Spline age 24-29	0.05	6.4	0.03	3.4	-0.01	-0.9	-0.01	-0.5
Spline age 29-34	0.04	4.9	0.03	3.1	0.07	8.6	0.08	9.2
Spline age 34-39	-0.10	-13.4	-0.08	-10.9	-0.09	-8.8	-0.08	-8.6
Spline age 39-49	-0.35	-16.4	-0.29	-13.5				
Wage (100 NOK)	-0.02	-0.1	0.41	1.4	-0.03	-0.1	-0.02	0.0
Wage squared	0.25	3.1	0.12	1.5	0.22	2.5	0.22	2.6
Non-labor inc. (100 NOK)	0.03	7.2			0.03	5.9		
Third birth								
Constant	-13.31	-16.4	-12.78	-15.8	-13.63	-11.9	-13.64	-11.9
Years since second birth (spline)								
0-2	4.50	19.8	4.54	20.0	4.50	16.7	4.50	16.7
2-4	0.27	12.7	0.28	13.2	0.36	14.4	0.36	14.4
4-6	-0.15	-7.3	-0.14	-6.9	-0.22	-8.4	-0.22	-8.4
6-8	-0.40	-14.1	-0.39	-13.7	-0.41	-9.5	-0.41	-9.5
8+	-0.36	-15.4	-0.35	-14.7	-0.25	-5.8	-0.25	-5.8
Spline age 19-24	-0.13	-0.9	-0.22	-1.6	-0.25	-1.2	-0.25	-1.2
Spline age 24-29	0.05	2.8	-0.04	-2.2	0.03	1.2	0.03	1.3
Spline age 29-34	-0.03	-3.6	-0.04	-4.2	-0.09	-8.4	-0.09	-8.6
Spline age 34-39	-0.16	-15.8	-0.18	-16.6	-0.08	-6.3	-0.08	-6.4
Spline age 39-49	-0.34	-12.5	-0.34	-12.5				
Wage (100 NOK)	0.85	2.0	1.39	3.3	1.59	3.1	1.61	3.1
Wage squared	0.13	1.1	-0.04	-0.3	-0.08	-0.6	-0.09	-0.7
Non-labor inc. (100 NOK)	-0.01	-2.6			0.00	-0.7		
Sigdelta	0.90	49.1	0.66	27.3	1.00	55.0	1.00	55.3
Log Likelihood	-547772		-225401		-479930		-479947	

[1] Dummies for birth region are also included in the estimations.

For the third birth the effect of introducing non-labour income is more ambiguous. Based on both the estimation results in this table and the previous ones, is seems to be more difficult to obtain precise estimates for the third birth, particularly compared to the first but also the second birth.

Specifications with education as a separate variable

Some researchers might argue that we should have included education in the specification of hazard functions. Education might influence preferences for the timing of births, in addition to its effect through wages. Besides it is an indicator of initial human capital accumulation and may reflect differential human capital appreciation and depreciation rates (Mincer and Polachek 1974, 1978; Mincer and Ofek 1982). We understand this argument, but would argue that there is a fundamental identification problem if one wants to include education both in the wage equation and in the hazard function. Therefore, we have chosen to ignore education as a separate argument in the hazard function. However, for those who are not entirely convinced about this argument, we present – in Table B1 in Appendix B, – the estimation results when education measured in years and a dummy for being a student are included in the specifications as separate arguments. Figures B1–B3 in the Appendix shows the corresponding wage effect curves.

From Figures B1–B3 we see that all the wage effect curves are still U-shaped by the introduction of education variables in the hazard function. The main pattern across cohorts within a particular birth is also unchanged by this modification. By comparing the estimated effects of wages and wages squared in Table 1 and Table B1 we find that most of the estimates are not very different. Thus, for first birth the estimated wage effect does not depend on whether or not we include education in the hazards model. For second birth, introducing education yields larger negative parameter estimates related to wages and also larger parameter estimates determining the effect of wages squared. The fact that the estimates related to wages squared are higher means that the U-shape determining the effect of wage changes becomes more pronounced. For third birth there is no systematic change in the estimates determining the effects of wage changes when including education in the hazard function.

Concluding remarks

We have studied the effect of wage changes on the hazard for first, second and third birth. The applied sample is a very large dataset covering all women born in Norway during 1974-2009, and we follow the potential mothers from age 19 to 49. By dividing the sample into four 5-years cohorts, we eliminate cohort effects in the estimation results. Another particular feature of our analysis is that we have access to data from the Norwegian Labour market survey covering the same period. By linking these data to other register data sets with information about wage incomes and education, we estimate wage equations separately for each year included in the analysis. Using these equations to predict wages for all women in the sample, we take into account that the shadow price of raising children is positive and not equal to zero for all women including those that do not participate in the labour market.

Economic theory postulates that a wage increase has both a positive income effect due to a larger family budget, and a negative substitution effect due to a higher "shadow price" of children. Since raising children is time-intensive and mothers are still the main care-givers we expect the substitution effect to dominate the income effect for most women. Thus, in line with New Home Economics (Becker 1960, 1965) we anticipated a negative effect of female wages on fertility.

For first birth, our findings are largely in accordance with predictions. However, the wage effect is not uniformly negative, and turns positive for wages above a certain level, which varies with cohort. By comparing these levels to women's average wages as reported in the National Accounts for the relevant period, we found that most women are on the downward sloping part of the wage effect curve. Hence, a wage increase reduces the first birth hazard for most women. The results indicate moreover that there is a substantial difference in the wage effect between the 1955-59 cohort and the three younger cohorts born from 1960 to 1974 as the U-shape is much more pronounced for the younger cohorts. Thus, a wage change had a larger effect on the first birth hazard for these cohorts than for the 1955-59 cohort. The higher wage sensitivity in the younger cohorts might be due to increased labour market participation rates among these cohorts.

For the second and third birth, most women are on the upward sloping part of the wage effect curve, indicating that the hazard for second and third birth often increases with increasing wage. Besides, the sensitivity of a wage change in different cohorts is the opposite of the pattern we observed for first birth. For second and third birth the wage effect on the log hazard is more pronounced for the oldest cohort compared to the three younger ones.

In line with previous authors (Heckman and Walker 1990a, Merrigan and St.-Pierre 1998) we thus find the strongest negative effects of wage increases on first birth. For second and third birth, the positive income effect seems to dominate the negative substitution effect, and this conclusion holds even when we control for non-labour income.

The reason why younger cohorts seem to respond less to wage changes than older cohorts when further childbearing is concerned may be associated with the expansion of family- and work-related policies in Norway. This has made the female wage an increasingly less accurate proxy for the opportunity cost of childbearing. The paid parental leave has for example been increased from about 4 months in the mid 1980s to more than a year in 2009, and to encourage fathers to get more involved in childcare, an increasing portion of the parental leave has been reserved fathers (from 4 weeks in 1993 to 10 weeks in 2009). Besides, the number of childcare centres has increased considerably, and now most pre-school children have access to formal childcare. A cash transfer introduced in the late 1990s to parents who do not use subsidised care increased the effective price of formal childcare, but later a system of maximum prices has greatly reduced the prices in care centres. The total effect of family policies on the shadow price of having children is not clear without more rigorous calculations, but it is quite clear that the policy expansions have reduced women's opportunity costs over time. To some extent this might explain why the wage effect profiles for second and third birth have become less pronounced for the younger cohorts compared to the older ones.

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Appendix A. Wage equations estimation

Wage rates are predicted using a set of estimated wage rate equations. Since data cover a fairly long period of time — from 1974 to 2009 — it might be assumed that the parameters determining return to schooling and education are not constant over time. Thus, the wage equation is estimated separately for each of the years 1974—2009. Heckman's selection model is used to consider potential selection bias in the estimation results. As an example of the results obtained, Table A1 reports the estimation results for a selected year, year 2000.

Table A1. Estimates of the parameters in the wage equation. Year 2000

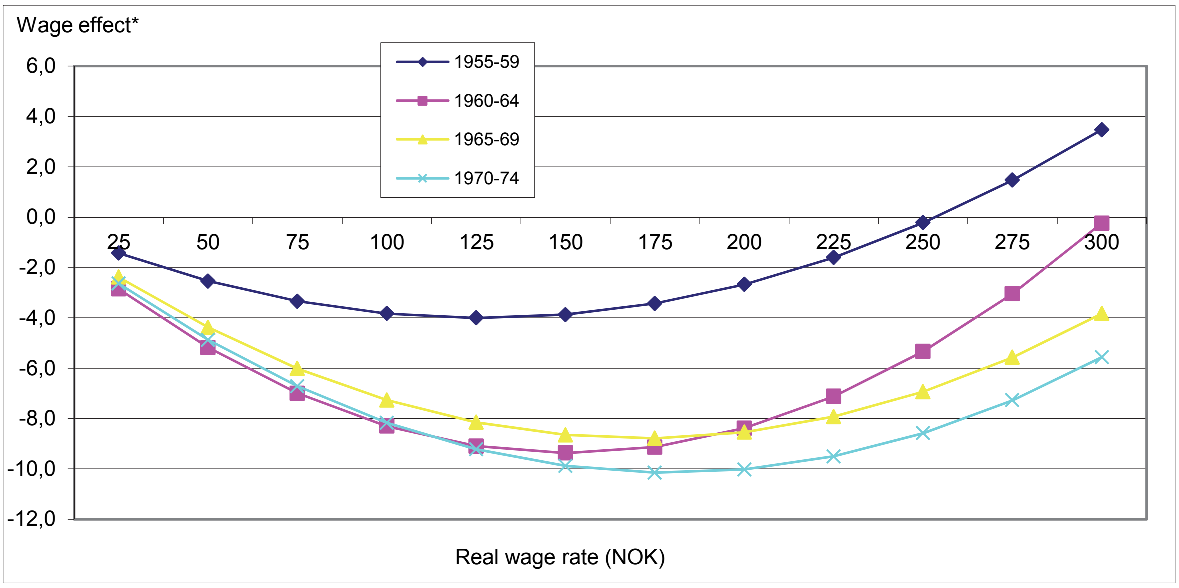
Parameter	Estimate	T-statistics
Log wagerate		
Intercept	3.90	161.4
Education (years)	0.045	27.6
Experience (years)	0.033	29.0
Experience squared	-0.0005	-19.0
Std. error of regression	0.37	104.8
Labour market participation		
Intercept	0.98	7.5
Age	-0.02	-8.5
Education	0.11	14.1
Nbr. children 0-2 years	-0.67	-16.3
Nbr. children 3-6 years	-0.38	-10.8
Nbr. Children 7-18 years	-0.10	-5.3
Non-labor income (NOK 1000)	0.00	-1.2
Non-western immigrant	-1.17	-7.7
Years of residence in Norway, immigr.	0.05	3.8
Correlation coefficient	-0.56	-14.4
Number of observations	10295.00	
Log likelihood	-6624	

Appendix B. Specifications including education variables

Table B1. Simultaneous estimation of transitions to first, second and third birth. Specifications including education. All females

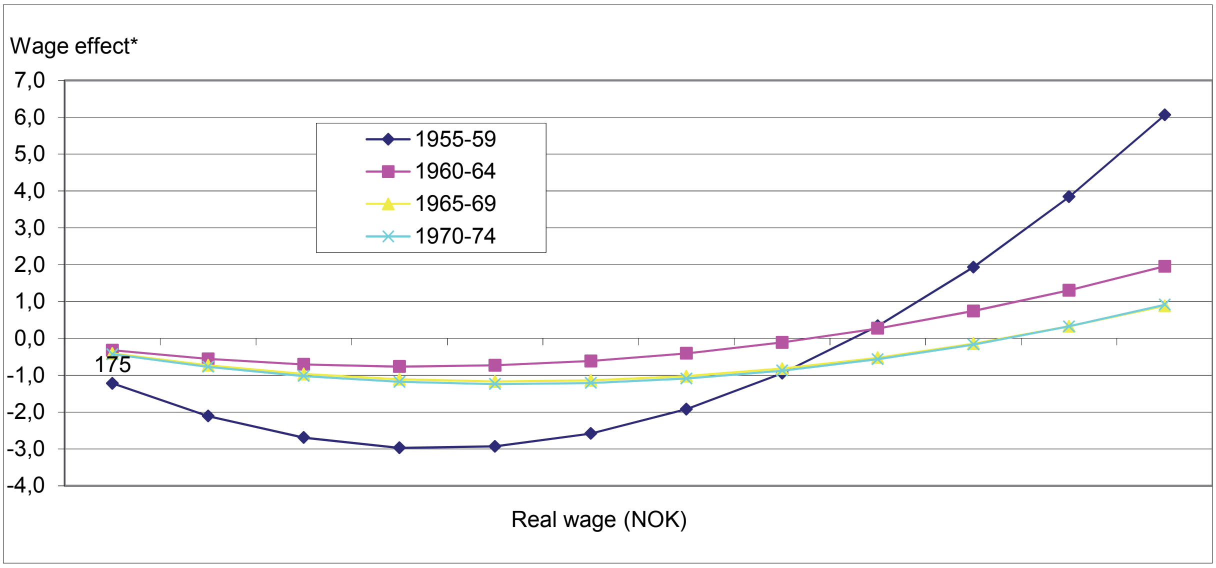
	Cohort 1955-59		Cohort 1960-64		Cohort 1965-69		Cohort 1970-74	
Covariate	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic	Estimate	T-statistic
First birth								
Constant	0.55	7.0	4.84	32.0	3.69	27.4	4.74	32.3
Spline age 19-22	0.25	30.2	0.25	34.5	0.39	47.1	0.39	44.2
Spline age 22-25	0.16	30.3	0.27	48.0	0.28	51.3	0.29	45.1
Spline age 25-29	0.06	13.4	0.18	39.5	0.16	38.2	0.23	47.7
Spline age 29-33	-0.06	-12.2	-0.05	-10.5	0.03	5.6	0.04	8.1
Spline age 33-37	-0.10	-15.8	-0.09	-14.0	-0.10	-17.5	-0.13	-21.4
Spline age 37-41	-0.24	-21.8	-0.27	-27.2	-0.25	-25.9	-0.18	-8.1
Spline age 41-49	-0.50	-28.5	-0.57	-32.5	-0.45	-13.0		
In education	-0.89	-112.0	-0.83	-105.6	-0.82	-113.4	-0.80	-110.1
Secondary educ.	0.16	12.2	0.39	23.8	0.35	20.9	0.45	23.1
University, short	0.29	11.2	0.63	22.7	0.61	22.3	0.61	20.0
University, long	0.19	3.5	0.36	7.8	0.41	10.0	0.05	1.1
Wage (100 NOK)	-6.33	-38.7	-12.42	-53.0	-10.26	-52.4	-11.33	-56.9
Wage squared	2.50	31.2	4.11	53.9	3.00	55.3	3.16	62.4
Second birth								
Constant	-7.57	-35.0	-10.50	-46.6	-10.57	-51.2	-11.19	-47.5
Years since first birth (spline)								
0-2	3.75	54.2	4.07	58.5	4.32	62.0	4.56	59.4
2-4	0.47	59.3	0.40	52.7	0.37	50.6	0.40	51.8
4-6	-0.27	-33.7	-0.30	-37.1	-0.33	-42.1	-0.33	-38.4
6-8	-0.29	-29.2	-0.22	-22.2	-0.19	-19.3	-0.19	-16.9
8+	-0.16	-39.7	-0.16	-38.8	-0.16	-36.1	-0.16	-27.0
Spline age 19-24	0.14	11.4	0.12	10.4	0.11	9.3	0.13	9.7
Spline age 24-29	0.09	21.6	0.11	23.5	0.12	26.2	0.12	21.2
Spline age 29-34	0.04	8.0	0.03	7.6	0.04	8.9	0.05	9.9
Spline age 34-39	-0.10	-22.1	-0.09	-17.2	-0.09	-17.7	-0.07	-12.3
Spline age 39-49	-0.37	-42.3	-0.37	-43.0	-0.33	-29.4		
In education	-0.67	-44.6	-0.61	-47.9	-0.60	-56.8	-0.57	-53.9
Secondary educ.	0.25	14.1	0.16	7.9	0.18	8.8	0.14	5.5
University, short	0.47	15.2	0.41	13.0	0.42	13.5	0.34	9.2
University, long	0.27	4.8	0.35	7.3	0.28	6.1	0.01	0.2
Wage (100 NOK)	-5.47	-16.2	-1.47	-5.2	-1.82	-7.9	-1.92	-7.5
Wage squared	2.32	17.9	0.71	7.8	0.70	11.1	0.74	11.1
Third birth								
Constant	-10.17	-20.3	-10.96	-23.0	-8.54	-19.7	-9.09	-17.3
Years since second birth (spline)								
0-2	3.37	32.3	3.61	35.4	3.80	36.9	3.69	33.3
2-4	0.12	9.1	0.10	8.1	0.09	7.4	0.12	8.9
4-6	-0.10	-8.3	-0.09	-7.7	-0.14	-12.5	-0.16	-12.0
6-8	-0.25	-20.5	-0.25	-20.2	-0.23	-17.6	-0.17	-10.8
8+	-0.15	-27.3	-0.14	-24.4	-0.12	-19.0	-0.11	-11.6
Spline age 19-24	0.01	0.2	0.10	1.4	-0.08	-1.2	0.07	0.8
Spline age 24-29	0.09	10.0	0.12	13.0	0.12	13.3	0.13	11.7
Spline age 29-34	0	7	0	3	0	6	0	2.6
Spline age 34-39	-0.11	-21.4	-0.12	-21.4	-0.13	-23.0	-0.11	-16.4
Spline age 39-49	-0.44	-41.7	-0.41	-40.3	-0.35	-27.4		
In education	-0.33	-14.2	-0.34	-17.9	-0.31	-18.2	-0.26	-14.5
Secondary educ.	-0.01	-0.5	0.07	2.5	0.12	4.0	0.09	2.1
University, short	0.33	7.1	0.51	11.8	0.58	12.9	0.53	8.9
University, long	0.28	3.6	0.62	9.6	0.63	9.8	0.47	5.5
Wage (100 NOK)	-1.41	-2.1	-1.00	-2.3	-3.07	-9.0	-3.14	-7.4
Wage squared	0.88	3.5	0.40	2.9	0.88	9.4	0.88	7.8
Sigdelta	0.72	51.7	0.78	60.1	0.77	60.8	0.82	62.2
Log Likelihood	-782200		-803260		-853017		-739451	

Figure B1. The shape of the wage effect on log hazard for the first birth, for four different cohorts. Specifications with education variables



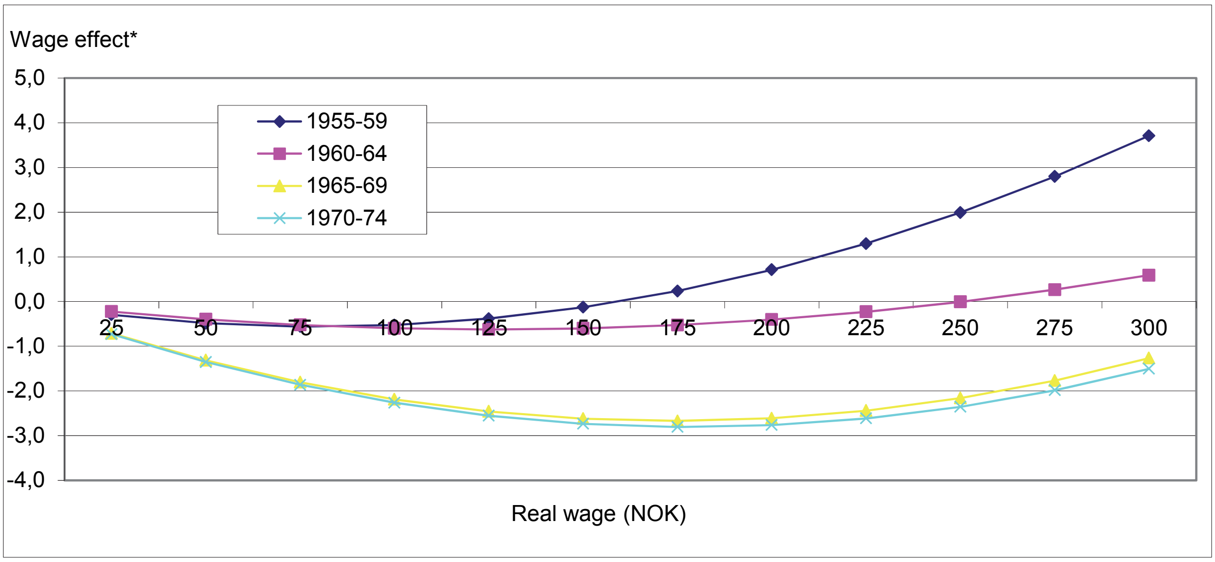
* Estimated as the contribution from wages through $\beta_j' X_{ij}(t)$ in Eq. 2).

Figure B2. The shape of the wage effect on log hazard for the second birth, for four different cohorts. Specifications with education variables.



* Estimated as the contribution from wages through $\beta_j' X_{ij}(t)$ in Eq. 2).

Figure B3. The shape of the wage effect on log hazard for the third birth, for four different cohorts. Specifications with education variables



* Estimated as the contribution from wages through $\beta_j^{'} X_{ij}(t)$ in Eq. 2).

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