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Marriage, Motherhood and Wages: Recent Evidence

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Abstract

Using data from the 1982 National Longitudinal Survey of Young Women (NLSYW) we replicate Korenman and Neumark's (K&N) (1992) study "Marriage, Motherhood and Wages" and obtain similar OLS estimates. Applying IV estimation to account for the endogeneity of experience and tenure, K&N find evidence in favor of the endogeneity and obtain more negative coefficients for the children coefficients. In the IV estimation we obtain experience and tenure coefficients which are not significantly different from the OLS estimates, hence rejecting the endogeneity. As in K&N's study, our first difference specification exhibits signs of heterogeneity bias. Sample selectivity bias is marginally evident in K&N's paper, wheras we do not find evidence in favor of it. We furhtermore examine the development of wage differentials over time by looking ar more recent NLSYW surveys (1987 and 1991). We find that as women age, expereience, tenure and education become better predictors for wage differentials among women with different marital and fertility status.

I. Introduction

The influence of marital and fertility status on earnings has been the subject of numerous studies. Researchers often compare the influences of marriage and children on wages between men and women. Children are generally found to have adverse effects on women's wages, while the same variables positively affect men's wages. Cain (1986) proposes two possible reasons as follows:

> (case 1) "children have no real effect on productivity of either men or women but that employers have a uniform preference for paying men with children more and women with children less" or,
>
> (case 2) "women with children are less committed to market work than women with no children, and men with no children are less committed to their jobs than men with children" This point is also made by Becker (1993).

Case 1 states discrimination as the reason for the observed outcome. Case 2 highlights heterogeneity, both between men and women as well as among women. Cain further explains the underlying rationale:

> "prior equality in tastes between men and women is often denied on the grounds that cultural and biological forces, which are presumed exogenous to the economic system (or, more narrowly, to the labour market), are the causes of a preference for market work relative to housework among men and vice versa for women".

Empirically, studies further specifying the phenomenon of gender specific wage differentials find three hypothesis for the earnings gap. Firstly, evidence has been found that employers have preferences for paying higher wages to married men [with children] than to married women or single men (Hill 1979; Bartlett and Calahan 1984). Secondly, it is hypothesized, by Becker (1981) and Nakosteen and Zimmer (1987), that men who marry are more productive than unmarried men or married women. The third hypothesis contradicts the previous approach: Becker (1985) and Kenny (1983) suggest that the act of marriage makes men change their behaviour and become more productive as opposed to unmarried men or married women. However, Korenman and Neumark (1990) conclude that, although finding “support for the hypothesis that marriage enhances men’s labour market productivity, a definitive judgement requires more evidence about the process or causal mechanisms linking marriage to...productivity”. In a British study, Greenhalgh (1980) finds evidence that married men earn more and married women earn less because of family specific behaviour. Hence, it is of interest how this specialisation influences women’s earnings. The consequences of the resultant gender role specialisation suggest that either women's behaviour is brought about by power relations in their interaction with men (McCrate 1987, 1988) or that women develop comparative advantage in household responsibilities. Gunderson and Riddell remark that "differences in household responsibilities can lead to a situation where women develop a comparative advantage in household tasks and men develop a comparative advantage in labour market tasks...whereby women become less prone to accumulate the continuous labour market experience that leads to higher earnings".

Looking at earning differentials among women, the same questions arise: are married women different than single women or does marriage make women exhibit a different behaviour. There are individual characteristics which are unobservable and vary according to the environment and individual preferences. These differences bring about differences in career orientation, ability and effort as well as in marriage and fertility decisions. This can lead to bias by unmeasured heterogeneity[1]: unmeasured characteristics make women self-select into different status’ of fertility (Hill 1979, Dolton and Makepeace 1987[2]), for example least career women rather select into having more children. These self selection decisions then produce (an upward) bias in OLS estimates of the effect of marriage and children on wages . Endogeneity bias also arises if labour supply and/or fertility decisions are not exogenous to the equation but depend on the level of wages. It is not clear whether the number of children is dependent on the obtained wage level or vice versa. In the case of endogenous fertility decisions the effect of children would be downward biased, e.g. overstated. Hence the implications from endogeneity bias lead effectively to the same predictions as in the case of heterogeneity bias[3]. Furthermore, a sample selection bias could exist: married women earning relatively high wages tend to select into employment as their opportunity cost of leisure is higher compared to low earning married women.

In a detailed study by Korenman and Neumark (1991) [K&N], they found that marriage has no significant influence on [log hourly] wages[4]. However, a strong negative relationship does arise when two or more children are present (with the coefficient of -.30). When controlling for education, experience and tenure this negative relationship is weakened but remains significantly negative (-.07). Before drawing conclusions from this attenuation, K&N examined whether experience and tenure were endogenous to the wage equation[5]. K&N find some (weak) evidence that experience and tenure are endogenous, proposing that the negative effects of having two or more children could be more pronounced. Furthermore, heterogeneity bias is detected by their first-differences estimates. Finally, selectivity corrections and the corresponding tests exhibit no clear signs of self-selection bias. As K&N point out however, it is difficult to establish any causal interpretation from these findings.

The centre piece of our work is to replicate Korenman and Neumark's result of the 1982 survey and to update and alter the specification through more recent data (1987 and 1991). Our research for 1982 confirms the basic results of K&N that suggest that there are forms of heterogeneity which lead to bias of the regression. Our instrumental variables estimation and also the selectivity correction gives no clear evidence for bias. In contrast, the first difference approach exhibits signs of heterogeneity biased OLS estimates. Throughout our calculations we found that the OLS and IV estimation techniques were sensitive to the sample exclusion restrictions. Additionally, the IV results and selectivity correction results were sensitive to slight changes in the IV's or participation predictors. Hence, one must be cautious in interpreting these results.

Utilizing a longitudinal survey allows us to investigate about the impacts and nature of marital status and fertility influences in greater detail. The issue is whether these factors will have a permanent or temporary effect on women's wages. If the influences are temporary, looking at several surveys would then show relatively higher wages before childbirth and a wage increase with increasing age of the child. On the other hand, permanent influences on wages could prevail because women's heterogeneity can lead to long-term interactions between the level of education, choice of number of children and the corresponding labour force attachment.

These long-term economic consequences can be seen as manifestations of heterogeneity among women. According to their preferences (or environment), and based on different expected stocks of human capital (as explained in Goldin and Polachek 1987), women set forth early on the path of lifetime wages by making an ex-ante education-marital status-fertility decision. We find evidence for group specific effects. For example, women with several children have relatively lower education and a disproportionably low experience level. This supports the early career package choice proposition. The data suggests that three main groups of women can be distinguished by their choices of "packages":

> Package 1 chosen by career oriented women:
> They have either chosen a high education level or are very attached to the labour market (highest experience level) and choose to have no children. A high fraction of these women are single[6].
> Package 2 chosen by less career oriented women:
> These women are mainly married, well educated and have one child[7]. The loss of years experience is proportionate to experience of child rearing.
> Package 3 chosen by least career oriented women:
> They are married, less educated and have primarily two or more children. Furthermore, they are less attached to the labour market and have therefore accumulated disproportionably few years of experience.

Our theory was augmented by the recent development of the literature, since some researchers have pointed out components of the above approach. It makes sense that heterogeneity unfolds very early in life. Although individuals underlie changes throughout life, the manifestation of their development and of their heterogeneity is always subject to their conditioning. Polachek and Kim (1994), who base their inferences on life-cycle models, find “that most individual-specific differences manifest themselves early in one's work career, probably even in the type of schooling received". For example, there is evidence that late child bearers (typically package 2), have significantly increased wages, as opposed to early child bearers (typically package 3). We see the reasons in women’s heterogeneity and not in the delay of childbirth *per se* (Chandler et al. 1994). Those who choose package 3 primarily have two or more children but their labour force attachment is disproportionably low. Even as children mature, these women do not become proportionally more attached. Nakamura and Nakamura (1994) find that labour supply behaviour is based on each woman’s labour supply history and that number of children can be a predictor for it.

Our theory underlies that these packages are chosen by different types of women. Thus the earnings function does not pick up the pure effect of the included variables, rather it picks up these different types of women who manifest themselves by choosing different types of career packages. Of course it is somewhat arbitrary to divide the whole population of women into three categories. We are aware that there are many forms of mixtures and smooth transitions[8]. Our aim is not to account for all the possible types of women but to give a convenient categorization of women’s heterogeneity and career choices.

The career package hypothesis makes it necessary to look at subsequent surveys as this data exhibits more evidence for the package-presumption. The full extent of such a package choice can only be evident after all components have fully developed over time - later in a woman’s life. Hence, examinations should focus on women who have primarily achieved their aimed education level, their ideal marital status and number of children and consequently their family role specialisation. In turn these achievements determine their attachment to the labour market [determining the level of experience, or level of effort and motivation]. However, the absolute wage gap between the specific marriage and fertility categories has been increasing over time, indicating that there are indeed long term effects of marital status and fertility decisions. We presume that these growing wage differentials are not due to the fact of marriage or children per se. Rather they are due to the occurring consequences of the one chosen package type[9].

Therefore, we will mainly focus our discussion on the latest wave of the survey (1991) as the permanent influences of heterogeneity and the resulting interaction of education, experience and number of children can be observed in 1991 more accurately. As the magnitude of these various influences is of interest, we will focus on regressions. Regressions allow for tests to the determination of heterogeneity, endogeneity and for the sensitivity of specifications. A further advantage of the following regression analysis is that the interactions of marital status and fertility variables, and of education and experience can be examined more carefully. For this purpose we account for structural changes over time by working with pooled data and taking longer differences to discover how changes in the regressors effect changes in log earnings over time.

II. Descriptive Statistics

In this section we will explore a table of mean earnings arranged by marital status and fertility categories (columns) and by variables like age, education and experience (rows). We want to find more explanation of the development of the wage differential over time and according to the specific categories. Statistically speaking, wage differentials across categories are approximately significant if they exceed 0.04.

Table of Mean Characteristics Arranged by Marital & Fertility Status, 1982

	married	**divorced/ separated**	**notmarried**	**nochild**	**onechild**	**two children**	**three children**
avg. age	32.69	33.04	31.70	31.65	32.00	33.02	33.78
avg. wage	6.33	6.40	6.54	6.55	6.43	6.32	6.19
avg. edu	13.31	13.18	14.99	14.77	13.60	13.08	12.62
avg. exp	7.58	7.88	8.85	8.67	8.42	7.49	6.88
avg. kids	.99	1.52	.06				

Table of Mean Characteristics Arranged by Marital & Fertility Status, 1982

	married	**divorced/ separated**	**notmarried**	**nochild**	**onechild**	**two children**	**three children**
observ.	1166	339	194	453	339	564	241

Summarizing the contents of the table, it can be said that women with children have lower average wages than childless women. Also, married women earn less than nonmarried women. Moreover the differentials across and within different categories of women appear to be stable over time (compared to the results of 1987 and 1991, see Appendix). These findings are not necessarily evidence of a simple story. Finding constant growth in wages does not mean that there is no absolute widening of the wage gap between the categories.

The question is whether differentials in wages are based on changes of the characteristics of the composition of each category. This would mean the results are not brought about by a time invariant relationship and therefore could not be described as stable. Rather, it could be due to a time variant, evolving pattern. For example, it is known that returns to experience are increasing over time. Therefore, we would expect an absolute and relative widening of the wage gap. As we do not observe a relative increase of the gap in the table, there must be offsetting movements. One can think about a diminishing influence of children as they mature. On the other hand, the wage gap does not decrease absolutely as the women who have raised children are less career oriented and are, because of this heterogeneity, less attached to the labour market, even as the children mature.

However, it is not possible to draw clear conclusions from these findings. They reflect a more complex evolution of each women's heterogeneity and its effects on earnings. Therefore, more aspects have to be taken into consideration. We presume that there are interactions at the education level, years of experience and age, simultaneously influencing wages. To broaden the basis of discussion, we take average education, experience and age into account in the subsequent section.

Given that child rearing leads to less labour force attachment we expect experience to decrease with the number of children. The fact that the average age increases with the number of children should increase (potential) experience. Examination of the data reveals that these counter movements do not offset each other. This means that women with more children tend to have disproportionably less experience than women with less or no children[10]. On the other hand, married women as opposed to nonmarried women have 2.26 years of less experience (21% wage loss) and per average 1.52 children. Even when taking the average number of children into account, married women have disproportionably less experience compared to single women. An explanation could be that married women who have relative little opportunity costs of staying home are less attached to the labour market. Even before marriage these women had relatively low expected wages. This is a sign for early selection of a woman's career package.

Looking at the education level and assuming that in general the education is completed in this age range, age differences across specific categories should not matter. The fact that women with several children have relatively lower education and a disproportionably low experience level supports evidence of early career package selection. Three types of women distinguished by their selection are discussed in the Introduction.

<u>Package 1 chosen by career oriented women</u>
primarily single, high education level or high(est) experience level, no children.
<u>Package 2 chosen by less career oriented women</u>
primarily married, well educated, one child.
<u>Package 3 chosen by least career oriented women</u>
married, less educated, mainly two or more children.

The absolute wage gap between the specific marital status and fertility categories has been increasing over time, showing that there are indeed long term effects of marital status and fertility decisions. We presume that these growing wage differentials are not due to the fact of marriage or children per se. Rather they are due to the evolving consequences of the once chosen package type. Furthermore we presume that there is a correlation between the education level and demographics. Therefore we will focus in our discussion on three waves of the survey (1982, 1987 and 1991) to examine the development of the permanent influences of heterogeneity and the resulting interaction of education, experience and number of children, say the evolution of a woman's career package.

III. Data and Empirical Findings

A. Data

The data used in this study are from the National Longitudinal Survey of Young Women (USA) which commenced in 1968 by observing 5159 women aged 14-24. In order to replicate K&N's research, we try to construct similar cross-sectional specifications[11] that are mainly from surveys in 1980 and 1982. Moreover, the data for 1985, 1987, 1988 and 1991 are used to update and modify the specifications in the models.

B. Empirical Findings

Table 1 to 4 are the replications of K&N's study which include the OLS regression, instrumental variables estimates, sample selection correction and fixed-effects estimates.

Part I

1.1 OLS Estimates for 1982

Table 1 reports the estimates from OLS cross-sectional log wage regressions. The results are the replication of part C (table 2) in K&N's paper.

Table 1

Wage Equation Estimates for White Working Women, 1982, Ordinary Least Squares[a]
Dependent Variable: Natural log of hourly earnings

	(1)	(2)
Married, spouse present	-0.006 (0.044)	0.051 (0.038)
Divorced or separated	0.019 (0.048)	0.106 (0.042)
One Child	-0.130 (0.038)	-0.031 (0.033)
Two+ Children	-0.313 (0.033)	-0.069 (0.031)
Education	---	0.060 (0.004)
Experience	---	0.024 (0.004)
Tenure	---	0.029 (0.003)
F-test[b]	0.0001	0.0155
R^2	0.0819	0.3072

a. There are 1,698 observations. Using the same criteria as in K&N's study, observations are included only if the wage reported is for the current job. However, Korenman and Neumark only have 1,207 observations. They might have excluded some of the observations according to other criteria. We tried more stringent exclusion restrictions for the construction of experience, resulting in a decrease of the experiment size and making the regressions even more sensitive to different specifications. However, our results with the bigger sample are similar to K&N's results. Never married and no children aged 28 are the reference categories and therefore contained in the intercept. Single-year age dummy variables are included in all specifications. Standard errors are reported in parentheses.
b. In all specifications controls of region of residence included [south via. non-south and urban vs. SMSA]
c. P-value for joint test of significance of marital status and fertility variables.

We confirm the finding of K&N. When including work experience and tenure in the regression the magnitude of the coefficients of fertility variables are reduced. Therefore, after controlling for experience and tenure, marriage and children have less association with wages. Our results in column (2) indicates the coefficient of two or more children becomes less negative with -0.07 and are consistent with K&N's finding. Therefore it can be concluded that differences in accumulation of experience and tenure are good proxies for a woman's heterogeneity and a good indicator of her career package selection.

Though more observable variables were added in the regression to eliminate the potential bias, K&N suggested that there may be certain

> "sources of bias associated with unobservables: endogeneity; heterogeneity arising from selection into different categories of marital status and number of children on the basis of unmeasured characteristics correlated with wages and selection into employment."

In the following section , we will pursue the procedures used by K&N to explore these sources of bias.

1.2 Instrumental Variables Estimates

Following Korenman and Neumark, the instrumental variables we used fall into two categories: family background variables and measures of attitudes and expectations. The variables are almost the same as those used by Korenman and Neumark[12].

Two stage least square is used to estimate the coefficients for the same regression in Table 2. Experience and tenure are treated as potentially correlated with the wage equation error[13].

Table 2

Wage Equation Estimates for White Working Women, 1982, Two Stage Least Squares[a]

Dependent Variable: natural log of hourly earnings

	OLS[b]	Experience and Tenure Endogenous	
Coefficients	**(1)**	**(2)[c]**	**(3)[d]**
Married, spouse present	0.051 (0.038)	0.037 (0.043)	0.063 (0.041)
Divorced or separated	0.106 (0.042)	0.082 (0.059)	0.135 (0.049)

	OLS[b]	Experience and Tenure Endogenous	
Coefficients	**(1)**	**(2)[c]**	**(3)[d]**
One child	-0.031	-0.046	-0.014
	(0.033)	(0.042)	(0.037)
Two+ children	-0.069	-0.139	-0.033
	(0.031)	(0.085)	(0.049)
Experience	0.024	0.004	0.014
	(0.004)	(0.029)	(0.024)
Tenure	0.029	0.016	0.056
	(0.003)	(0.035)	(0.024)
F-test[e]	0.0155	0.0417	0.0409
Family background as IVS	No	Yes	Yes
Expectations/attitudes as IVS	No	No	Yes
Test for Overidentification[f]	---	0.6858	0.072
Hausman Test[g]: Experience	---	0.463	0.177
: Tenure	---	0.141	1.312

a. There are 1,698 observations. Standard errors are reported in parentheses. Single-year age dummy variables and controls for region of residence were included in all specifications.
b. The OLS regression is the same as in Table 1, column (2).
c. Using family background variables as instruments.
d. Using family background - and attitudinal/expectational variables as instruments.
e. P-value for joint test of significance of marital and fertility status variables.
f. P-value of F-test for overidentifying restrictions.
g. Statistics of Hausman test (Chi-squared distribution).

The corresponding first stage regression for column (2) (see Appendix, table 12) shows that experience is significantly lower for women with two or more children. Among the instruments, number of siblings and the estimate of whether the mother worked when respondent was aged 14, are deemed significant. The remaining four IV's are insignificant predicting experience. When using tenure as an independent variable, again the estimate of two or more children is significantly negative. None of the instruments are significant. Additional use of family background variables as IV's in

Table 2, column (2) indicates that the first stage regressions have the same significant coefficients as the previous specifications and none of the additional instruments are significant in predicting experience or tenure. Turning to the second stage for the family background instruments, the p-value of overidentification test is 0.68 (greater than the critical value of 0.05). This implies that the validity of the instruments is not rejected. The results of the Hausman tests show that the IV estimates of Experience and Tenure are not different from the OLS estimates. Hence we cannot reject the null hypothesis that there is no mis-specification of the OLS regression**.** This result differs from K& N's study which rejected the exogeneity of experience and tenure. This may be due to a slightly different sample. When both family background and expectations/attitudes are used as the IVS, the p-value for the overidentification test is 0.06. The finding is quite significant and is closer to K&N's result than in column (2)[14]. In both studies, the coefficients for experience and tenure are insignificant. However, we find no evidence of the endogeneity of experience and tenure.

1.3 Selectivity Correction

Our estimation, based on a higher proportion of high earning women than there are in the population, could be upward biased. If this is the case, then employment decisions of women are correlated to their marriage and fertility decisions (ie. career package selection). Correcting the selection bias of employment should decrease the impact of the marital and fertility variables on wages.

We employed Heckman's (1979) selectivity correction method and a linear Probit model is used to estimate the unobservable heterogeneity among women.

Table 3

Wage Equation Estimates for White Working Women, 1982 Sample Selectivity-Corrected (SSC) Estimates[s]
Dependent Variable: natural log of hourly earnings

	OLS	SSC[b,d]
	(1)	**(2)**
Married, spouse present	0.064 (0.040)	0.064 (0.044)
Divorced or separated	0.074 (0.044)	0.074 (0.046)
One child	-0.066 (0.036)	-0.067 (0.045)
Two+ children	-0.169 (0.033)	-0.170 (0.055)
λ	---	0.002 (0.197)
F-test[c]	0.0001	0.0025

a. There are 1,868 observations in the Probit procedure and 1,471 observations in the OLS estimation.The employment Probit includes all variables used K&N's study (except for number of weeks the husband was unemployed the previous year since only very small portion of respondents answered this question in the survey) to estimate the value for the inverse of Mill's ratio (λ), which then is used as a regressor in the Probit model.
Standard errors are reported in parentheses. Single-year age dummy variables and controls for region of residence and education were included in all specifications.
b. Variables in the employment Probit include all variables in Table1, column (2), as well as dummy variables indicating whether a respondent changed marital or fertility states between 1980 and 82 and measures of husband's income, sum of income from alimony and child support (set to 0 for never-married women). K&N did include the variable of weeks husband spent unemployed in 1982 which we excluded from the Probit procedure because it causes a serious decrease in the number of observations.
c. P-value for joint test of significance of marital status and fertility variables.
d. Value of Log Likelihood for normal: -911.31

The coefficient of λ is very small (0.002) and insignificant (t-statistic of 0.008) which does not support that the OLS estimates are biased by self-selection. Moreover, our OLS and SSC regressions give approximately the same coefficient estimates for other independent variables, showing that we do not need to include the inverse Mill's ratio in subsequent specifications. K&N found λ as 0.13 and marginally significant, (t-statistic of 1.86) and that the estimates of the fertility variables' coefficients become more negative when selectivity was taken into account. The differences in results propose that the estimate of λ is sensitive according to the specification and sample size used.

1.4 First Differences Estimates

It may be the case that women are less attached to the labour market not because of marriage/ divorce or children *per se* but because of unobserved heterogeneity. To examine this hypothesis, first differences are applied.

Table 4

Wage Equation Estimates for White Working Women, First-Difference Specification[a]
Dependent Variable: levels and changes of natural log of hourly earnings

	1982[b]	**82-80 change**[c,f]
Coefficients	**(1)**	**(2)**
Married, spouse present	0.050 (0.041)	-0.039 (0.044)
Divorced or separated	0.070 (0.044)	0.028 (0.036)
One child	-0.033 (0.036)	-0.027 (0.041)
Two+ children	-0.112 (0.032)	-0.010 (0.029)
F-test[e]	0.0034	0.7235

a. There are 1,396 observations. Only women, for whom data in the years 1982 and 1980 are available, were included in this test, decreasing our sample size. Change in experience and tenure is not included in the equation as it is only focused on women who are employed in the two consecutive years 1980 and 1982. Standard errors are reported in parentheses. Controls for changes in education and changes in region of residence were included.
b. OLS regressions using same variables as Table 1, column (2), without experience and tenure.
c. Standard fixed-effects estimator.
e. P-value for joint test of significance of marital status and fertility variables (column 1) or of significance of changes in marital status and fertility status (column 2).
f. A specification test rejected selection into employment on basis of wage growth. See Appendix, Table 14.

All estimates of the difference specification were found to be insignificant. It is concluded that controlling for unobserved heterogeneity different marital and fertility states should not have influence per se on women's wages. Rather, the adverse effects exhibited by OLS could be caused, among others, by heterogeneity among women.

In Part 1, we sought to replicate K&N's study by constructing the same variables and procedures. Having obtained almost identical estimates for the OLS specification, K&N's results for IV estimation exhibit signs of endogeneity of experience and tenure, whereas we did not find evidence in favor of it. Both, K&N's and our first difference specification show signs of heterogeneity bias. Our results for the sample selectivity did not exhibit evidence of biased OLS estimates, whereas K&N find some evidence in favor of it. One has to be careful in drawing conclusions from these differences in results, as the specifications were sensitive to the sample exclusion restrictions used. In part, it questions the application of these estimation techniques. In order to further investigate the impacts of endogeneity and heterogeneity among women and to find evidence for our career package hypothesis, we modified the specifications and use more recent waves of the survey for analyses in the following sections.

Part 2: Survey of 1991

2.1 OLS Estimates for 1991

Looking at the 1991 survey, we hope to determine the route of this story. The same regressions in Table 1 are run using the data for 1991 wave and the results are reported in the following table.

Table 5

Wage Equation Estimates for White Working Women, 1991, Ordinary Least Squares[a]

Dependent Variable: Natural log of hourly earnings

	(1)	**(2)**
Married, spouse present	0.009 (0.064)	0.067 (0.051)
Divorced or separated	0.003 (0.067)	0.092 (0.054)
One Child	-0.121 (0.053)	0.013 (0.044)

	(1)	(2)
Two+ Children	-0.279 (0.045)	-0.031 (0.038)
Education[b]	---	0.094 (0.005)
Experience	---	0.024 (0.003)
Tenure	---	0.015 (0.002)
F-test[c]	0.0001	0.3456
R^2	0.0476	0.3800

a. There are 1,266 observations. Observations are included only if the wage reported is for the current job. Never married and no children aged 37 are the reference categories and therefore contained in the intercept. Single-year age dummy variables and controls for region of residence are included in all specifications. Standard errors are reported in parentheses.
b. Education level as of 1987.
c. P-value for joint test of significance of marital status and fertility variables.

As shown in column (2), the joint impacts of marital and fertility status on earnings had been decreasing over time (jointly zero on a 35% level). One possible reason for this may be that the marital and fertility variables themselves do not effect the wages of women, rather they reflect the preference for a specific career package among women. It is then the level of education, experience and tenure implied by the package that exerts influences on wages.

Compared to 1982 (positive divorce premium present) and to 1987 (positive marriage premium present) (see Appendix, Tables 17 - 21), the marriage coefficients have both become insignificant. This means that more heterogeneity has been taken away by labour market controls, especially by the education coefficient. It can be taken as a sign that package3 choosers marry earlier than package 2 choosers, such that the marriage coefficient had become positive in 1987 compared to 1982. Comparing 1991 to 1982 the fertility variables also become insignificant. One proposition is that as the children mature, career unoriented women participate relatively more in the labour market. A counter proposition is that as consequences and effects of package selection unfold, the education, experience and tenure variables capture more of the heterogeneity of the specific package choice. The impact of education changes over time which becomes more positive. This may due to the increasing dispersion of return to schooling. However, it may also be due to the sample selection bias that career oriented women tend to have more education and higher wages over time. On the other hand, the results in 1991 are quite different from the results in 1982 and 1987, which indicates that there may be structural changes within the sample over time.

2.2 Pool 1982 - 1991

Table 6
Wage Equation Estimates for White Working Women, Pooled Data 1982 & 1991[a]
Dependent Variable: natural log of hourly earnings

	(1) **Unrestricted**	**(2)** **Restricted**
Married, spouse present	0.059 (0.031)	0.510 (0.039)
Divorced or separated	0.102 (0.034)	0.106 (0.042)
One child	-0.013 (0.027)	-0.031 (0.034)
Two+ children	-0.030 (0.240)	-0.069 (0.032)
Education	0.074 (0.003)	0.060 (0.004)
Experience	0.031 (0.002)	0.024 (0.004)
Tenure	0.191 (0.001)	0.029 (0.003)
d[b]	---	-0.028 (0.135)
d(Married, spouse present)	---	0.016 (0.064)

	(1) Unrestricted	(2) Restricted
d(Divorced or separated)	---	-0.014 (0.068)
d(One child)	---	0.044 (0.054)
d(Two+ children)	---	0.038 (0.049)
d(Education)	---	0.033 (0.007)
d(Experience)	---	0.000 (0.005)
d(Tenure)	---	-0.014 (0.003)
F-test[c]	0.0275	0.0001
R^2	0.4873	0.5058

a. There are 2,964 observations. Standard errors are reported in parentheses.Controls for region of residence and single-year age dummy variables were included in all specifications for wage levels. We included 18 age dummies, of which there are 9 dummies for the age groups in the specific year.
b. We pooled 1991 with 1982 by including dummies which were set to equal one in the year 1991 and to zero in 1982. If these dummies turn out to be jointly different from zero, we find evidence for a structural change. Contrary to a Chow Test, we can specify where this break comes from by looking at the t-statistics of the specific dummy. This will allow us to conclude how the influence of certain variables evolves over time.
c. P-value for joint test of significance of marital status and fertility variables for the unrestricted regression P-value for joint test of significance of all change dummies of the restricted regression.

As can be seen from table 6, the negative influences of children have been weakened in 1991 compared to 1982 (when running a regression only for 1982), which has the effect of reducing the wage gap. This can result from older children, who need less time spent on their rearing. In 1987 (see Appendix), the influence of one child begins to decline, in contrast the negative effects of two of more children become more pronounced. Moreover, the returns to education have been steadily increasing over time, indicating that the effects of a higher education become more apparent. This makes the wage differentials more pronounced. In addition, returns to experience had increased in

1987, but remain constant in 1991 compared to 1982. This can result from a slight downward bias of the experience coefficient in 1991 implying package 3 selectors face declining returns to experience. The additional returns for tenure also seem to decrease over time. All these findings suggest that there are complex interactions between the fertility variables and education and experience which are not time consistent. These changes in influences make it necessary to run different regressions for each survey and show that the effect of heterogeneity is indeed time variant. However, all this is evidence for the hypothesis that there are long-term effects of a chosen career package.

Although we think that the 1991 OLS estimates do not suffer from bias, we conduct the corresponding tests since we know about the sensitivity of the OLS estimates to sample size and exclusion restrictions.

2. 3 Instrumental Variables Estimates

Table 7
Wage Equation Estimates for White Working Women, 1991, Two Stage Least Squares [a]
Dependent Variable: natural log of hourly earnings

	OLS[b]	**Experience and Tenure Endogenous**	
Coefficients	**(1)**	**(2)**[c]	**(3)**[d]
Married, spouse present	0.067 (0.051)	0.048 (0.066)	0.059 (0.053)
Divorced or separated	0.092 (0.054)	0.038 (0.081)	0.064 (0.063)
One child	0.013 (0.044)	-0.063 (0.065)	0.006 (0.048)
Two+ children	-0.031 (0.038)	-0.271 (0.104)	-0.078 (0.059)
Experience	0.024 (0.003)	-0.060 (0.035)	0.014 (0.017)
Tenure	0.015 (0.002)	0.010 (0.018)	0.004 (0.013)

	OLS[b]	Experience and Tenure Endogenous	
Coefficients	**(1)**	**(2)[c]**	**(3)[d]**
F-test[e]	0.3456	0.0599	0.2904
Family background as IVS	No	Yes	Yes
Expectations/attitudes as IVS	No	No	Yes
Test for Overidentification[f]	---	0.0675	0.0003
Hausman Test[g] :Experience	---	5.890	0.384
:Tenure	---	0.252	0.659

a. There are 1,266 observations. Standard errors are reported in parentheses. Single-year age dummy variables and controls for region of residencewere included in all specifications. Same family background and attitudes/expectations IV's as in 1982 are used here.
b. The OLS regression is the same as in Table 5, column (2).
c. Using family background variables as IV's.
d. Using family background- and attitudinal/expectational variables as IV's.
e.P-value for joint test of significance of marital and fertility status variables.
f.P-value of F-test for overidentifying restrictions.
g. Statistics of Hausman test (Chi-squared distribution).

In 1982 and 1987 the overidentification test did not reject the validity of family background instruments. In 1991 they are only accepted on a 7% level. Moreover, in 1982 the enodogeneity of both experience and tenure was rejected, in 1987 tenure was marginally found to be endogenous and in 1991 experience is found endogenous. This demonstrates that either these specifications are sensitive to the sample used or that there are structural changes occurring. When using family background variables as IV's, the first stage regression for experience (see Appendix, Table 13) shows that one child leads to a significantly lower experience level (-0.96 years) and having two or more children decreases experience by -2.4 years. Among the instruments, only the dummy for whether the mother worked when the respondent was aged 14 is significantly positive. Using tenure as an independent variable in the first stage regression, we find that women with higher levels of tenure also appear to have higher levels of education. From the instrument list only the dummy for living with father and mother when aged 14 is significant, but surprisingly, negative. It is worth remarking that beginning in 1982 the R^2 of the first stage regressions has been constantly decreasing, indicating that the quality of the instruments decreased.

Given that family background variables are valid instruments to control for the endogeneity of experience and tenure, as indicated by the overidentification test, it can be inferred from Table 7 that the OLS estimate for Two or more children is upward biased, i.e., the negative impact of this variable is understated. In fact, all estimates are not significantly different from the OLS estimates except for this two plus children coefficient. It proposes, when instrumenting experience and tenure, that we are again adding heterogeneity to the regression. This also becomes clear when looking at how the experience variable has changed. In the OLS regression we found positive returns to experience, whereas after instrumenting it, the estimate becomes insignificant. We conclude therefore, that experience is probably endogenous, a good predictor and indicator of a woman's heterogeneity and career package choice. When adding attitudinal IV's to the instrument list, the overidentification test rejects their validity, as experienced in 1987 and 1982 (only marginally). This could be due to the fact that these IV's are the attitudes and expectations of the respondents at early ages and the predictive power of these variables over present work experience and tenure may be diminishing over time.

2.4. Selectivity Correction

In order to examine whether or not there is selection bias for employment decision in 1991, we again apply Heckman's selectivity sample correction and the results are shown in the following table.

Table 8

Wage Equation Estimates for White Working Women, 1991 Sample Selectivity-Corrected (SSC)Estimates [a]
Dependent Variable: natural log of hourly earnings

	OLS	**SSC[b,d]**
	(1)	**(2)**
Married, spouse present	0.050 (0.057)	0.100 (0.063)

	OLS	SSC[b,d]
	(1)	**(2)**
Divorced or separated	0.060 (0.060)	0.040 (0.061)
One child	0.001 (0.049)	-0.002 (0.049)
Two+ children	-0.120 (0.042)	-0.095 (0.044)
λ	---	-0.590 (0.322)
F-test[c]	0.0036	0.0202

a. There are 1,364 observations in Probit procedure and 1,175 observations in the OLS estimates. Standard errors are reported in parentheses. Single-year age dummy variables amd controls for region of residene were included in all specifications.
b. Variables in the employment Probit include all variables in Table 5, column (3), as well as dummy variables indicating whether a respondent changed marital or fertility states between 1988 and 91. Measures of husband's income, sum of income from alimony and child support (set to 0 for never-married women). As in Table 4, the variable of weeks husband spent unemployed in 1991 is excluded from the Probit procedure.
c. P-value for joint test of significance of marital status and fertility variables.
d. Log likelihood for normal: -522.86

The results show that the coefficients of Married and One child are insignificant after correcting for the selectivity bias and hence have no impact on wages. The coefficient of Two or more children becomes more negative and remains significant, suggesting that the negative impact of Two or more children is being overstated in the OLS estimate. This implies that there may be some negative selection among women, i.e., women with more children tend to participate in the labour market and earn lower wages, e.g. if they have to work for financial reasons[15]. This finding contradicts the 1982 result, where no signs of selectivity were found, but the results confirm the finding of 1987, where λ turned out being significant with an estimate of -0.684. The change in the outcome is probably due to a change in the composition of the sample which started after 1982 (more and more women were ending up in their position implied by the career package). In this 1991's specification, the λ becomes insignificant, but as the Two plus children coefficient changes it makes sense taking the inverse Mill's ratio into account. The question arises as to whether this SSC approach should be given much weight in our analysis. Firstly, the portion of non-working women in the sample has become very small (14%) as opposed to previous years (1982: 21 %; 1987: 15 %) Secondly, the variables used in the Probit procedure do not have good predictive power on the labour participation decision. Therefore, there is no concrete evidence to support that the OLS regression does suffer from self-selection bias of employment.

2.5. First Difference Estimates 1991 - 1987

Table 9
Wage Equation Estimates for White Working Women, First-Difference Specifications[a,e]
Dependent Variable: levels and changes of natural log of hourly earnings

	1991[b]	**91-88 change**[c]
Coefficients	**(1)**	**(2)**
Married, spouse present	0.024 (0.063)	-0.053 (0.057)
Divorced or separated	-0.024 (0.067)	-0.005 (0.053)
One child	-0.094 (0.052)	0.141 (0.111)
Two+ children	-0.222 (0.044)	-0.028 (0.059)
F-test[d]	0.0001	0.6149

a. There are 1,047 observations. Standard errors are reported in parentheses.Only women for whom data in the years 1991 and 1988 are available, were included in this test, decreasing our sample size. Changes in experience and tenure are not included in the equation as it is only focused on women who are employed in the two consecutive survey years 1988 and 1991. Controls for changes in region of residence were included in the specifications.
b. OLS regressions using same variables as Table 1, column (2) but without experience and tenure.
c. Standard fixed-effects estimator.
d. P-value for joint test of significance of marital status and fertility variables (column 1) or of significance of changes in marital status and fertility variables (column 2).
e. A spcification test rejected selection into employment on basis of wage growth (see Appendix, Table 14).

The first differences confirm once more the presence of heterogeneity bias, as all coefficients become insignificant (t-stat < I1I). In 1982 we obtained a similar result, whereas in 1987 we obtained a significantly negative marriage premium. We contributed this to the possibility that in 1987 package 3 choosers who have relatively lower wages had downward biased the marriage coefficient. However, we can conclude that changes in earnings are brought about to a large extent by unmeasured heterogeneity: by career package selection and not the influence of marital status and fertility variables itself. The joint test of the marital and fertility variables suggests, on a 0.61 level, that these variables are jointly zero[16].

In all previously considered surveys we have found signs of heterogeneity bias. It is interesting to explore how these earnings develop over several years after a change in marital status or fertility has occurred. Conclusions may then be drawn about the long-term effects of these changes. We expect to find more evidence for or against our hypothesis.

Due to our findings that education and experience influence wages significantly in determining women's wages, it is of interest to examine how long-term changes impact on wage growth. Therefore we now examine the difference from 1991 to the basis year 1982.

2.6. Nine Year Difference 1991 - 1982

Table 10
Wage Equation Estimates for White Working Women, First Difference Specifications[a]
Dependent Variable: levels and changes of natural log of hourly earnings

	1991[b]	**91-82 change[c]**
Coefficients	**(1)**	**(2)**
Married, spouse present	0.069 (0.051)	0.031 (0.042)
Divorced or separated	0.071 (0.054)	0.082 (0.043)

	1991[b]	91-82 change[c]
Coefficients	**(1)**	**(2)**
One child	0.024 (0.043)	-0.036 (0.069)
Two+ children	-0.014 (0.038)	0.025 (0.038)
Education	0.096 (0.005)	0.069 (0.032)
Experience	0.023 (0.003)	0.012 (0.008)
Tenure	0.013 (0.002)	0.01 (0.002)
F-test[d]	0.4930	0.3425

a. There are 1,135 observations. Standard errors are reported in parentheses. Single-year age dummy variables were included in all specifications for wage levels. Other independent variables are controls for region of residence.
b. OLS regressions using same variables as Table 1, column (3).
c. Standard fixed-effects estimator.
d. P-value for joint test of significance of marital status and fertility variables (column 1)
P-value for joint test of significance of changes of marital status and fertility variables (column 2).

As we now consider a longer time-span, we also include variables for changes in education and experience. When estimating a nine year difference[17], again we find fixed effect bias, proposing that cross sectional OLS estimates suffered from heterogeneity bias. The differenced marital and fertility variables are jointly zero on a .34 significance level. The only significant estimates are the differences in education and tenure, showing tenure matters more than experience. As almost all women have differences in experience and tenure since 1982, the decreased influence of experience and tenure can also be a sign that the OLS coefficient had been upward biased. This resulted as women, who receive higher earnings and returns to experience because of their career orientation, tend also to have higher levels of experience. The proposition is that these variables (experience, tenure and education) are endogenous to the difference equation as well as good predictors of the career package selection. The OLS education coefficient could have been upward biased for the same reason. In the case of education, the difference estimates can still be upward biased, as it is strongly correlated with the decision of career oriented women who choose higher levels of education[18].

Another problem arises with this longer difference estimation: The fixed component which is assumed to be different out refers to the effect that this individual component has on wages. Reasoning with our package hypothesis this means that women choose in an early age their specific package. This choice, in general, is fixed. However, the consequences of this choice change over time, as proposed by the descriptive statistics and the pooled data. The effects of such a choice must therefore be considered in a long-run, dynamic framework. This means that longer difference estimates actually do not cancel out the individual's effect entirely, for example when differences in growth rates are varying. Therefore, shorter differences are more adequate in terms of cancelling out fixed effects (as they do not change rapidly) but they do not contribute much to our discussion if the story evolves in a dynamic way. In this case inferences can only be drawn when considering wage determinants in a broader or more holistic way. The reason for using the long difference is that we can control for changes in the independent variables, such as education and experience, that could not be controlled before using a short difference. Therefore, we expect the long difference to absorb a considerable portion of the heterogeneity. This benefit should balance the above mentioned possibility of partly biased difference variables. Nevertheless, we will instrument the difference approach[19].

2.7. Nine Year Difference Instrumental Variable Estimation

Table 11

Wage Equation Estimates for White Working Women, 1991, Two Stage Least Squares [a]

Dependent Variable: change in natural log of hourly earnings between 1982 and 1991

	Nine Year Difference[d]	Experience and Tenure Endogenous	
Coefficients	**(1)**	**(2) Whole set of IV's[b]**	**(3) Subset of IV's[c]**
d Married, spouse present	.031	.0001	.062
	(.042)	(.059)	(.066)
d Divorced or separated	.082	.049	.122
	(.043)	(.062)	(.071)
d One child	-.036	-.013	-.086
	(.069)	(.093)	(.096)
d Two+ children	.025	-.107	.056
	(.038)	(.061)	(.064)
d Experience	.012	-.219	.095
	(.008)	(.095)	(.116)
d Tenure	.01	-.008	.041
	(.002)	(.020)	(.035)
d Education	.069	.044	.095
	(.032)	(.046)	(.052)
F-test[e]	.3425	.3826	.2422
Family background as IVS	No	Yes	Yes
Test for Overidentification[f]	---	0.0011	.0002
Hausman Test[g] :Experience	---	4.811	.512
:Tenure	---	0.813	.772

a. There are 1,135 observations. Standard errors are reported in parentheses. Controls for changes for region of residence were included in all specifications.
b.Same family background IV's as in previous specifications are used here.
c. Using the following instruments only: father's education, parent's education goal, mother worked. See overid equation in the Appendix, table 16.

d. The Regression is the same as in Table 10, column (2).
e. P-value for joint test of significance of marital and fertility status variables.
f. P-value of F-test for over identifying restrictions.
g. Statistics of Hausman test (Chi-squared distribution).

We are using family background variables as instruments only, as expectational variables were rejected by instrumenting the OLS approach. In the first stage regression (see Appendix, Table 15) with difference in experience as independent variable we find that having only one child over the time period 1982-1991 had no significant effect on experience. This is what we expected, as women with one child were classified as career oriented and attached to the labour market (package 2 selection). In contrast, women with two or more children significantly accumulated disproportionably less experience which is a sign for package 3 selection. We can conclude that it is more the effects of working vs. nonworking, captured with the difference in experience first stage, that contributes to our theory[20].

The second stage, based on family background instrumental variables, exhibits evidence of a marginally different dExperience coefficient by conducting a Hausman test. The coefficient of dTenure is found to be exogenous. However, as the overidentification test rejects the validity of the instruments used (p-value .01%) it suggests that the two stage least squares coefficients are inconsistent and do not contribute much to our story. Hence, we cannot conclude that dTenure and dExperience were endogenous to the first difference equation. Although applying IV estimation, the influence of the marital status and fertility variables is not changing. This indicates our difference equation is robust to these various estimation techniques.

To find out which of the instruments causes the rejection of the validity, we constructed the overidentification equation (see Appendix, Table 16) by regressing the residuals of the second stage on the set of instruments. Through use of significant variables only as instruments in the subsequent second stage equation, the overidentification test rejects the validity of this new set of instruments even more strongly. This suggests that we have lost predictive power by excluding the insignificant instruments, such that the new instruments as a whole are less efficient than before[21].

To summarize: we found no evidence of the bias of the dExperience and dTenure estimate in the long difference estimation. This confirms our presumption of upward biased OLS coefficients. From the presence of this bias we derive evidence for our package hypothesis as it indicates that mainly career oriented women with no child or one child accumulate more experience and higher returns than non-career oriented women with several children.

IV Conclusion

Having confirmed the basic results underlying K&N's paper the further purpose of this paper is to specify, with recent data from 1991, three possible effects that marital status and fertility variables can have on wages. Firstly, these variables can have independent effects on wages (effects *per se*). Secondly, they can affect the acquisition of human capital which would lead, via lower experience levels, to a relative reduction in wages. Thirdly, women's marital status and fertility decisions can be correlated to unobservables. Different types of women choose different marital and fertility states (choice of career package) such that differences in wages are a function of the underlying type of woman, and not of her marital or fertility status per se.

In regressing wages on marital status and fertility variables we find that both, in 1982 and 1991, only fertility decisions have significant influence on wages. The R^2 of the regression decreases from .0819 in 1982 to .0476 in 1991 showing that the predictive power of the regressors diminishes over time.

The children variables become less negative and the influence of one child on wages becomes insignificant when we include labour market controls, experience and tenure, and education, in 1982. In addition the marriage coefficient is also insignificant. Due to an uneven age distribution across the different marital status and fertility categories we conclude that high transition (that can lead to survey year specific bias, i.e. a significant 10% divorce premium) across categories is in evidence. Therefore representative inferences cannot be drawn from the 1982 survey. The presumed heterogeneity (or endogneity) downward bias is partly absorbed by the inclusion of experience, tenure and education.

In 1991 the marital status and fertility variables are insignificant in themselves and jointly at a 35% level. Apparently experience, tenure and education have considerably absorbed the previous bias of these variables. Hence, wage differentials are due to differences in experience, tenure and education. We consider these variables as good proxies for a woman's heterogeneity. The R^2 has increased from 0.30 in 1982 to 0.38 in 1991 illustrating that these variables become better predictors when women have completed the career position implied by package selection, say later in life.

In 1991, first differences confirm the zero influence of the marital and fertility variables on wages. Applying IV estimation and conducting an overidentification test we find that the instruments lose predicitive power over time such that we cannot draw further inference from this specification. A sample selectivity correction gives no clear evidence in favor of sample selection bias.

These findings for 1991 are consistent with the outcomes predicted by our career package hypothesis. Furthermore, we conclude that there is no influence of marital status and fertility variables per se on wages. Rather, having children decreases the accumulation of human capital to a magnitude of approximately -.7 years per child and affects wages only via a relatively lower experience level. The prevailing wage differentials can be largely explained by heterogeneity. Consequently the earnings function does not only pick up the effect of the regressors but specifically the effect of the underlying type of woman according to her package choice. Hence, our work should be understood as an extension of the model underlying K&N's findings: we confirm their basic results for the year 1982 and conclude that it is rather necessary to take the timevariant evolution of wage determinants into account, by looking at several survey years.

Appendix

I. Further Tables

Table of Mean Characteristics Arranged by Marital & Fertility Status, 1987

	married	divorced/ separated	notmarried	nochild	onechild	two children	three children
avg. age	37.63	38.12	36.65	36.94	37.23	37.76	38.44
avg. wage	6.68	6.72	6.89	6.90	6.80	6.64	6.60
avg. edu	13.63	13.38	15.14	14.78	13.86	13.41	13.32
avg. exp	12.45	11.97	13.25	13.34	12.99	11.58	11.32
avg. kids	1.95	1.67	0.17	---	---	---	---
Observ.	994	299	162	317	257	522	221

Table of Mean Characteristics Arranged by Marital & Fertility Status, 1991

	married	divorced/ separated	notmarried	nochild	onechild	two children	three children
avg. age	41.57	42.16	41.39	41.16	41.27	41.69	42.19
avg. wage	6.94	6.97	7.14	7.16	7.04	6.94	6.85
avg. edu	13.64	13.58	15.05	14.89	13.71	13.66	13.27
avg. exp	15.57	15.79	17.68	17.59	16.53	15.41	15.22
avg. kids	2.03	1.81	0.14	---	---	---	---
observ.	892	263	111	246	213	470	216

Table 12

First Stage of Instrumental Variable Estimation of Wage Equation Estimates for White Working Women,OLS Specification[a b,] , for the year 1982

dependent variable:	experience[c]	tenure[c]	experience[d]	tenure[d]

dependent variable:	experience[c]	tenure[c]	experience[d]	tenure[d]
married, spouse present	-.122 (.248)	-.389 (.360)	.087 (.252)	-.360 (.368)
divorced or separated	-.051 (.269)	-.961 (.390)	.162 (.275)	-.928 (.400)
one child	-.126 (.214)	-.542 (.310)	.002 (.215)	-.488 (.313)
twoplus children	-1.575 (.197)	-.891 (.291)	-1.312 (.202)	-.781 (.298)
education	-.002 (.032)	.008 (.047)	-.071 (.045)	-.048 (.066)
experience	--- ---	.577 (.032)	--- ---	.566 (.033)
tenure	.275 (.015)	--- ---	.265 (.015)	--- ---
siblings[e]	.117 (.040)	-.105 (.059)	.100 (.041)	-.100 (.059)
father's education	-.030 (.025)	-.057 (.036)	-.030 (.025)	-.058 (.037)
mother's education	.024 (.029)	-.016 (.043)	.010 (.029)	-.023 (.043)
parents goal	-.011 (.020)	-.022 (.029)	-.012 (.020)	-.021 (.029)
mother worked when respondent aged 14	.398 (.141)	-.284 (.205)	.418 (.141)	-.253 (.206)
respondent lived with father and mother when 14	.510 (.489)	-.183 (.709)	.552 (.492)	-.222 (.719)
hdisagr	--- ---	--- ---	.766 (.841)	-.250 (1.228)
hagr	--- ---	--- ---	1.027 (.849)	.026 (1.240)
agem	--- ---	--- ---	.046 (.035)	-.065 (.051)

dependent variable:	experience[c]	tenure[c]	experience[d]	tenure[d]
idkids	---	---	-.120	-.147
	---	---	(.096)	(.140)
eduexp	---	---	.087	.026
	---	---	(.049)	(.071)
egoal	---	---	-.053	.051
	---	---	(.050)	(.073)

a.there are 1698 observations
b.single year age-dummies and controls for region of residence are included in all specifications
c.with family background variables as IVs
d.with family background variables and attitudinal/expectational variables as IVs
e. see endnote #10 for explanatnions about the construction of the IVs

Table 13

First Stage of Instrumental Variable Estimation of Log Wage Equation Estimates for White Working Women,OLS Specification[a ,b] , for the year 1991

dependent variable:	experience[c]	tenure[c]	experience[d]	tenure[d]

dependent variable:	experience[c]	tenure[c]	experience[d]	tenure[d]
married, spouse present	-.112 (.433)	-.361 (.820)	.186 (.440)	-.396 (.843)
divorced or separated	-.073 (.458)	-1.960 (.867)	.249 (.467)	-1.999 (.894)
one child	-.965 (.366)	.707 (.695)	-.699 (.368)	.694 (.707)
twoplus children	-2.398 (.316)	-.531 (.613)	-2.041 (.322)	-.534 (.627)
education	.043 (.051)	.310 (.095)	-.047 (.072)	.309 (.137)
experience	--- ---	.507 (.052)	--- ---	.513 (.053)
tenure	.141 (.014)	--- ---	.139 (.014)	--- ---
siblings[e]	.100 (.065)	.034 (.123)	.085 (.065)	.040 (.125)
father's education	-.053 (.039)	-.024 (.074)	-.053 (.039)	-.027 (.076)
mother's education	.051 (.047)	.004 (.089)	.035 (.047)	.004 (.090)
parents goal	-.046 (.033)	-.064 (.062)	-.045 (.033)	-.061 (.063)
mother worked when respondent aged 14	.526 (.224)	-.606 (.425)	.516 (.223)	-.585 (.428)
respondent lived with father and mother when 14	1.529 (.784)	-4.712 (1.481)	.1.679 (.788)	-4.867 (1.507)
hdis	--- ---	--- ---	1.222 (1.280)	-2.274 (2.455)
hagr	--- ---	--- ---	1.399 (1.292)	-2.084 (2.477)
agem	--- ---	--- ---	.138 (.057)	-.150 (.110)

dependent variable:	experience[c]	tenure[c]	experience[d]	tenure[d]
idkids	--- ---	--- ---	-.177 (.147)	-.106 (.282)
eduexp	--- ---	--- ---	.007 (.094)	.085 (.180)
egoal	--- ---	--- ---	.017 (.087)	-.036 (.166)

a.there are 1266 observations
b.single year age-dummies and controls for region of residence are included in all specifications
c.with family background variables as IVs
d.with family background variables and attitudinal/expectational variables as IVs
e.see endnote #10 about the construction of the IVs

Table 14
Wage Equation Estimates for White Working Women, First-Difference Estimation[a] on Basis of Wage Growth[a] Dependent Variable: levels and changes of natural log of hourly earnings

	1982 - 1980	**1991-1988**
Difference Coefficients	**(1)**	**(2)**
d Married, spouse present	-0.077 (0.047)	-0.053 (0.057)
d Divorced or separated	0.025 (0.039)	-0.006 (0.053)
d One child	-0.011 (0.042)	0.147 (0.111)
d Two+ children	0.025 (0.031)	-0.025 (0.058)
Early Wage[b]	-0.040 (.022)	-0.036 (.024)

a. There are 1,185 observations for column (1). There are 1047 observations for column (2). Only women, for whom data in the two consecutive difference-years and on the early wage are available, were included in this test, decreasing our sample size. Standard errors are reported in parentheses..

b. Wage as of 1978 used in column (1), and wage as of 1987 used in column (2).

Table 15

First Stage of Family Background Instrumental Variable Estimation of Wage Equation Estimates for White Working Women, Nine Year Difference Specification[a,b]

dependent variable:	**d experience**	**d tenure**
dmarried, spouse present	0.097	-1.068
	(.159)	(.611)
ddivorced or separated	.066	-1.373
	(.162)	(.62)
d one child	-.108	1.371
	(.261)	(1)
d twoplus children	-.433	.166
	(.144)	(.556)
d education	.086	-.945
	(.120)	(.640)
d experience	---	.641
	---	(.113)
d tenure	.043	---
	(.008)	---
siblings[c]	-.001	.045
	(.027)	(.104)
father's education	.011	.031
	(.019)	(.072)
mother's education	.039	.055
	(.383)	(.083)
parents goal	-.012	.024
	(.015)	(.058)
mother worked when respondent aged 14	.042 (.107)	-.357 (.412)
respondent lived with father and mother when 14	.272 (.384)	-5.289 (1.)

a. There are 1135 observations.

b. Controls for changes of region of residence are included in all specifications.

c. See endnote #10 about construction of the IVs

Table 16

Overidentification Equation[a], referring to Wage Equation Estimates for White Working Women, Nine Year Difference Specification

dependent variable: residuals	
number of siblings	.003 (.009)
father's education*[b]	.013 (.006)
mother's education	-.002 (.007)
parents goal*	.011 (.005)
mother worked when women was aged 14*	-.051 (.036)
women lived with father and mother when aged 14	.085 (.129)

a. There are 1135 observations.
b. * indicates the significance on a <5% level, ** on a approx. 15% level, and hence these instruments were maintained. Corresponding two stage least square regressions in table 11, column 3.

II. Regression Results for 1987

OLS Estimates for 1987 Data

Table 17
Wage Equation Estimates for White Working Women, 1987, Ordinary Least Squares[a]
Dependent Variable: Natural log of hourly earnings

	(1)	**(2)**
Married, spouse present	0.072 (0.057)	0.104 (0.048)
Divorced or separated	0.078 (0.060)	0.175 (0.051)

	(1)	(2)
One Child	-0.122 (0.047)	-0.027 (0.040)
Two+ Children	-0.334 (0.041)	-0.083 (0.036)
Education		0.072 (0.005)
Experience		0.027 (0.004)
Tenure		0.024 (0.002)
F-test[b]	0.0001	0.001
R^2	0.0696	0.3529

a. There are 1,424 observations. Observations are included only if the wage reported is for the current job. Never married and no children aged 33 are the reference categories and therefore contained in the intercept. Single-year age dummy variables and controls for region of residence are included in all specifications. Standard errors are reported in parentheses.
b. P-value for joint test of significance of marital status and fertility variables.

Pooled Data

Table 18
Wage Equation Estimates for White Working Women, Pooled Data 1982-1987[a]
Dependent Variable: natural log of hourly earnings

	(1) Unrestricted	(2) Restricted
Married, spouse present	0.073 (0.03)	0.51 (0.039)
Divorced or separated	0.135 (0.032)	0.106 (0.043)
One child	-0.029 (0.026)	-0.031 (0.034)
Two+ children	-0.066 (0.23)	-0.069 (0.032)

	(1) Unrestricted	(2) Restricted
Education	0.066 (0.003)	0.06 (0.004)
Experience	0.028 (0.003)	0.024 (0.004)
Tenure	0.026 (0.002)	0.029 (0.003)
d[b]	---	-0.018 (0.126)
d(Married, spouse present)	---	0.054 (0.061)
d(Divorced or separated)	---	0.069 (0.065)
d(One child)	---	0.004 (0.051)
d(Two+ children)	---	-0.014 (0.047)
d(Education)	---	0.012 (0.006)
d(Experience)	---	0.003 (0.005)
d(Tenure)	---	-0.005 (0.003)
F-test[c]	0.0001	0.3987
R^2	0.3885	0.3935

a. There are 3,122 observations. Standard errors are reported in parentheses.Controls for region of residence and single-year age dummy variables were included in all specifications for wage levels. We included 18 age dummies, of which there are 9 dummies for the age groups in the specific year.
b. We pooled 1987 with 1982 by includeing dummies which were set to equal one in the year 1987 and to zero in 19821. If these dummies turn out to be jointly different from zero, we find evidence for a structural change. Contrary to a Chow Test, we can specify where this break comes from by looking at the t-statistics of the specific dummy. This will allow us to conclude how the influence of certain variables evolves over time
c. P-value for joint test of significance of marital status and fertility variables for the unrestricted regression P-value for joint test of significance of all change dummies of the restricted regression.

Instrumental Variable Estimates for 1987

Table 19

Wage Equation Estimates for White Working Women, 1987, Two Stage Least Squares[a]
Dependent Variable: natural log of hourly earnings

	OLS[b]	Experience and Tenure Endogenous	
Coefficients	**(1)**	**(2)**	**(3)**
Married, spouse present	0.104 (0.048)	0.084 (0.058)	0.095 (0.050)
Divorced or separated	0.175 (0.051)	0.100 (0.069)	0.139 (0.057)
One child	-0.027 (0.040)	-0.062 (0.050)	-0.044 (0.043)
Two+ children	-0.083 (0.036)	-0.237 (0.090)	-0.158 (0.058)
Experience	0.027 (0.004)	0.003 (0.039)	0.015 (0.024)
Tenure	0.024 (0.002)	-0.018 (0.024)	0.004 (0.018)
F-test[c]	0.0013	0.0210	0.0037
Family background as IVS	No	Yes	Yes
Expectations/attitudes as IVS	No	No	Yes
Test for Overidentification[d]	---	0.7098	0.0101
Hausman Test[e]: Experience	---	0.385	0.270
: Tenure	---	3.012	1.205

a. There are 1,424 observations. Standard errors are reported in parentheses. Single-year age dummy variables and controls for region of residence were included in all specifications. Same family background and attitudinal/expectational IV's as in 1982 are used here.
b. The OLS regression is the same as in Table 12, column (2).
c. P-value for joint test of significance of marital and fertility status variables.
d. P-value of F-test for overidentifying restrictions.
e. Statistics of Hausman test (Chi-squared distribution).

Selectivity Correction 1987

Table 20

Wage Equation Estimates for White Working Women, 1987 Sample Selectivity-Corrected (SSC) Estimates [s]
Dependent Variable: natural log of hourly earnings

	OLS	SSC[b,d]
	(1)	(2)
Married, spouse present	0.084 (0.055)	0.038 (0.059)
Divorced or separated	0.121 (0.061)	0.055 (0.068)
One child	-0.039 (0.047)	-0.003 (0.050)
Two+ children	-0.203 (0.042)	-0.094 (0.067)
λ	---	-0.684 (0.326)
F-test[c]	0.0001	0.4948

a. There are 1,477 observations in Probit procedure and 1,250 observations in the OLS estimates. Standard errors are reported in parentheses. Single-year age dummy variables and controls for region of residence and education were included in all specifications.
b. Variables in the employment Probit include all variables in Table5, column (3), as well as dummy variables indicating whether a respondent changed marital or fertility states between 1985 and 87. Measures of husband's income, sum of income from alimony and child support (set to 0 for never-married women). As in Table 4, the variable of weeks husband spent unemployed in 1987 is excluded from the Probit procedure.
c. P-value for joint test of significance of marital status and fertility variables.
d. Log likelihood for normal is -612.83

First Difference Estimation 1987

Table 21

Wage Equation Estimates for White Working Women, First-Difference Specifications[a]
Dependent Variable: levels and changes of natural log of hourly earnings

	1987[b]	87-85 change[c]
Coefficients	**(1)**	**(2)**
Married, spouse present	0.028 (0.055)	-0.110 (0.052)
Divorced or separated	0.031 (0.060)	0.005 (0.049)

	1987[b]	87-85 change[c]
Coefficients	**(1)**	**(2)**
One child	-0.038 (0.044)	0.022 (0.064)
Two+ children	-0.242 (0.037)	0.031 (0.053)
F-test[e]	0.0001	0.3071

a. There are 1,090 observations. Standard errors are reported in parentheses.Other independent variables are controls for region of residence and controls for changes in education.
b. OLS regressions using same variables as Table 1, column (2).
c. Standard fixed-effects estimator.
d. Natural log of hourly earnings in 1983.
e. P-value for joint test of significance of marital status and fertility variables (column 1) and
P-value for joint test of significance of changes of marital status and fertility variables (column 2).

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Endnotes

1. Kim and Polachek (1994) discuss different sources of bias in their paper.

2. This is also pointed out by Dolton and Makepeace (1987): "... some aspects of a woman's labour force attachment, such as career aspirations, occupational choices, attitudes to work, commitments to particular jobs ... cannot be measured but are related to her child rearing plans. The presence of a child may be a good indicator of these unobservable characteristics."

3. Due to the virtual zero effect of children on wages in the 1991 OLS, we can conclude that it is not [unexpected] wage growth that influences fertility decisions.

4. When bias correcting the effects of marriage on wages, K&N found in 1994 that there is a positive marriage premium.

5. Experience and tenure might be correlated with the wage equation error. These acquired attributes (usually observable) might be correlated with innate characteristics (usually unobservables), contained in the error term.

6. Since a fraction of women in the no child category is married, these women are in a state of having no children yet. This means that the lower average of education and experience could be caused by these women who have just become married and are in their relative short time period of having no children.

7. This type is hard to detect in the data as the marriage category also contains a big fraction of women who are Less career oriented and who have more children. On the other hand, the one child category is also influenced by less career oriented women who are on their way on having two or more children.

8. Polachek and Kim (1994) develop individual-specific slope models to account for heterogeneity among women.

9. Over time we expect to find either an increased marriage premium (as package 2 choosers who marry when older have higher wages than package 3 choosers who had married earlier) and a higher standard error, as the dispersion of

wages of married women increases. In addition, the estimate of the one child coefficient is presumed to become less negative over time as the coefficient is more and more dominated by package 2 choosers. Moreover, we expect that the coefficient of two or more children will decrease by a small number, as children need less parental effort, these women can become more attached to the labour market. Finally, we expect that, although probably endogenous, labour market controls experience, tenure and education are good indicators of a woman's career package and of her heterogeneity. This should remove most of the significance of the marital status and fertility variables.

10. Reduction in years of experience by number of children, after controlling for differences in age: 1: -0.7 ; 2: -2.55 ; 3: -3.92 .

11. Small deviations in results may be caused by lightly different construction and exclusion restrictions of some variables. In certain cases, slightly different specifications are used in the regressions and are reported under each section. Example: experience/education is constructed such that no missing values are generated in order to maintain a higher sample size.

12. We use the same family background variables as K&N used. Family background variables used as IVS include: father's education; mother's education; parents' educational goal for respondent at age 14; number of siblings; a dummy variable equal to one if the respondents' mother worked when respondent was age 14; a dummy variable equal to 1 if the respondent lived with both a father and mother at age 14; and dummy variables corresponding to each of these variables, equal to 1 when the variable was missing (in which case the variables were set equal to 0).

However, we cannot construct the same expectational and attitudinal variables as K&N used. The expectational/attitudinal variables we used as IVS include: a dummy variable set equal to 1 if respondent disagreed or strongly disagreed with the statement that it is alright for a woman to work even if her husband disagrees, asked in 1972 (K&N used the same variable for 1971, however, we cannot find this question in the 1971 survey); and a dummy variable set equal to 1 if respondent agreed or strongly agreed with this statement, in 1971; ideal age at marriage reported by respondent in 1968 (set equal to 0, with a dummy variable set equal to 1 if response was never to marry); ideal number of children , in 1970 (K&N used expected number of children but we cannot find this question in 1970's survey); educational expectations, in 1970; educational goal, in 1970; and dummy variables corresponding to each of these variables, equal to 1 when the variable was missing (in which case the variables were set equal to 0).

13. although we know that education is probably endogenous we could not find adequate instruments to pursue this approach. Family background and attitudinal/expectational variables were rejected by various override tests.

14. In fact, the result for using both two groups of IVS should be different from the result in K&N's paper since some of the IVS we use for expectations and attitudes are slightly different.

15. Nakamura and Nakamura (1994) also find that women with more children become more attached to the labour market, probably for financial reasons.

16. We will again apply a specification test to see whether we can keep the first difference approach for further examination. Including the wage level of 1987 in the list of regressors, we find that this estimate is insignificant, and as the other coefficients virtually do not change, and are not important in the regression. We conclude that the first difference specification is not rejected in favour for a regression allowing for wage growth. See the appendix for regression table.

17. this specification was tested against regressions with inclusion of wage coefficients of 1978 and of 1987 and 1980. In both regression they obtained insignificant estimates. Therefore the use of level differences was confirmed.

18. From our earlier discussion we know that primarily women with higher education determine this coefficient because package 3 selectors, by assumption, had completed their education in 1982. Finding increasing returns to schooling in the 80's is consistent with the results by Blackburn et al. (1994) and Grogger and Eiger who offer, as a possible reason, the acquisition of more valuable degrees.

19. Again, we would need to instrument education. Having tried several sets of instruments, the validity of the IV's had always been rejected.

20. The first stage regression with difference in tenure as independent variable has the only interesting finding that differences in education are significantly and negatively correlated with differences in experience.

21. We also applied the instrumental variables estimation with both, family background and expectational variables. The first stage equations correspond our previous first stage in its interpretation. In the second stage, the validity of the instruments is rejected more strongly. Excluding the insignificant instruments, the F-value of the overidentification test increases even more.