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The Impact of Recent Changes in Family Assistance on Partnering and Women's Employment in New Zealand*

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This paper estimates the effects of recent changes to Family Assistance tax credits on the partnering and employment outcomes for New Zealand women. We use a difference-in-differences approach to control for economic and other confounding factors. Specifically, we investigate differences in partnering, employment and work hours over time across

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1. Introduction

Recently enacted changes in Family Assistance, by offering more generous benefits to families, are expected to reduce child poverty by one-third (MSD 2006). As side effects, the changes alter the relevant benefit of both partnering (marriage or cohabitation) and work for low- income families. The new rules offer expanded benefits which increase the number of families eligible, and reduced abatement rates which raise net wages and could affect labour supply. Since benefits are abated based on helps control for background factors when estimating policy impacts. We compare changes in partnering and work behaviour over time for demographic groups that differ in their eligibility or likely use of the programmes. We use three variations of this approach in identifying these policy effects. Firstly, families with children receive additional benefits that can be contrasted to childless families. Secondly, women with low education are more likely to qualify for these benefits than women with high education. Finally, women with low wages are more likely to be eligible for these programmes than those with high wages. We also estimate these impacts using the preferred child/no child identifier with the sample restricted to the low wage or low education subgroup.

2. Changes in Family Assistance

Aid to low-income families in New Zealand is largely administered through the tax system in the Family Assistance set of tax credits: family support, child tax credit, family tax credit and parental tax credit. Family support is not conditioned on work, but the remaining credits are only available to working families. The Working for Families programme altered a number of the dimensions of Family Assistance. We focus on changes to family support (renamed Family Tax Credit), a benefit available to income qualified families with children, and the child tax credit, a benefit available to families that have children and meet a work test. In April 2005, Family Support Rates increased by $25 per week for the first child and $15 per week for additional children. In April 2006, the child tax credit was replaced by the In-Work Payment, and amounts were raised for families qualifying by working at least 30 hours per week for a couple or 20 hours for a sole parent. The new payment provides $60 per week for up to three children, and then $15 for each additional child. For a family with one child, this amounts to an increase of $45 per week from $15 to $60, or an annual increase of $2,340. For families with two children, the increase is $30 per week from $30 to $60, or an increase of $1,560 annually. The abatement schedule was also changed, increasing the no abatement range to $35,000 (from $20,356), removing the 18 percent benefit reduction rate, and lowering the 30 percent rate to 20 percent for income above this threshold. This expands eligibility to more families by raising the breakeven level. In addition, the Family Tax Credit, a guaranteed income now called the Minimum Family Tax Credit, was raised from $15,080 to $17,680. Given the size of these changes beginning in 2005, we choose this year as the start date from which we might potentially see behavioural changes.

Other changes in policy have taken place over our study period. These may have affected employment for low-income women and thus should be

The expansion of benefits has made the partnering penalties larger. Consider the same example as above using the 2004 rules. The combined gross benefit for the custodial parent would have been $5,668 ($4,108 in family support, $1,560 in child tax credits), but abatement would have reduced this benefit to $3,662. If the pair were to partner, the combined income of $60,000 would have fully abated the benefit. Thus, this particular penalty for partnering was substantially lower in 2004 ($3,662) than in 2006 ($5,000).

A Family Assistance marriage bonus could occur in some circumstances. If a woman with children had less than twenty hours of work per week, she would not have qualified for the 2006 In-Work payment of $3,120 (annually) for up to three children. If she partnered with a man such that their total work hours exceeded 30 per week, they would have received the In-Work payment (although it would have been abated by the additional partner's income if joint income exceeded $35,000). In another example, if an unpartnered woman worked 20 hours per week but was ineligible for the In-Work payment because she received a part benefit from DPB, she could have an incentive to partner to get access to the In-Work payment even though she would forgo her part benefit. The sizes of the bonuses or penalties depend on incomes and number and ages of children.

Partnering incentives tell only part of the story. Couples decide to partner for many reasons. We argue that financial incentives can have an impact on propensity to partner for some, as is evident from the literature. As discussed below, other factors such as gains from economies of scale of the larger partnered family, labour supply adjustments, and income stability could also play roles in this decision. Increased income stability of sole parents may result in increased marriage rates in the future. The increase in family incomes of those on Family Assistance could potentially stabilize incomes and thus stabilize marriage (more on this possibility below). Our method is not able to tease apart these separate influences, and estimates only the total effect.

3. Background Literature

For labour supply, the impact of welfare and tax credit programmes in the US on work effort by low-income persons is well surveyed elsewhere (e.g., see Moffitt 1992, Hotz and Scholz 2003 and Moffitt 2003) and we discuss it only briefly here. One theme that emerges in this literature is that the work participation decision is more sensitive to policy than the hours of work decision (Meyer 2002). Our work focuses on employment outcomes as well as hours of work. The earned income tax credit, in particular, has been found to have a positive and substantial impact on labour participation of single mothers (Meyer and Rosenbaum 2001, Dickert, Houser et al. 1995). It also appears to have a modest negative effect on the hours worked by second earners in couples (Eissa and Hoynes 2004).

Two New Zealand studies have investigated the impact of the Working for Families Programme on work effort (Kalb and Scutella 2003 and Kalb, Cai et al. 2005). These studies employ a structural simulation methodology, predicting the impacts based on labour supply elasticities from tax changes in the 1990s. They predicted a small increase in labour force participation for sole parents by two percent over the next several years. Our methodology is quite different, using actual data from before and after these policy changes.

To gauge the potential impacts of partnering penalties from past literature, we can look at US studies of the Earned Income Tax Credit (EITC) which is also a tax credit for low-income working families. Since EITC benefits depend on combined family income, it also creates marriage penalties or bonuses in different situations. Single mothers with no earnings can reap a bonus by marrying a man with earnings so that the couple qualifies for the EITC. On the other hand, a single mother with moderate earnings who qualifies for the credit could lose the credit if she married a man with enough earnings to put the married couple beyond the earnings limit. The size of these penalties and bonuses have been documented by Dickert-Conlin and Houser (1998) and Holtzblatt and Rebelein (2000).

To help clarify this issue, consider a simple conceptual model. Suppose a single woman decides whether or not to partner by comparing her expected utility if married to her expected utility if she remains single. She chooses the higher valued option. The value of each option would depend on expected income, taking into account labour supply adjustments, leisure, government taxes and benefits, and tastes. The difference in utility between the married and unmarried states is generally modelled as a function of the incomes in the two states and demographic characteristics. A key issue is that couples are not usually observed in both the married and unmarried state and thus income differences must be predicted for the marital state not observed.

Within this framework, Eissa and Hoynes (1999) look at marriage by single mothers. They model incomes in each marital state based on current earnings and a tax/transfer function that calculates taxes and transfers including the EITC based on assumptions about household makeup after a split. They find small or nonexistent effects on family formation. Dickert-Conlin and Houser (2002) also model the impact of EITC using this approach. They use panel data on individuals to get an initial distribution of earnings, and then compute benefits over time allowing the benefit rules to change but holding the distribution of earnings fixed. They conclude that the EITC expansions of the 1990s had little impact on marriage. Overall, the EITC literature from the US suggests that marriage responses are likely to be either nonexistent or small in magnitude.

Studies of the U.S. experience are not directly applicable to NZ for several reasons. Firstly, the EITC benefits in the US are structured differently than Family Assistance. The EITC has an initial earnings subsidy component that increases net wages. It then abates away, but the EITC does not have hours-of-work thresholds like the NZ In-Work payment. Secondly, the social stigma of unwed motherhood and marriage customs vary between the countries. Cohabitation is less stigmatized in NZ and benefit rules explicitly allow for partner benefits as long as there is a marriage-like relationship. In our study, we use a broad definition of partnering that includes reported cohabitation and marriage. Furthermore, non-working low-income married couples in the US do not receive cash benefits apart from unemployment insurance (if applicable), but rather in-kind aid such as food stamps, housing assistance, energy assistance, and subsidised child care. The broader availability of cash aid in NZ complicates any comparison.

4. Empirical Models: Reduced Form Difference-in-Differences Approach

With the availability of annual cross sectional data on individuals from before and after the expansion of Family Assistance, we develop empirical models of the partnering and employment behaviour of women. We begin with a descriptive analysis of the trends in the proportions partnered and employed for various groups. We then jointly estimate the propensity to partner and the propensity to work. Joint estimation by bivariate probit allows for correlation in the unobservables across the equations and should improve precision of parameter estimates. Women with a high unobserved propensity to work (more productive in the market), for example, may be more sensitive to financial marriage incentives. This model is similar in form to that of Buffeteau and Echevin (2003). We estimate hours worked for those employed and the probability that a family’s hours of work will exceed the hours threshold for the In-Work payment.

Our model attempts to isolate the impact of the Family Assistance changes that took effect beginning in 2005. To control for other time varying factors in the economy that confound with changes in benefits and tax credits, our model includes the local unemployment rate, a measure of the strength of the labour market which varies by region over time. Annual time dummies or a linear time trend are also included to absorb other general trends in partnering and employment outcomes that are unrelated to policy changes.

To better control for non-policy influences we use a difference-in-differences approach. As one example, part of the variation in Family Assistance is due to the presence (and ages) of children because families without children are not eligible. The decisions of childless individuals to partner will be affected by other changes in the economic and social environment over time. Thus 4.4(u)-4rs 4 ecuent op.5(t o)-4(t o)-4.3.7(o)1.6(n)

$$m_{it} = \alpha_0 + \alpha_1 T_i + \alpha_2 R_t + \alpha_3 T_i R_t + \alpha_z z_{it} + \alpha_t t + \varepsilon_{it}$$

and her propensity to work be expressed as:

$$p_{it} = \beta_0 + \beta_1 T_i + \beta_2 R_t + \beta_3 T_i R_t + \beta_z z_{it} + \beta_t t + \eta_{it}$$

where T_i = 1 indicates that the woman is in the treatment group (with children) and T_i

common time effects, but identify the policy effects of the *changes* in behaviour between the two groups after the policy is implemented. We condition on background characteristics to make the treatment and controls more similar, conditional on z_{it}. These background characteristics can include child variables (like age of youngest child) as long as these effects are constant over time (in the absence of policy change) after conditioning on other variables (for further discussion on this point, see Hotz and Scholz (2003) or Moffitt and Willhelm (2000)).

Furthermore, we estimate the impact of the reforms on hours worked for the subsample of women who work. We estimate a linear regression of weekly hours of work on background characteristics including the time and treatment indicators and interactions. We estimate these hours equations separately for single and partnered women. In each case, we must control for selection into the working sample, as well as selection into the partnered or single sample. A bivariate selection model allows for separate but correlated treatment of the two selections, based on computing Heckman-type selection correction terms from the estimated bivariate probit coefficients (Ham, 1982). The double-selection model assumes a correlation between the regression error and the errors in the two selection equations, and results in the addition of two selection-correction terms in a two-step correction procedure. For the selected sample of those employed and partnered, we estimate the following:

$$E\left(Hours_{it} \middle| x_{it}, P_{it} = 1, M_{it} = 1\right) = \gamma_x x_{it} + \lambda_1 s_1 + \lambda_2 s_2$$

where s_1 and s_2 are the selection correction terms from Ham, which are computed using output from the bivariate probit estimation. We calculate robust standard errors to allow for possible heteroskedasticity from the selection model.[2]

Finally, as noted earlier, the In-Work payment sets up a jump in benefits if the family meets the work hours threshold of at least 20 hours per week if unpartnered, and 30 hours per week (combined) if partnered. To investa 7 TD-0.0017 Tc0.02(e p)4.6h-0.0017tesh-0.0017f acorr selection due to partner status. That is, we estimate the probability of exceeding the hours threshold conditional on partner status, and partner status is estimated jointly as a selection equation. This allows for correlation in unobservables between the equations, with identification coming from additional age-related variables on children in the partnering probit.[3]

including both cohabitation and possibly some married women living apart from their spouses.

The second dependent variable is employment. We define a woman as employed if her usual hours worked per week in the survey period are positive. We use usual hours of work both as the dependent variable in the hours regressions and to define a binary variable for whether the In-Work payment hours threshold was exceeded.

Part of the analysis requires forming groups based on predicted wages. We predicted wages using a selection-corrected regression, with selection into employment allowing for different coefficients between partnered and unpartnered women. The models used covariates of age, age squared, educational qualification indicators, ethnicity indicators, region of residence, regional unemployment rate, and year indicators. The probit model for probability of selection additionally included several variables for number of children of various ages to aid in identification. A wage was predicted for each woman in the sample, including an adjustment for sample selection. For the employed women we estimated:

$$E\left(Wage_{it} \middle| x_{it}, P_{it} = 1\right) = \beta_x x_{it} + E\left(\varepsilon_{it} \middle| x_{it}, P_{it} = 1\right)$$

where the last term is a Heckman selection term (Mills ratio) for employment. For those not employed we estimated:

$$E\left(Wage_{it} \middle| x_{it}, P_{it} = 0\right) = \beta_x x_{it} + E\left(\varepsilon_{it} \middle| x_{it}, P_{it} = 0\right)$$

using a different Heckman selection

Table 2: Difference-in-Differences Estimates of WFF Impact on Partnering

A. Probability of Partnering Probit: Column labels are group identifiers

Variable	**Has Child**	**Has Child**	**Low Education**	**Low Education**	**Low Wage**	**Low Wage**
Year	-0.001	-0.003***	0.002	0.000	-0.005***	-0.007***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
D05	-0.000	0.007	0.025**	0.016	0.003	0.023***
	(0.007)	(0.007)	(0.010)	(0.011)	(0.008)	(0.007)
Identifier	0.193***	0.114***	-0.049***	-0.081***	-0.294***	-0.357***
	(0.004)	(0.004)	(0.006)	(0.007)	(0.005)	(0.008)
Identifier • D05	**0.007**	**0.003**	**-0.042*****	**-0.025****	**0.025****	**0.014**
	(0.008)	**(0.007)**	**(0.011)**	**(0.012)**	**(0.010)**	**(0.010)**
Has Covariates?	No	Yes	No	Yes	No	Yes
Log Likelihood	-45,274.4	-42,427.2	-18,199.2	-16,534.5	-28,206.3	-25,907.8
Sample Size	81,627	81,627	30,236	30,236	54,415	54,415

Notes: Merged HLFS-IS data 1997-2007 for women aged 22 to 50 weighted by Statistics New Zealand sample weights. Covariates include age, age-squared, ethnicity indicators, region indicators, age of youngest child. The ‘Has Child’ specification includes education qualification indicators. Coefficients are partial derivatives of probabilities, and standard errors are included in parentheses. The reported results on the interaction terms (Identifier • D05) are corrected difference-in-difference probabilities as

C. Probability of Employment Probit ─ Unpartnered: Column labels are group identifiers

Variable	Has Child	Has Child	Low Education	Low Education	Low Wage	Low Wage
Year	0.007*** (0.002)	0.002 .002)				

6.1 A. Presence of Children as a Policy Identifier Partnering

Figure A1 in the Appendix plots the proportion of women partnered between 1997 and 2007. There appears to be a slight dip in this propensity after 2005. When we plot this proportion separately for women with and without children, we can see some differences. While childless women show a slight dip, those with children show a slight rise in the proportion partnered. This suggests that the Family Assistance expansion did not reduce partnering in absolute levels.

To test this hypothesis, we ran a probit model for partnering, which included a time trend, an indicator for presence of children, an indicator for the policy change years 2005, 2006 and 2007 (D05), and the interaction between presence of children and this policy dummy. The interaction term tells the tale. These results are shown in the first two columns in Panel A of Table 2.

The first column of Table 2 shows that the partnering change is not statistically significant at conventional levels. The second column reports on a specification that adds several background covariates (estimated coefficients not shown). This makes the groups more comparable and controls for potential changes in partnering caused by modifications in the composition of the samples over time. The second column results condition on qualifications (and other things) so that any change in education will not drive the partnering results. The results in this second column tell the same story as the first. We conclude that adding additional covariates does not matter and the policy does not appear to have influenced partnering over the period from 2005 to 2007.

6.2 Employment and Hours of Work

Figures A2 and A3 in the Appendix display the separate time trends for the employment of partnered and unpartnered women. For both partnered and unpartnered women, employment has been rising primarily among women with children. The employment propensities for women without children have been relatively stable over this time perio12rih9vely-on qualifsbe

We ran a joint model of marriage and employment that allows correlation of the error terms for the two equations shown in Section 4. The advantage of the joint model is that unobservable influences likely correlate across the two dependent variables and the bivariate probit uses that correlation to improve the precision of the estimated coefficients. All models include the background characteristics listed in the notes at the bottom of each table. Results from the full model are shown in Panel A of Table A1 in the Appendix. The estimated correlation in unobservables between the partnering and employment equations is 0.234 and statistically significant at better than a 1% level. This positive correlation indicates that women who have unobservable traits making them more likely to be employed are also women with unobservables that make them more likely to be partnered. This argues against the Becker notion that

Notes: Models also include age, age squared, ethnicity indicators, region indicators, education qualifications, unemployment rate and two selection correction terms. Robust Standard Errors are reported. Bivariate probit estimation on partnering and employment was used to produce Heckman correction terms for significantly different from zero at a 1% level for partnered women and only for this specification. We discount the result because it flows from a *rise* in partnering among more educated, with no obvious explanation since they are not affected by the policy, rather than a fall for the less educated, and because low education is clearly a coarse proxy for eligibility.

Turning to employment propensities, we see quite different policy effects for partnered and unpartnered women. These differences highlight the advantages of estimating employment responses for the two groups. The employment trends in Figures A2 and A3 show an increase in employment for the less educated women relative to the more educated women after 2005. These effects are statistically significant in both sets of probit results in Panel B of Table 2. Using our preferred specification that

Table 4: Hours Regressions: Low and High Education Interaction Coefficients for Difference-in-Differences

A. Hours of Work for Employed Women: Selection Corrected Regression

	Partnered Women		Unpartnered Women	
Variable	**Hours Coefficients**	**Standard Errors**	**Hours Coefficients**	**Standard Errors**
Year	-0.160*	0.083	-0.109	0.125
D05	-1.858***	0.466	0.576	0.697
Low Education	-0.459	0.580	-5.970***	0.670
D05 • Low Education	**3.129*****	**0.524**	**1.920****	**0.783**
Root MSE	14.04		13.39	
R Squared	0.094		0.123	
F for all Zero Coefficients	58.06		32.44	
Sample Size	13,687		5,639	

Notes: Models also include age, age squared, ethnicity indicators, region indicators, unemployment rate and two selection correction terms. Robust Standard Errors are reported. Bivariate probit estimation on partnering and employment was used to produce Heckman correction terms for these regressions with additional age of children variables as identifiers. These bivariate probit results are not reported.

B. Probits for Hours Thresholds: Selection Corrected for Partner Status

	Partnered Women 30 or More Combined Work Hours for couple		Unpartnered Women 20 or More Work Hours	
Variable	**Coefficients**	**Standard Errors**	**Coefficients**	**Standard Errors**
Year	-0.004	0.009	0.011	0.008
D05	0.089	0.057	0.096*	0.060
Low Education	-0.507***	0.031	-0.869***	0.039
D05 • Low Education	**0.001**	**0.015**	**0.012**	**0.017**
Rho	0.649	0.077	-0.851	0.025
Chi Squared	1,363.78		1,621.34	
Sample Size (uncensored)	20,249		9,987	

Notes: Models also include age, age squared, ethnicity indicators, region indicators, education qualifications, and unemployment rate (for employment probit). Selection on partnering model adds number of children less than age 6, number of children 6-12 and number of children 13-15 to the regression. The reported results on the interaction terms (D05 • Low Education) are corrected difference-in-difference probabilities as derived in Norton et al. (2004) evaluated at the mean value of regressors, and the standard errors are computed using the appropriate delta method (Corneliβen and Sonderhof, 2008).

Significance levels:

Table 5: Hours Regressions: Low and High Wages
Interaction Coefficients for Difference-in-Differences

A. Hours of Work for Employed Women: Selection Corrected Regression

	Partnered Women		Unpartnered Women	
Variable	**Hours Coefficients**	**Standard Errors**	**Hours Coefficients**	**Standard Errors**
Year	0.098*	0.059	-0.232**	0.092
D05	-1.791***	0.316	-0.064	0.614
Low Wage	5.831***	0.393	-0.635	0.827
D05 • Low Wage	**1.529***	**0.407**	**1.227***	**0.638**
Root MSE	13.755		13.402	
R Squared	0.114		0.089	
F for all Zero Coefficients	115.17		35.66	
Sample Size	26,456		10,125	

Notes: Models also include age, age squared, ethnicity indicators, region indicators, education indicators, unemployment rate and two selection correction terms. Robust Standard Errors are reported. Bivariate probit estimation on partnering and employment was used to produce Heckman correction terms for these regressions with additional age of children variables as identifiers. These bivariate probit results are not reported.

B. Probits for Hours Thresholds: Selection Corrected for Partner Status

	Partnered Women 30 or More Combined Work Hours for couple		Unpartnered Women 20 or More Work Hours	
Variable	**Coefficients**	**Standard Errors**	**Coefficients**	**Standard Errors**
Year	-0.007	0.007	-0.012*	0.007
D05	0.063	0.042	0.111**	0.053
Low Wage	-0.277***	0.081	-0.551***	0.060
D05 • Low Wage	**0.089***	**0.051**	**-0.041**	**0.055**
Rho	0.379	0.107	-0.776	0.028
Chi Squared	2,272.78		2,351.18	
Sample Size (uncensored)	37,937		16,478	

Notes: Models also include age, age squared, ethnicity indicators, region indicators, education qualifications and unemployment rate (for employment probit). Selection on partnering model adds number of children less than age 6, number of children 6-12 and number of children 13-15 to the regression The reported results on the interaction terms (D05 • Low Wage) are corrected difference-in-difference probabilities as derived in Norton et al. (2004) evaluated at the mean value of regressors, and the standard errors are computed using the appropriate delta method (Corneliβen and Sonderhof, 2008).

Significance levels: *** 1%, ** 5% and * 10%

6.5 D. Additional Specification Tests

To sharpen contrasts further, we also explored the use of children/no

B. Probits for Hours Thresholds: Selection Corrected for Partner Status

	Partnered Women 30 or More Combined Work Hours for couple	Unpartnered Women 20 or More Work Hours

smaller in some specifications. We conclude that there is no qualitative difference in the results. In addition, we re-estimated the initial education group and wage group models excluding all children variables as potentially endogenous. Again, the results did not qualitatively change.

7. Discussion and Future Directions

We conducted difference-in-differences analyses using three alternative identifying assumptions. The results from our different assumptions gave somewhat mixed results. These results are summarised in Table 8. We favour the comparison of women with and without children as the best policy identifier. For partnering, we found no significant policy effects from changes in Family Assistance for the groups with children or facing low predicted wages. For the low education group there was a possible partnering effect in the hypothesised direction, but we discount it for reasons given above.

Employment rose significantly for women with children relative to childless women. This effect was somewhat larger among unpartnered women, so the policy may be producing a larger employment response among sole mothers. When low education and low predicted wages were used as policy indentifies, the employment responses appear to be concentrated among partnered women. We consistently found positive effects on hours worked among those employed in almost all specifications. In addition, the proportion of couples with combined hours of work that exceed the In-Work payment threshold of 30 hours per week increased in the policy period for those with children. A similar effect is found for unpartnered women with children whose hours of work exceeded the 20-hour threshold for the In-Work payment. There is little evidence of these hours threshold effects in other specifications. Overall, our analysis finds little evidence to suggest that these policies influenced partnering, but we do find evidence of positive effects on employment and hours of work. Yet, several caveats should be borne in mind[5].

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Table 8: Summary of Interaction Effects for Difference-in-Differences Estimation

				For Employed Women			
		Probability of Employment		Hours of Work		Hours Thresholds	
Policy Identifiers	**Probability of Partnering**	**Partnered**	**Unpartnered**	**Partnered**	**Unpartnered**	**Partnered 30 or more**	**Unpartnered 20 or More**
Has Child	0.003	0.016*	0.044***	1.002***	1.001***	0.035***	0.028**
Low Education	-0.025**	0.041**	0.008	3.129***	1.920**	0.001	0.012
Low Wage	0.014	0.025**	-0.022	1.529***	1.227*	0.089*	-0.041
Low Wage/Has Child	---	---	---	-0.219	0.407	0.092	0.036
Low Education/Has Child	---	---	---	2.457***	3.511**	0.038	0.092

Notes: Results reported in the first three columns are taken from Table 2 and pertain to our preferred regrenf358.4as

Firstly, we have a relatively short period over which to observe any behavioural adjustments to these policy changes. This is particularly true for partnering decisions. Families have to learn about the new rules and learn about how they would be personally affected. Families in our data have not had long to react to the Family Assistance changes. We might expect larger impacts over a longer time horizon, consistent with evidence of Family Assistance on labour force participation and unemployment. Since it could be argued that these policies were designed partly to encourage unemployed women into work, this issue might merit further analysis. Another avenue would be to move beyond the descriptive difference-in-differences approach, and develop a joint structural estimation of discrete labour supply and partnering status wherein a woman chooses both her partner status and her work hours, based on income expectations that depend on tax and benefit policies. Such a model could estimate responses to changes in benefit amounts.

At this point, subject to the caveats above, we provide some evidence of employment increases and more solid evidence of work hours increases for those working due to the Family Assistance policy changes. Evidence on partnering is more elusive but there are certainly no large impacts currently.

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Appendix – Table A1
Bivariate Probits and Hours Regression: With and Without Children
Full Set of Coefficients

A. Bivariate Probit Coefficients for Partnered and Employed

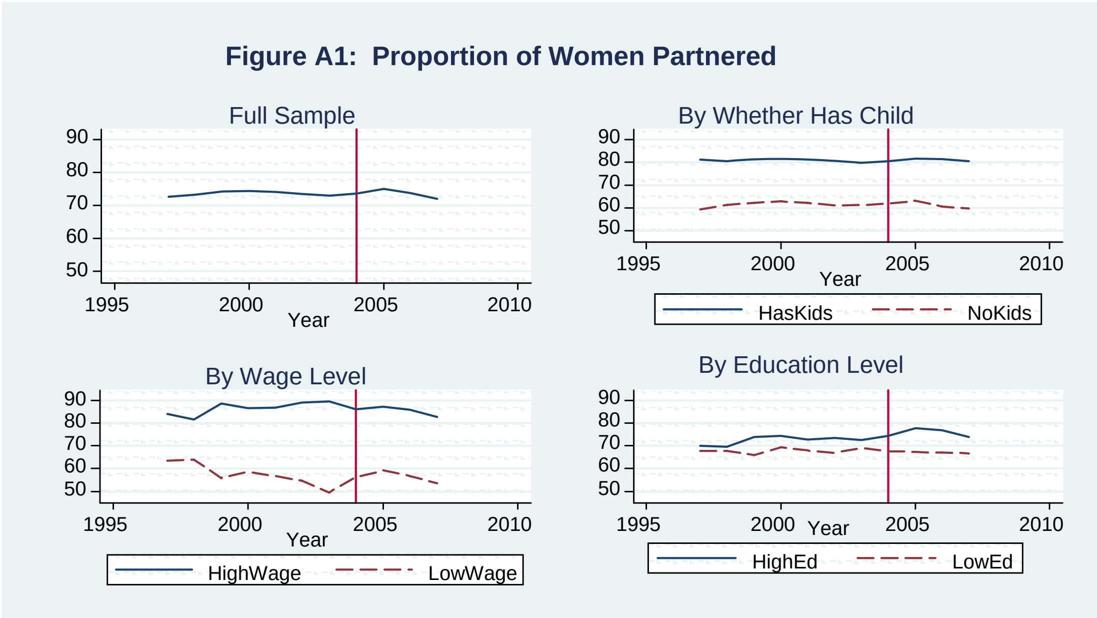
	Partnered Equation		Employment Equation	
Variable	Coefficients	Standard Errors	Coefficients	Standard Errors
Year	-0.007***	0.002	0.002	0.003
D05	0.020	0.022	-0.013	0.024
Has kids	0.359***	0.013	-0.279***	0.014
D05 • Has Kids	0.012	0.022	0.068***	0.024
Age	0.104***	0.006	0.067***	0.007
Age Squared	-0.001***	0.000	-0.001**	0.000
Maori	-0.499***	0.014	-0.296***	0.014
Pacific Island	-0.178***	0.021	-0.219***	0.021
Asian	0.317***	0.027	-0.513***	0.024
Other	0.150***	0.027	-0.599***	0.025
None	0.023	0.160	-0.336***	0.153
Num. Kids < age 6	0.275***	0.009	-0.447***	0.008
Region 2	-0.085***	0.026	-0.038	0.029
Region 3	0.003	0.028	-0.005	0.031
Region 4	0.014	0.030	0.051*	0.031
Region 5	-0.021	0.029	-0.006	0.032
Region 6	-0.016	0.032	0.025	0.036
Region 7	-0.043	0.030	-0.049	0.033
Region 8	-0.073***	0.027	0.050	0.032
Region 9	0.086***	0.030	-0.005	0.036
Region 10	-0.045*	0.027	0.031	0.031
Region 11	0.012	0.031	0.040	0.034
Region 12	0.122***	0.033	0.002	0.039
Primary/School Cert	0.237***	0.017	0.377***	0.017
Sixth Form/Bursary	0.312***	0.017	0.410***	0.017

Vocational	0.186***	0.013	0.532***	0.013
Bachelors	0.309***	0.019	0.704***	0.019
Post Grad Degree	0.195***	0.027	0.801***	0.029
Unemployment Rate	---	---	-0.028***	0.005
Constant	12.281***	4.739	-5.233	6.711

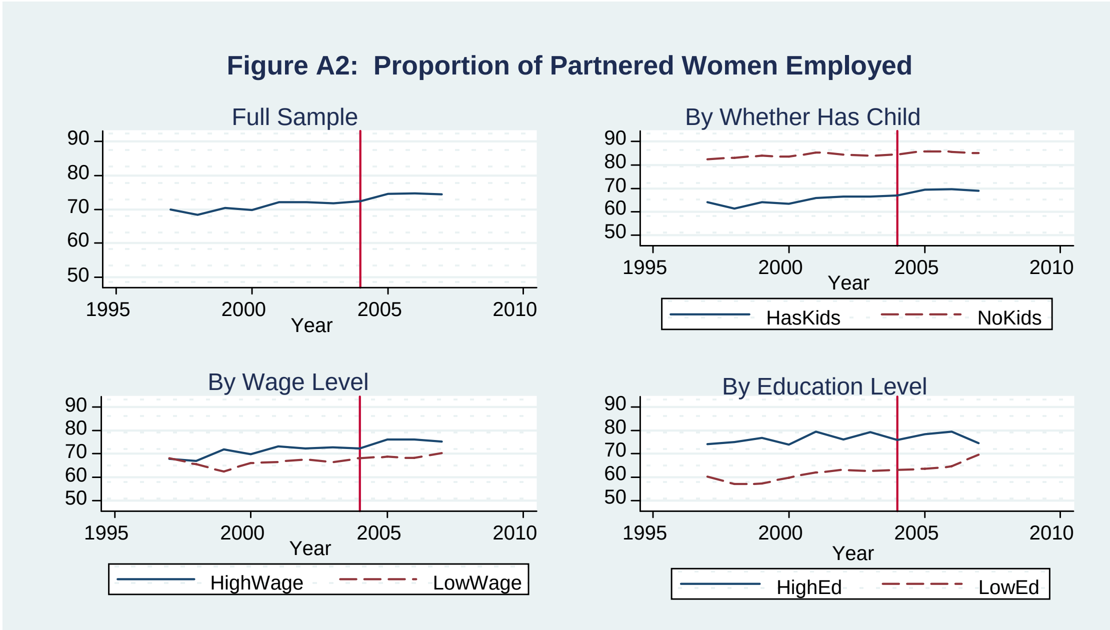
Region 5	-0.315	0.504	0.952	0.800
Region 6	-0.225	0.552	-0.650	0.871
Region 7	0.062	0.505	-0.733	0.829
Region 8	0.904*	0.480	1.663**	0.764
Region 9	-0.321	0.547	0.550	0.883
Region 10	-1.470***	0.479	0.302	0.762
Region 11	-1.110**	0.507	-1.801**	0.806
Region 12	-0.456	0.590	0.478	0.993
Primary/School Cert	-0.452	0.263	1.949***	0.417
Sixth Form/Bursary	0.175	0.278	1.757***	0.419
Vocational	-0.346	0.237	2.687***	0.357
Bachelors	1.172***	0.366	5.856***	0.506
Post Grad Degree	0.234	0.501	6.899***	0.889
Unemployment Rate	0.050	0.075	-0.030	0.118
Lambda Partnered	0.105***	0.007	-0.003	0.008
Lambda Employed	-0.039***	0.005	0.104***	0.027
Constant	32.759	93.349	23.383	147.717
Root MSE	13.625		13.061	
R Squared	0.113		0.134	
F for all Zero Coefficients	180.72		77.16	
Sample Size	40,596		15,207	

Notes: Models also include age, age squared, ethnicity indicators, region indicators, education qualifications, unemployment rate, and two selection correction terms. Robust Standard Errors are reported.

Significance levels: *** 1%, ** 5% and * 10%



Notes: Merged HLFS-IS data 1997-2007 for women aged 22 to 50 weighted by Statistics New Zealand sample weights.



Notes: Merged HLFS-IS data 1997-2007 for women aged 22 to 50 weighted by Statistics New Zealand sample weights.