# Estimating the labour supply response to the earned income tax credit for single parents in the Netherlands

CPB Achtergronddocument,

behorend bij CPB Policy Brief 2011/14 'Hoe prikkelbaar zijn Nederlanders'

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22 december 2011

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December 22, 2011

**Abstract**

In 2002 the Dutch government extended the eligibility conditions of the earned income tax credit for single parents to single parents with a child 12 to 15 years old. We use this natural experiment to study the labour supply response of single parents. We find that the policy reform increased the participation rate of single parents with a youngest child 12 to 15 years old by 1.7–2.0 percentage points, and increased their average number of hours worked by 0.8 hours per week. The implicit wage elasticity of participation is 0.4–0.5, and the implicit wage elasticity of total hours worked is 0.6.

**JEL codes**: C21, H24, J22
**Keywords**: Policy evaluation, difference-in-differences, regression-discontinuity, labour participation, hours worked, single parents

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# 1 Introduction

Up to 2001 working single parents received a tax credit, the *Aanvullende Alleenstaande Ouderkorting* ('Additional Credit for Single Parents'), if the youngest child was younger than 12 years old. In 2002 this age limit was raised to 16 years old, to stimulate the labour participation of single parents (Ministry of Finance, 2001). The current government has plans to reverse the policy change of 2002, to reduce the budget deficit (CPB, 2010), but for the moment these plans are on hold.

We use the change in the age limit as a 'natural experiment' to evaluate the impact of the tax credit on the labour supply of single parents. The focus is on a difference-in-differences (DD) approach to estimate the effect on the participation rate and on hours worked.[1] To the best of our knowledge, our analysis is the first to use a DD approach to study the effect of a change in a tax credit on participation and hours worked in the Netherlands.[2,3] Single parents are an interesting group to study because their number is on the rise, from 360,000 in 1995 to a projected 500,000 in 2011 in the Netherlands[4], and because they receive special attention by policymakers as evidenced by the various specific subsidies and tax credits targeted at single parents.

We expect the policy reform to increase the participation rate of single parents since the net income of the employed increases relative to the net income of the non-employed. We also expect a positive impact on hours worked of single parents. Not only does net income from working increase, but also net income from working more hours. For the major part of working single parents, the tax credit is increasing with income, and we expect the positive substitution effect to dominate the smaller negative income effect.

The estimation results are in line with expectations. In our preferred DD estimate the policy reform increased the participation rate of single parents with a child 12 to 15 years old by 1.7 percentage points. As a sensitivity check we also consider the effect using a regression-discontinuity approach, and obtain a value of 2.0 percentage points. Our preferred DD estimate for the effect on the average number of hours worked is 0.8 hours per week. The implicit wage elasticity of participation is 0.4–0.5, and the implicit wage

[1] The application of DD estimators in labour supply analyses is reviewed in Blundell and MaCurdy (1999), Blundell and Dias (2009), Imbens and Woolridge (2009) and Meghir and Phillips (2010).

[2] Euwals (2008) uses a DD setup to study the overall impact of the 2001 tax reform on the participation rate of married women, and uses single women as a control group. However, he does not include controls, nor does he check for placebo reforms. This is unfortunate as the treatment and control group are presumably quite different in many relevant ways.

[3] For an overview of labour supply responses in the Netherlands using structural models see Mastrogiacomo et al. (2011).

[4] See Figure 4 in the Appendix.

elasticity of total hours worked is 0.6. Part of the response in hours is due to an increase in hours worked by individuals already working.

Our results are broadly in line with the findings of related studies on single parents using DD abroad. There is an extensive literature on the impact of the *Earned Income Tax Credit* (EITC) in the US, introduced in 1975. This literature is reviewed in Hotz and Scholz (2003). They conclude that the EITC appears to have had substantial, positive effects on the participation of single parents. Eissa and Liebman (1996) is one of the earlier studies that has applied the DD methodology to labour supply responses. They estimate the impact of the EITC-expansion in 1987, combined with other elements of the Tax Reform Act of 1986, by comparing the change in labour supply of single women with children to the change of single women without children.[5] They find that the EITC-expansion explains an increase in participation by 2.8 percentage points (from a base of 74 percentage points). According to Meyer and Rosenbaum (2001), income taxes of working, single mothers fell by 491$ over the period 1984-1988, while their earnings were 17,250$ in 1984.[6] This implies an elasticity of participation with respect to net-of-tax income of $3.8/2.8 = 1.3$. They estimate that the reform had no significant effect on the working hours of single mothers who were already working.[7] Meyer and Rosenbaum (2001) examine the effects of changes in both welfare and tax policies during the 1984-1996 period on the labour supply of single mothers. DD estimates suggest that the policy changes over this period have raised the participation of single mothers by 7.1 percentage points relative to single women without children. Supplemented by estimates of a structural model, they show that changes of the EITC explain 62% of the total increase in participation. Tax reforms, mostly consisting of EITC changes, have increased net-of-tax earnings by around 1000$ for single mothers, who earned 18165$, on average. As a result, participation is raised by $0.62 \cdot 7.1 = 4.4$ percentage points, starting from a level of 75.9%. This yields an elasticity of participation with respect to net income of $5.8/5.5 = 1.1$.

Several studies have examined the introduction of the *Working Families' Tax Credit* (WFTC) in the UK, introduced in 1999. Brewer and Browne (2006) compare the findings of Gregg and Harkness (2003), Francesconi and van der Klaauw (2004), Leigh (2005) and Blundell et al. (2005). The increase in the participation rate varies from an insignificant 0.6 percentage points in Leigh (2005) to 7 percentage points in Francesconi and van der Klaauw (2004). Blundell et al. (2005) report an impact of 3.6 percentage points for single

[5]The use of this control group is being criticized by Blundell and MaCurdy (1999) and Meghir and Phillips (2010).

[6]See Appendix 2 in Meyer and Rosenbaum (2001); amounts are expressed in 1996 dollars.

[7]Eissa and Hoynes (2004) estimate the impact of the EITC-expansion in 1993 on the participation rate of married couples.

mothers and 4.6 percentage points for single fathers. According to Brewer et al. (2006), WFTC raised mean net incomes of working lone parents by 16.24 pound a week from a base of 250.22 pound. An increase in the participation rate by 3.6 percentage points, starting from a base of 47 percentage points, implies an elasticity of 7.7/6.5 = 1.2. Estimates of the increase in hours worked by single mothers that were already working range from 1.2 hours per week in Gregg and Harkness (2003) to (an insignificant) 1.7 hours per week in Leigh (2005).[8]

Stancanelli (2008) studies the impact of the *Prime Pour l'Emploi* ('Work Premium') in France, introduced in 1991. She finds no significant effect of the reform on single mothers when compared to single women without children.[9] She attributes the insignificant results partly to the poor specification of the control group as, in particular, the average participation rate of the control group is much higher than the treatment group.

The paper is structured as follows. Section 2 discusses the policy reform that we use in our empirical strategy. Section 3 outlines the estimation methods. Section 4 describes the data used in the analysis, and gives some descriptive statistics for the treatment and control groups. Section 5 presents the estimation results and various robustness checks. Section 6 concludes.

## 2 The natural experiment

Working single parents in the Netherlands receive an additional tax credit, the *Aanvullende Alleenstaande Ouderkorting* ('additional credit for single parents'), next to the general tax credit for all working individuals, the *Arbeidskorting* ('working credit'). Up to 2001, only single parents with a (dependent) youngest child younger than 12 years old[10] received this additional tax credit. In 2002 this age limit was raised to 16 years old, to promote labour participation of single parents (Ministry of Finance, 2001).

The tax credit for working single parents is income dependent and amounts to 4.3% of gross income up to a maximum credit of 1,301 euro in 2002, see Figure 1.[11] As can be

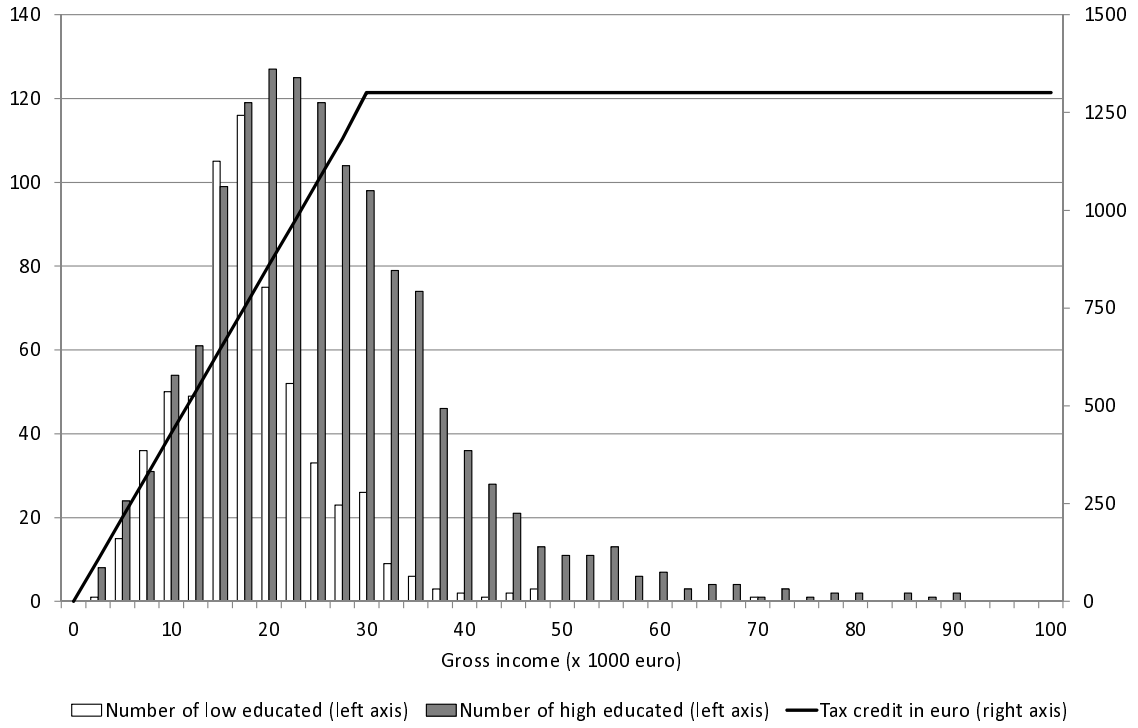
[8] Brewer et al. (2006) use a structural model to evaluate the WFTC. Model simulations yield results comparable to the DD studies. Simulation results suggest that due to the WFTC the participation rate of single mothers increased by 0.6 percentage points in 2000 and 3.7 percentage points in 2002. The simulation results further suggest that the WFTC increased average hours worked by 0.4 hours per week in 2000 and 0.8 hours per week in 2002.

[9] Furthermore, Stancanelli (2008) finds that the participation rate of married women fell by 3 percentage points due to the tax credit, whereas the participation rate for cohabiting women increased by 6 percentage points.

[10] On the 1st of January.

[11] The annual increase in the tax credit is linked to the growth of average gross wages, rising to 1401

Figure 1: Tax credit and distribution of working parents in treatment group in 2002




seen, the phase-in is up to a gross income of about 30,000 euro, which is about twice the minimum wage. The credit is not phased out.

Figure 1 also shows the distribution of working single parents with a youngest child aged 12 to 15 years old, the treatment group. We show the distribution of low and high educated single parents in this group. Most of the low educated single parents are in the phase-in range of the credit, and the majority of high educated single parents is also in the phase-in range of the credit. So, for most working single parents the tax credit increases the marginal net income from working more hours.

# 3 Empirical methodology

We use a difference-in-differences (DD) approach to estimate the impact of the modified tax credit (see *e.g.* Imbens and Woolridge (2009) and Blundell and Dias (2009)). This approach identifies the labour supply response of a tax reform by combining two types of comparisons. First, we consider the difference in labour supply before and after the reform

euro in 2005.

for the treatment group (the first difference). Second, we subtract the change in labour supply before and after the reform for the control group (the second difference). By taking a double difference, the approach removes common time effects and time-invariant group effects.

The treatment group in our study consists of single parents with a youngest child aged 12 to 15 years old. The control group consists of single parents with a youngest child aged 16 to 21 years old. We motivate this choice below.

Below we outline the methodology for estimating the effect of the policy reform on the participation rate and average number of hours worked of single parents in the treatment group.

### 3.1 Participation

To explain the participation rate we estimate a linear probablity model (see *e.g.* Angrist and Pischke (2009))

$$P(part_{it} = 1) = \beta_0 + \beta_y year_t + \beta_g group_{it} + \delta treatment_{it} + \beta_X X_{it} + \epsilon_{it} \qquad (1)$$

using ordinary least squares. The outcome variable equals 1 if the parent works and equals 0 otherwise. The vector *year* consists of year dummies, the dummy variable *group* controls for a constant difference between the treatment and control group, the *treatment* variable equals one if the youngest child is aged 12 to 15 years old and the year is in the interval [2002,2005]. We are primarily interested in the treatment coefficient $\delta$. We also add a number of additional controls $X$ to control for differences in age, education, gender and ethnicity between groups and over time (see the data section below). Since we have panel data, we also run the model with fixed effects, to control for unobserved characteristics. We follow the specification in Blundell and MaCurdy (1999), where we add a fixed effect next to a group dummy. The number of single parents is on the rise. This may affect the composition of the treatment and/or control group not corrected for by observable characteristics. Below we will see that it is important to control for unobserved characteristics.

As a robustness check we also consider the treatment effect using a regression-discontinuity (RD) approach, see e.g. van der Klaauw (2008) and Imbens and Woolridge (2009). A RD approach identifies the average effect of the tax credit based on the cut-off age that determines the entitlement to the credit. The idea is that in the absence of the tax credit, participation and working hours are a smooth function (in our case a second or third order polynomial) in the age of the youngest child, and the targeted tax credit introduces a discontinuity in this relation. We estimate the following linear probability model for

participation using the RD approach

$$P(part_{it} = 1) = \beta_0 + \beta_y year_t + \beta_{a1}(ageyoungestchild)_{it} + \beta_{a2}(ageyoungestchild)^2_{it} + \delta treatmentRD_{it} + \beta_X X_{it} + \epsilon_{it}, \quad (2)$$

where $treatmentRD = 1$ if the parent is entitled to the tax credit. Notice that the RD approach estimates another effect than the DD approach. The RD approach estimates the effect on all single parents that receive the tax credit, not just single parents with a youngest child 12 to 15 years old as in the DD approach.

### 3.2 Monthly number of hours worked

We are also interested in the effect on the number of hours worked. Following Angrist and Pischke (2009), we estimate the effect on the average number of hours worked using ordinary least squares. Again, we also estimate a fixed effects model. In all specifications we measure the effect on the average number of hours worked for all single parents in the treatment group, hence including the zeros for the non-employed. The specification for hours is

$$h_{it} = \beta_0 + \beta_y year_t + \beta_g group_{it} + \delta treatment_{it} + \beta_X X_{it} + \epsilon_{it}, \quad (3)$$

where $h_{it}$ is the monthly number of hours worked by the individual.

We are also interested in how much of the response in average hours worked is due to additional participation by former non-employed and how much is due to additional hours worked by single parents that were already working. To this end we estimate (3) on single parents in the treatment and control groups that were always working, and are observed before and after the reform.

We also estimated the effect on average hours worked using a RD approach. Unfortunately, this did not result in significant parameter estimates for the treatment dummy.

## 4 Data

We use data from the *Arbeidsmarktpanel* ('Labour Market Panel'), see Statistics Netherlands (2009). This dataset combines the data from the *Sociaal Statistisch Bestand* ('Social Statistical Database') and the *Enquete Beroepsbevolking* ('Labour Force Survey') to create a large administrative panel dataset for households for the period 1999–2005. With the policy reform in 2002, the dataset contains observations up to three years prior to the reform and up to four years after the reform.

The dataset contains detailed administrative information on socio-economic and demographic characteristics of households. In particular, we have the employment status,

Table 1: Descriptive statistics: treatment and control group

| | Treatment group | | Control group | |
|---|---|---|---|---|
| | Youngest child 12-15 | | Youngest child 16-21 | |
| | Mean | SD | Mean | SD |
| Age of youngest child | 13.5 | 1.1 | 18.1 | 1.6 |
| Age | 44.4 | 5.0 | 48.1 | 4.3 |
| Female | .86 | .35 | .79 | .41 |
| Male | .14 | .35 | .21 | .41 |
| Lower educated | .38 | .48 | .36 | .48 |
| Higher educated | .62 | .48 | .64 | .48 |
| Native | .73 | .44 | .76 | .42 |
| Western immigrant | .16 | .37 | .12 | .33 |
| Non-Western immigr. | .11 | .31 | .11 | .32 |
| >3 children | .02 | .12 | .00 | .05 |
| Employed | .79 | .41 | .85 | .35 |
| Hours per month | 104.2 | 35.5 | 119.0 | 58.0 |
| Observations | 16171 | | 17953 | |

the monthly number of working hours and individual characteristics like age, gender, education and ethnicity. Furthermore, we have the number and ages of the children in the household. From this dataset we select single parents aged 20-57 years old who have a (dependent) child aged 0-27 years old.

Table 1 gives some descriptive statistics of the treatment and control group. The treatment group consists of single parents with a youngest child aged 12 to 15 years old. The control group consists of single parents with a youngest child aged 16 to 21 years old. We have 16171 observations in the treatment group and 17953 observations in the control group.

As expected, the youngest children in the treatment group are on average younger than the children in the control group. Furthermore, also their parents are on average younger than the parents in the control group. Most of the single parents in the treatment group are women, and the same is true for the control group. The share of women in the treatment group is somewhat higher than in the control group (a composition effect). The majority of single parents in both the treatment and control group are higher educated. The differences in the shares of low and high educated across the two groups are small. Most of the single parents are 'natives' (both their parents are from the Netherlands),

Figure 2: Participation pre and post reform period: treatment group and control groups

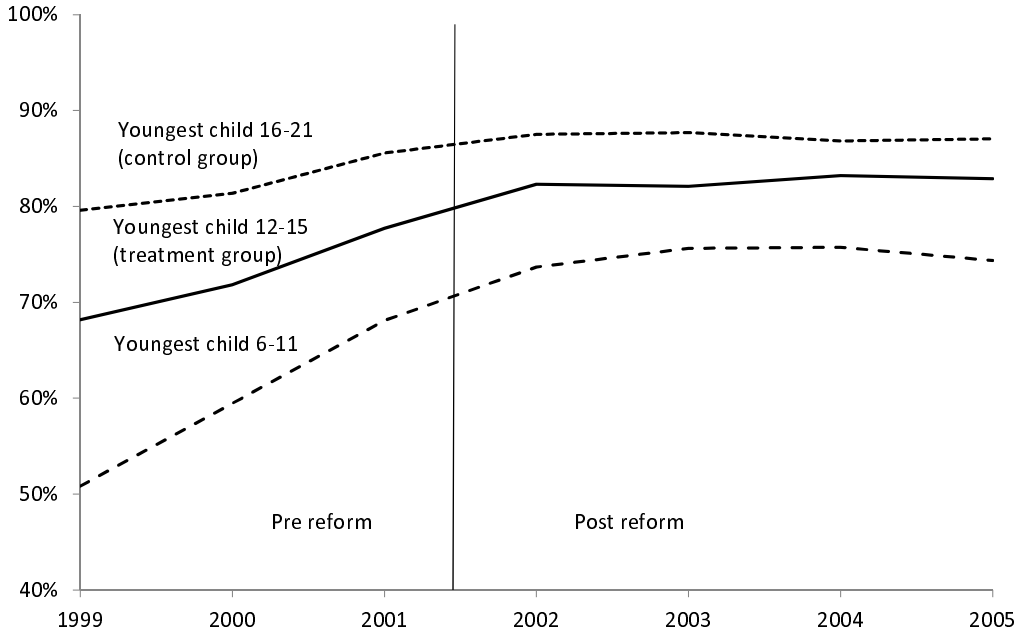
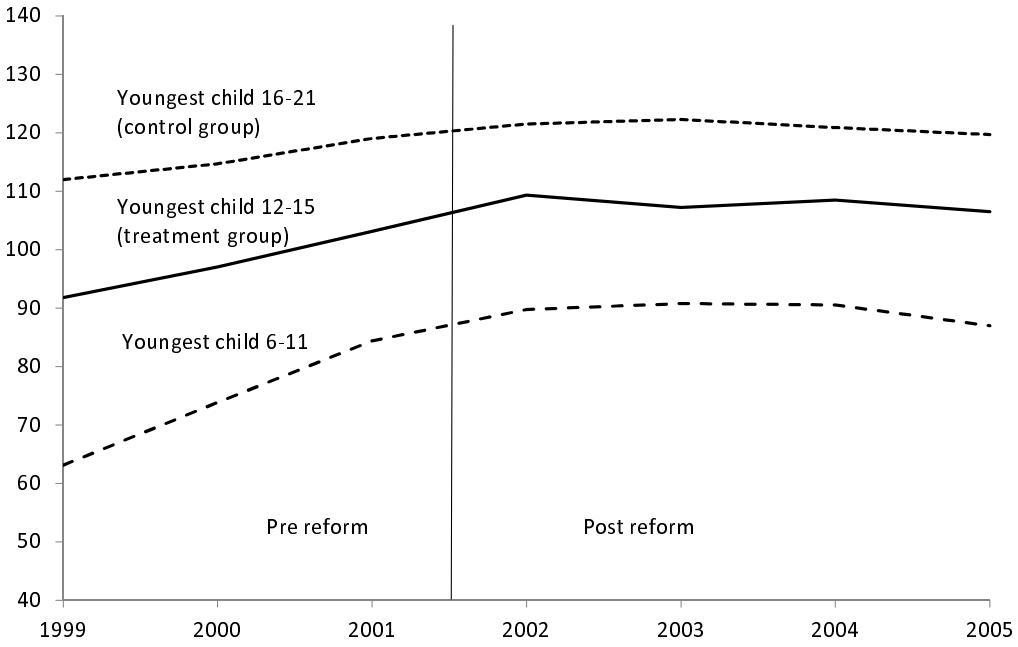



Figure 3: Average hours worked pre and post reform period: treatment group and control groups



but some have parents from abroad. The shares of natives, Western immigrants and Non-Western immigrants differ somewhat between the treatment and control group. Finally, there are more big families (here defined as having more than 3 children living at home) in the treatment group than the control group.

Regarding the outcome variables, we see that the participation rate of the treatment group is lower than the control group. Hence, single parents with younger children are less likely to work. The average number of hours worked per month in the treatment group is also lower than the control group. This reflects both the difference in participation and the average number of working hours of those parents that are working. The difference in average working hours is bigger in percentage terms than the difference in participation rates, so working single parents in the treatment group work less hours than working single parents in the control group.

We choose this control group because of the similarity in the development of participation and average working hours before the reform. Figure 2 gives the participation rates for the treatment and control group over the period 1999-2005, and Figure 3 gives the same for average hours worked. An eyeball test suggests that the control group has a similar pattern to the treatment group in both cases (which is not true for single parents with a youngest child aged 6 to 11 years old). However, this is not controlling for differences in *e.g.* age and education of the parents. Below we will test whether a placebo treatment dummy for a period before the reform (2000) is significantly different from zero (bad news) and significantly different from the treatment effect (good news).

# 5 Results

We first consider the estimated treatment effect on the participation rate, and subsequently consider the effect on working hours. We present baseline results and robustness analyses.

## 5.1 Participation

Table 2 gives the results for participation. The first column gives the treatment effect when we only include a group dummy and year dummies, but none of the other controls (full regression results can be found in the Appendix). We find an effect of 4.7 percentage points. When we control for observable characteristics, column 2, the treatment effect drops to 3.8 percentage points. Finally, when we control for unobservable characteristics via fixed effects, the treatment effect drops to 1.7 percentage points. The fixed effects estimates indicate that the composition of either the control or treatment group changes over time not captured by observable characteristics.

Table 2: Effect on participation: DD

| | Without controls | With controls | With fixed effects |
|---|---|---|---|
| Treatment | 0.047*** | 0.038*** | 0.017** |
| | (0.012) | (0.012) | (0.007) |
| Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1. | | | |

Table 3: Effect placebo on participation: DD

| | With fixed effects |
|---|---|
| Placebo | -0.005 |
| | (0.007) |
| Treatment | 0.016** |
| | (0.007) |
| Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1. | |

The estimate including fixed effects is our preferred estimate. The average participation rate in the period 2002-2005 was 83 percentage points. Hence, the policy reform led to an increase in the participation rate of 1.7/(83-1.7) = 2.1%.

To check whether we capture the effect of the policy change as opposed to some other difference in the growth of participation in the treatment or control group we do a placebo test. Specifically, we include a separate treatment dummy for the year 2000. The coefficient on this dummy should not be significantly different from zero and should be significantly different from the genuine treatment dummy.

Table 3 gives the result. The placebo dummy is indeed not significantly different from zero but with a p-value of 0.0222 we can reject that the placebo dummy is equal to the treatment dummy. As another robustness check we consider the results of a regression-discontinuity approach, equation (2). As noted above, the RD-design identifies the average effect of the tax credit for all working parents with a youngest child younger than 16. We present results for 2002-2005.

Table 4 gives the results using the RD approach. With a second order polynomial we find a treatment effect of 3.3 percentage points for parents with a youngest child 0-12 years old, but with a (significant) third order polynomial this effect drops to 1.9 percentage points. When we zoom in on the dicontinuity the estimate drops to 2 percentage points (the third order term is no longer significant when we zoom in). The RD-results are quite close to the DD results. 2 percentage points is an increase in the participation rate due to the reform in 2002-2005 of 2.5%.

Table 4: Effect on participation: RD

| Youngest child | 0-27 | 0-27 | 12-20 | 14-18 |
|---|---|---|---|---|
| TreatmentRD | 0.0332*** | 0.0192** | 0.0197** | 0.0202** |
| | (0.00801) | (0.00902) | (0.00846) | (0.00928) |
| Age youngest child | 0.0219*** | 0.0311*** | -0.0298* | -0.0778 |
| | (0.00180) | (0.00379) | (0.0160) | (0.0538) |
| (Age youngest child)$^2$ | -0.000382*** | -0.00132*** | 0.00119** | 0.00273* |
| | (6.36e-05) | (0.000334) | (0.000494) | (0.00165) |
| (Age youngest child)$^3$ | | 2.38e-05*** | | |
| | | (8.11e-06) | | |
| Observations | 52135 | 52135 | 20539 | 11933 |

Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1.

We can also determine the implicit participation elasticity. A modal single parent has gross income of 28,000 euro in 2002. Her or his net disposable income is 23,326 euro. The tax credit for this parent is 1,204 euro, and is proportional to gross income, which suggests an increase in net income of about 5.2%. Participation increases by 2.1% (DD) to 2.5% (RD). Hence, the implicit wage elasticity of participation is 0.4–0.5.

### 5.2 Monthly number of hours worked

We are also interested in the effect on the number of hours worked. Table 5 gives the results for the treatment dummy in the DD model in equation (3). Again, the full estimation results including the estimated coefficients for the controls, can be found in the Appendix. Excluding controls, we find a treatment effect on the average number of hours worked per month by the treatment group (including the unemployed) by 4.1 hours per month. Including controls, this number drop to 3.3 hours per month. Including also fixed effects, the treatment effect becomes 3.5 hours per month (=0.8 hours per week). This is 3.3% of the average number of hours worked by the treatment group over the period 2002-2005 (minus the treatment effect).

Table 6 then again considers the effect of a placebo treatment dummy for 2000. As with participation, we find that the placebo dummy is not significantly different from zero, and it is significantly different from the genuine treatment dummy with a p-value of 0.0032.

We can again calculate the implicit labour supply elasticity, now for total hours worked. Net income rises by 5.2%. Hours worked rises by 3.3%. The implicit total hours worked

Table 5: Effect on monthly hours worked: DD

| | Without controls | With controls | With fixed effects |
|---|---|---|---|
| Treatment | 4.09** | 3.34** | 3.52*** |
| | (1.88) | (1.79) | (1.02) |
| Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1. | | | |

Table 6: Effect placebo on monthly hours worked: DD

| | With fixed effects |
|---|---|
| Placebo | -0.61 |
| | (1.06) |
| Treatment | 3.35*** |
| | (1.02) |
| Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1. | |

Table 7: Effect on monthly hours worked, individuals always working and observed before and after policy change: DD

| | Without controls | With controls | With fixed effects |
|---|---|---|---|
| Treatment | -2.50** | -2.15* | 2.62*** |
| | (1.17) | (1.10) | (0.80) |
| Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1. | | | |

Table 8: Effect placebo on monthly hours worked, individuals always working and observed before and after policy change: DD

| | With fixed effects |
|---|---|
| Placebo | 0.91 |
| | (0.90) |
| Treatment | 2.87*** |
| | (0.83) |
| Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1. | |

elasticity is then 0.6.

Finally, we are also interested in whether the change in total hours worked is due to both an increase in participation, and an increase in working hours by individuals that would work anyway. To explore this question, we estimate model (3) only on individuals that are i) working in all periods that they are observed, and ii) observed both before and after the reform (to prevent selection bias due to individuals that start working after the reform due to the reform).

Table 7 gives the results for this group. Surprisingly, we find negative numbers without controls and with controls. However, once we control for unobservable characteristics using fixed individual effects, we do find a significant positive effect on hours worked of 2.6 hours per week for individuals already working. Hence, part of the response in total hours worked seems due to an increase in hours worked per week.

Table 8 gives the corresponding placebo test. The placebo dummy is again not significantly different from zero, but with a p-value of 0.0645 is is only borderline significantly different from the genuine treatment dummy.

# 6 Conclusion

In 2002 the eligibility conditions of the tax credit for working single parents were relaxed to stimulate labour force participation. We consider the increase of the maximum eligible age of the youngest child as an attractive natural experiment to evaluate the impact of the tax credit. We apply a difference-in-differences approach and a regression-discontinuity approach to estimate the effects on labour supply of single parents.

We find that the policy change contributes to an increase in the employment rate by 1.7–2.0 percentage points. Because we have panel data we are able to control for observed as well as unobserved individual characteristics. This turns out to be important for the result. If we ignore the unobserved characteristics we find larger effects for participation, which would overstate the impact of the reform. Average working hours increase by 0.8 hours per week. Part of this increase is due to an increase in the number of hours worked by individuals that would work anyway (the intensive margin) These results remain politically relevant as the current government has plans to reverse the increase in the age limit.

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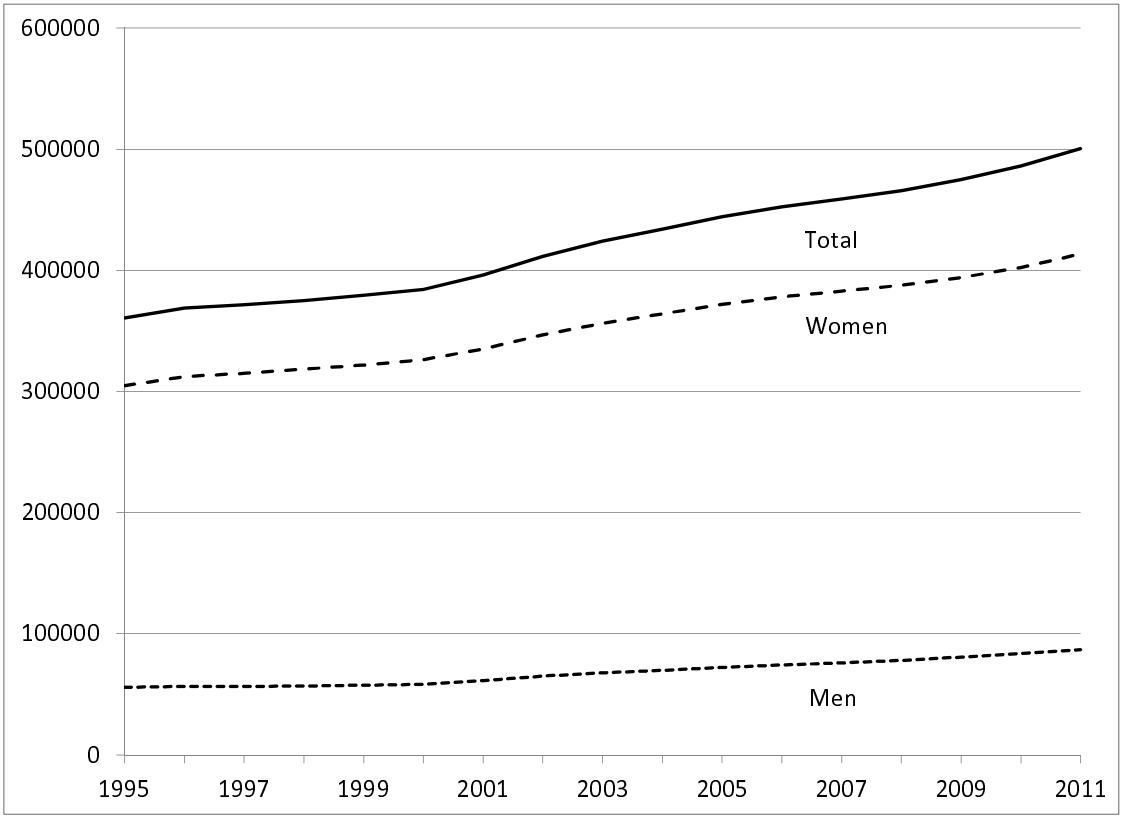
# A Single parents over time

Figure 4: Single parents 1995-2011




Source: Statistics Netherlands.

# B Full regression results

Table 9: Effect on participation rate: DD

| | Without controls | With controls | With fixed effects |
|---|---|---|---|
| Treatment | 0.0465*** | 0.0384*** | 0.0167** |
| | (0.0120) | (0.0117) | (0.00659) |
| Group | -0.0939*** | -0.0562*** | -0.00604 |
| | (0.00995) | (0.0100) | (0.00644) |
| Age | | 0.0899*** | 0.0584*** |
| | | (0.0104) | (0.0120) |
| $(\text{Age})^2$ | | -0.000924*** | -0.000440*** |
| | | (0.000111) | (0.000125) |
| Male | | 0.0872*** | |
| | | (0.00742) | |
| Western immigrant | | -0.0999*** | |
| | | (0.0131) | |
| Non-Western immigrant | | -0.0625*** | |
| | | (0.0127) | |
| Higher educated | | 0.120*** | |
| | | (0.00684) | |
| Year 2000 | 0.0207*** | 0.0202*** | 0.00581 |
| | (0.00613) | (0.00605) | (0.00376) |
| Year 2001 | 0.0697*** | 0.0674*** | 0.0273*** |
| | (0.00731) | (0.00724) | (0.00405) |
| Year 2002 | 0.0799*** | 0.0806*** | 0.0274*** |
| | (0.00970) | (0.00954) | (0.00413) |
| Year 2003 | 0.0800*** | 0.0803*** | 0.0209*** |
| | (0.00974) | (0.00963) | (0.00329) |
| Year 2004 | 0.0819*** | 0.0806*** | 0.0120*** |
| | (0.00963) | (0.00958) | (0.00213) |
| Year 2005 | 0.0829*** | 0.0805*** | |
| | (0.00968) | (0.00967) | |
| Observations | 32780 | 32780 | 32780 |

Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1.

Table 10: Effect placebo on participation rate: DD

| | With fixed effects |
|---|---|
| Placebo | -0.00448 |
| | (0.00672) |
| Treatment | 0.0155** |
| | (0.00658) |
| Group | -0.00470 |
| | (0.00658) |
| Age | 0.0586*** |
| | (0.0120) |
| $(\text{Age})^2$ | -0.000441*** |
| | (0.000126) |
| Year 2000 | 0.00776* |
| | (0.00396) |
| Year 2001 | 0.0270*** |
| | (0.00410) |
| Year 2002 | 0.0276*** |
| | (0.00410) |
| Year 2003 | 0.0211*** |
| | (0.00326) |
| Year 2004 | 0.0121*** |
| | (0.00212) |
| Observations | 32780 |
| Robust standard errors in parentheses, ***p<0.01, ** p<0.05, * p<0.1. | |

Table 11: Effect on participation rate: RD

| Youngest child | 0-27 | 0-27 | 12-20 | 14-18 |
|---|---|---|---|---|
| TreatmentRD | 0.0332*** | 0.0192** | 0.0197** | 0.0202** |
| | (0.00801) | (0.00902) | (0.00846) | (0.00928) |
| Age youngest child | 0.0219*** | 0.0311*** | -0.0298* | -0.0778 |
| | (0.00180) | (0.00379) | (0.0160) | (0.0538) |
| (Age youngest child)$^2$ | -0.000382*** | -0.00132*** | 0.00119** | 0.00273* |
| | (6.36e-05) | (0.000334) | (0.000494) | (0.00165) |
| (Age youngest child)$^3$ | | 2.38e-05*** | | |
| | | (8.11e-06) | | |
| Age | 0.0662*** | 0.0652*** | 0.0773*** | 0.0927*** |
| | (0.00315) | (0.00320) | (0.0122) | (0.0175) |
| (Age)$^2$ | -0.000693*** | -0.000680*** | -0.000790*** | -0.000955*** |
| | (3.76e-05) | (3.82e-05) | (0.000130) | (0.000185) |
| Male | 0.0693*** | 0.0696*** | 0.0596*** | 0.0567*** |
| | (0.00643) | (0.00643) | (0.00787) | (0.0102) |
| Western immigrant | -0.0952*** | -0.0947*** | -0.0981*** | -0.104*** |
| | (0.00855) | (0.00855) | (0.0139) | (0.0175) |
| Non-Western immigrant | -0.0526*** | -0.0525*** | -0.0618*** | -0.0528*** |
| | (0.00942) | (0.00941) | (0.0137) | (0.0174) |
| Lower educated | -0.207*** | -0.207*** | -0.165*** | -0.160*** |
| | (0.00647) | (0.00647) | (0.00920) | (0.0116) |
| More than 3 children | -0.143*** | -0.143*** | -0.0402 | -0.0304 |
| | (0.0197) | (0.0197) | (0.0429) | (0.0573) |
| Year 2002 | | | 0.00197 | -0.00171 |
| | | | (0.00432) | (0.00601) |
| Year 2003 | 0.00419 | 0.00421 | | |
| | (0.00285) | (0.00285) | | |
| Year 2004 | -0.000215 | -0.000220 | 0.00121 | -0.00378 |
| | (0.00339) | (0.00339) | (0.00411) | (0.00575) |
| Year 2005 | 0.00101 | 0.00105 | -0.00196 | -0.00882 |
| | (0.00379) | (0.00379) | (0.00509) | (0.00721) |
| Observations | 52135 | 52135 | 20539 | 11933 |

Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1.

Table 12: Effect on monthly hours: DD

| | Without controls | With controls | With fixed effects |
|---|---|---|---|
| Treatment | 4.093** | 3.335* | 3.520*** |
| | (1.877) | (1.790) | (1.018) |
| Group | -17.52*** | -12.80*** | -3.122*** |
| | (1.538) | (1.521) | (0.980) |
| Age | | 12.25*** | 7.207*** |
| | | (1.575) | (1.735) |
| $(\text{Age})^2$ | | -0.131*** | -0.0523*** |
| | | (0.0169) | (0.0182) |
| Male | | 37.41*** | |
| | | (1.283) | |
| Western immigrant | | -4.737** | |
| | | (2.044) | |
| Non-Western immigrant | | -5.985*** | |
| | | (1.931) | |
| Higher educated | | 18.95*** | |
| | | (1.180) | |
| Year 2000 | 3.138*** | 3.677*** | 1.707*** |
| | (0.947) | (0.920) | (0.577) |
| Year 2001 | 7.901*** | 8.319*** | 3.762*** |
| | (1.123) | (1.090) | (0.616) |
| Year 2002 | 10.47*** | 10.93*** | 3.206*** |
| | (1.528) | (1.473) | (0.647) |
| Year 2003 | 9.804*** | 10.25*** | 2.616*** |
| | (1.536) | (1.487) | (0.515) |
| Year 2004 | 9.747*** | 9.928*** | 1.611*** |
| | (1.527) | (1.486) | (0.375) |
| Year 2005 | 8.495*** | 8.553*** | |
| | (1.532) | (1.495) | |
| Observations | 32780 | 32780 | 32780 |

Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1.

Table 13: Effect placebo on monthly hours: DD

| | With fixed effects |
|---|---|
| Placebo | -0.614 |
| | (1.060) |
| Treatment | 3.346*** |
| | (1.024) |
| Group | -2.937*** |
| | (1.011) |
| Age | 7.231*** |
| | (1.739) |
| $(\text{Age})^2$ | -0.0524*** |
| | (0.0182) |
| Year 2000 | 1.974*** |
| | (0.658) |
| Year 2001 | 3.722*** |
| | (0.622) |
| Year 2002 | 3.239*** |
| | (0.642) |
| Year 2003 | 2.637*** |
| | (0.511) |
| Year 2004 | 1.622*** |
| | (0.374) |
| Observations | 32780 |
| Robust standard errors in parentheses, *** p<0.01, ** p<0.05, * p<0.1. | |