Appendix A: Summary tables of the articles dealing with the antipoverty and employment effects of various policies

In the following tables, some abbreviations have been used in order to reduce the size of paragraphs:

w/ → with, w/o → without, b/w →between, ppts → percentage points, pvty → poverty, min → minimum, CC→ childcare, hh→ household, ctrl →control.

Table A1: Employment effects of the minimum wage

Author(s), country. Evaluated program, period	**Method, independent variable(s)**	**Employment effects**	**Comments**
Sabia (2008), USA, effective minimum wage, 1992-2005	Regression model: employment, weekly hours, weeks last year, annual hours regressed on the ln(effective min wage), state and year fixed effects, a state-specific time effect (quadratic), state economic controls (average wage rate, unemployment rate, ln(GDP), and welfare variables (welfare waivers, ln(max AFDC+food stamps)), **Single mothers**, for various educational levels.	No impact on single mothers overall, even in specifications including government transfers and the EITC; however, reduces weekly and annual hours and weeks last year for high school dropouts. HS dropouts: Increase 10% in min wage reduces employment by 8.8%, weekly hours -9.2%, weeks last year -11.6%, annual hours -11.8%	Elasticities are quite large compared to literature on teenagers , maybe due to the fact that before PRWORA there was strong disincentive to work; in addition min wage increases may shift employment away from low-skilled adults
Bazen (2000), France, Belgium, the Netherlands, Germany, Italy, Denmark, UK, US, minimum wage and collective bargaining	Correlations; Remark: min wage 60% in F, 50% in BE and in NL, poverty line = 50% average income	There is no one-to-one mapping between the presence of mechanisms to regulate low wages and labor market performance	EC: purely descriptive, **not included in vote-counting procedure.**

Vedder, Gallaway (2002), USA, state & federal minimum wage, 1953-1998	Regression model (ARIMA), impact of min wage deflated with CPI on hours worked and overtime hours, controlling for the unemployment rate and GDP growth, **for all full-time year-round workers in non agricultural sector.**	Significant negative relationship b/w hours worked and overtime hours and the real minimum wage, full-time year-round workers. A 10% increase in the minimum wage reduces the number of hours worked by 1.3%; a 1$ (real) increase reduces weekly hours in manufacturing sector by 0.816 hrs and overtime hours by 0.349 hr.	Reason for choosing full-time year-round workers: they usually do not lose their jobs as a consequence of minimum wage hikes, yet they may have a reduction in hours worked. Indeed, FTYR workers tend to work a bit less when min wage increases, i.e. the employment effect of min wage is probably understated in most studies (only unemployment is considered).
Neumark, Wascher (2002), USA, state & federal minimum wage, 1986-1995	Regression model (multinomial logit), probability of more workers/same number/fewer workers regressed on real minimum wage (deflated w/ CPI), a state dummy variable, fixed year effects, a vector of controls (unemployment, quartiles of earnings distribution, and welfare policies (AFDC waivers, max AFDC benefits). **All families.**	Effect on number of workers in family: total effect has a significant positive impact on the likelihood of having fewer workers in year 2; only significant for nonpoor families (in year 1). A $1 increase in real min wage increases the likelihood of having fewer workers in year 2 than in year 1 by 0.021 for nonpoor families (0.015 for all observations, significant at 10% level)	A higher minimum wage generates trade-offs: some families gain and escape poverty and others slip into poverty, due in part to negative employment effects. On balance, no compelling evidence that minimum wages help in the fight against poverty. Various trade-offs more closely resemble income redistribution among low-income families than income redistribution for high- to low-income families.

Watson (2000), UK, if all workers' wage=w(p), i.e wage that could be earned given their human capital; i.e. without underpayment, what would happen ? 1985-1993	Simulation, calculation of full earnings given the number of hrs worked. H0: workers cannot change contractual hours & H0': if min wage <=w(p), no decrease in employment. In order to determine underpayment, ln(wage) is regressed on a sum of human capital characteristics, and unobserved individual characteristics to calculate the wage person should get.	Wage capacity rates differ substantially across industries (agriculture, energy, metal extraction, engineering, other manufacturing, construction, distribution, transport, bank/finance, other services), difference b/w union and non-union. The wage capacity ranking is as follow (increasing): agriculture, distribution, other manufacturing, transport and communication, construction, other services, engineering, metal extraction, bank/finance, energy. As the degree of underpayment increases in a group, the minimum wage rate that can be introduced without causing unemployment will also increase.	As wage capacity rates differ substantially, a Wage Councils orientation is preferable to a national minimum wage. EC: does not provide an estimate of employment effect, hence **not included in vote-counting procedure.**
Neumark, Adams (2003), USA, living wage legislation + minimum wage, 1996-2000	Regression model, impact of presence of living wage ordinance and minimum wage. **Probability of employment in various ranges of the wage distribution (lowest decile, b/w 10th and 25th percentile, between 25th and 50th percentile**, and b/w 50th and 75th) is regressed on year, month and city fixed effects, on max(ln(wage ordinance, min wage)), on ln(min wage) and a vector of individual characteristics →Workers in cities that have adopted living wage legislations are compared to workers in metropolitan areas that haven't ; lags of 6 and 12 months are tested	Living wage has a significant negative impact (w/ 12-month lag) on probability of being employed for workers below lowest decile of wage distribution, and positive impact b/w 25th and 50th (contemporaneous and 6 months) and b/w 50th and 75th percentile (contemporaneous and 12 months). Minimum wage has no significant impact. Lowest-wage workers (lowest decile) : a 10% increase in LW lowers probability of employment by 0.56 percentage points	LWO: Disemployment effects appear moderate. There is some evidence of positive employment effects for workers in the higher percentiles of the wage distribution.
Portugal, Cardoso (2006), Portugal, minimum wage, 1986-1989	Regression model (Poisson regression/ firm-specific random effects), impact of large increase in minimum wages for teenagers in 1987. Gross flows (accessions and separations are analyzed separately). Number of **teenage workers** (hired/ separated/dismissed) regressed on year dummies (88 and 89), firm size and hiring rates and market concentration. Also simulations w/ 3 alternative scenarios for teenage shares: actual 1986, the 1988 & 1989 share	Companies significantly decreased the share of teenagers among their newly hired workforce; however, also a significant negative impact on share of teens in job separations. New firms ('88 and '89) recruited a significantly lower share of teenagers than those set up before min wage hike of 1987; in addition, teens overrepresented in firms going out of business. In 1988.in 1988 & 1989, the share of teenagers in overall job accessions to continuing firms was 4% lower than in 1986; however, the share of teenagers in job separations was 15% lower in 1988, and 14% lower in '89. New firms (set up in '88 & '89) hired a 4% lower share of teens. Simulations: the change in min wage was responsible for an increase in total employment of 0.41 percentage pts in	Therefore, the decline in separations has clearly driven the rising teenage employment level. Another specification measures the impact of retention rates, which increased by 38%, which points to the relevance of supply-side factors. It should be noted that Portugal's labor market has one of the most stringent employment

	that would prevail given the estimated impact (regression parameters) and actual shares in '88 & '89	'88 and 0.32 percentage pts in '89.	protection legislation
Skedinger (2004), Sweden, minimum wage through collective bargaining, 1979-1999	Regression model, situation in Swedish **hotels and restaurants**, gross flows (hirings and separations) limited to **unskilled workers**. Sample is divided into two parts: observations with increasing or with decreasing min wages. Each is divided into four groups: wage lower than min wage, treatment group with wage b/w old and new min wage, control group w/ wage b/w new min wage and 1,05*new min wage, and above. Separate estimations for 1979-1991 and 1992-1999. Logit: probability of separated in the next period (min wage increase) or accessed the job in current period (min wage decrease) explained by group dummies*real increase/decrease in min wage and controls (real wage, age, number of employees, gender, type of contract, occupation).	Min wage increases and decreases contribute significantly to job separations and accessions, respectively, except for teenagers 1993-1988; job separations increase when min wage increases, but the evidence is less conclusive for accessions.1979-91: an increase in min wage increases job separations in treatment group by 1.332 % more than in the "high" wage group, by 0.756% for control group. The difference 1.332-0.756 = treatment effect = 0.576%. For decreasing min wages, the difference in elasticities is 0.843%. 1992-1999: the effect is insignificant for workers aged 20-59 in the full sample, but significant for two-year panel of firms; for teens the impact is insignificant	The effects are not dramatic, but of non-negligible magnitude. Another specification leads to the conclusion that the effects of decreasing min wages on job separations are insignificant and smaller than the effects of increasing min wages; however, accessions are at least as large when minimum wages increase as when they decrease. Differs from Card and Kruger, maybe due to the fact that the min wage bite is larger in Sweden
Campolieti, Gunderson, Riddell (2006), Canada, minimum wage, 1981-1997	Empirical estimate, new research design by Neumark, where researcher has to precomit to a methodology, specifications, and data set prior to estimating employment effects; method applied here to Canadian data. Regression explaining the employment to population ratio of **youth (16-19 & 20-24).** Baseline model: employment-to-population ratio for a given age group regressed on ratio of min wage to average wage of workers 16-64 in given region (both year t and t-1), regional and year dummies, and controls. Another specification uses the "fraction below new minimum wage" instead of the minimum wage itself.	The sum of current and lagged effect: statistically significant negative effect on employment for teens (16-19, elasticity -0.282) and young adults (20-24, elasticity -0.155). Using "fraction below" as min wage variable, also significant negative impact. Then, two-stage regression model, modeling demand elasticities for teens, in which the wage of the group under study is regressed on the minimum wage and control variables, and then in the second stage the employment rate is regressed on this instrumented wage rate and control variables. This 2-stage model leads to elasticities most of which are insignificant for teens, but significant for 20-24. For 16-24, specification with minimum wage index, elasticity of -0.256 (current + lagged), and lagged elasticity is larger than contemporaneous. The adverse effect is larger for 16-19, elasticity = -0.282. For specification w/ "fraction below"	(i) There may be publication bias, but also "author bias" where authors run alternative specifications until they get the results they want. (ii) in Canada, no federal min wage b/c under provincial jurisdiction --> considerable cross-sectional and time series variation, hence particularly appropriate (iii) substantial adverse employment effects for

		variable, results are fairly similar.	youths, however, estimates for demand elasticities are mixed.
Ragacs (2007), Austria, minimum wage through collective bargaining, 1967-1995	Empirical estimate, growth rate of employment = f(time trend, growth rate of output, lagged labor productivity, ln(real minimum wage)), also lagged employment is added as a predictor. **Aggregated data for Austrian industry, average employment**	Contrary to the "textbook model", no significant negative effect of minimum wages on employment could be found.	Austrian labor market may be organized differently than described in textbook model, and unions' wage claim may be very modest, due to productivity-oriented wage settings
Doucouliagos, Stanley (2009), USA, minimum wage	**Meta-regression** analysis, **1,474 estimates** from 64 US studies, accounting for publication bias by regressing the estimated elasticities on their standard error. There's evidence of publication bias, i.e. selection for negative employment effects. Selection bias included in regression model.	Estimates of the empirical effect corrected for publication bias show that there's no evidence of a genuine employment effect when using Card and Kruger data set. Using the authors' much extended data set and meta-regression analysis techniques, conclusions are similar: there's strong evidence of publication selection; once this publication selection is filtered, no evidence of a minimum-wage effect remains. The publication bias appears rather severe. publication bias -> elasticity of employment effect = f(standard error): the intercept of regression model lies between -2 and -3 for the specifications preferred by the authors or the articles, -1.3 to -2.63 for all estimates. The all-set meta-regression finds, contrary to the best-set one, a very small but significant negative minimum-wage effect --> a 10% increase in min wage decreases employment by 0.1%. Conclusion: the average publication bias for the minimum-wage literature is -0.231 (multivariate meta-regression analysis) or -0.273 (simple multiple regression). Subtracting the estimated publication bias, the average employment effect of -0.19 is converted to +0.041, but too small to be of practical import.	The elasticity estimates are getting 0.14 larger (or less negative) every decade, not as a result of falling real minimum wage EC: meta-analysis, hence **not included in own meta-analysis, but analyzed in text.**

Sabia (2009 a), USA, effective minimum wage, 1979-2004	Regression model, impact of ln(min wage) on **teenage** employment-to-population ratio and hours of work (unconditional and conditional hours), with state effects and a set of state-level time-varying controls (economic + demographic). Appropriateness of including year effects, which may capture much of the variation in the min wage measure, but this is hotly contended. First a model w/ months dummies, then another specification w/ year & months dummies.	For period 1979-1997: Without year effects, stat. significant negative impact on employment/population ratio for teens, whereas when year effects are included, the estimated elasticity becomes insignificant. For period 1979-2004: the impact is statist significant w/ or w/o year effects. Hours effect: consistent evidence of decrease in average weekly hours. Evidence for conditional hours = inconclusive. For 1979-1997: w/o year effects, a 10% increase in min wage is associated with a 2.8-3% decline in teenage employment ratio. 1979-2004: w/o year effects, -2%, and w/ year effects -0.3%. When lagged effects included (t and t+1), for 1979-2004, the elasticity is slightly larger and significant: a 10% increase in min wage associated with a "long-run" 2.5-3.3% decrease in ratio of employed teens. (Unconditional) hours effects: "long-run" elasticities range from -0.37 to -0.51	
Kalenkoski, Lacombe (2007), USA, Federal & state minimum wage, 2000	Regression model (OLS) at county level, b/c geographical variables are important, e.g. Card & Kruger noted changes induced by the inclusion of regional dummies. Log of employment-to-population ratio for **16-19 yrs old** at the county level, explained by log of effective min wage, income/capita, weekly wage, unemployment rate, and demographic controls, accounting for spatial autocorrelation, notably neighboring counties. 3,605 counties in sample	Increase in the effective minimum wage has significant impact on teenage employment/population ratio. Increase of 10% leads to a statistically significant decrease in employment/population ratio of teens of 2.5%. With the spatial autocorrelation model, a 10% increase causes a 3.2% decrease in youth employment to population ratio	Accounting for spatial autocorrelation leads to a coefficient that is 28% higher than the OLS estimate
Addison, Blackburn, Cotti (2009), USA, federal & state minimum wage, 1990-2005	Regression model at country level (payroll data) for the **retail-trade sector nationwide**, ln(employment) is regressed on ln(min wage), county and quarter fixed effects, and controls (population, total employment, weekly earnings, unemployment and enrollment rate); the model incorporates a county-specific time trend in the error term, as in Sabia 2008. Results are sensitive to the inclusion of state-specific trends, impact of the "effective" minimum wage.	W/o time trend: 3 out of 5 coefficients are positive but insignificant (1), (2) & (4); only significant effect is negative in subsector (3). With county-specific trends, 4 out of 5 coefficients are stat significant and positive (whereas for sector (3) at 10% level only). Another model accounting for border counties leads to a largely similar general pattern of results as that obtained w/ detrended data. W/ county-specific trends: a 10% increase in min wage leads to a 1 to 2% increase in employment in 4 out of 5 subsectors of retail trade	Explanation may be efficiency wage or monopsony, but authors not convinced. Could be the impact of increased earnings for minimum-wage workers (increased demand).

	5 subsectors: (1)food & beverage stores, (2)supermarket/grocery stores, (3)convenience stores, (4)specialty food stores, (5)beer/wine/liquor stores		
Fang, Gunderson (2009), Canada, Minimum wage, 1993-1999. In Canada, min wage is under provincial jurisdiction	Regression model, probability of being employed (probit model) in subsequent year depending on whether in treatment (wage b/w old & new min wage at time t-1) or control group, controlling for individual characteristics, and including time and region dummies. The focus is on **older workers**, contrary to most of the literature, namely workers **50+.** Various ctrl groups: ranging from 0-0.25 Canadian $ above min wage, up to $4 above. Authors prefer 0-1 $ above min wage as ctrl group. Other control groups also include wages below old minimum wage. 22 regressions were carried out (11 comparisons groups, w/ or w/o wages below min wage).	Model for preferred control group, i.e. within $1 above the min wage: minimum wages have a statistically significant and positive impact on affected workers aged 50+. When all 22 regressions are considered, overall, the effects are significant and positive; no coefficient is negative. Model w/ control group within $1: workers 50+ affected by min wage increase were 14 % points more likely to be employed the next year compared with otherwise similar low-wage workers not affected by theses increases	Results for older workers are completely at odds with the results for youth found in the recent Canadian literature, which found substantial negative effect
Wessels (2007), USA, Minimum wage, 1990-1991 and 1996-1997	Reevaluation of Card and Kruger (1995) difference-in-difference model, with the main minimum wage indicator being the proportion of workers affected by the minimum wage hike, but dependent variable changed. Log (employment-to-population ratio 1992/e-t-p ratio 1989), regressed on fraction of affected workers + ln(change in adult employment) as control. **Teenagers.**	Re-evaluate C & K: when controlling for business cycle (adult employment change), the regression coefficient becomes insignificant - same conclusion as C&K. However, second model for 1996-1997 min wage hike leads to a significantly negative coefficient. Regression rerun for states with the federal min wage only: for 1996-97, coefficient is stat significant and negative. Elasticities for 1996-1997: regression coefficient is-0.4112 for fraction of affected workers, and the fraction averaged 44.9% for teenagers, hence elasticity should be -0.4112*0.449=-0.18, as the equation is of the log-linear type. When only states with federal min wage in the equation, the coefficient is -0.7576 (but % of affected workers not indicated hence no elasticity)	C&K were correct with their model for 1990-1991; however, for the 1996-1997 hike, the impact was significantly negative, maybe due to a higher number of workers affected by the hike.

Feldmann (2009), 73 countries, minimum wage, 2000-2003	Regression model (random effects) based on 2004 Executive Opinion Survey, based on 60-70 executives pro country, who gave answers on scales between 0 and 1 (agree/disagree), on 4 topics: minimum wage has little impact b/c low or not obeyed; hiring and firing are determined by private contract; the share of wages set by collective bargaining is low; unemployment benefits preserve incentives to work + 1 objective measure: duration of conscription, rescales --> 0 to 1. Dependent variable is the unemployment rate. A set of controls is added: business regulation, tax burden, GDP growth, % of children, ethnic fractionalization, whether the country is landlocked, wars, transition economy, and the model includes year dummies.	Coefficient for variable "little impact of minimum wage" has not a significant impact on unemployment rate	EC: not quite reliable (opinion questions), hence **not included in meta-analysis**
Böckerman, Uusitalo (2009), Finland, Minimum wage through collective bargaining 1991-1996, reduction in min wage for workers under 25 over period June 1993- June 1995 (80% of the lowest task- and region-specific tariffs)	Regression model (difference-in-difference) →effect of this decrease based on payroll records. Time (before/during/after the reform) and treatment group (Trainee, under 25) dummies and their interactions to explain share of workers in control and treatment group in each firm (employment share + share of hours worked), control group =workers under 30 w/ a maximum of 2 years' work experience. **Workers under 25.**	Contradictory findings: there was a decrease in employment in the affected group both when min wage exceptions were introduced and then removed. Potential endogeneity problem as subminimum wage is reaction to growing unemployment, hence inclusion of interaction term (business cycle x treatment group) and GDP growth or unemployment rate. Authors also narrowed the age range, and compare 24- w/ 26-year-olds. Specifications with interaction terms --> treatment*business cycle (real GDP, unemployment) lead to negative estimates very close to zero and insignificant. W/ narrower age range, slight increase when min wage decreased and slight decrease when exceptions removed, but not significant. Conclusion: No statistically or economically significant effects of changes in youth minimum wage.	i) Due to a severe recession in early 1990s unemployment skyrocketed, and unions signed agreements w/ employer organizations to relax wage regulations for young workers. ii) Out of 86 minimum wage studies surveyed by Neumark and Wascher (2007), only Skedinger (2006) analyzes the effects of union-negotiated min wages

Thompson (2009), USA, minimum wage, 1996-2000, in 15 states covered in dataset based on unemployment insurance wage records and affected by changes in the min wage in the mid-1990s	Much of research showing small effects relies on state-level panels => here regression model: impact of the 1996 and 1997 min wage rises on **teenage employment** share w/ county level data, counties are classified as "low-impact" and "high-impact" based on regional earnings variation, either with thirds or fifths (top and bottom) of teenage earnings distribution. Model (difference-in-difference): teen employment is regressed on impact variable (low vs. high impact), time indicator (=1 after reform) and controls(unemployment, teens as a % of working-age population, adult earnings)	Membership in the high-impact group following the 1996 increase has a stat significant negative impact on teenage employment, the addition of states fixed effects has no impact. Interestingly, county-level data show a stat significant negative effect, while state-level data yields insignificant effects. W/ usual covariates, 1996: membership in high-impact group (bottom third) is associated with a 3.1% reduction in teenage employment (-3.0% if states fixed effects are included); -3.4% if high-impact is bottom fifth; that is elasticities of -0.26 and -0.29 respectively; for the 1997 increase, elasticities are -0.29 and -0.37, but when high-impact defined by bottom third, the results narrowly miss significance. A stable/transitory (stable= lasting at least the full quarter) tenure variable is added, which shows min wage increases only have an impact on transitory teenage employment.	(i) No significant impact on 19-22 years old in most regressions in 1996 and 1997 (ii) Impact was stronger is small counties i.e. with total employment below 10,000 (elasticites in the range 0.59-0.67 in '96 and 0.38-0.54 in '97)
Orrenius, Zavodny (2008), USA, Effective minimum wage, 1994-2005	Regression model (Huber White std errors and correction for heteroskedasticity): log(employment-to-population ratio) of **low-skilled immigrants** is regressed on log(min wage), business cycle controls (log GDP/capita, unemployment, and construction permits), and state and time fixed effects, w/ state-level data. **3 groups: 20-54 natives w/o high school diploma, same group but foreign-born and not US citizen at birth, and all teens.** Alternative independent variable is ln(minimum wage/real average wage of 20-54 yrs old), same model for hours worked. Another model includes mobility variables, w/ % of low-skilled immigrants and average yrs of education among immigrants as dependent variables.	Real min wage has no significant effect but for teens when controlling for state-level economic conditions (significant for male but not for female teens). No significant employment losses for immigrants. Same conclusions for relative minimum wage. Same for hours worked, only significant effect is for teens; however, hours increase for male teens, and decrease for female teens (which explains why no employment effet for female teens). Finally, results suggest that min wages influenced low-skilled immigrants' location patterns. Teens: a 10% increase in the min wage reduces teen employment by 1.8%, a result driven by male teens, among whom a 10% increase reduces employment by 1.9%; non stat significant effect for female teens. The estimated responses to this min wage increase are 1% increase in male teens hours worked, and a 1.3% decrease in female teens' average hours. Regarding mobility patterns: raising the minimum wage has a stat significant negative impact on the fraction of low-skilled immigrants at state level	(i)Including a one-year lag of the real minimum wage is not significantly negatively associated with employment rates or average hours, except for employment among low-skilled adult immigrant women. Hence, results are broadly unaffected by this addition. (ii) the gender asymmetry in teen results is striking, may depend on incidence of tipped jobs and higher enrollment in school of young women

Sabia (2009 b), USA, Effective minimum wage, 1979-2004	Regression model of **retail industry**-wide employment effects with a state-month panel, effect on **16-64 as well as teenagers (16-19).** Dependant variable is ratio of retail employment to total population 16-64 regressed on ln(min wage) , state, month, and year fixed effects, and economic controls (ln wage rate of workers 25-54, unemployment rate), average weekly hours equations are also specified.	Minimum wage increase leads to significant but modest reduction in retail employment. However, when state-specific linear month trends are included, adverse retail employment effects disappear. With Kaitz index as independent variable, not stat different from zero. Hours effects: significant negative impact on hours worked, but again quite sensitive to the inclusion of state-specific time trends. Elasticities: min wage hikes seem to have larger effects on retail employment rather than on conditional hours worked. Effects are stronger for teenagers. W/o time trend, employment: a 10% increase in min wage is associated with a 1.1% decline in employment. Hours worked: a 10% increase reduces average retail hours worked by about 1%. Teenagers, employment effect: a 10% increase in min wage associated with a 3.4% decrease in teen retail employment, and a 3.8% to 4.1% decrease in average teenage hours worked in retail trade	State trends may be capturing retail employment variation that the model seeks to explain
Hyslop, Stillman (2005), New Zealand, Large reform of minimum wage affecting youth workers in 2001, period: 1997-2003. Before 16-19 yrs old min wage = 60% of min wage, after 18-19 yrs old get adult min wage + youth min wage 80% of adult min wage	Regression model (difference-in-difference estimates, robust standard errors) comparing employment rate for **16-17 yrs old and 18-19 yrs old** to those of young adults (20-25), before and after the 2001 reform: employment is regressed on age dummies interacted with post-2001 dummy and on single-age dummies, quarter-specific dummies, and on controls (gender, ethnicity, marital status, NZ born, urbanicity, region, and the size of each age group).	Descriptive stat: min wage changes that occurred after 2001 do not appear to have had a substantial effect, at best weak effects. Regression results: the estimated employment effects are negative but not stat significant. Control for announcement effect with a dummy variable (announcement in April 2000, dummy=1 for Q2, Q3 and Q4 of 2000): announcement effects are quite important in 2003. Main model: interact quarterly and age-specific dummy variables and also add 26-49 yr-olds unemployment rates; the estimated impacts are non significant in 2001 & 2002, but significantly negative in 2003. For hours worked, most effects are insignificant. Estimates of the impact on unemployment, results are not plausible. Accounting for announcement effects, there is a significant negative impact on youth employment in 2003 (-2.6% to -2.9%). With the more complete models, including age-specific business cycles, there is a decline in 2003 b/w 2% and 4% of 16-17 and 18-19 year-olds	no evidence of adverse effects on youth unemployment immediately following the reform; some weak evidence of employment loss by 2003; but also an increase in hours worked, depending on specification adopted

Kawaguchi, Yamada (2007), Japan, Min wage is Japan is rather low (as a share of median earnings). All Japanese prefectures are divided into four categories. Period 1993-1999	Regression model based on a panel survey of consumers. 5 groups are defined, observations for 2 consecutive years, the full sample (n=1438): workers with wage <110% min wage (sample A, n=236), sample A w/ wage below 110% for 2 yrs or more (sample A', n=152), sample A' but w/o workers w/ wage below min wage (sample B, n=148), and sample B in 2 yrs or more (sample B', n=96). Linear probability model: Employment dummy regressed on treatment group dummy (wage b/w old and new minimum) and controls (age, children, education, marital status, job openings) and prefecture and time dummies	Results are not statistically significant. The preferred samples, namely A and B, lead to negative regression coefficients, suggesting disemployment effects. However, none is significant. This may be due to very small sample sizes.	EC: small sample sizes, hence **not included in meta-analysis.**
Kawaguchi, Mori (2009), Japan, min wage set at the prefecture level, years 1982, 1987, 1992, 1997, 2002	Regression model (weighted least square), the change in employment rate regressed on the fraction affected over a 5-year period, proportion in each category (7 categories: **male/female teens, male/female young adults (20-24), male/female 60+, and 20-59 women**), year dummies, unemployment, and average wages. Workers working less than 200 days in a year are dropped.	Significant reduction of employment rate among male teenagers, and among women in the 20-59 age group, but not for young adults aged 20-59. Male teens: A 1% increase in the fraction affected reduces employment by 0.2 percentage point. Less clear for female teens. Women aged 20 to 59: A 1% increase in the fraction affected decreases the employment rate by 0.4-0.8 percentage point	Min wage in Japan is rather low (as a share of median earnings)
Abowd, Kramarz, Margolis, Philippon, France, min wage, 1990-1998	Regression model (logit): Treatment t and control groups are defined by the log real minimum wage in year t and t+1 (control group’s upper bound = log (1.1) + ln(real min wage)). Exit models: Probability of being in employment in year t+1 conditional on being in employment in year t and on an increase in the min wage is regressed on interaction treatment*difference b/w new and old min wage (amount), on interaction control*difference, and on a vector of controls (age, sex, seniority, type of contract, education, and year). Two separate model for increasing and decreasing real minimum wages.	Difference-in-diff estimator of exit based on both increases and decreases: the elasticity is -0.404 for French men and -0.2983 for French women, and the impact is statistically significant. The probability of being in employment in year t+1 significantly decreases when the real minimum wage increases. When increases and decreases are analyzed separately, elasticities are very high: approximately -2 for men and -1.5 for women. Regarding entry probabilities: the effects are essentially zero. Hence, an increase in minimum wage strongly and negatively affects exit probabilities in France, while it has virtually no	

	There are also entry models, similarly specified. **Overall employment, broken down by gender.**	effect on entry rates.	
Grogger (2003), USA, EITC, 1979-2000	Based on state-level variation in welfare reform policy. Employment and weeks worked of **female-headed families** are regressed on log (minimum wage), the introduction of time limits, a dummy for the introduction of welfare reform, demographic controls and the generosity of the maximum credit of the EITC. The model includes states dummies and state-specific quadratic trends.	Log(minimum wage) has no significant impact on the employment rate and on the number of weeks worked among female-headed families.	

Table A2: Antipoverty effects of the minimum wage

Author(s), country. Evaluated program, period	**Method, independent variable(s)**	**Antipoverty effect**	**Comments**
Neumark, Wascher (2001), USA, federal & state minimum wage, 1986-1995	Regression model: linear probability model using first-difference estimators (robust standard errors)→ probability of being nonpoor in year 2 if poor in year 1 (*pre-tax* poverty) regressed on real minimum wage, controlling for the unemployment rate, # of children under 18 (1, 2, 3+), with state and year dummies. **Poor and low-income families**. similar specifications to examine effects on changes in the income-to-needs ratio (official poverty line)	No significant effect overall, however positive impact on families with children, with or without workers in year one (lagged effect if no worker in year 1). Significantly increases pre-tax income-to-needs ratio of families below poverty line in year one. An average increase in the real minimum wage (+ $0.2) increases the probability of escaping poverty by 0.013 for families with kids, 0.02 for families w/ at least 1 worker in year 1. Income-to-needs ratio increase = 0.018, which is weak	(i) 23% of families in sample are pretax poor (ii) Families w/o worker in year one +0.012; families w/ at least one worker +0.02
Sabia (2008), USA, effective minimum wage, 1992-2005	Regression model: the dummy variable poverty (income is below the official poverty line) is regressed on the ln(effective min wage), state and year fixed effects, a state-specific time effect	No significant impact on single mothers' poverty (whether they work or not and whatever their educational level)	

	(quadratic), state economic controls (average wage rate, unemployment rate, ln(GDP), and welfare variables (welfare waivers, ln(max AFDC+food stamps). **Single mothers.**		
Joassart-Marcelli (2004), California, state minimum wage, 1998-2000	Simulation, official poverty line, no employment effect	Minimum wage of 8.45$ would be necessary so that a full-time full-year worker can provide the basic needs of a family of 4 i.e. much higher than minimum wage b/w 98 and 2000 ($5.75); if poverty line = 50% of median in Los Angeles, then minimum wage should amount to $12.80	EC: does not provide any estimate of the antipoverty effect, hence **not included in meta-analysis**
Vedder, Gallaway (2002), state & federal minimum wage, 1953-1998	Regression model (ARIMA), impact of real minimum wage (deflated with CPI) on official poverty rate, controlling for unemployment, GDP growth, GDP/capita, and quadratic term for real transfers/capita. **Full-time year-round workers in nonagricultural sector.**	No statistically significant relationship b/w min wage and poverty rate, not even for full-time year-round workers, and not in 3 out of 4 broad census regions: South, Midwest, and West ; in Northeast, significant positive impact: higher min wage increases poverty (for f-t y-r workers). Increase of 1$ in real min wage leads to 0.36 percentage point increase in official poverty rate in the Northeast (+1.56 for Nonwhites)	Reason for choosing full-time year-round workers: they usually do not lose their jobs as a consequence of minimum wage hikes, yet they may have a reduction in hours worked. Indeed, f-t y-r workers tend to work a bit less when min wage increases, i.e. the employment effect of min wage is probably understated in most studies (only unemployment is considered).
Neumark, Wascher (2002), USA state & federal minimum wage, 1986-1995	Regression model (logit), impact of real minimum wage on probability of escaping poverty (official poverty line), remaining in poverty and slipping into poverty and impact on needs-to-income ratio (n-to-i ratio <1 = poor), similar specification as described above.	Full model: real minimum wage has a significant negative impact on probability on staying in poverty in year 2 (contemporaneous effect); however, no significant lagged effect. Total effect on poor population is not significant. For nonpoor persons in HH with at least 1 worker in year 1, minimum wage significantly increases probability of slipping into poverty; no effect for HH w/o a worker. Another model	A higher minimum wage generates trade-offs: some families gain and escape poverty and others slip into poverty, due in part to negative employment effects. On

	All families.	(multinomial logit): effect of min wage on various income-to-needs categories: <=1[poor], 1-1.5, 1.5-2, >=2). Total effect (contemporaneous +lagged) are generally not significant, but increased likelihood to slip into poverty for persons in inc-to-needs ratio >1.5. For all poor HH, regardless of the number of workers: 1$ increase decreases probability of staying in poverty by 0.094 (-0.096 for HH w/ at least 1 workers, -0.081 for HH w/o a worker). For nonpoor HH, a 1$ increase increases probability of slipping into poverty by 0.02 (0.024 for HH w/ at least 1 worker, not significant for HH w/o a worker). Second model shows that for families entering poverty, a $1 higher minimum wage reduces the income-to-needs ratio by 0.08; for families remaining in poverty, however, the contemporaneous effect of a $1 increase is an increase of 0.072 of the income-to-needs ratio, which is nontrivial given that these HH have a ratio <1 (by definition)	balance, no compelling evidence that minimum wages help in the fight against poverty. Various trade-offs more closely resemble income redistribution among low-income families than income redistribution for high- to low-income families.
Heller Clain (2007); USA, living wage legislation + state minimum wage at county level, two time periods: ca. 1990 and ca. 2003	Regression model, impact of presence or absence of both living wage ordinance and state minimum wage above federal minimum on poverty rate. Poverty rate is regressed on a vector of behavioral and demographic controls (race, age, family structure, education, labor force status), unemployment rate, unionization rate, a dummy for state min wage above federal min wage and dummy for the presence of local living wage ordinance. But endogeneity problem, hence 2-stage estimation to estimate instrument variables for the local economic activity and state and local wage policies. **Overall poverty rate.**	Living wage ordinances have a statistically significant impact: they reduce poverty rates; on the contrary, there is no significant evidence that state minimum wages reduce poverty. Presence of a living wage ordinance in at least one municipality of significant size reduces poverty rate by 1.821 - 1.965 ppts (depending on specifications), i.e. for the average county with a population of 720,273 in 2000, b/w 13,115 and 14,155 individuals lifted out of poverty	No evaluation of employment effect, but the favorable effect of living wage ordinances is probably due to the absence of disemployment effects: Coverage is much narrower - approx. 1-2% of workers in lowest quartile of wage distribution - and the employers of low-wage workers targeted by legislation sell services to local governments, who intentionally maintain their demand and absorb the higher costs; the redistribution occurs b/w taxpayers in general and low-wage

			workers
Watson (2000), UK, if all workers' wage=w(p), i.e wage that could be earned given their human capital; i.e. without underpayment, what would happen ? Poverty line = 140% of a single person's supplementary benefit/income support allowance rate. 1985-1993	Simulation, ln(wage) = function of human capital, unobserved characteristics and error term. If the latter = 0 => potential wage. Calculation of full earnings and wage capacity earnings given the number of hrs worked. H0: workers cannot change contractual hours & H0': if min wage <=w(p), no decrease in employment. Impact on disposable equivalized income.	The payment of household income capacity (wage capacity earnings) reduces poverty, but not sufficiently to eradicate it. Payment of household income capacity (full earnings capacity given contractual hours) reduces working poverty by 39.4%	Theory of dynamic monopsony: due to job search costs, workers will accept wages below their marginal productivity; as wage capacity rates differ substantially, Wage Councils orientation is preferable to a national minimum wage. EC: no estimate of antipoverty effect, hence **not included in meta-analysis.**
Burkhauser, Sabia (2007), USA federal & state minimum wage, 1988-2003	Regression model at state level: ln(pvty rate) is regressed on ln(min wage), ln(% of adult males unemployed) and ln(mean wage rate), some specifications w/ year and state fixed effects, and Prais-Winsten GLS to deal w/ autocorrelation and heteroskedasticity across states. **Total population / single mothers.** Simulation of 96-97 increase $4.25 to 5.15 and proposed increase $5.15 to 7.25, H0: no employment effects and no decrease in hours worked (overall + single mothers), sample of workers aged 16-64.	Across specifications there is little evidence of a significant relationship b/w min wage increases and overall state poverty rate; no evidence that min wage decreases poverty rate among single mothers. Regression coefficients have negative sign, but never statistically significant. Simulation: 1997 increase only helped 27.3% of poor workers and only 14.7% of beneficiaries lived in poverty; however, 55.6% of single mothers who benefited from min wage increase lived in poverty; simulation of proposed increase: would help 31.1% of poor workers; only 13.4% of beneficiaries live in poor households (53.4% among single mothers)	(i) In the presence of monopsonies, minimum-wage hikes could have positive employment effects (ii) increase from $4.25 to $5.15: 60.6% of benefits went to workers in households with income-to-needs ratio greater than 2.
Morgan, Kickham (2001), USA, Effective minimum wage 1987-1996	Regression model: **child poverty** rate regressed on minimum wage, average EITC and food stamps benefits, indicators of child support effort and collection, and controls (# of births to unmarried women, single-parent families, % of African Americans, average pay, unemployment rate), w/ and w/o state dummies.	Minimum wage has a significant negative impact on child poverty rate. If real minimum wage decreases by an inflation-adjusted $1, child poverty rate increases by 2.51 % points	The minimum wage's employment effects and its impact on poverty are less than commonly believed, the poverty-fighting potential has more support

Gundersen, Ziliak (2004), USA, effective minimum wage ; 1981-2000	Regression model, log pre-tax and post-tax Foster-Greer-Thorbecke indicator w/ α=0 regressed on ln(effective min wage), a lagged term (poverty in previous year), other policy variables (dummy before and after introduction of TANF, log(food stamps+AFDC), log effective EITC), macroeconomic indicators (unemployment, employment growth rates, state median wage & wage inequality), aggregate and state-specific fixed time effects, a state-specific time trend →two dependent variables: poverty rate and squared poverty gap (α=2) for **all families, female-headed families, married-couple families, white families, black families.**	The higher the state minimum wage relative to the federal level, the lower the poverty head count, before as well as after tax; also a significantly negative impact on pre-tax squared poverty gap, but no significant impact on post-tax squared poverty gap. A 10% increase in the state min wage lowers pre-tax poverty by only 0.5%; conclusion changed little after tax liabilities netted out and EITC credits added. Also reduced pretax squared poverty gap, and regression coefficient is larger than in poverty equation (-0.036 vs. -0.027)	
Sawhill and Thomas (2001) in Bartik (2004), USA, federal minimum wage increase	Simulation of increase from \$5.15 to \$6.15 (simulation parameters not indicated)	This increase would reduce poverty by 4%	EC: indirect account w/o details hence **not included in meta-analysis**
Neumark, Adams (2003), USA, living wage legislation + minimum wage, 1996-2000	Regression model, impact on workers living in cities w/ living wage ordinance vs. workers in cities w/o such ordinance on poverty, depending on level of living wage. Also impact of minimum wage on probability of being poor. For both policies, lags of 6 and 12 months are tested. For more details, see above. **All families.**	Living wage has significant negative impact on probability of income below poverty line only w/ 12-month lag. Minimum wage has a contemporaneous impact (significant and negative), but no significant lagged effect. A 10% increase in the living wage reduces the probability that a family lives in poverty by 0.0033 to 0.0039, i.e. 1/3 of a percentage pt. A 10% increase (contemporaneous) in min wage decreases probability of income below poverty line by 0.9 to 1.38 percentage points	Disemployment effects appear moderate. There is some evidence of positive employment effects for workers in the higher percentiles of the wage distribution.
Leigh (2007), Australia, Minimum wages 1994-2003, real min wages increased by 9% over the period	Simulation using a range of plausible elasticities; Elasticity of wages b/w 0 and 1, and elasticity of labor demand b/w 0 and -1. Three scenarios: (1,0); (0;-1), (1;-1), for hourly wages and labor demand, estimates are averaged over 50 replications of the simulation. Simulation of the effect of giving minimum-wage workers a 10% pay rise, firing 10% of minimum-wage workers, or both.	Estimated impact on the share below half the median pretax equivalized income; income inequality rises under the three scenarios, i.e. relative poverty. These are extreme cases. Among workers, the correlation b/w hourly wages and family income is 0.43. Simulation results: Status quo, pre-tax poverty rate = 21.7%, scenario (1,0) poverty rate=22.4%, scenario (0,-1) pr=22.0%; scenario (1,-1) pr=22.4%. Assuming an hourly wage elasticity of 1, a min wage rise will lower inequality (Gini of pretax income) if elasticity of labor demand < -0.4; if this elasticity is 0.5, min wage rise will lower inequality if	(i) Stigler wrote in 1946:' the connection b/w hourly wages and the standard of living of the family is… remote and fuzzy.' (ii) the typical minimum-wage worker is likely to live in a middle-income household (iii) Australia

		elasticity of labor demand <-0.2	has relatively high min wage, 58% of mdn vs. 34% in the US, subminimum wages exist under 21 (iv) these estimates do not take account of income support and taxation EC: no estimate of antipoverty effect→**not included in meta-analysis**
Hyslop, Stillman (2005), New Zealand. Large reform of minimum wage affecting youth workers in 2001, period: 1997-2003. Before: 16-19 yrs old min wage = 60% of min wage, after: 18-19 yrs old get adult min wage + youth min wage 80% of adult min wage	Regression model comparing the outcomes (log weekly income) for **16-17 yrs old and 18-19 yrs old** to those of young adults (20-25), before and after the 2001 reform, with age dummies and post-2001 dummy and covariates. For more details on specifications: see above.	W/o age-specific business cycle controls, significant positive impact in 2002; w/ business cycle controls, income impact is insignificant. W/o business cycle controls, for 16-17 yr-olds: significant 16 log points increase in individuals' total weekly income in 2002, and for 18-19 yr-olds about 10 log pts increase in 2002. When business controls are added, no significant effect.	
Neumark, Schweitzer, Wascher (2005), effective minimum wage, 1986-1995	Regression model with non parametric technique, difference-in-difference estimators of the effect of minimum-wage increases on the density at each income-to-needs ratio. Availability of 2 consecutive years for each family. Treatment = in states in which min wage increased b/w years 1 and 2, control = min wage remained constant. Does not rely on the linearity of any relationship. Lagged effects are accounted for. **All families in various income-to-needs ratio categories**.	Contemporaneous effect: share of income-to-needs (i-t-n) ratio b/w 0 & 0.6 decreases, share of i-t-n 0.6 to 1.5 increases and proportion w/ i-t-n b/w 1.5 and 2.7 decreases. Lagged effect: unambiguously increases share w/ i-t-n ratios b/w 0 and 1.3 and reduces the share above 1.3. Total effect: essentially no change for i-t-n below 0.3, a marked increase for i-t-n b/w 0.3 and 1.4 and decrease for i-t-n b/w 1-4 and 3.3. Not significant for i-t-n ratios b/w 0 and 0.5 but significant b/w 0.5 and 1. An increase in the min wage has no effect for income-to-needs ratio below 0.5. In contrast, increase of 0.0079 in the proportion of families w/ i-t-n b/w 0.5 and 1; that is an increase of 0.0083 (0.83 % point) in the share of the population w/ an income-to-needs ratio b/w 0 and 1, i.e.	The share of nearpoor families also increases, but only significant at 10%-level

		increase in poverty; as the poverty rate is 18%, this change represents a 4.6% increase in the poverty rate	
Vedder, Gallaway (2001), USA, federal minimum wage, 1953-1998	Regression model: poverty rate regressed on real min wage, unemployment, and real transfer payments/capita (w/ quadratic term), aggregate data at national level; overall 8 models for 9 cohorts. First-difference approach, which solves many econometric issues. Evidence reviewed based on changes in the poverty rate. Also cross-sectional analysis using state data over period 1996-1999. **Overall poverty rate.**	Rem: in 1999, only 12% of the poor worked full-time. Most evidence suggests no stat significant relationship (127 out of 144 models). First-difference regressions: in all 72 regressions the observed relationship b/w poverty rate and real min wage was not stat different from zero. Same conclusions with another poverty indicator and another price index. Analysis by region (Northeast, South, Midwest, West), no significant effect. Regressions for nonwhite workers: results are the same. Then use of state cross-sectional state-level data, average poverty rate 1996-1998 regressed on number of times state min wage was changed b/w 96 and 99, state unemployment rate, and level of income per capita. No significant impact on poverty. Identical analysis w/ data 1991-1993, same conclusions.	(i) The literature on the poverty effects of the min wage is surprisingly modest compared with that on the employment effects (ii) Joseph Stiglitz wrote "a higher min wage does not seem to be a particularly useful way to help the poor".
Müller, Steiner (2008), Germany, minimum wage, 2008 (2005 income data adjusted with average growth rates)	Simulation of the introduction of a min wage of 7.5 €, using a model accounting for the complexity of the German tax-benefits system: means-tested schemes, exemptions from social security contributions for "mini jobs", joint income taxation for married couples. Poverty line = 50% of median income and Foster-Greer-Thorbecke. Changes in labor supply and demand are deemed negligible, based on Neumark & Wascher (2007), hence no effect simulated. **All workers**.	Overall, the income change would amount to roughly 1.5 billion € a year, about 40% of the total increase in net household income would go to East Germany (around 20% of population). There would be less people affected in the lowest income decile than in 2nd, 3rd, 4th and 5th. Very little effect on the incidence and depth of poverty, decrease in the poverty rate would be weak, from 11,75% to 11,56%, with no change in West Germany and a decrease from 21,25% to 20,28 in East Germany. Results are fairly robust to the definition of the poverty line. The poverty gap remains virtually unchanged.	There has been little research on the question to what extent minimum wages affect the income distribution and may serve as an instrument to reduce poverty. Relationship b/w lower wages and low incomes is rather weak.
Kawaguchi, Mori (2009), Japan, minimum wage, 1982, 1987, 1997, 2002	Descriptive stat on profile of min wage workers: empirical first-difference estimates of the impact of the "fraction below" on employment rates of various groups at the prefecture level allow authors to draw conclusion on efficiency of minimum wage as an antipoverty tool. People working less than 200 days/year on irregular schedules are dropped.	70% of minimum-wage workers are not household heads, and around 50% of min wage workers belong to middle- to high-income families as nonhead of household. More than half of min wage workers are middle-aged women (30-59). 7 categories are defined: male 15-19, male 20-24, female 15-19, female 20-24, elderly women (60+) and 25-59 married females. Fraction below: significant reduction in male teen employment (+10% increase in min wage => employment -	EC: no estimate of the antipoverty impact, hence **not included in meta-analysis**

		2%). For married women 25-59%, a 10% increase in min wage decreases employment by 4 to 8% (also decreases employment for teens and young adults). Overall, a rising minimum wage does not seem to be an efficient policy to alleviate poverty in Japan, because it is not well targeted to the poor and reduces the employment of less-skilled workers.	
Sutherland (2001), UK, national minimum wage (NMW) combined with Working Family Tax Credit (WFTC) and other benefits, 2000/2001	Simulation, poverty=60% of median, using POLIMOD, simulation program of tax and benefits. No employment effects simulated. Counter factual: the policy that would have prevailed had Labour not come to power. Situation in 2000/1 with tax and benefits unchanged since April 1997; poverty rates are higher than those observed. **Overall poverty/ singles/couples w/ children.**	Reform package proposed by government (NMW + WFTC + increases in some benefits) reduced overall poverty by 23%, from 18.6% to 14.4%. Of these 23.0 percentage points, 21.8% attributable to tax-benefit changes and only 1.2% to the NMW. Of families affected by the NMW, only 16% are poor	
Giannarelli, Morton, Wheaton (2007), EITC, USA, 2004	Microsimulation model TRIM3, which contains detailed state-specific modeling of the rules of tax and transfer programs. Employment effects are also simulated using estimates derived from US evaluations. Several variables were imputed: monthly transfers, child care expenses, housing expenses. Simulation based on income definition that is broader than the official definition, but poverty lines set at a level that produces headcount ratios very similar to the official ones. Model simulates an increase in the federal minimum wage to $7.25 per hour (from 2010 to 2003 dollars), b/c new minimum wage set to be implemented in 2010. Assumption: a min wage worker's probability of losing his or her job equals 0.06 times the percentage increase in the min wage, and indirect wage gains for those slightly above and below min wage (spillover effects). **Workers**	Assuming no employment or wage effects, the increase in the minimum wage lifts approximately 200,000 people out of poverty. Assuming the expected job loss and indirect wage gains, poverty falls by 475,000 people.	

Bargain (2009), income support, WFTC and other reforms in the UK , 1998-2001.	Microsimulation and decomposition into three effects i) changes in tax-benefit policy ii) adjustments of tax-benefits monetary parameters according to market income growth iii) changes in market income inequality, by calculating counterfactuals based on the EUROMOD tax-benefit calculator. Poverty line fixed at 60% of median income. During this period, Income Suport was increased, WFTC more generous than its predecessor and introduction of the National Minimum Wage (NMW). **Overall poverty rate.**	Over the period, total poverty decreased 2.4 ppt as measured by the FGT indicator w/ $\propto=0$ (headcount), and by 0.1 ppt with the FGT indiactor with $\propto= 1$ (poverty gap). The severity of poverty FGT ($\propto= 2$) increased by 0.1 ppt. The changes in tax-benefit policy amounts to -2.5/-2.4, -0.6/-0.1 and -0.2/0.1 respectively, which means that these changes more than offset slight increases in market income inequality. More specifically, of the -2.4 ppt overall effect, -0.2 ppt can be attributed to the introduction of the NMW, i.e. 8% of the decrease in poverty attributed to the tax-benefit system (NMW, WFTC, Income Support, and other changes)	
Grogger (2003), USA, EITC, 1979-2000	Based on state-level variation in welfare reform policy. Income and log(income) of **female-headed families** are regressed on the introduction of time limits, a dummy for the introduction of welfare reform, demographic controls and the generosity of the maximum credit of the EITC. The model includes states dummies and state-specific quadratic trends.	Log(minimum wage) has no significant effect on income and log(income) among female-headed families.	
Gerfin, Leu, Brun, Tschöpe (2002), Switzerland, EITC and WFTC (simulated)	Simulation: introduction of a minimum wage set at 3’000 Swiss Francs gross/net, no employment impact simulated. Results are based on families in which **household members together work at least 40 hours a week** (i.e. 1 full-time job at the HH level).	Effect is positive w/ a ½ ppt decrease in poverty among workers who live in a household whose members work at least 40 hours a week in total.	

Table A3: Employment effects of tax credits

Author(s), country. Evaluated program, period	Method, independent variable(s)	Employment effects	Comments
Ellwood (2000), USA, EITC, 1975-1999	First, runs a wage equation for 1998 for women aged 18-44 who worked 26+ weeks, and then can predict a potential 1998 wage for all women 18-44 in each sample from 1975-1978 and defines quartiles → it obviates the need to do regression-corrected estimates (thanks to consistent wage/skill quartiles). Predicts incentives with effective tax rate for median earnings for women who work more than 26 weeks. Then compares change in employment between 1986 and 1999 across quartiles for unmarried women with children and then for married mothers, and checks whether differences are statistically significant. **Mothers 18-44, single or married.**	The difference-in-difference estimated impacts are statistically significant for unmarried mothers, and amount to: . Increase in bottom quartile – increase in highest quartile = +18 percentage points (ppts) . Increase in bottom quartile – increase in 3rd quartile =+13 ppts). For married mothers, . Increase in bottom quartile – increase in 3rd quartile = -3 ppts, not significant . Increase in bottom quartile – increase in 2nd quartile = -5 ppts, significant	
Neumark, Wascher (2001), USA state & federal EITC, 1986-1995	Regression model, impact of changes in EITC credit rate + other policies (minimum wage) on probability of adding an adult worker in year 2 if no worker in year 1 or 1 worker in year 1, and impact on change in hours worked in poor **families with kids.** Unemployment and other controls. Average EITC federal credit rate = 14,8% for families with children, 4,8% for state credit	Both federal and state credit have positive impact on P(add a worker if no worker in year 1), which improves likelihood by 11,2 and 14,1 ppts respectively, but no impact on P(add a worker if one worker in year 1) . For hours worked, significant impact of state credit on both families with or without workers in year 1, positive impact for families w/o worker (+205.39 hours) and negative for families w/ workers (-162.64). Federal EITC has essentially no effect on hours worked.	The absence of effect of federal EITC on hours worked is at odds with the large positive employment effect. Lack of consistency of the results for the federal credit, while state credit has sizable positive effects.
Eissa, Hoynes (2004), USA, EITC, 1985-1997	1. Regression model (probit), difference-in-difference estimates, Labor force participation (LFP) of married couples is regressed on fixed group effect*time effect, post 1992*2+kids fixed effects, time and group effects and individual characteristics. A 2nd model adds year dummies*any child 2. Reduced form labor force participation equation (probit), LFP regressed on net nonwage income, gross wage and tax rate and controls, then parameters estimated are used for a	1. All mothers were 1.5 ppt less likely to be working after the EITC expansions, and mothers with larger families (2+ kids) were additionally 3.6 ppts. Total effect for mothers of 2+ kids is a 5 ppts decrease 2. Simulation: married mothers were 1.1 ppt less likey to work in 1996, while fathers were only 0.2 ppt more likely to work. For married women whose husband's wage is in the lowest decile, probability of employment is reduced by 1.7 ppt, and in the second decile 1.6 ppt. For married men the increase is always less than 0.6 ppt. In the phase-in range, married women have an increase in employment	

	simulation of the employment impact. **Married couples.**	probability (+1.1 ppt), a decrease in flat range (-1.5 ppt) and in the phase-out range (-2.1 ppt). For men, impact is always positive and very small, whatever the region of the tax credit	
Grogger (2003), USA, EITC, 1979-2000	Regression model: Based on state-level variation in welfare reform policy. Employment rate and weeks worked among **female-headed families** are regressed on the introduction of time limits, a dummy for the introduction of welfare reform, demographic controls and the generosity of the maximum credit of the EITC. The model includes states dummies and state-specific quadratic trends.	The maximum credit has a significant impact on both the employment rate and weeks of work. A $1000 increase in the maximum credit leads to a 3.6 ppt increase in employment for single-mothers and to an increase of about 1.2 weeks of work. The effect of the EITC explains 34% of the increase in employment among single mothers and 27% of the increase in weeks worked.	
Trampe (2007), USA, EITC, 1993 expansion	Regression model: Focuses on the phase-out range of the credit, 200 households in the phase-out range in 1993, 1994, and 2006 and calculates the phase-out rate that applies to each household (all households, not only single mothers as is often the case). Hours worked are regressed on the EITC phase-out range, age, gender, number of children, educational level, marital status, and school enrollment of parents.	The phase-out rate has a small but statistically significant negative impact. The 1993 expansion (a 7.2 ppt increase in the phase-out range) caused those in the phase-out range to reduce their hours of work by 2.7 hrs/week if they have 2 kids and by 1.1 hours/week for those with 1 kid.	Hoynes: If women who enter the workforce work less than those already in the labor force, there is an endogeneity problem. In addition, no control for year fixed effects, no controls for macroeconomic trends, and the analysis pools single and married women. EC: very small sample. => **Not included in meta-analysis**
Herbst (2008), USA, EITC, 1986-2005	Regression model: **Single mothers with at least one child 0-18**. Probit: probability of various outcomes (any work, work and no welfare, full-time full-year work) as a function of social policy reforms → federal and state maximum EITC credit, child care subsidies, welfare benefits (+work requirements, sanctions, time limits), state and year fixed effects, state unemployment rate, demographic and human capital controls and time trends. Another specification decomposes the impact across quartiles of unemployment levels (heterogeneous effects model)	EITC has a significant impact on any employment and full-time year round employment. Average effects model: a $1,000 increase in the max credit is expected to increase any employment by 1.1 ppt, but decrease full-time full-year employment by a similar amount. Heterogeneous effects model: same magnitude (+1.1 ppt and -1.1 ppt respectively). The impact is very similar across unemployment quartiles, including for single mothers with a high school diploma or less and for non-white single mothers.	75% of EITC dollars are paid to single-parent families (48% of all claimants)

Noonan, Smith, Corcoran (2007), USA, EITC, 1991-2003	Regression and simulation: **focus on single mothers, differences b/w black and white mothers**. Multilevel logistic regression: Variable “employed in the week prior to the survey” regressed on maximum federal and state EITC, welfare policies (any waiver, TANF, sanctions, benefits) and sociodemographic controls. Simulation: 1991, 2000 and 2003 samples→ thanks to regression coefficients, probability of being employed is calculated; then, each women is assigned the conditions (labor market and welfare policies) she would have faced in 2000, and also the EITC she would have received (for the 1991 sample). Finally, this counterfactual probability of being employed is compared to observed values.	Full regression model: the max EITC benefit has a significant impact on employment. For each $1,000 increase in the maximum EITC, the odds of employment are 9.1% greater. Simulation: between 1991 and 2000, changes in the EITC explain 20-25% of the increase in employment for black single mothers and 23-31% for white single mothers (19-25% for all single mothers). The decrease in employment after 2000 -2% is completely attributable to the increase in unemployment.	Black single mothers are younger, more likely to be never married and high school dropouts, and have more children on average.
Meyer, Rosenbaum (2001), USA, EITC, 1984-1996	Regression with focus on **all single mothers**. Probit employment estimates (worked in reference week), independent variables are: a vector of demographic and economic controls (state, age, race & ethnicity, education, marital status and kids, unemployment, unearned income, central city, month), year dummies and an interaction term year*any children. The coefficient of the latter variable gives the difference-in-difference estimates.	Taxes have a significant impact on employment: a $1,000 reduction in income taxes (or increase in tax credit) increases employment last week by 2.7 ppt, and increases employment last year by 4.5 ppt. The effects are larger for the less educated: for high school dropouts the corresponding increases are +4.2 ppt and +8.8 ppt respectively. For hours worked per year conditioning on positive hours, the policy variables have the same signs, but smaller and less significant effects. Using parameter estimates, the EITC explains 62% of the increase in weekly employment over the 1984-1996 period, but only 27% b/w 1992 and 1996. For annual employment, EITC explains 61% of annual employment.	Single parents received about two thirds of EITC dollars; EITC credits increased fifteenfold.
Blundell (2006), United Kingdom, WFTC, 1996-2003	Regression models and simulation to evaluate the impact on **single mothers under 45**: First a structural model→ discrete choices from a small subset of hours (0, 1-15, 16-22, 23-29, 30-36, 37+) w/ usual sociodemographic and economic controls as well as child care demand as a function of hours worked and fixed costs of work. Then parameters are used for a simulation. 2nd regression model used as a quasi-experimental approach: difference-in-difference estimate of the	Difference-in-difference results: the impact was a 3.5 to 4 ppt increase in single mothers labor supply attributable to the WFTC policy (using 2 different surveys), the response was slightly larger for the lower education group. Significant impact. Simulation: moving from the Family Credit to the WFTC →increase in single mothers’ employment rate +5.95 ppt (about 7.5 ppt for children aged 3 to 10). The same simulation including all reforms directed to single mothers	A pronounced puzzle: the UK policy appears twice as generous as the US policy. Yet the impact looks to be half what is was among similar groups in the US.

	impact of the WFTC on employment (comparison group = single women w/o children)	shows an increase of 3.86 ppt: the contemporaneous increase in Income Support dulled the positive labor supply of the WFTC	
Brewer, Duncan, Shephard, Suarez (2006), UK, WFTC, 1999-2003	Simulation based on parameters of structural equation→ Tobit regression model, sample of **all families with children in Great Britain, but separate regression for single mothers and couples**. Probability of choosing from a subset of working hours (same one as Blundell (2006)) as a function of demographic and other household characteristics. Wage equation also specified.	Simulation shows that replacing Family Credit with WFTC lead to a statistically significant 5.11 ppt rise in the proportion working among lone mothers. Hours worked are estimated to increase by 14%, with average weekly hours worked by those working increasing by 2.7% (0.75 hrs/worker). The non-WFTC reforms reduced the positive employment impact by 1.45 ppt. Among couples with children, there was a slight increase for women whose partner doesn't work (+0.06 ppt), but a decrease amongst women whose partner is in work of -0.64 ppt. The overall decrease: -0.57 ppt. For men in couples, the WFTC increased employment by +0.75 ppt.	
Blundell (2000), UK, WFTC, 1994-1996	Simulation using discrete choice structural labor supply model similar to the one already described above, simulation for samples of **single parents and married couples** (including *de facto* married) with children, excluding self-employed from the 1994-1995 and 1995-1996 family resources survey; simulation allowing child care demand to vary w/ hours worked, fixed cost of work, and stigma associated with welfare.	WFTC simulations show an increase of 2.2% in the number of single parents who work (+34,000), 1.32% for women with a nonworking partner (+11,000), a decrease of 0.57% for women whose partner is in work (-20,000) → an overall increase among mothers +24,000. In addition, an increase for men with a nonworking partner of 0.37% and for men with a working partner of 0.3%. The overall decrease in workless families amounts to 57,000 families.	
Francesconi, Rainer, van der Klaauw (2009), UK, WFTC, 1991-2002 Francesconi, van der Klaauw (2007), WFTC, UK	Regression model, **married and cohabiting couples**. If man works 16+ hours and is in top quartile of earnings distribution → excluded from sample. Quasi-experimental: employment is regressed on dummy experimental vs. control group, a time trend interacted with this dummy, the difference b/w any year and the year of the reform, a vector of individual characteristics and an individual fixed effect. Lone mothers and women in couples.	The overall impact on women in couples with children is not significant, but heterogeneity in responses. Strong effect among women whose partner did not work or worked fewer than 16 hours: increase in employment by 3 ppt, with an increase in full-time employment rate by 2 ppt. Very similar findings in Francesconi and van der Klaauw (2007) for single mothers. For women with a partner who works 16+ hours, no significant effect. No statistically significant difference between low-education sample and the rest of the sample. Among women whose partner works less than 16 hours or not at all, significant impact of WFTC on persistence probability (i.e. staying in employment) as well as entry probability.	

		No significant impact on men whatever their partner's labor force attachment.	
Bargain, Orsini (2005), France, Germany, Finland	Simulation of the introduction of the WFTC in these 3 countries (but extended to childless singles and couples), as well as that of a hypothetical low-wage subsidy (LWA) using the lowest decile of the wage distribution as a threshold, those earning between 1 and 1.4 times this amount getting the subsidy, Simulation uses the 1998 Income Distribution Survey for Finland, the 1998 Socio-Economic Panel for Germany and the 1994 Household Budget Survey for France. The parameters stem from a structural model, a discrete-choice logit model similar to that developed by Blundell, including disposable income, the costs of work, a vector of socio-demographic characteristics. Wages are predicted for non-participants. **Lone mothers and married women.**	WFTC single women: 1.8% of single women in Germany and in Finland enter the labor force, while the effect is smaller in France (0.51%). 80% of these movers are single mothers. As for women in couples, 4.3% in France, 1.43% in Germany and 1.17% in Finland leave their job. Almost all negative responses concern women with children. Overall, the disincentive effect for married women prevails so that the net effect on employment is negative. LWS: The impact on single women is lower, between half and 2/3 of what is found with the WTC. For married women the impact is positive: 3.1% in France, 0.99% in Germany and 0.34% in Finland →% of the sample who enters the labor market.	Contrary to the US and the UK, the overall effect is negative. Why? A smaller share is in the phase-out range in the UK and the US. In addition, these countries have a wider wage distribution and lower level of taxation. Moreover, the poverty rate of lone mothers is lower in Continental Europe and Scandinavia; however, it is as high in Germany as in the UK. More fundamentally, the employment level of single mothers is higher in Continental Europe and even higher in Scandinavia.
Scarth, Tang (2008), Canada, Working Income Tax Benefit	Simulation, based on a nine-equation system, based on some important assumptions: globalization constraint (if government raises taxes, owners of capital can relocate their factors of production, there is involuntary unemployment among the unskilled, government budget constraint –how the program is financed, the rich receive 2/3 of income, the WITC benefits 10% of workers, all poor have same benefits – although there are a phase-in and a phase-out range, a the take-up rate is 100%. Options: an income tax credit financed by higher tax on "rich" population or a cut in spending (upper half) or by a cut in spending (both amount to 1 ppt), allowing for changes in labor force participation or not. **Poorest decile of the population.**	Poorest decile of the population who work at minimum-wage levels of remuneration, allowing for changes in the labor force participation: Increase in taxes: the unemployment rate decreases by approximately 0.17 ppt Decrease in spending: the unemployment rate decreases by approximately 0.2 ppt.	

Gregg, Harkness, Smith (2009), UK, WFTC, 1998 and 2003, and employment rates 1993-2003 (labor force survey panel)	Detailed evidence on entry and exit rates, dynamic approach. Difference-in-difference approach w/ 2 control groups 1) women w/ kids in couples 2) single childless women. Probit regression model: probability of being employed regressed on an interaction lone mother*post-reform, a lone mother dummy, a post-reform dummy, and controls (age, education, age of child, ethnicity, region, and interaction terms). Dynamic aspects→ probit model probability of job loss after break-up for partnered women in year t-1 (become single mother in year t), conditioning on education, age, and other differences. Another probit model (difference-in-difference) for entry/exit with an interaction lone mother*post-reform, a lone-mother dummy, and same controls as 1st model described above, w/ and w/o poor health control. Fixed-effects regression for weekly hours, similar controls, with two groups: < 16 hrs/week and >= 16 hrs/week. **Lone parents.**	Lone mothers, employment: compared to single childless women: probability increased by +5.2% (significant), compared to mothers in couples +3.8%(significant) Lone parents, employment: compared to single childless adults +4.1%(significant), compared to parents in couples + 3.8%(significant) Dynamic model of exit rates: in 1993-1999, becoming a lone mother increased the exit probability by a stat significant 9.5% and in 1999-2003 the impact was insignificant. Regarding the probability of entering the labor market, when single mothers are compared to childless single women, the difference is significant (+4.8%), however when a poor health control is included the result is insignificant. There is no difference when the comparison is made w/ mothers in couple. Regarding exit probabilities: significant w/o poor health control (-3.7%), not significant when this control is included. No difference with mothers in couples. Hours worked (mothers in couple for comparison): + 3.06 hours among part-timers but -1.33 hours among full-timers	The increase in generosity of out-of-work benefits reduced some of the WFTC's pro-employment effects. The increase in lone mothers' employment has come largely from a sharp increase in the share of mothers becoming lone parents holding on to work at the point of transition into lone motherhood.
Haan, Myck (2007), Germany, introduction of UK's WTC and CTC, 2005	Simulation w/ 2003 data, based on discrete choice labor supply estimation, similar model as Bargain and Orsini (2006), extended to women *and* men, and uses 2005 parameters to account for the Hartz reforms. The structure and generosity of the simulation are based on the 2005 system in the UK: working tax credit (WTC) and child tax credit (CTC) introduced in 2003, childless individuals becoming eligible (contrary to WFTC). H_0: income from tax credits is included in the means test for income support, which is withdrawn at a rate of 100%. Use of the STSM microsimulation model of the German tax and benefit system	Labor supply, restricted to households where both spouses are aged b/w 25 and 59 not in education and not self-employed: the overall employment of single women increases by more than 95,000 or +2.9%, almost exclusively borne by lone mothers. The effects on single men are modest, namely +10,000 (+0.3%). Total employment among women in couples decreases by more than 55,000, i.e. -0.8%; for men in couples the effect is also negative but smaller, namely -13,000 or -0.2%. The negative impact is highly concentrated among two-earner couples (-53,800 among women and -29,900 among men), while there is an increase among workless households (+8,500 and +26,100 respectively). Put differently, many no-earner and two-earner couples become one-earner couples.	These estimations call for a high degree of caution as far as 'importing' UK-style tax credits to Germany is concerned. A solution could come in the form of an individual tax credit combine w/ addressing the problem of supply shortages of childcare places.

	Lone parents and couples.	The overall effect is slightly positive (+35,000), contrary to Bargain and Orsini's simulation findings for Germany.	
Shannon (2009), Canada, changes in various provinces, 1976-2001	Article follows Meyer and Rosenbaum (2001), with probit employment equations, the probability of employment being regressed on the effect of interaction term having children* year, with the following controls: age, education, province, marital status, unemployment rate, age of children. Comparison of **single mothers** w/ unattached women (=in single-person households). Policy effects are measured by 1) classifying provinces by their degree of aggressiveness 2) calculate an aggressiveness indicator based on cuts in real welfare benefits, presence and level of income support program for the working poor at the province level. Ontario and Alberta are the most aggressive reformers, Newfoundland and New Brunswick the least aggressive. Regression model, employment probits, with employment regressed on province classified by degree of aggressiveness, maximum welfare benefit available, full-time earnings at minimum wage level, and a "welfare wall" variable including income supplements for the working poor and their tax treatment. Some specifications include aggressiveness and welfare generosity dummies. When possible, the whole sample, otherwise the 1989-2001 sample. Simulation using the coefficients with welfare pre-reform values and unemployment rates at their 1992 value.	The negative impact on employment of having a child diminished markedly after 1994: By 2001, the effect of a child on weekly employment had diminished from 16.6% to 4.7% and that on annual employment from 14.8% to 3.3%. Restricting the sample to younger women (20-34 years) made little difference, as did an analysis of the difference b/w single and nonsingle. In Canada, unattached women are also eligible to collect welfare, contrary to the US. By contrast, the differences by skill level are striking, the effect being, as in the US, strongest for the least educated. Changes in income support policies do not explain much of the rise in lone mothers' employment in Canada, increased by 3.1-3.5 ppt. Policy changes explain at most 10-20% of the rise in Canadian lone mother's employment during the 1990s.	Canadian welfare program changes went less far than those in the US. However, in the UK, in contrast to Canada, welfare benefits to nonworking lone mothers rose significantly. Yet, the UK reforms appear to explain a larger share of the increase of lone mothers' employment. Maybe the use of other datasets and techniques will uncover more significant results. EC: does not really provide an estimate of employment effect, hence **excluded from meta-analysis.**
Stancanelli (2008), France, "employment premium" (PPE)	Difference-in-difference approach: difference b/w the employment probabilities of women in the treatment group and that of women in the control group. 3 different "treatments": 1) potential eligibility (conditional on earnings and household income if married) 2) Comparison of married and cohabiting women 3) Comparison of lone mothers	Treatment 1: eligibility →No significant effect on overall female employment. For married women, negative and significant impact (-3 ppts) in fixed-effects model, but insignificant in random-effect model. For cohabiting women, the effect is significant at the 10% level and positive (+6 to 6.9 ppts). The effect is not significant for all single women.	This is the 1st evaluation of the employment effects of the French tax credit using non-experimental methods.

	and single childless women. Logit of employment dummy regressed on treatment dummy, fixed group and year effects, and a vector of controls (experience, education, number of children, region, nationality, wage rate) , with robust standard errors→autocorrelation). A random-effect model is also specified (term c_i added for individual unobserved effects). **Women: Married, cohabiting, single.**	Treatment 2: marital status → significant negative impact (-3 to 3.5 ppts) for married women compared to cohabiting women. Treatment 3: lone parents vs. single childless women → the impact is not significant. Maybe due to the fact that childless singles are also eligible for the tax credit. Net impact on total female employment is very small, about 2,000 jobs.	
Bloemen, Stancanelli (2007), France, "employment premium" (PPE), 1999-2002	Estimation of the employment effects, accounting for potential endogeneity: Eligibility depends on earnings and so does the employment decision in theory, wage rates and employment are potentially correlated → this may introduce a bias in the usual difference-in-difference estimates. First, a wage equation is estimated, so that the probability distribution of eligibility can be obtained: log earnings are regressed on individual characteristics, and then a probit specification is used to estimate P(eligibility). Then sophisticated regression models are specified that estimate the conditional probability of employment whether the person is eligible or not, and the probability of non-employment, whatever the eligibility. Hence, P(employment) is regressed on eligibility, eligibility*policy year, and controls, with a specific joint distribution of residuals. Robust standard errors are estimated to control for serial correlation. The policy year is 2002. **Sample of women: single, married and cohabiting**. Self-employed are dropped, retired women and full-time students as well. Those who have a retired or self-employed husband too.	W/o controls: overall significant (at the 10% level) negative impact on employment. Insignificant effect on married women, cohabiting women and single women. When controls are included, w/ and w/o corrections for potential endogeneity, all effects are insignificant. Hence, there is no evidence of any positive effect of the tax credit on employment. Regarding working hours, the effect isn't significant either. The PPE did not affect women's working hours.	

Table A4: Antipoverty effects of tax credits

Author(s), country. Evaluated program, period	**Method, independent variable(s)**	**Antipoverty effects**	**Comments**
Gundersen, Ziliak (2004), USA, state and federal EITC, 1981-2000	Regression model, impact of macroeconomic factors and policy factors (TANF, waivers, food stamps, EITC and minimum wage) on logarithm of pre-tax and post-tax Foster-Greer-Thorbecke (FGT) indicator + squared poverty gap (alpha=2), for various groups (**all families, female-headed families, married-couple families, white families, black families**), with time and state fixed effect and state-specific time trends + macroeconomic and policy controls (described above). A lag is included in the specification, namely FGT poverty in year t-1 EITC variable = ln(state-federal EITC)	EITC increases pre-tax poverty and squared poverty gap, no significant impact on post-tax poverty and squared poverty gap; however, the 1990s trend-break variables (after 1990, after 1992, after 1995) have stat significant and higher coefficients in post-tax models suggesting a positive role of the EITC in eradicating post-tax poverty in the 1990s	
Morgan, Kickham (2001), USA, EITC, 1987-1996	Regression model, pooled times series (OLS with state fixed effects), impact of rise of the maximum tax credit eligibility (beginning of phase-out range) on **child poverty** rate.	In specifications only with significant state dummies, the EITC variable has a stat significant impact. As the threshold for maximum tax credit eligibility rises by $1000, child poverty declines by 0.18 percentage points.	
Neumark, Wascher (2001), USA state & federal EITC, 1986-1995	Regression model, impact of on **pre-tax** poverty among poor families with kids, probability of being nonpoor in year 2 if poor in year 1, and impact of income-to-needs ratio (official poverty line) of changes in EITC credit rate + other policies (minimum wage). Unemployment and other controls. Average EITC federal credit rate = 14,8% for families with children, 4,8% for state credit. **Poor families.**	An average change in state credit rate (0.04) has a significant impact on P(nonpoor in year 2/poor in year 1) which is increased by 7 ppts, and on change in income-to-needs ratio (+0.076) for families with children, federal EITC has no significant effect. For families with children and no adult worker in year 1, the state EITC significantly increases P(nonpoor in year 2) by 10.9 ppts, but no effect on families w/ kids who already have a worker. The same pattern applies for income-to-needs ratio (significant +0.1 if no worker, no increase otherwise). Federal EITC has no effect.	i) Probability of increasing **earned** income → impact on pretax poverty. The increase in total resources would be more pronounced if one considered the additional income received from the credit itself ii) It is not clear why the incentives posed by the federal and state tax systems should differ.
Grogger (2003), USA, EITC, 1979-2000	Based on state-level variation in welfare reform policy. Income and log(income) of **female-headed families** are regressed on the introduction of time limits, a dummy for the introduction of	Surprisingly, the results from income and log(income) regressions suggest that the EITC has no net effect on income.	The absence of a significant effect on income may be due by the offsetting effects of a decrease in welfare use and

	welfare reform, demographic controls and the generosity of the maximum credit of the EITC. The model includes states dummies and state-specific quadratic trends.		an increase in work and earnings. However, the author underlines that EITC income is not reported at all in the March Current Population Survey
Keegan Eamon, Wu, Zhang (2009), USA, EITC 1996-1998 and 2003-2005	Measure used is disposable family income obtained by adding the EITC and near cash government benefits to the family's income and subtracting federal, state and payroll taxes. Authors carried out simple calculations comparing the poverty rate before and after the EITC benefit is added to the disposable family income measure for 2004 and 2005. For previous years, they used figures from other authors' articles based on the same database (Current Population Survey) and similar indicators. Child poverty.	Reduction in child poverty rate thanks to the EITC: 1996 -14.5% for all children (-30% for children with working parents), 1997 -15.6% (-27.2%), 1998 -17.9% (-27.7%), 2003 -19% (n/a), 2004 -18.3% (n/a) and 2005 -19.5% (n/a)	"the method might overstate the poverty reduction effectiveness…First…the EITC can serve as a disincentive for employed parents to work more hours…or for married women to enter the work force. Second… earning more income increases taxes and decreases eligibility for or the amount of other means-tested benefits" (p.924) EC: simple pretax/transfer – posttax/transfer comparison, hence **not included in meta-analysis.**
Bargain, Orsini (2005), France, Germany, Finland	Same simulation as described above: Ignoring behavioral responses, there is obviously a reduction in the poverty rate, but also when behavioral responses are accounted for. Poverty line set at 50% of median equivalized income (also 40% and 60% for comparisons). **Overall poverty** rate.	For the WTC: France: poverty rate decreases from 7.03% to 6.35%, Germany 5.65%→5.41 and Finland 3.75 →3.71%, when behavioral responses are included in the simulation. The impact of the LWS is surprisingly similar to that of the WTC, even though smaller amounts are distributed to many more working families, including those in high income brackets with 6.45% (France), 5.5% (Germany) and 3.66% Finland	

Gerfin, Leu, Brun, Tschöpe (2002), Switzerland, EITC and WFTC (simulated)	Simulation: the EITC and WFTC parameters are adjusted using purchasing power parities. A structural model allows predictions of family labor market participation, following Blundell's discrete-choice model described above. Results are based on **families in which household members together work at least 40 hours a week** (i.e. 1 full-time job at the HH level).	The EITC has no impact, while there is a very slight increase (+0.1 ppt) in the working poor rate with the WFTC, because some households decide to work full-time but remain poor nonetheless.	
Scarth, Tang (2008), Canada, Working Income Tax Benefit	Simulation, based on a nine-equation system, based on some important assumptions: globalization constraint (if government raises taxes, owners of capital can relocate their factors of production, there is involuntary unemployment among the unskilled, government budget constraint –how the program is financed, the rich receive 2/3 of income, the WITC benefits 10% of workers, all poor have same benefits – although there are a phase-in and a phase-out range, a the take-up rate is 100%. Two options are tested: an income tax credit or an individual wage subsidy as the one proposed by Phelps; both programs can be either financed by a 1 ppt increase in the tax rate of by a 1 ppt cut in other spending. **Poorest decile of the population.**	Poorest decile of the population who work at minimum-wage levels of remuneration: Increase in taxes, no change in labor force participation: average income goes up by 3.9%, with change in labor force participation by 4.3% Decrease in spending, no change in labor force participation: average income increases by 8.1%, with changes in labor force participation by 8.6%	EC: antipoverty impact depends on the average income gap (in terms of incidence)
Haan, Myck (2007), Germany, introduction of UK's WTC and CTC, 2005	Simulation w/ 2003 data, based on discrete choice labor supply estimation, similar model as Bargain and Orsini (2006), extended to women *and* men, and uses 2005 parameters to account for the Hartz reforms. The structure and generosity of the simulation are based on the 2005 system in the UK: working tax credit (WTC) and child tax credit (CTC) introduced in 2003, childless individuals becoming eligible (contrary to WFTC). H_0: income from tax credits is included in the means test for income support, which is withdrawn at a rate of 100%.	Distributional impacts: families in the 2nd decile (+ €52.10/week i.e. +4%) and in the 3rd decile (+€ 60 i.e. +3.8%) would gain the most, while families in lowest decile would gain € 25.80 i.e. +3.4%	EC: antipoverty impact depends on the average income gap (in terms of incidence). Behavioral impact not accounted for, even though authors show the employment effect is negative for couples. **Not included in meta-analysis.**

Giannarelli, Morton, Wheaton (2007), US, EITC, 2004	Microsimulation model TRIM3, which contains detailed state-specific modeling of the rules of tax and transfer programs. Employment effects are also simulated using estimates derived from US evaluations. Several variables were imputed: monthly transfers, child care expenses, housing expenses. Simulation based on income definition that is broader than the official definition, but poverty lines set at a level that produces headcount ratios very similar to the official ones. **Workers.** 3 types of EITC expansions simulated. For childless workers, phase-in rate of 20% instead of 7.65%, and phase-out rate 0f 16%. Further, extension to childless workers aged 18-24 who aren't full-time students. Moreover, the provision for married couples excludes ½ of the earnings of a lower-earning spouse if it would result in larger EITC, and, third, for families w/ 3+ kids, phase-in rate=45% and phase-out rate=23.69%.	Assuming no employment effects, the packet of EITC changes reduces poverty by approximately 2 million individuals. Assuming higher employment among childless workers, poverty would decline by 2.2 million individuals.	
Bargain (2009), UK, income support, WFTC and other reforms , 1998-2001.	Microsimulation and decomposition into three effects i) changes in tax-benefit policy ii) adjustments of tax-benefits monetary parameters according to market income growth iii) changes in market income inequality, by calculating counterfactuals based on the EUROMOD tax-benefit calculator. Poverty line fixed at 60% of median income. During this period, Income Suport was increased, WFTC more generous than its predecessor and introduction of the National Minimum Wage (NMW). **Overall poverty rate.**	Over the period, total poverty decreased 2.4 ppt as measured by the FGT indicator w/ $\propto$=0 (headcount), and by 0.1 ppt with the FGT indiactor with $\propto = 1$ (poverty gap). The severity of poverty FGT ($\propto = 2$) increased by 0.1 ppt. The changes in tax-benefit policy accounts for -2.5/-2.4, -0.6/-0.1 and -0.2/0.1 respectively, which means that these changes more than offset slight increases in market income inequality. More specifically, of the -2.4 ppt overall effect, -0.8 ppt can be attributed to the change from Family Credit to the WFTC, i.e. 1/3 of the changes in total poverty attributable to the tax-benefit system(NMW, WFTC, Income Support, and other changes).	

Bargain, Terraz (2003), France, "employment premium" (PPE), 2002	Microsimulation based on SYSIFF98, used for static simulations of the tax-benefit system, i.e. employment effects aren't accounted for, based on Household Budget Survey 2002. Impact of the reform of the initial PPE introduced in 2001 w/ increased credit for part-time workers (in force from 2003 onwards, hereafter named "PPE Raffarin", named after the Prime Minister who was in charge at that time). It is a 4.4% credit rate w/ an increased credit for part-time workers (corresponding to approx. 6.6%). Two poverty lines: 50% and 60% of median income. **Overall poverty rate.**	PPE Raffarin has an extremely weak impact on poverty despite the increased credit for part-time workers: the poverty rate decreases from 6.51% to 6.47% (poverty line=50% of median) resp. from 12.97% to 12.94% (poverty line=60% of median).	EC: Accounting for employment effects wouldn't change much, as these effects are usually estimated to be very weak.

Table A5: Employment effects of family cash benefits

Author(s), country. Evaluated program, period	**Method, independent variable(s)**	**Employment effects**	**Comments**
Del Boca, Pasqua, Pronzato (2008), Italy, Spain, France, Belgium, the Netherland, 1999 Family allowances.	Use the large variations across EU countries in terms of CC slots and opening hours. Bivariate probit model, i.e. the probability of choosing to work and to have children are modeled jointly. The probability of working and of having children are regressed on the woman's age (and its square), educational level, other hh income, the age of the youngest child, part-time jobs' availability and quality, CC availability, family allowances, length of optional leave, and cluster dummies (pro-natalist, pro-traditional, pro-egalitarian, on-interventionist). **HH w/ women aged 21-45, married or living w/ a partner**	Family allowances have a statistically significant negative impact on the probability to be in work. Regressions are carried out separately for women w/ tertiary education and women w/ less than a tertiary education. Family allowances only have a statistically significant negative employment impact for women w/ less than tertiary education. For women w/ tertiary education, the impact is also negative but much smaller and insignificant.	

Sánchez-Mangas, Sánchez-Marcos (2008), Spain, 1996-2004. Introduction of €100 monthly cash benefit for working mothers of children under three.	Introduction in Spain in 2003 of a monthly cash benefit of €100 per child under three for working mothers. Difference-in-difference-in-differences approach (DDD), based on a probit model which regresses the likelihood of labor market participation on after-the-reform dummy, a treatment dummy (mothers of kids under 3 as opposed to the comparison group: mothers of kids aged 3-6), an interaction after*treatment, a time trend, an interaction time*treatment, educational attainment, age and age squared, number of kids, employment status in previous year, and employment status of spouse. Fertility is assumed to be exogenous. **Married women under 45.**	The policy variable, namely after*treatment, has a statistically significant and positive effect on employment. The estimated change that results from this reform is a 2.93 ppts increase in labor market participation, which represents approximately 5% of the labor market participation of the treatment group in 2002. Estimates by educational level are the following: the reform has a positive effect on each educational level; however, it is only significant for women with an intermediate educational level (as opposed to those with a primary or tertiary level).	EC: employment-conditional benefit
Berninger (2009), 21 European countries, 2004-2005. Family cash benefits	Situation in Belgium, Denmark, Germany, Finland, France, Greece, UK, Ireland, Iceland, Italy, Luxembourg, the Netherlands, Norway, Austria, Poland, Portugal, Sweden, Slovakia, Spain, Czech Republic, and Hungary. Multilevel regression models: maternal employment rates regressed on individual characteristics (age, age squared, youngest kid's age, number of kids, educational level, marital status, personal representation of mother's role) and macro characteristics (supply of CC for kids under 3, CC supply for kids aged 3-5, family cash benefits as a % of GDP, number of weeks of maternal leave, employment of childless women, and national representation of mothers' role). The representation of the maternal role is measured w/ the question: "A woman should be prepared to down on her paid work for the sake of her family" **Mothers of children under 16, aged 25-60**	In the model that includes all individual and macro characteristics, income transfers have a negative but insignificant effect on maternal employment. Indeed, the log odds amount to 0.977, indicating a very small decrease in the odds of being in employment.	The insignificant effect of cash benefits could be due to lack of differentiation of the operationalization, which includes universal child benefits, benefits that provide incentives to stay at home, and employment-conditional benefits.

Milligan, Stabile (2007), Canada, 1996-2000. Introduction of National Child Benefit.	1998: introduction of the National Child Benefit (NCB), w/ national benefit, also for nonworking parents + provincial benefit that is employment-conditional. Some provinces subtract NCB from welfare payments ("clawback states") while others don't. Clawback states increase incentives to work through child benefit. The provincially-run earned income supplements provide more incentive to join the labor force. Regression: linear probability model or OLS, both w/ instrument variable b/c of the endogeneity of benefits. NCB depends on income, which in turn depends on earnings and other income sources. **Welfare recipients, mothers aged 18-50, married women are excluded.**	The interaction term Clawback state*NCB has a significant positive impact on single mothers' probability to have positive earnings, and also for all single women. This interaction term has a positive but insignificant impact on the number of weeks worked; however, its impact on hours worked is statistically significant and positive for single mothers. For all single women, however, the impact on both variables is insignificant.	EC: employment-conditional benefits
Naz (2004), Norway, 1998-1999. Cash-for-care benefit.	In 1998 the Norwegian government introduced cash benefits up to approx. €400/month for parents of 1-to-3 years old kids who don't utilize state-subsidized day-care facilities. Difference-in-difference estimator, treatment = parents of kids under 3, control = parents of kids 3-6. Outcome variables are: market intensity (husband's working hours+ his wife's), specialization (husband's working hours – his wife's), wife's working hours and husband's working hours. 1st specification: each outcome is regressed on child-under-3 dummy, a before/after dummy, and an interaction child under 3*after the reform. 2nd specification adds a tertiary education dummy variable and interaction terms w/ previous dummies. 3rd specification adds a vector of control variables and interactions w/ the same dummies (child under 3, after, child under three*after). **Married and cohabiting couples who have children**	Policy variable is the interaction term Child under 3*after the reform. It has a statistically significant and positive impact on specialization (+3.28 hours), a significantly negative impact on market intensity (-2.42 hours) and on wife's working hours (-2.85 hours), and an insignificant (but positive) impact on husband's working hours (+0.43 hour). Adding the educational level shows that the difference b/w the reform's effects for the two types of hh (tertiary vs. below tertiary educational level) is stat insignificant. If the wife has a university degree, the increase in specialization amounts to 5.06 hours (significant), in other hh by 2.43 hours (insignificant). For market intensity, the decrease amounts to 3.49 for women w/ a tertiary educational level, and -1.9 hours otherwise, but these decreases, as well as the difference, are not significant. Wives' working hours are decreased in both groups, but only significantly so for women w/ a higher educational level. The impact on husbands' working hours is positive but insignificant whatever the educational level.	Surprisingly, bigger impact on wives w/ a university degree. Reason could be that the use of subsidized childcare is higher among higher-educated mothers.

	at least one-year-old		
Brink, Nordblom, Wahlberg (2007), Sweden, 1999. Simulated child benefit reform.	In 2002, a maximum was introduced in Sweden. Simulations are carried out, based on parameters obtained from a sophisticated structural labor supply model. This model allows estimating pre-reform labor supply and disposable income. Then, the maximum fee is applied (3% of gross hh income for the 1st child, 2% for the 2nd and 1% for the 3rd, for family incomes below 38,000 SEK → this fee is applied to all hh). A tax-benefit simulation model from Statistics Sweden. Information about CC is simulated: for single mothers, number of working hours = use of CC; for couples, time in CC = working hours of the parent who works the least. This policy is compared with a theoretical child-benefit increase that gives the same budgetary implications as the maximum fee reform. **Single mothers and couples w/ children born b/w 1994 and 1998.**	Simulated child benefit reform, which amounts to a 5,500 SEK/year increase, i.e. a 61% increase: for single mothers, employment decreases by 1% overall and 6.5% in the lower income quartile, whereas hours of work decrease by 2.4% overall and 5.4% in the lower income quartile. For two-parent families, husband's and wife's employment rate and hours of work remain unchanged.	
Van Damme, Kalmijn, Uunk (2009), 13 countries, Denmark, UK; Belgium, Netherlands, Austria, Germany, France, Ireland, Finland, Italy, Portugal, Spain, Greece, 1994-2001. Family cash benefits	Measuring the impact of family benefits and CC policy on the odds of after-separation employment. Simple and multinomial logit models: Outcomes are regressed on cash benefits in PPPs as the sum of 3 allowances (basic welfare + single-parent allowance + child allowance), on the number of public CC slots/100 kids under 3, macro-level controls (female unemployment rate, incidence of part-time work, and gender role values derived from a scale), and individual variables (married before separation, ex partners' income quartile, living w/ adult family e.g. mother's parents, education, a dummy for repartnered mothers, duration of inactivity for mothers who didn't work before break-up, dummies kids 0-6 and kids 7-	<u>Women who didn't work before separation:</u> The net monthly allowance for single-parents has a negative and significant impact in the baseline model, in the model controlling the impact of gender role values and in the model where the impact of CC is interacted w/ dummies kids 0-6 and kid 7-15. It is negative and insignificant in the model w/ the interaction term net allowance*1st quartile. Baseline: an increase of 100 PPP in allowances reduces the entry odds by 9%. <u>Women who worked before separation (3 outcomes: increase in working hours, decrease, and exit):</u> The net monthly allowance has a positive but insignificant	

	15, dummies for # of years after separation. The allowance variable is interacted w/ 1st income quartile of ex partner and CC variable interacted w/ dummies kid 0-6 and kid 7-15. **Women aged 18-60 at the time of separation who experienced a separation during the panel period**	impact on the odds of exit (compared to stability, the reference category), i.e. negative employment impact, but nonsignificant.	
Cho (2006), Korea, 1998-2003 Child allowances	Complex simulation process: first, a wage equation is calculated, and then a non-maternal income equation. Then, a certain number of parameters are estimated from the dataset, based on an initial guess. The estimated values are compared to existing values (employment, hours of work, labor force participation within 5 years since birth, income share of expenditure on kids, number of kids, age at 1st birth), and if necessary, parameters are adjusted. **Women aged 20-40 w/ at least 1 child**	The introduction of child allowances would decrease labor force participation by 5.4% (from 57.6% to 54.5%). For women within 6 years since birth, employment is reduced by 13,9% (from 34.6% to 29.8%).	Interesting, b/c only one type of family policy exists in Korea, namely CC subsidies
Jaeger (2010), 10 countries, 1995-2000: Australia, Canada, Czech Republic, Finland, Germany, New Zealand, Norway, Spain, Sweden, and the US. Family cash benefits	Probit regression model: probability of employment is regressed on family cash benefits as % of GDP, benefits in kind as a % of GDP, a dummy variable strong religious ties (based on 2 questions), and 2 interaction terms (cash benefits*religious ties and in-kind benefits*religious ties), and a set of sociodemographic control variables (age, family status, # of kids, education, living area, denomination, social class, chief-earner dummy, living w/ parents, female unemployment rate, GDP/capita, and growth rate). **Mothers aged 25-40 or 25-54.**	Model w/o interaction terms: family cash benefits have a negative but insignificant impact on labor force participation, whereas the impact is stat significant and negative for full-time employment. The same conclusions apply to the models that include interaction terms.	

Table A6: Antipoverty effects of family cash benefits

Author(s), country. Evaluated program, period	Method, independent variable(s)	Antipoverty effect	Comments
Milligan, Stabile (2007), Canada, 1996-2000. Introduction of National Child Benefit.	1998: introduction of the National Child Benefit (NCB), w/ national benefit, also for nonworking parents + provincial benefit that is employment-conditional. Some provinces subtract NCB from welfare payments ("clawback states") while others don't. Clawback states increase incentives to work through child benefit. The provincially-run earned income supplements provide more incentive to join the labor force. Regression: linear probability model or OLS, both w/ instrument variable b/c of the endogeneity of benefits. NCB depends on income, which in turn depends on earnings and other income sources. **Welfare recipients, mothers aged 18-50, married women are excluded.**	The interaction term Clawback state*NCB has a positive, yet insignificant effect on single mothers' total income. On the contrary, it has a statistically significant and positive effect on total income for all single women.	EC: employment-conditional benefits
Matsanganis, Levy, Mercader-Prats, Toso, O'Donoghue, Coromaldi, Farinha Rodrigues, Tsakoglu (2005) Spain, Portugal, Greece, Italy, 1995-1996 Family cash benefits.	Static simulation, does not account for behavioral responses; as it appears that means-tested or universal benefits tend to reduce maternal labor force participation, the estimates are overoptimistic. Microsimulation based on EUROMOD, tax and benefit simulator for 1998. Data 1995-1996, adjusted for 1998. Impact of family allowances and non-refundable tax credits. **Child poverty**	The impact on the child pvty rate / pvty gap : Greece: -8.1% (-1.5 ppt) / -11.4% Italy: -19.0%(-6.2 ppts) / -28.2% Spain: -7.7% (-1.8 ppts) / -12.1 Portugal: -20.9% (-6.1 ppts)/ -36.7% The impact on Foster-Greer-Thorbecke w/ α=2 : Greece: -13.2% Italy: -30.8%	The most striking finding is that the overall value of family transfers in southern Europe is extremely low. **EC: doesn't account for behavioral responses, not included in vote count.**

		Spain: -14.0% Portugal: -44.0%	
Brink, Nordblom, Wahlberg (2007), Sweden, 1999. Maximum fees reform.	In 2002, a maximum was introduced in Sweden. Simulations are carried out, based on parameters obtained from a sophisticated structural labor supply model. This model allows estimating pre-reform labor supply and disposable income. Then, the maximum fee is applied (3% of gross hh income for the 1st child, 2% for the 2nd and 1% for the 3rd, for family incomes below 38,000 SEK → this fee is applied to all hh). A tax-benefit simulation model from Statistics Sweden. Information about CC are simulated: for single mothers, number of working hours = use of CC; for couples, time in CC = working hours of the parent who works the least. This policy is compared with a theoretical child-benefit increase that gives the same budgetary implications as the maximum fee reform. **Single mothers and couples w/ children born b/w 1994 and 1998.**	Simulated child benefit reform, which amounts to a 5,500 SEK/year increase, i.e. a 61% increase, increases disposable income by 4.6% overall, and by 4.3% in the lower quartile for single mothers. For two-parent hh, disposable income increases by 1.5% overall and 2.3% in the lower quartile. Distributional effects: the reforms decrease inequality (Gini coefficient) by 3.4% and P90/P10 by 1.22%.	
Misra, Moller, Budig (2007), 11 countries: Belgium, Finland, France, Germany, Luxembourg, Netherlands, Norway, Sweden, Canada, UK, USA. Mid-1990s → early 2000s. Family benefits.	Logistic regressions, w/ robust estimator (Huber-White for heteroskedasticity). Probability of poverty is regressed on family benefits (% of social insurance), on the % of 1-2-year olds in formal CC, paid leave and family leave (including family leave^2) and controls: age, part-time and full-time employment, education, partnered or not, parent or not, partnered*parent. **Women aged 25-59**	W/o control for paid and family leaves, family benefits have a significant and positive antipoverty effect (poverty reduced by 1.9%). Same result when paid leave is included (-1.2%). When family leave and its square are entered, CC availability has a positive but insignificant effect on poverty.	Misra, Moller, Budig (2007), 11 countries: Belgium, Finland, France, Germany, Luxembourg, Netherlands, Norway, Sweden, Canada, UK, USA. Mid-1990s → early 2000s.

Bäckman, Ferrarini (2009), 21 countries, 2000: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Estonia, Finland, France, Germany, Hungary, Ireland, Italy, The Netherland, Norway, Poland, Slovenia, Sweden, Switzerland, UK, US child allowances, lump-sum maternity grants, tax deductions	Multilevel, random intercept regression model: odds that children are poor are regressed on "traditional" policies (child allowances, tax deductions, lump sum grants), dual-earner policies (parental insurance transfers), age of family head, labor market attachment i.e. the # of earners, a female-head dummy variable, and interaction terms. In some specifications, the public CC coverage of kids aged 0-3 is added. Moreover, the coefficients are estimated w/ the whole sample or w/ the sample w/o postcommunist countries (and w/o Denmark which displays very specific patterns). **HH with pre-school children**	Whole sample: The "traditional" benefits variable has a positive but insignificant antipoverty effect; controlling for the number of earners, the gender of the hh head, and interaction terms doesn't affect these conclusions. The odds ratio in the 1st model is 0.19; surprisingly enough, this large change in the odds is not stat significant. When the public CC coverage for kids aged 0-3 is added, the impact becomes stat significant and remains positive (odds ratio=0.26). When the sample excludes postcommunist countries (Czech Republic, Estonia, Hungary, Poland and Slovenia), the conclusions are the same: traditional programs only have a significant impact when public CC coverage is included in the model.	
Frick (2007), 15 countries, 1994-1998: Italy, Greece, Spain, Portugal, UK, Ireland, Luxembourg, Netherlands, France, Germany, Austria, Denmark, Finland, Poland and Hungary.	First, usual pretransfer and posttransfer comparison to calculate a "poverty reduction effect" due to the receipt of family-related transfers (FRT) expressed in %. Poverty is measured w/ the Foster-Greer-Thorbecke indicator w/ $\propto$=2 (severity of poverty). The "poverty reduction effect" is regressed on family benefits as a % of GDP, length of maternity leave, female employment, country dummies, welfare-regime dummies, and on socioeconomic characteristics: educational level within hh, # of siblings, age structure of hh, health status of adults, single-parenthood dummy, immigrant status dummy, unemployment dummy, and an inactivity dummy. FRT = sum of family related cash benefits, social assistance and housing allowance. **Children identified as poor in a fictitious world w/o FRT**	Family cash benefits as a % of GDP has a stat significant and positive effect on poverty reduction in the model that includes country dummies as well as in the model that contains welfare-regime dummies. W/ the Corporatist regime as a reference category, the social-democratic dummy has a positive and significant impact, *ceteris paribus*, whereas the Southern regime and the liberal regime have a negative and significant impact on poverty severity reduction. The negative impact is much stronger for the Southern dummy than for the liberal.	

Table A7: Employment effects of childcare (CC) availability and cost

Author(s), country. Evaluated program, period	Method, independent variable(s)	Employment effects	Comments
Del Boca, Pasqua, Pronzato (2008), Italy, Spain, France, Belgium, the Netherland, 1999 CC availability.	Use the large variations across EU countries in terms of CC slots and opening hours. Bivariate probit model, i.e. the probability of choosing to work and to have children are modeled jointly. The probability of working and of having children are regressed on the woman's age (and its square), educational level, other hh income, the age of the youngest child, part-time jobs' availability and quality, CC availability, family allowances, length of optional leave, and cluster dummies (pro-natalist, pro-traditional, pro-egalitarian, on-interventionist). **HH w/ womend aged 21-45, married or living w/ a partner**	CC availability (% of children aged 0-2 using CC facilities) has a significant positive effect on the probability of working for the entire sample. Regressions are carried out separately for women w/ tertiary education and women w/ less than a tertiary education. CC availability is significantly positive for both groups of women, but the coefficient is higher for lower-skilled women.	
Kornstad, Thoresen (2006), Norway, 1998/2003. Reform aiming at reducing queues and fees.	The Norwegian parliament has passed a resolution to end queues in childcare centers and to reduce fees. Simulation based on parameters derived from a joint labor supply and childcare choice decision model, from a finite set of jobs and childcare arrangements (3 modes of care and 4 brackets of working hours). A multinomial logit regression model is used: the probability of choosing 1 of the 12 combinations is explained by disposable income, the number of hours worked, number of children, the number of opportunities and of jobs available. The simulation is based on 1998 data projected to 2003; the tax and benefit system in 2003 serves as a baseline. **Married and cohabiting parents with at least one child aged 1-5.**	Increasing the number of places at CC centers up to a point where there are no more queues has a strong impact on female labor supply. The probability of working 38+ hours increases from 31 to 37%, whereas the probability of working b/w 28 and 38 hours increases from 14% to 17% (approx.). The likelihood also increases in others work hours brackets, hence overall increase in employment. In the lowest and 3rd deciles of the income distribution, there is a one-hour increase in mothers weekly working time. The increase is even higher in the 2nd lowest decile (approx. 1.2 hours). The increase is slightly less than 1 hour in the upper two quintiles. Overall, expected hours of work for married or cohabiting mothers of pre-schoolers increase by 4% A 50% reduction in CC fees increases labor supply by about 2	**EC: only 768 observations, hence not included in vote-counting procedure.** Only small number of families has access to informal cares by others, especially grandparents. 70% of kids aged 1-5 attented childcare centers in 2003 in Norway.

		hours a week, i.e. about 8%. When queues are suppressed and fees reduced by 50%, labor supply increases by 13%.	
Sánchez-Mangas, Sánchez-Marcos (2008), Spain, 1996-2004. Introduction of €100 monthly cash benefit for working mothers of children under three.	Introduction in Spain in 2003 of a monthly cash benefit of €100 per child under three for working mothers. Difference-in-difference-in-differences approach (DDD), based on a probit model which regresses the likelihood of labor market participation on after-the-reform dummy, a treatment dummy (mothers of kids under 3 as opposed to the comparison group: mothers of kids aged 3-6), an interaction after*treatment, a time trend, an interaction time*treatment, educational attainment, age and age squared, number of kids, employment status in previous year, and employment status of spouse. Fertility is assumed to be exogenous. **Married women under 45.**	As indicated above, the introduction of a €100 benefit per child under 3 for working mothers had a significantly positive impact on maternal employment, namely around 3 ppts. Since the cash benefit represents 39% of the average child care cost, the implied elasticity of labor market decision of mothers of children under 3 with respect to CC prices is -0.07, assuming that there is no effect of the cash benefit on the equilibrium price of CC services.	
Berninger (2009), 21 European countries, 2004-2005 Belgium, Denmark, Germany, Finland, France, Greece, UK, Ireland, Iceland, Italy, Luxembourg, the Netherlands, Norway, Austria, Poland, Portugal, Sweden, Slovakia, Spain, Czech Republic, and Hungary. CC availability.	Multilevel regression models: maternal employment rates regressed on individual characteristics (age, age squared, youngest kid's age, number of kids, educational level, marital status, personal representation of mother's role) and macro characteristics (supply of CC for kids under 3, CC supply for kids aged 3-5, family cash benefits as a % of GDP, number of weeks of maternal leave, employment of childless women, and national representation of mothers' role). The representation of the maternal role is measured w/ the question: "A woman should be prepared to down on her paid work for the sake of her family". **Mothers of children under 16, aged 25-60**	In the model w/ all individual and macro variables, CC availability for kids under 3 has a significant and positive impact on maternal employment, whereas rather small effect as the odds increase by 3.6 percent → $e(\beta)=1.036$. In the 2 models w/ all individual variables but only significant macro variables, 2 interaction terms are added. Conclusions: CC availability has a significant positive impact and the odds ratio increases by 3 to 4.7 percent. Interaction term CC availability*age of youngest child shows that the negative impact of the age of the youngest child significantly decreases when CC availability increases.	

Baker, Gruber, Milligan (2008), Quebec/Canada, 1994-2003, Québec's new CC policy	Quebec introduced a major family-policy reform: began in 1997 w/ the extension of full-time kindergarten to all 5-year-olds, and the provision of CC at an out-of-pocket price of $5.00 for all children 0-4. Program was phased in, starting w/ the 4-year-olds, than 3-year-olds in '98, all 2-year-olds in '99, and all kids under 2 in 2000. Moreover, the number of spaces doubled b/w '97 and '05. Difference-in-difference: %of mothers in employment is regressed on a policy eligibility dummy (=1 if kid is under 5 and lives in the province of Quebec), province and year dummies, as well as a vector of parents' characteristics, city size, number of siblings, age and sex of child. Panel survey of **children** w/ details on parental situation.	For two-parent families: There is a statistically significant rise in employment of women in Quebec, relative to the rest of Canada, of 7.7 ppts, that is, 14.5% of baseline participation. Including an economic control (male unemployment rate) slightly decreases this coefficient: 7.4 instead of 7.7. Alternative samples: for single mothers, the impact is positive, but insignificant. For mothers in couples w/ kids aged 0-2, the impact is positive and significant (+0.09), as well as for kids aged 3-4 (+0.055). For mothers in couples high school or less, the impact is positive but insignificant, contrary to mother w/ some post-high school education (+0.095).	Before the 1997 reform, CC policies targeted low-income families → the reform brought little gain for the lowest-income families, which explains why impact is much larger for women in couples and w/ higher education level.
Brink, Nordblom, Wahlberg (2007), Sweden, 1999, impact of the maximum fees reform.	In 2002, a maximum was introduced in Sweden. Simulations are carried out, based on parameters obtained from a sophisticated structural labor supply model. This model allows estimating pre-reform labor supply and disposable income. Then, the maximum fee is applied (3% of gross hh income for the 1st child, 2% for the 2nd and 1% for the 3rd, for family incomes below 38,000 SEK → this fee is applied to all hh). A tax-benefit simulation model from Statistics Sweden. Information about CC is simulated: for single mothers, number of working hours = use of CC; for couples, time in CC = working hours of the parent who works the least. This policy is compared with a theoretical child-benefit increase that gives the same budgetary implications as the maximum fee reform. **Single mothers and couples w/ children born b/w 1994 and 1998.**	Single mothers: the maximum fee reform increases average labor force participation by 0.7%, and by 4.6% in the lower income quartile. The increase in working hours amounts to 1.4% overall, and 16.5% in the lower quartile. Two-parent households: the wife's labor market participation increases by 0.4% and working hours by 0.5%, and for the husbands the increases amount to 0.2% and 0% respectively. In the lower income quintile, wife's employment increases by 2.5% and working hours by 3.1%; for husbands, these increases amount to 1.3% and 0.3%.	

Del Boca, Vuri (2007), Italy, 1998 Availability and cost of CC.	Regression: bivariate probit model to jointly estimate the probability of maternal employment and of CC use. Both outcomes are regressed on hourly CC costs, on a dummy variable for regions in which the provision of CC is highest (Emilia Romana, Lomardia and Veneto), and sociodemographic controls (wife's and husband's age and education, dummy grandparent still alive, husband's labor income and hh's nonlabor income, family transfers, # of kids 4-5, # of kids 6-13) and labor market variables (% part-time jobs, unemployment rate). A second model includes an interaction hourly CC costs*dummy(Emilia Romana, Lombardia, Veneto). In addition the estimates are used for a simulation of the absence of rationing and of two subsidies: 100% and 50% of CC costs. **Married adults w/ the youngest kid under 3.**	Simulation: W/ no rationing, a 50% CC subsidy increases employment by 15.5% and a 100% subsidy increases employment by 26.5% (from a baseline of 61.5%). Model I: hourly CC costs have a negative but insignificant effect on P(mother works). Living in in a region in which the provision of CC services is highest significantly increases this probability. Model II, w/ interaction term: hourly CC costs still have a negative but nonsignificant impact; living in one of the 3 "high CC provision" regions has a stronger impact. The interaction term has a stat significant and negative impact on P(mother works), which demonstrates that CC cost do have an impact, but only if CC availability is not heavily constrained.	EC: simulating the absence of rationing and a 100% subsidy is not realistic→ **simulation results not included in vote count.** - An important component: the distance b/w the family's house, the workplaces and the location of CC facilities. - In the North, 15-20% of kids under 3 are in public CC, while in the South only 1-2%.
Bub, McCartney (2004), USA, children born in 1991 followed until 1st grade,	**Families were recruited after the birth of a child in 1991, the follow-up interviews until kid is in 1st grade**, i.e. 6-7 years old. Families recruited in 10 cities across the country. Three measures: 24 months after birth, 36 months and 54 months. Maternal employment hours are regressed (OLS) on hours of childcare, a dummy continuous employment, pre-birth maternal employment, partner employment status, income-to-needs ratio dummy (below or above 2*official poverty line), interactions hours in CC*maternal education and hours in CC*poverty status, and control variables (child gender and ethnicity, maternal partner status and maternal education).	The hours spent in CC at 24 months have a positive and stat significant impact on maternal employment when the kid is in 1st grade, and so do hours in CC at 36 months and at 54 months, w/ coefficients of 0.39, 0.43, and 0.5 respectively. Interaction terms do not have a significant impact, but poverty status*Hours in CC at 54 months. Signs for maternal education are negative, signs for poverty status are positive. So the impact does not vary by maternal educational level nor by hh's income level; however, hours in CC at 54 months have a bigger impact for higher income families.	EC: sample not representative, sample size is small, and attrition is not at random --> The final sample: highly educated overrepresented, partnered women overrepresented, African American underrepresented. **Not included in vote-counting procedure.**
Blau, Tekin (2007), USA, 1999, substantial increase in funding of CC subsidy programs	Welfare reform in 1996 consolidated four existing CC subsidy programs into a single block grant, the Childcare and Development Fund (CCDF), which substantially increased funding. Parents must be	4 specifications: w/ or w/o lagged variables, welfare reciept since January 1997 (dummy) and CC assistance after welfare since January 1997. Either one OLS regression (2nd equation) or 2-stage least square model (both equations) 2SLS --> model	EC: results from models III and IV are disputable. And criticized by the authors themselves. →

	employed, in training, or in school. However, there aren't enough funds approx 12-15% of those eligible are served. **A subsample of 13 oversampled states, all hh headed by an unmarried mothers w/ at least 1 kid under 13.** Pair of linear probability equations: 1) subsidy receipt (dummy) is regressed on family characteristics and policy variables 2) employment (and other outcomes) is regressed on this estimated subsidy dummy, the same hh characteristics and other policy variables. Rationing defined at the county level --> county dummies (i.e. the instrument variables) included in the set of policy variables in the 1st equation, but not in the second. State fixed effects are included in both equations. HH characteristics are: age, age^2, black/white/Hispanic dummies, a good health dummy, 12-15 years of education and education > 15 years, nonwage income, family size and its square, kids 0-5 only and kids 6-12 only --> used in the first equation. Then, estimated likelihood of subsidy receipt included in 2nd equation. **Single mothers of children under 13.**	1= OLS w/o lagged variables, model 2= OLS w/ lagged variables, model 3 = 2SLS model w/o lagged variables, model 4 = 2SLS model w/ lagged variables. CC subsidy receipt has a significant positive impact in all 4 models: likelihood of employment increases by 13 ppts in model I, by 12.5 ppts in model II, by 17 ppts in model III, and by 32.9 ppts in model IV. However, in model III the impact is not significant, b/c standard errors are very large in the 2SLS model. Equations III and IV: the validity of county dummies as instruments for CC subsidy is disputable.	**2SLS not included in vote-counting procedure, only OLS results.**
Fitzpatrick (2010), US, 2000, universal pre-kindergarten in Georgia and Oklahoma.	Georgia and Oklahoma introduced universal prekindergarten for all 4-year-olds→ kids who turned 4 by September 1, 2000. Program take-up is high: 50 to 60% of all 4-year-olds. Identification of eligible and ineligible kids in the 1999-2000 school year, **kids born within 100 days before the cutoff date and 100 days after the reform.** Census data. Outcomes are regressed on the difference b/w date of	The policy variable (eligibility) has a negative but insignificant impact on labor market participation (worked last week & worked during the year prior to the interview) and hours of work per week. The impact on weeks worked is positive but insignificant. The cutoff variable (child born before cutoff date) also has a negative and insignificant impact on labor force participation and hours per week. The impact on weeks worked is negative	Conclusion: labor supply of mothers didn't increase following the introduction of pre-K for all 4-year-olds. First, female labor supply elasticities seem to be declining; moreover, many women in the comparison group may already be receiving

	birth (DOB) and cutoff date (number of days), on a dummy "born before cutoff date", an interaction term number of days*cutoff, and an eligibility term (born in Oklahoma and Georgia before the cutoff date) and a set of demographic controls. For the "number of days" variable and the interaction term "number of days*cutoff", a polynomial is used (cubic). Outcomes pertain to **maternal labor supply**: employment over a year, employment last week (Probit), and hours/weeks of work (OLS).	but insignificant. Generally, it is not possible to reject the assumption that the estimated effects are the same across groups of mothers (single or married, w/ or w/o a younger child). The effect is never significant. One exception: for weeks worked, the "cutoff variable" has a significant negative impact for single mothers who don't have younger kids or for married mothers who do.	subsidies.
Pettit, Hook (2005), 19 countries, mid-1990s: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Hungary, Italy, Luxembourg, Netherlands, Norway, Poland, Russian Federation, Sweden, UK, USA. Public provision of CC.	Multilevel model: For each country, a dummy "woman employed or not" is regressed on individual-level characteristics: age, age^2, marital status, education, # of kids, having a kid under 3, youngest child 4-6. Then, the coefficients of these country-level regressions are used as dependent variables and regressed on national characteristics: Maternity leave weeks, parental leave weeks, parental leave^2, public provision of CC, unemployment, service sector growth, and the share of women in parliament. Both OLS and random-effects regression are carried out. **Women 18-65**	OLS specification: public provision of CC has a positive impact for married women and women w/ children 0-3 and children 4-6. Another specification leads to identical conclusions. The random-effect model leads to similar conclusion, but the impact is only significant for children aged 4-6.	
Lundin, Mörk, Öckert (2008), Sweden, 2001 and 2003, maximum fees reform in Sweden	Analyze the 2002 maximum CC fees reform. Employment dummy and share of full-time conditional on employment are regressed on the price of CC and household characteristics. A 2nd specification includes hh type fixed effects as well as hh type*municipality fixed effects, and the 3rd (preferred by the authors) also further includes a hh type*time effect. **Two-parent hh w/ at least one child aged 1-9.**	Model w/ hh characteristics: The maximum fee has a statistically significant and negative impact on the probability of being in employment. However, in the preferred specifications w/ hh type fixed effects, interacted w/ municipality and time, the impact of maximum fees is negative but insignificant. Regarding the impact on the share of full-time mothers among those already in the labor force, the impact is always negative but insignificant.	Highly subsidized CC already existed in Sweden before the reform, hence the difference w/ previous studies.

Lefebvre, Merrigan, Verstraete (2009), Canada, 1999-2004 Québec's new CC policy	Québec's new CC policy w/ large increase in the number of subsidized spaces and low fees provides a Canadian "natural experiment". A difference-in-difference (DD) model is specified: Various outcomes are regressed on a Québec dummy, a dummy for mothers having benefited from the program, and a sum of year dummies multiplied by the Québec dummy, plus a set of control variables. Both a model w/ equal effects (the coefficients of the policy variable are constrained to be identical every year) and w/ unequal effects are specified. The article focuses on unequal effects. A DD-in-difference (DDD) model is also specified, mothers w/ kids aged 12-17 and no kind under 12 are added, but only as a robustness check. **Mothers w/ at least 1 kid aged 6-11 and no kid under 6,** a group that potentially benefited from the new policy when their kid was under 6.	Labor force participation (LFP): coefficients for 2002, 2003 and 2004 are positive and have a significant effect and LFP increased by 3.4, 6.1, and 5.7 ppts respectively. For women w/ high school diploma or less, the effect is significantly positive and very large: 8.3, 16.1 and 17.3 ppts, but insignificant for mothers w/ a higher educational level. Weeks of work/year: coefficients for 2002, 2003 and 2004 are positive and have a significant effect; +2.07, +2.76 and +3.98 weeks respectively. For mothers w/ HS diploma or less, effect is significantly positive and large: +5.63, 9.29, and 10 weeks respectively. For mother w/ a higher educational level, the impact is insignificant. Hours of work/year: coefficients for 2002, 2003 and 2004 are positive and have a significant effect; +95, +92 and +147 weeks respectively. For mothers w/ a HS diploma or less, effect is significantly positive and large: +114, 266, and 318 hours respectively. For mother w/ a higher educational level, the impact is insignificant. It is noticeable that there is an increasing pattern.	
Kalb, Wang-Sheng Lee (2008), Australia, 2002 Cost and availability of formal and informal CC	Sophisticated microsimulation based on regression models and a tax/benefit simulation program. A simultaneous bivariate tobit model predicts formal CC hours and informal CC cost. The average price of formal CC is available from an external source; hence the cost of formal CC can be predicted. Another bivariate model predicts the demand for both formal and informal CC. The budget constraint is based on a tax and transfer simulator (MITTS). Finally a labor supply model is specified based on wage rates, nonlabor income other than taxes and transfers, number and ages of kids, age and education of each parent. Two scenarios are simulated : a 10% increase in net costs (=cost of formal CC – CC subsidies + cost of informal CC) and a 10% increase	A 10% increase in the net costs of CC: - Lone parents: hours of labor decrease by 1.4% and participation decreases by 1.5% - Mothers in couples: hours of labor decrease by 0.3% and participation by 0.2% - Fathers in couples: both indicators remain unchanged. A 10% increase in gross hourly prices w/ adjustment in demand: - Lone parents: hours of labor decrease by 1.6% and participation decreases by 1.9%	

	in gross hourly prices allowing for adjustment in demand. **Families w/ kids younger than 12 years of age.**	- Mothers in couples: both indicators remain unchanged - Fathers in couples: both indicators remain unchanged.	
Lefebvre, Merrigan (2008), Canada, 1999-2002, Québec's new CC policy	September 1, 1997, the government of Québec implemented a new policy of day-care subsidies. On September 1, 2000, all kids aged less than 59 months were eligible for reduced contribution spaces. Difference-in-difference procedure (DD), comparison group is made up of women w/ kids under 6 in other Canadian provinces. Outcomes are regressed on a Québec dummy, an after-the-reform dummy, a sum of time-specific Québec dummies, and a vector of controls (mother's age (& age squared), years of education (& education squared), dummy mother born abroad, single-mother dummy, number of kids under 5, dummy kid under 6 and dummy kid under 3, and earned income from other sources). **All mothers aged 18-56 w/ at least 1 kid younger than 6.**	<u>For all mothers w/ kid under 5:</u> Participation: all coefficients positive and significant for period 1999-2002, w/ increases by 7.6, 5.3, 8.3, and 8.1 ppts respectively. Annual hours: all coefficients positive, but only significant for 2001 and 2002: + 84, 64, 169, and 231 hours respectively Annual weeks: all coefficients positive and significant for period 1999-2002: + 3.8, 3.29, 5.09, and 5.17 weeks respectively <u>For mothers w/ educational level <= high school</u> Participation: all coefficients positive, however only one significant at 10%-level (1999) Annual hours: all coefficients positive, however only one significant at 10%-level (2002) Annual weeks: all coefficients positive, however only one significant at 10%-level (2002) <u>For mothers w/ educational level> high school</u> Participation: all coefficients positive, all significant but in 2001 Annual hours: all coefficients positive, significant in 2001 and 2002 Annual weeks: all coefficients positive and significant	The price reduction was larger for high and middle-income families, because CC subsidies for low-income families existed prior to the reform.
Uunk, Kalmijn, Muffels (2005), 13 EU countries, 1994-1999, Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, The Netherlands, Portugal, Spain, and the UK. CC	Impact of the availability of CC, but also of the GDP level and of gender values (measured w/ European Values Survey questions). For each year, assess whether a child was born since the prior wave. Working hours are measured 1 year before birth and 2 years after. Multilevel regression model (random intercept model): changes in women's working hours after 1st childbirth are regressed on a CC indicator (number of public spaces per child aged 0-3), GDP in	1st specification w/ the CC variable and individual controls only: a 10% increase in CC availability increases mother's afterbirth labor supply by 2.9 hours (significant increase). Adding the impact of GDP in a 2nd specification leads to the conclusion that the impact is stat significant and positive; it is larger than in the 1st model: +4.72 hours. A 3rd model also includes gender values: the impact of CC availability grows even bigger +5.44%.	GDP has a stat significant and negative impact on mother's employment when controlling for the availability of CC slots. Gender values have an insignificant impact once controlling for CC availability.

availability	US$ at price levels and exchange rates of 1995, scale results (1-4) derived from 2 attitudinal questions on gender roles, and individual control variables (mother's age, education, partner's hours, and hh income/1000 in PPPs). **Women aged 20-40 in married or unmarried cohabitation, focus on 1st childbirth.**		EC: small sample 1004 observations, but OK for vote-counting procedure.
Van Ham, Mulder (2005), The Netherlands, 1998. CC availability	Availability of CC, including geographical factors → depending on distance b/w home and CC facilities. Logistic regression: whether or not in paid employment for more than 12 hours is regressed on the number of institutionalized CC slots/100 kids aged 0-4 within 10 minutes' travel from the residence, the number of jobs by jobs level that can be reached by car in 15 minutes, and sociodemographic controls (age, age^2, education, a more-than-one-kid dummy, a single-mother dummy, additional income, a dummies for immigrant mothers and religious mothers). **Mothers w/ children 0-6**	Having good access to CC has a positive and significant impact on the odds of maternal employment: One extra CC slot per 100 kids increases the odds of a mother being in paid employment by 2.2%.	
Van Damme, Kalmijn, Uunk (2009), 13 countries, Denmark, UK; Belgium, Netherlands, Austria, Germany, France, Ireland, Finland, Italy, Portugal, Spain, Greece, 1994-2001. CC availability	Measuring the impact of family benefits and CC policy on the odds of after-separation employment. Simple and multinomial logit models: Outcomes are regressed on cash benefits in PPPs as the sum of 3 allowances (basic welfare + single-parent allowance + child allowance), on the number of public CC slots/100 kids under 3, macro-level controls (female unemployment rate, incidence of part-time work, and gender role values derived from a scale), and individual variables (married before separation, ex partners' income quartile, living w/ adult family e.g. mother's parents, education, a dummy for repartnered mothers, duration of inactivity for mothers who didn't work before break-up, dummies kids 0-6 and kids 7-15, dummies for # of years after separation. The allowance variable is interacted w/ 1st income quartile of ex partner and CC variable interacted w/ dummies	<u>Women who didn't work before separation:</u>. Public CC provision has a positive and significant impact on entry odds an additional CC place increases the odds by 1.6% in the baseline model, by 1.4% in the model controlling the impact of gender role values, by 0.9% in the model where the impact of CC is interacted w/ dummies kids 0-6 and kid 7-15, and by 1.3% in the model w/ the interaction term net allowance*1st quartile <u>Women who worked before separation (3 outcomes: increase in working hours, decrease, and exit):</u> Public CC provision significantly increases the odds of increased working hours but also of decreased hours, as compared to stability. The odds of exit increase, but the effect	The model for women who were in employment before separation: effects are weaker and less in line w/ expectations than the effects on labor market entry.

	kid 0-6 and kid 7-15. **Women aged 18-60 at the time of separation who experienced a separation during the panel period**.	is insignificant.	
Stähli, Le Goff, Levy, Widmer (2009), Switzerland, 1998-99, CC availability	The reduction in labor activity (never reduced/reduced at some moment/stopped for kids but now active again/has completely stopped working) is regressed on the type of CC (collective day nursery, nanny & day mother, informal, …), on geographical context (metropolitan, small center, periurban, etc.), age, national origin, occupational position (manager, professionals, etc.), # of kids, age of youngest kid, partner's income, feminized job (> 70% women) or masculinized or mixed, lifestyle preferences (opinion questions→ home-centered, work-centered or adaptative), kind of network (sparse, friendship, interfering, etc.). Multinomial logistic regression, the reference category of the dependent variable being "homemaker". **Mothers in married and unmarried couples w/ at least 1 child of any age**.	Having a child in a collective day nursery has a positive and significant impact on the odds of never having reduced labor force participation (rather than leaving the labor market and still being out of it) and of having interrupted employment participation but being back in the labor market. The impact on the odds of having reduced at some moment rather than completely stopped is positive but insignificant. Having a kid minded by a nanny or a "day mother" significantly increases the odds of never having reduced labor force participation or having temporarily reduced it rather than completely leaving the labor market. The impact is positive but insignificant on the odds of having temporarily interrupted labor market participation rather than leaving the labor market.	
Rammohan, Whelan (2007), Australia, 2002, CC cost	Dependent variable: odds of being in full-time employment, part-time employment, and not working. Model is a structural ordered probit regression. First, labor force participation is regressed on age, age^2, experience, experience^2, a dummy partner works regular shifts, education, immigration status, region of residence, number and age of kids. A wage equation is calculated using the same variables. The cost of CC equations are calculated using age, partner works regular shifts, migration status, region, number and age of kids. Then, the odds of labor outcomes are regressed on predicted wage, predicted CC costs, region, and number and age of kids.	The (predicted) cost of CC has a negative but insignificant impact on the odds of being in part-time employment and in full-time employment. Hence, no stat significant impact on maternal employment status. Restricting the sample to mothers of children of preschool age yields similar results.	

	Mothers w/ kids aged less than 15		
Tekin (2007), USA, 1997, CC cost and CC development fund	Structural model: a multinomial logit model and two OLS equations are modeled jointly. The price of CC is regressed on state dummies and a vector of individual characteristics (mother's age, education, nonwage income, race, ethnicity, health, region of residence, presence of kids by age); the logarithm of wage rates for full-time or part-time employment is regressed on state dummies and the same set of socio-demographic characteristics. In the multinomial logit model, the outcomes are regressed on these estimated prices of CC and wage rates and on the same set of individual characteristics, plus a control for labor demand factors (unemployment rate). The outcomes are full-time work w/ paid care, and part-time work w/ paid care; the reference category is no employment and no CC. Based on regression coefficients, a a $1 decrease in CC costs is simulated. **Single mothers w/ kids younger than 13.**	The price of CC has a negative but insignificant impact on part-time employment w/ paid care (vs. no employment & no CC) w/ a CC price elasticity of -0.068, and a negative and significant impact on full-time employment w/ CC (vs. no employment, no CC), w/ a CC price elasticity of -0.139. A simulated decrease in hourly CC price of $1, which corresponds to an annual subsidy of $2,080, leads to an increase in overall employment of 3.7 ppts, i.e. 5.2%.	A higher price of CC is a stronger deterrent to full-time employment than it is for part-time employment

Table A8: Antipoverty effects of childcare (CC) availability and cost

Author(s), country. Evaluated program, period	Method, independent variable(s)	Antipoverty effect	Comments
Kornstad, Thoresen (2006), Norway, 1998/2003. Reform aiming at reducing queues and fees.	The Norwegian parliament has passed a resolution to end queues in childcare centers and to reduce fees. Simulation based on parameters derived from a joint labor supply and childcare choice decision model, from a finite set of jobs and childcare arrangements (3 modes of care and 4 brackets of working hours). A multinomial logit regression model is used: the probability of choosing 1 of the 12 combinations is explained by disposable income, the number of hours worked, number of children, the number of opportunities and of jobs available. The simulation is based on 1998 data projected to 2003; the tax and benefit system in 2003 serves as a baseline. **Married and cohabiting parents with at least one child aged 1-5.**	The change in post-tax equivalent income due to fee reductions of around 50% amounts to approx. 3,000 Norwegian crowns (NOK) in the lowest 3 quintiles and around 4000 NOK in upper two quintiles, i.e. around US$ 415 and $550 respectively in 2003, from a baseline of 326,200 NOK (unequivalized), i.e. an equivalized disposable income of 163,100 given that the mean number of children is 2 and the equivalence scale used is the square root of the number of household members. Overall, an increase of 3,600 NOK, around +2.2% in disposable equivalized income. The impact of increasing the number of spaces at childcare centers on disposable income is much smaller, namely less than 1,000 NOK in the lower 2 quintiles and virtually zero in the other three quintiles, due to the moderate impact on working hours.	**EC: only 768 observations, hence not included in vote-counting procedure.** Only small number of families has access to informal cares by others, especially grandparents. 70% of kids aged 1-5 attended childcare centers in 2003 in Norway.
Brink, Nordblom, Wahlberg (2007), Sweden, 1999. Maximum fees reform.	In 2002, a maximum was introduced in Sweden. Simulations are carried out, based on parameters obtained from a sophisticated structural labor supply model. This model allows estimating pre-reform labor supply and disposable income. Then, the maximum fee is applied (3% of gross hh income for the 1st child, 2% for the 2nd and 1% for the 3rd, for family incomes below 38,000 SEK → this fee is applied to all hh). A tax-benefit simulation model from Statistics Sweden. Information about CC are simulated: for single mothers, number of working hours = use of CC; for couples, time in CC = working hours of the parent who works the least.	Single mothers: Disposable income increases by 3.7% overall, and by 3.2% in the lower income decile. Two-parent families: disposable income increases by 2.7%; in the lower income quartile, it increases by 1.7%. Distributional impact: two-parent families gain more than single mothers and high-income more than low-income families. The maximum fee reform keeps the Gini coefficient constant, but the P90/P10 ratio increases by 1,8% → reform enlarges the gap b/w low and high-income families.	The low cap makes the fee reduction larger for high-income families.

	This policy is compared with a theoretical child-benefit increase that gives the same budgetary implications as the maximum fee reform. **Single mothers and couples w/ children born b/w 1994 and 1998.**		
Fitzpatrick (2010), US, 2000. Universal pre-kindergarten in Georgia and Oklahoma.	Georgia and Oklahoma introduced universal prekindergarten for all 4-year-olds→ kids who turned 4 by September 1, 2000. Program take-up is high: 50 to 60% of all 4-year-olds. Identification of eligible and ineligible kids in the 1999-2000 school year, **kids born within 100 days before the cutoff date and 100 days after the reform.** Census data. Outcomes are regressed on the difference b/w date of birth (DOB) and cutoff date (number of days), on a dummy "born before cutoff date", an interaction term number of days*cutoff, and an eligibility term (born in Oklahoma and Georgia before the cutoff date) and a set of demographic controls. For the "number of days" variable and the interaction term "number of days*cutoff", a polynomial is used (cubic). Outcomes pertain to **maternal labor supply**: employment over a year, employment last week (Probit), and hours/weeks of work (OLS).	The policy variable (eligibility) has a positive but insignificant impact wage and salary. The cutoff variable (child born before cutoff date) has a significantly negative impact (-$1,672). Analyzing the effect across groups of mothers (single or married, w/ or w/o a younger child), one result is significant: The policy variable reduces the wage of single mothers who also have younger children	**EC: impact on wage, not on income, hence not included in vote count**
Misra, Moller, Budig (2007), 11 countries: Belgium, Finland, France, Germany, Luxembourg, Netherlands, Norway, Sweden, Canada, UK, USA. Mid-1990s → early 2000s.	Logistic regressions, w/ robust estimator (Huber-White for heteroskedasticity). Probability of poverty is regressed on family benefits (% of social insurance), on the % of 1-2-year olds in formal CC, paid leave and family leave (including family leave[2]) and controls: age, part-time and full-time employment, education, partnered or not, parent or not, partnered*parent.	W/o control for paid and family leaves, CC availability has a significant and positive antipoverty effect (poverty reduced by 0.5%). Same result when paid leave is included (-0.4%). When family leave and its square are entered, CC availability has a positive but insignificant effect on poverty.	

CC availability	**Women aged 25-59**		
Bäckman, Ferrarini (2009), 21 countries, 2000: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Estonia, Finland, France, Germany, Hungary, Ireland, Italy, The Netherland, Norway, Poland, Slovenia, Sweden, Switzerland, UK, US. Public CC coverage of kids under 3.	Multilevel, random intercept regression model: odds that children are poor are regressed on "traditional" policies (child allowances, tax deductions, lump sum grants), dual-earner policies (parental insurance transfers), age of family head, labor market attachment i.e. the # of earners, a female-head dummy variable, and interaction terms. In some specifications, the public CC coverage of kids aged 0-3 is added. Moreover, the coefficients are estimated w/ the whole sample or w/ the sample w/o postcommunist countries (and w/o Denmark which displays very specific patterns). **HH with pre-school children**	For the whole sample: an increase in public CC coverage has a stat significant positive antipoverty impact; conclusions are not affected if postcommunist countries are not taken into account, nor when both postcommunist countries and Denmark aren't included in the sample.	
Kreyenfeld, Spiess, Wagner (2000), Germany, 1996. CC fees	Day care fees are expressed as a percentage of household equivalized net income, for each income quintile. **Families w/ kids 0-11 years old.**	HH in the bottom quintile spend 4.1% of their income on day care fees, while this share amounts to 3.3 in the 2^{nd} quintile, 3.3 in the 3^{rd}, 3.0 in the 4^{th}, and 2.3 in the 5^{th}, although day care fees increase w/ income. Hence these fees could increase relative poverty as they seem to increase income inequality.	From the 1920s onward, municipalities started funding day care for the working class. However, slots in public funded centers are scarce.