Income, the Earned Income Tax Credit, and Infant Health: A Comment

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Hoynes, Miller and Simon (2015), henceforth HMS, report the national expansion of the Earned Income Tax Credit at point in time is associated with decreases in low birth weight. We eliminate any associations between multiple expansions of the EITC and low birth weight across race and ethnicity with simple controls for trends in parity. We question HMS's difference-in-difference analysis of the 1993 increase in EITC due to differential trends by parity prior to expansion. Because variation in the EITC is based on parity at the national level, we collapse their data by year and parity and find standard errors increase. (JEL H24, I12, I38, J13)

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Hoynes, Miller and Simon (2015, henceforth HMS) use the expansion of the Earned Income Tax Credit (EITC) in 1993 to test the effect of an income transfer to working families on birth outcomes. The EITC is considered a highly successful anti-poverty program with expenditures of over 63 billion dollars and 26 million recipients in 2013 (Nichols and Rothstein 2016). The maximum EITC in 2017 was $510 for workers with no children, $3,400 for workers with one child and $6,318 for workers with two or more children (Tax Policy Center 2017). HMS exploit the relative changes in the differences in tax credits by family size to identify effects of income on infant health.

The question of whether income transfers can improve birth outcomes is fundamental to U.S. social policy. First, preterm birth (< 37 weeks gestation) is the most important predictor of infant mortality. Over two-thirds of all low birth weight births (< 2500 grams) are preterm. Second, there is a clear inverse gradient between adverse birth outcomes and socio-economic status. Third, the long-term effects of low birth weight on adult health appear significant (Almond and Currie 2011).

HMS follow a large literature that evaluates the EITC on primarily employment by using a difference-in-difference (DD) design to compare the birth outcomes of single, less educated women who are having a first child with those having second and higher order births (Eissa and Liebman 1996; Meyer and Rosenbaum 2000; Meyer 2002; Eissa and Hoynes 2004; Eissa and Nichols 2005). The largest increase in the EITC occurred with Omnibus Budget Reconciliation Act of 1993 (OBRA93). HMS find that the change in the generosity of the EITC after 1993 lowers the incidence of low birth weight by 0.35 percentage points over a mean of 10.2 percent or a decline of 9 grams over a mean of 3206 grams by comparing single women with children to those without. HMS report associations between EITC and more prenatal care and less prenatal smoking as evidence of plausible mechanisms. HMS then extend the analysis to include the EITC expansions in 1986, 1990 and 1993. They find a $1000 increase the

maximum available benefit lowers the rate of low birth weight by 0.30 percentage points, a finding consistent with the DD analysis from 1993.

The brief summary of HMS's results does not do justice to their exhaustive data work, the clear presentation of their methods and the ease with which we could replicate their results. HMS provide an excellent review of the EITC, its employment incentives and the data used in their analysis. We also do not comment on the effect of OBRA93 on employment, earnings and other sources of income as presented by HMS.

Our focus is on the effect of the EITC on birth outcomes.[1] Despite HMS's careful analysis, we believe there is insufficient evidence to interpret their findings as causal. HMS analyze national changes in the EITC on birth outcomes. Exposure to the EITC varies only by parity. We show that differences in trends by parity in low birth weight prior to the 1993 expansion of the EITC violate the parallel trend assumption of their difference-in-difference (DD) design. Differential trends in low birth weight by parity are also apparent when we reanalyze HMS's multiple expansions of the EITC. Inclusion of a quadratic trend by parity at the national level, a modest addition to their specification, eliminates any association between multiple expansions of EITC and low birth weight.

Our re-analysis underscores the difficulty of estimating the causal effect of a national policy that lacks variation across time and space. HMS aggregate the census of individual birth certificates into cells by state, year, parity race, age, ethnicity, and education. However, the EITC varied only by year and parity. We demonstrate that thousands of their cells are superfluous. Following Bertrand, Duflo and

[1] The paper by HMS is the first published analysis of the effect of the 1993 expansion of the EITC on infant health. Strully, Renkhop and Xuan (2010) use variation in state EITCs to analyze their association with birth weight. However, by 1993 only 4 states had a refundable tax credit (Maryland, Minnesota, Vermont and Wisconsin) while Rhode Island had a nonrefundable credit (Hotz and Scholz 2001). Moreover, the size of the state tax credits, were roughly one-fifth the magnitude of the federal tax credits and yet, Strully, Renkhop and Xuan (2010) report estimated effects on birthweight that are twice as large as HMS.

Mullainathan (2004) we collapse their data by year and parity and compare changes pre-and-post the 1993 expansion. We replicate their point estimates almost exactly in the more aggregated specification, but standard errors increase substantially. To increase power at this level of aggregation, we use Donald and Lang's (2007) two-step procedure but find little support for the conclusion that the EITC improved infant health. Our results highlight the challenge of identifying clinically small effects from a point-in-time change with observational data.

II. Analyzing Multiple Expansions of the EITC

The EITC expanded in 1986, 1990 and 1993. As Figure 1 shows the number of tax filers for the EITC and the average tax credit received per family grew roughly linearly between 1986 and 1997. To evaluate the impact of all three expansions, HMS aggregate individual-level birth certificate data to cells defined by state, year, parity, maternal education, race, ethnicity and age from 1983 to 1998 (HMS, Appendix B). To account for the multiple expansions HMS estimate the following equation:

$$(2)\ Y_{pjst} = \alpha + \delta Maxcredit_{pt} + \beta X_{st} + \rho_p + (\rho_i * T) + \varphi_j + \lambda_s + \delta_t + \epsilon_{pjst}$$

Y_{pjst} is the rate of low birth by parity (ρ_p), demographic groups (φ_j), state (λ_s) and year (δ_t). X is set of state policies: Medicaid/SCHIP, welfare reform and the state unemployment rate. $Maxcredit_{pt}$ is the maximum tax credit available to eligible filers that varies by parity and year. HMS include an additional term (p_i* T) to control for linear trends by parity. The concern is well-founded. Figure 2 replicates HMS's figures which shows the rate of low birth weight among births to white and black women by parity from 1983 to 1998.[2] The rate of low birth weight is presented as deviations from the rate of low birth weight in 1993. The black line is the rate of low birth weight among first births (parity 1). These women were ineligible for the EITC until a small increase in 1991 and they serve as the comparison

[2] HMS provide these plots in Appendix Figure 1b and 1c of their online Appendix.

group. Their trends are approximately linear. The trends among women of higher parity are non-linear. This is particularly true for black women of parity 3+, the group for whom HMS report the largest and most robust effects.

The importance of correctly adjusting for the time-series pattern in low birth weight becomes apparent in Table 1. In column (1) we replicate HMS's results from the specification that does not include a linear trend in parity. The coefficient for all women indicates that a $1000 increase in the maximum available credit is associated with -0.307 percentage point decline in the rate of low birth weight. The effects for whites are substantially smaller [-0.119, column (4)], while those for blacks are much larger [-0.518, column (7)]. Inclusion of the linear trend term in parity alters the estimates substantively (Table 1, columns 2, 5 and 8). In this specification there is no association for white women whereas the coefficient for black women increases by 160 percent. According to these estimates, a $1000 increase in maximum available credit lowers the rate of low birth weight among black women by a sizeable 1.4 percentage points over a mean of 14.8. We next add an interaction of parity and time squared to equation (1) given the curvilinear relationship to low birth weight. All associations between *Maxcredit* and low birth weight are eliminated (Table 1, columns 3, 6 and 9).

One argument against this approach is that we have overfitted the data. This is a possibility, but the trend in low birth weight among black women of parity 3+ is unique within their sample. It rises rapidly in the late 1980s, peaks around 1990 and then regresses to its 1983 level. The dramatic pattern raises the possibility of confounding from time-varying factors. Another reason to appropriately adjust for trends in low birth weight by parity is because HMS's results are limited to this group of multiparous black women. Why an income transfer would only affect black women and not whites or Hispanics of the same SES and parity is difficult to explain. Lastly, HMS's identifying variation comes from differences in infant health by parity overtime at the national level. The relatively smooth growth in EITC benefits

and participants since 1986 makes identifying small effects of the program from underlying trends even more challenging (see Figure 1). The EITC may have improved infant health among subgroups of participants, but the absence of an association for any group with relatively simple controls for trend should weaken confidence in interpreting HMS's results as causal.

III. Violations of the Parallel Trend Assumption in the DD

HMS primary focus is on the 1993 expansion of the EITC. As with the analysis of the multiple expansions of the EITC, HMS limit their sample to single women with no more than a high school education. They use a difference-in-differences (DD) design and compare changes in infant health three years before and five years after the 1993 expansion for women with no previous live births (parity 1) relative to women having a second or higher order birth (parity 2+). They also provide separate comparisons between women of parity 2 versus parity 1 and parity 3+ versus parity 1. HMS find that the rate of low birth weight falls by 0.354 percentage points or 3.5 percent among women of parity 2+ in this high-impact sample relative to single women of parity 1 in the years after the 1993 EITC expansion relative to three years prior (HMS, Table 2, column 2). The difference in low birth weight of women of parity 2 versus parity 1 is smaller, 0.164 percentage points. These are small effects of questionable clinical relevance.[3]

The DD findings are consistent with those they report of the multiple expansions of the EITC. As with the latter, however, pre-EITC trends in low birth weight by parity threaten the internal validity of their findings. A key assumption of the DD is that trends among the exposed and comparison groups are parallel prior to the intervention. As supporting evidence, HMS use an event-study framework. They estimate regressions similar to those characterized by their equation (2), but allow for differences in low

[3] To provide some perspective on the magnitude of these effects, the EITC is associated with a decline of 6 grams in mean birth weight between parity 2 and parity 1. This is the weight of a standard letter envelop with nothing in it. In contrast, the effect of prenatal smoking on mean birth weight is estimated at between 150 and 250 grams and a doubling of rate of low birth weight.

birth weight by parity in 1991 and 1992 relative to 1993, the reference year (HMS Figure 3, Panels A-B).[4] They limit the pre-period to three years to exclude effect from the EITC expansion in 1990 (HMS p. 187). However, the 1990 expansion was phased in over three years much like the 1993 expansion. In fact, the year-to-year increase in tax filers for the EITC and the average tax credit per family from 1991 to 1993 are similar to those from 1994 to 1996 (see Figure 1).[5] Not only is 1991-1993 a limited period with which to determine pre-existing trends in low birth weight, but the smooth growth in EITC participants and tax credit refunds attenuates the discontinuity in 1993 and jeopardizes the parallel trend assumption in the pre-period.

Trends in low birth weight by parity from 1983-1998 provide a more complete picture of the time-series patterns in the DD analysis. This is the same study period HMS use in their analysis of multiple EITC expansions. Figure 3 shows the differential trends in low birth weight among single women with no more than a high school education by parity. Panel A presents the differential trend in low birth weight between parity 2+ relative to 1; Panel B compares parity 2 to parity 1; Panel C contrasts parity 3+ to parity 1 and Panel D compares parity 3+ to parity 2. The right vertical axis in Panels B-D shows the maximum credit available by parity in 1995 dollars.

In each figure, we plot the estimated coefficients of the interactions between parity and year fixed effects using the same specification as HMS in their Figure 3. As is evident in Panel A, there is a sharp rise in low birth weight among women of parity 2+ relative to parity 1 from 1983 to 1987 after which the series declines continuously to 1998. If we limited the pre-period to 1987-1992, we decisively

[4] The 1993 expansion became effective in 1994. HMS use 1993 as the reference period in the event study specification.

[5] Between 1989 and 1993 the average tax credit per family increased by $464 ($1028-$564) and the number of tax filers grew by 3,421 million. The changes were almost the same between 1993 and 1997. The average tax credit per family increased by $537 ($1567-$1028) and the number tax filers grew by 4,217 million (http://www.taxpolicycenter.org/statistics/eitc-recipients).

reject the null hypothesis of no differential trend in low birth by parity ($F_{6,50}$= 9.07). The same pattern is apparent in Panels B and C comparing changes in low birth weight between parity 2 and parity 1 (Panel B) and parity 3+ versus parity 1. How many time periods prior to an intervention should be included in tests of the parallel trends is unclear. However, by only estimating coefficients from 1991 and 1992 in their event study (HMS, Figure 3A, HMS), HMS neglect considerable differences in the time series patterns of low birth weight by parity in the run up to the 1993 EITC expansion.

HMS also estimate the effect of the 1993 EITC expansion on low birth weight by race and ethnicity. They find small and marginally significant declines in low birth for whites and Hispanics (0.130 percentage points), but larger declines for blacks (0.73 percentage points, HMS Table 3). As in the analysis of all women, we extend the pre-period and test for differences in trends in low birth weight by parity separately for whites, blacks and Hispanics. As we show in the Appendix the more extensive pre-period suggests that the decline reported by HMS after 1993 is a continuation of a trend that began earlier (See Appendix Figures A1-A3). As with the analysis of all women, we reject the null of no pre-trend differences in all cases. We repeat this analysis for other birth outcomes (See Appendix Figures A4-A8). The same pattern persists: visually clear and statistically significant differential trends in birth outcomes between women of parity 1 versus parity 2, parity 2+ and parity 3+ that precede the 1993 expansion of the EITC.

Birth Outcomes: Parity 2 vs Parity 3+

Prior to the 1993 expansion in the EITC, the maximum tax credit available to women with at least one qualifying child (parity 2+) was the same regardless of the number of children. After 1993 women of parity 3+ became eligible for a larger tax credit than those of parity 2 and the difference in available tax credits between 3+ versus parity 1 became even larger. As we show in Figure 3, Panel C, differences in the trend in low birth weight between parity 3+ and parity 1 are substantial and

statistically significant. However, the differences between those of parity 3+ and parity 2 are roughly similar from 1987-1992 (Figure 3, Panel D).[6] HMS find that rates of low birth weight fell 0.340 percentage points more among women of parity 3+ relative to women of parity 2 in the high-impact subsample after 1993. The lack of differential trends from 1987-1992 provides support for HMS's findings. But the association between the EITC and low birth weight among women of parity 2 versus parity 3+ is limited to births to black women. HMS report no association between the 1993 EITC and low birth weight among white or Hispanic women of parity 2 relative to parity 3+ (HMS Table 3, Panel C). These finding match the analysis of multiple EITC expansions. As we show in Table 1, however, the association between expansions in the EITC and low birth weight is lost for black women once we control for trends in low birth weight by parity. We believe the DD results from 1993 expansion are similarly contaminated by pre-intervention trends. In Figure 4, we show the differential trend in low birth weight between parity 3+ versus parity 2 among single black women with no more than a high school education. The rate of low birth weight of women of parity 3+ is rising much faster than those of parity 2 through the 1980s. In 1990, the trend reverses and the rate of low birth weight declines three years prior to the 1993 expansion.[7]

[6] The 1990 EITC instituted a very small differential tax credit between parity 2 and party 3 plus that was phased in prior to the 1993 expansion (See HMS Figure 2).

[7] In this comment, we do not discuss HMS's findings regarding possible mechanisms that might explain their reported effect of the EITC on infant health. HMS suggest that increases in prenatal care and decreases in prenatal smoking associated with EITC may account for their findings. We consider both mechanisms questionable. The relative changes in prenatal care visits and the percent of women who begin care before the third trimester are less than 1 percent, too small to be meaningful. Second, the trends in prenatal care prior to 1993 violate the parallel trend assumption for all except for parity 3+ versus 2 (see Appendix Figures A-9-A-10). Prenatal smoking, by contrast, is strongly related to low birth weight. Yet prenatal smoking has been declining faster among women of parity 2+ than parity 1 since 1987 (see Appendix Figures A11-A14).[7] Finally, smoking needs to be an inferior good among low income women of parity 2+ relative to parity 1 for there to be an inverse relationship between the income tax credit and prenatal smoking. One of the authors (Hoynes) has noted in an analysis of the food stamp program that smoking among low income women could be a normal good. "The increased transfer income could also encourage behaviors that could harm infant health such as smoking and drinking" (Almond, Hoynes and Whitmore Schanzenabach 2011, p. 391.

IV. The Level of Aggregation in the DD

HMS aggregate individual-level birth certificates to cells defined by state, year, parity, maternal education, race, ethnicity and age (HMS, Appendix B). They focus on a high-impact subsample of single women with no more than a high school education (n=47,687). In all their regressions they cluster the standard errors by state.[8] The EITC, however, is a national policy and its effect on low birth weight is identified by variation in parity with at most 4 clusters. This precludes standard procedures to correct for within group serial correlation. However, the EITC affected all groups at the same point in time. This enables us to collapse the time-series into a simple before and after design and mitigate the serial correlation (Bertrand, Duflo and Mullainathan 2004). To do so, we regress the rate of low birth on all the covariates used by HMS less year and parity fixed effects. We then aggregate the residuals into 8 cells that contain the mean rate of low birth weight by 2 periods (1991-1993 and 1994-1998) and 4 levels of parity (1, 2, 3 and 4 plus). We weight the cells by the count of births in each. We regress the rate of low birth weight on dummies for time, parity and an interaction of time and parity with these 8 observations. The procedure guards against Type I errors due to serial correlation. The results are displayed in Panel A of Table 1. The coefficients in column (1) replicate HMS's estimates comparing party 2+ versus parity 1 for all women using their specification and the full sample of 47,687 cells with the standard errors clustered by state. Estimates in column (2) are from the regressions with 8 observations. The point estimates in column (2) are very close to those in column (1), but the standard errors are roughly four times larger. In columns (3) and (4) we contrast the coefficients for parity 2 and

[8] We show that many of these 47,867 cells are unnecessary and artificially inflate the degrees of freedom. For instance, we re-ran their regressions and excluded all state-year variables and fixed effects except those for year, parity and race. We were able to replicate their point estimates almost exactly (results available upon request). This is not surprising given that identification comes exclusively from year-parity interactions. One reason to include state fixed effects is to control for the 5 states that had a state EITC by 1993 (see footnote 1) or to control for time-varying state differences in social safety net policies. HMS explicitly control for changes in welfare reform, Medicaid/SHIP expansions and the unemployment rate at the state level, but as they show in their Table 2, the time-varying state controls have absolutely no effect on their estimates.

parity 3 relative to parity 1 obtained from HMS's full sample and our 8 cells. The point estimates are almost identical but standard errors are much larger. The last two columns are from regressions contrasting the effect of the 1993 EITC on low birth weight between women of parity 3+ versus parity 2. Again we find no association in the more aggregated sample. In panels B and C we show results from the same exercise but for blacks and whites separately. The key takeaway is the reported effect of the EITC on the rate of low birth weight among black women is lost when we aggregate to the national level.[9]

Aggregation by parity pre- and post the 1993 EITC reduces the likelihood of Type I errors but at the expense of statistical power. As an alternative, we use Donald and Lang's (2007) two-step procedure. We regress the first difference-in-difference on a dummy variable that is one for the first period of the post-period and zero otherwise.[10] The results are displayed in Table 3. We estimate regressions for two sample periods (1991-1998 and 1987-1998) and for each period we show estimates from regressions with and without a constant. None of the estimates are statistically significant. The standard errors are smaller in magnitude than those from pre-post aggregation in Table 2 and we have more degrees of freedom. The estimates are consistent with a lack an association between the EITC and low birth weight for both white and black women.

[9] Surprisingly the standard errors in the regressions for white women become smaller and some estimates become statistically significant. We consider these inconsequential given the very small magnitude of the effects and the lack of any association between parity 3+ and parity 2. It should also be noted that where we do still find an effect the experiment is contaminated by visually clear and significant pre-trends (See Figure 3: Panel A, B, and C).

[10] As in the aggregation by parity to the pre-post level, we first regress low birth weight in the full sample (n-47,687) on HMS's full set of covariates less year and parity. We then aggregate the residuals to the year-parity level from 1987 to 1998. We show the results from this aggregation in Appendix Table A-1 to A-3. In column (3) we show the difference in low birth weight between parity 2+ and parity 1 within each year. In column (4) we take the first difference of the differences in column (3). We regress the differences in column (4) on the dummy variable described above. We estimate the same regressions for the differences in column (8) of Appendix Table X for the comparisons between parity 3+ and parity 2.

V. The Etiology of Low Birth Weight and Preterm Birth

We find the lack of an association between the EITC and low birth weight unsurprising given the clinical literature. In a review article in the *New England Journal of Medicine* on the prevention of preterm birth, the authors conclude that, ",.....substantial reductions in preterm birth are unlikely to be achieved (Goldenberg and Rouse 1998, p. 318). Almost 10 years later in an exhaustive review of the literature on preterm birth, the Institute of Medicine (IOM) states in the abstract, "The current methods for diagnosis and treatment of preterm labor are currently based on an inadequate literature, and little is known how preterm birth can be prevented" (Institute of Medicine 2007, p.2). HMS report that the EITC is protective against an increase in preterm birth but warn that gestational age is not well-measured and not reported for some state/year cells (HMS's Table 6). They also report that EITC lowers the percent of births less than 2000 grams by approximately 1 percent. Birth weight is well measured. In 2015, 94 percent of all births less than 2000 grams were preterm.[11] Based on the clinical literature, we question whether a tax refund delivered primarily in February and spent soon thereafter on mostly durables and transportation provides a credible mechanism by which to prevent declines in preterm birth of clinically meaningful magnitudes (Goodman-Bacon and McGranahan 2008). [12]

[11] Authors calculations based on Table 23 in https://www.cdc.gov/nchs/data/nvsr/nvsr66/nvsr66_01.pdf

[12] Despite the lack of interventions to prevent preterm birth, the trends in low birth weight reported by HMS and documented here vary significantly by race and parity. We can only speculate as to the factors driving these trends, but the pattern of low birth weight among black women of higher parity is consistent with the impact of the crack-cocaine epidemic of the 1980s. The use of crack-cocaine spread rapidly after 1983 in large urban areas (Evans, Garthwaite and Moore 2016). The epidemic is associated with huge increases in homicide rates among young, black males in urban areas (Blumstein, Alfred, Frederick Rivara and Richard Rosenfeld 2000). Use of cocaine and its attendant lifestyle also had a profound impact on black women giving birth. Rates of low birth weight among *all black* women in New York City rose 2.6 percentage points between 1984 and 1988 from 10.5 to 13.1 percent, reversing a 21-year decline (Joyce and Racine 1993). In the largest population prevalence study ever undertaken, 29,494 women were tested for perinatal substances at 202 California hospitals in 1992. The percent of women exposed to cocaine at delivery was 13 times greater among black non-Hispanics (7.79 percent) than white, non-Hispanics (0.60 percent) and Hispanics (0.55 percent) (Vega et al. 1993). Hospital-based studies of prenatal cocaine exposure found the average age of users was between 26 and 29 and the average parity was 3 as

VI. Conclusion

We end by praising HMS for trying to test the effect of an exogenous income transfer on infant health. The EITC represents a sizable increase in income among working single women with children. But attempts to identify small potential effects of a national change at a single point in time is exceedingly challenging. Without variation in the timing of the transfer across groups and geographical units, we contend there is too much confounding from other events and policies to identity small effects of the EITC on birth outcomes. The increase in standard errors as we aggregated the data to the level of the intervention most likely reflects the inherent uncertainty in the exercise.

computed by HMS (McCalla et al. 1991; Phibbs, Bateman and Schwartz 1991; Singer et al. 2002). A meta-analysis of 30 studies on the association between prenatal cocaine use and adverse birth outcomes reported that mean birth weight was 491 grams lower and that the odds of low birth weight was 3.4 times greater among women exposed relative to unexposed to cocaine. The authors emphasize that tobacco was a frequent concomitant risk factor (Gouin et al. 2011). Again, we can only conjecture, but HMS's finding that the EITC was limited essentially to black women of higher parity suggests the decline in low birth weight in the early 1990s represents regression to the mean as the epidemic waned.

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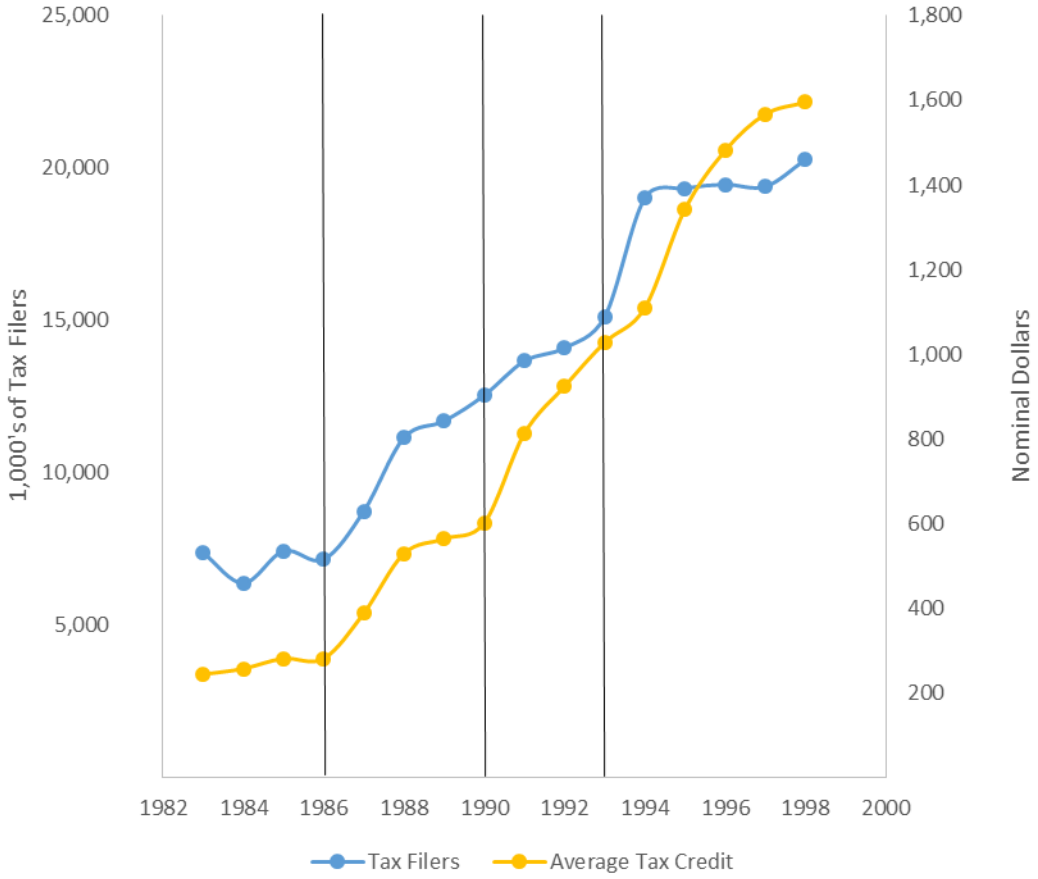
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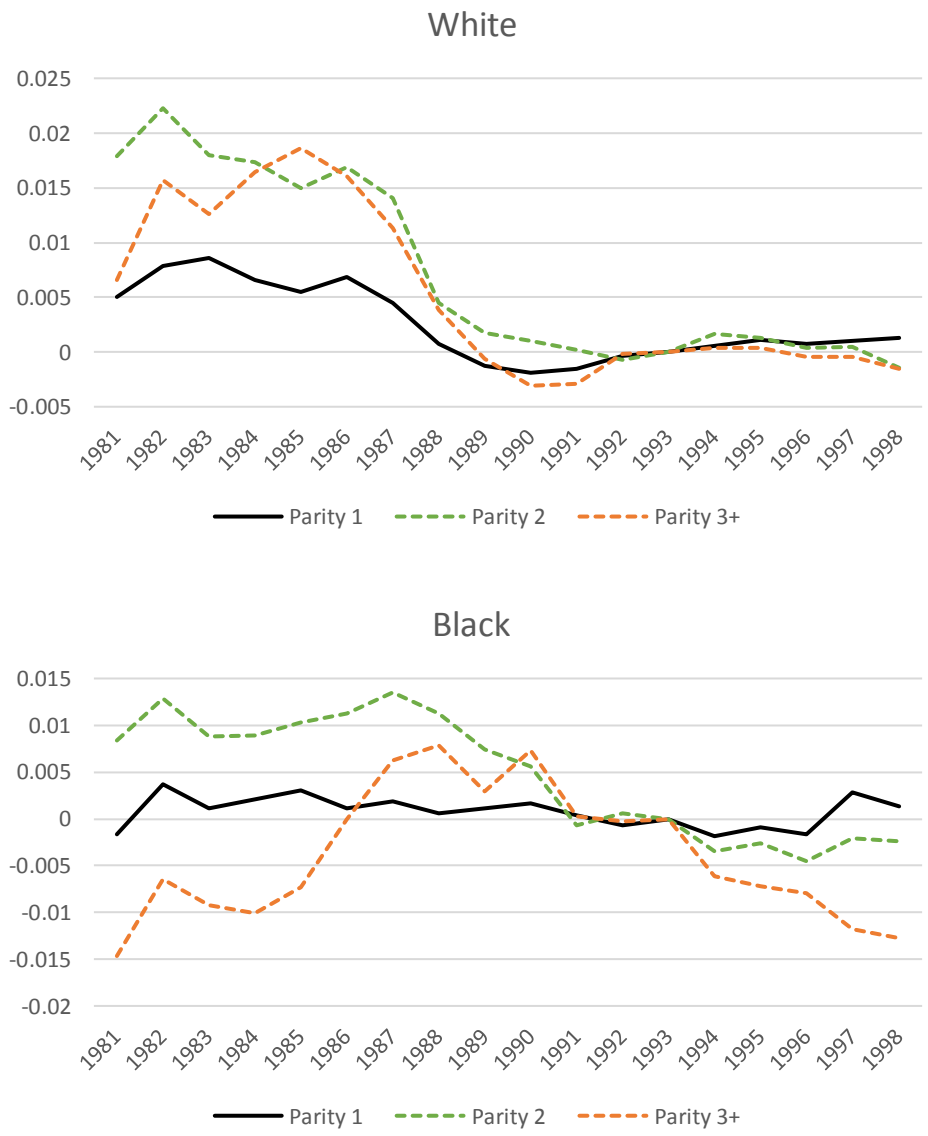
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Figure 1: Number of EITC Tax Filers and Average Tax Credit Received Per Family 1983-1998



Note: Source, 2015: IRS, Statistics of Income Division, Publication 1304, September 2017. Vertical dark lines represent expansions to the EITC.

Figure 2: Trends in Low Birth Weight relative to 1993 by Race and Parity



Notes: Low birth weight estimates refer to deviations as compared to low birth weight in 1993.

Table 1- Maximum Credit Estimates of EITC on Low Birth Weight, Single Women with a HS by Race

Model	All			White			Black		
	(1, HMS)	(2, HMS)	(3)	(4, HMS)	(5, HMS)	(6)	(7, HMS)	(8, HMS)	(9)
Maximum Credit ($1,000 of 95$)	-0.307***	-0.774***	-0.0733	-0.119**	-0.0842	0.0051	-0.518***	-1.359***	-0.0214
	(0.0659)	(0.127)	(0.150)	(0.0528)	(0.107)	(0.159)	(0.115)	(0.179)	(0.220)
Parity X linear time		X	X		X	X		X	X
Parity X quadratic time			X			X			X
Mean of the dependent variable	11.21	11.21	11.21	8.81	8.81	8.81	14.76	14.76	14.76
Observation	81,782	81,782	81,782	37,335	37,335	37,335	23,746	23,746	23,746

Note: Columns (1,2,4,5, 7,8) replicate HMS results (HMS Table 5). All regressions include controls for age, ethnicity, education, interactions of age*ethnicity, age*ed, state FEs, and time FEs. Columns 2,5 and 8) add a linear time trend interacted with paity. Columns (3,6, and 9) add the square of a time trend interacted with parity.

Figure 3: Event Time Estimates of Parity on Low Birth Weight, Single Women with a High School Diploma, 1983-1998

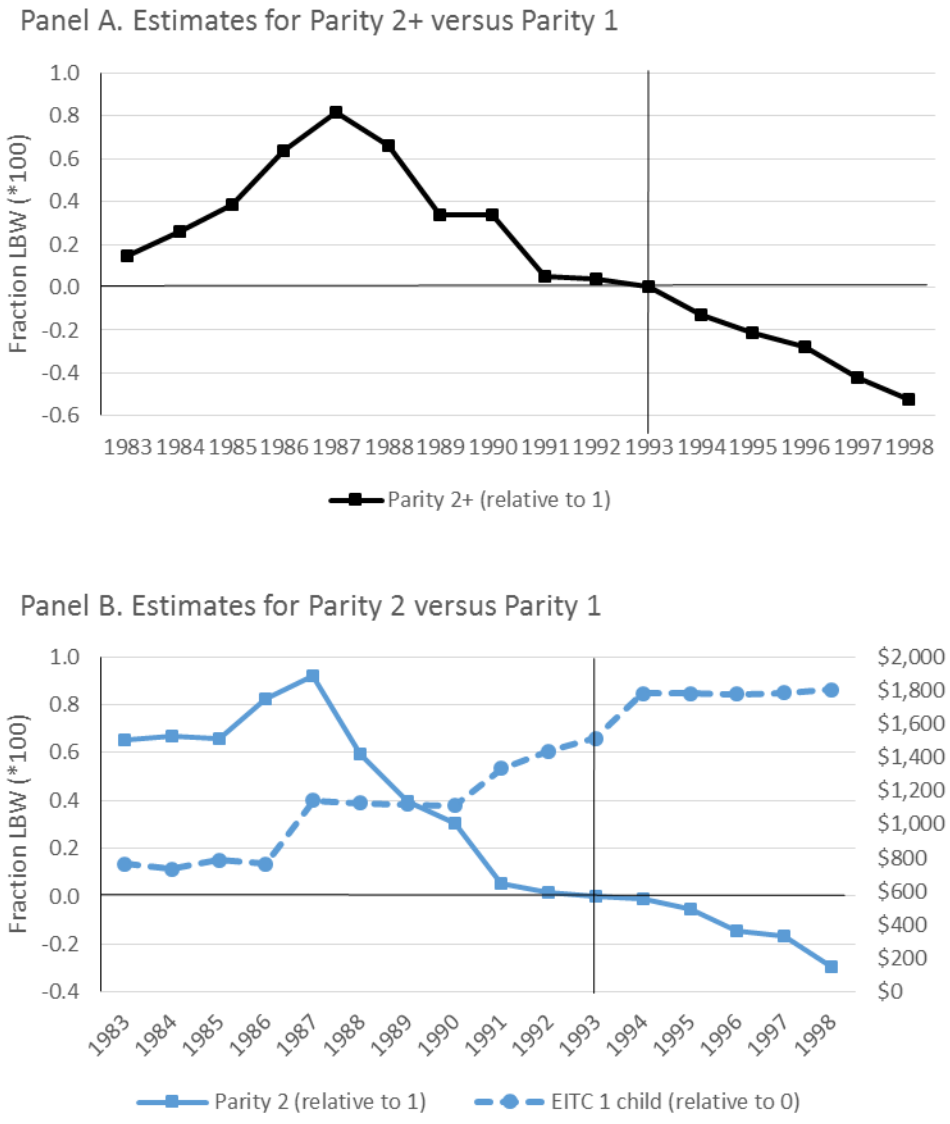
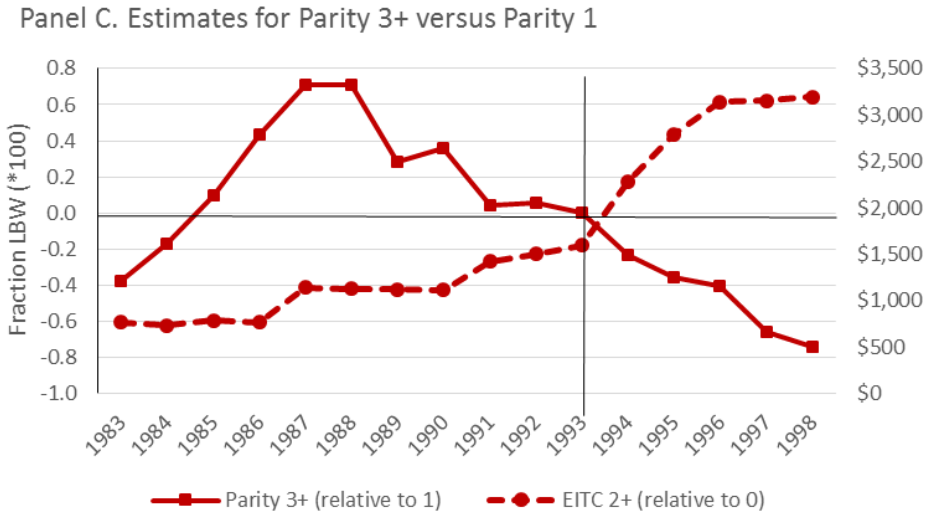
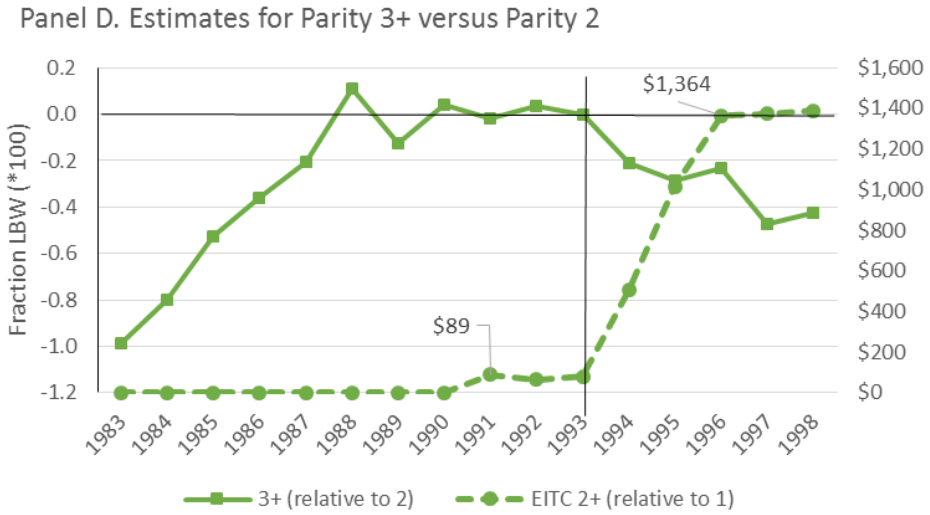


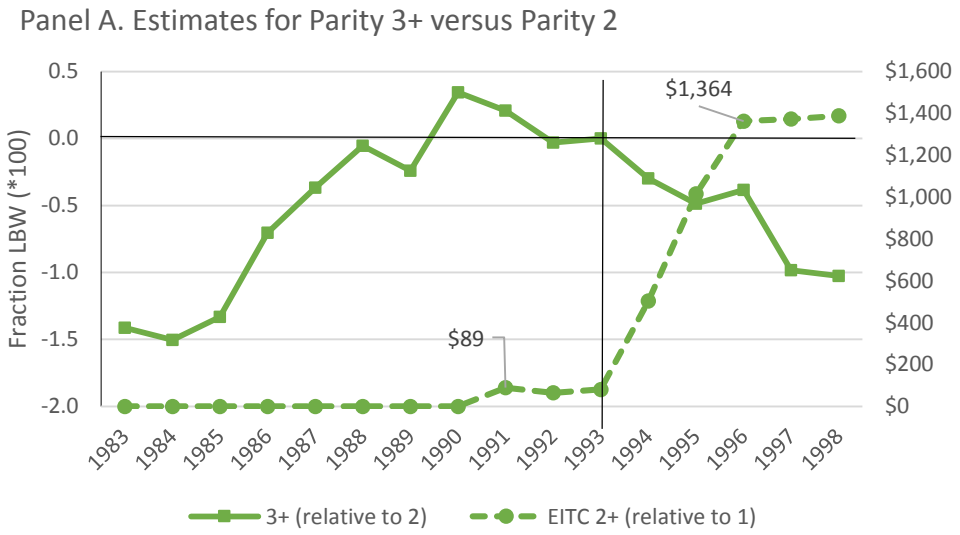
Figure 3, Continued: Event Time Estimates of Parity on Low Birth Weight, Single Women with a High School Diploma, 1983-1998





Note: Similar to HMS Figure 3, each figure plots coefficients from an event-study analysis where each point represents coefficients of time-period and parity interactions. Similar to HMS, the specifications include controls for year, state, parity, demographic group and state-year covariates for Medicaid/SCHIP, welfare reform, and unemployment rates. Panels B, C, and D include relative maximum EITC amounts in each year in 1995 $'s. The p-values from a joint F-test on interactions of corresponding treatment parities with year dummies for 1987-1992 for Panels A, B, C, and D is 0.000, 0.000, 0.000, and 0.065 respectively.

Figure 4: Event Time Estimates of Parity on Low Birth Weight-Black Single Women with a High School Diploma, 1983-1998



Note: Similar to HMS Figure 3, each figure plots coefficients from an event-study analysis for births to Black women where each point represents coefficients of time-period and parity interactions. Similar to HMS, the specifications include controls for year, state, parity, demographic group and state-year covariates for Medicaid/SCHIP, welfare reform, and unemployment rates. Panels B, C, and D include relative maximum EITC amounts in each year in 1995 $'s. The p-values from a joint F-test on interactions of parity 3+ relative to parity 2 with year dummies for 1987-1992 is 0.014.

Table 2- Difference-in-differences estimates of OBRA93 on Low Birth Weight among Single Women with a High School Education or Less

Model	Parity 2+ versus 1		Parity 2, 3+ versus 1		Parity 3+ versus 2	
	(1, HMS)	(2)	(3, HMS)	(4)	(5, HMS)	(6)
Panel A: All Women						
Parity2+ X After	-0.354***	-0.412				
	(0.074)	(0.304)				
Parity2 X After			-0.164**	-0.159		
			(0.072)	(0.293)		
Parity3+ X After			-0.529***	-0.538	-0.342***	-0.362
			(0.091)	(0.254)	(0.069)	(0.249)
Mean of the dependent variable	10.2	10.2	10.2	10.2	10.7	10.7
Observation	47,687	8	47,687	8	35,467	6
Panel B: Black						
Parity2+ X After	-0.728***	-0.798				
	(0.143)	(0.570)				
Parity2 X After			-0.310**	-0.303		
			(0.144)	(0.490)		
Parity3+ X After			-1.040***	-1.05	-0.715***	-0.731
			(0.160)	(0.425)	(0.121)	(0.424)
Mean of the dependent variable	14.43	14.43	14.43	14.43	14.92	14.92
Observation	13,780	8	13,780	8	10,273	6
Panel C: White						
Parity2+ X After	-0.132*	-0.134				
	(0.072)	(0.029)**				
Parity2 X After			-0.114*	-0.112		
			(0.065)	(0.034)		
Parity3+ X After			-0.151	-0.146	-0.0231	-0.034
			(0.093)	(0.029)	(0.071)	(0.029)
Mean of the dependent variable	8.14	8.14	8.14	8.14	8.23	8.23
Observation	21,775	8	21,775	8	16,247	6

Note: Columns (1), (3), and (5) replicate the results from HMS. They include fixed effects for age, ethnicity, education, interactions of age*ethnicity, age*ed, state FEs, and time FEs. Columns (2), (4), and (6) are obtained by first regressing on all of the same controls and then aggregating by parity and before/after the 1993 EITC expansion. Parity2+ includes all women delivery a second or higher order birth. Parity2 refers to women delivery a second birth only and parity 3+ is for women delivering a third or higher order birth.

Table 3- Donald and Lang Procedure, Difference-in-difference estimates of OBRA93 on Low Birth Weight Single Women with a High School Education or Less among all women and by race

Model	2+ versus 1				3+ versus 2			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
All	-0.049	-0.133	-0.018	-0.133	-0.168	-0.213	-0.213	-0.213
	(0.0494)	(0.0939)	(0.109)	(0.151)	(0.127)	(0.116)	(0.191)	(0.173)
White	0.077	0.014	0.106	0.014	-0.216	-0.153	-0.199	-0.153
	(0.0588)	(0.0803)	(0.0747)	(0.114)	(0.167)	(0.154)	(0.207)	(0.192)
Black	-0.198	-0.325	-0.170	-0.325	-0.111	-0.284	-0.250	-0.284
	(0.239)	(0.239)	(0.302)	(0.314)	(0.273)	(0.289)	(0.357)	(0.325)
Constant	X		X		X		X	
Observation	7	7	11	11	7	7	11	11

Note: We first obtain residuals where we control for all controls in HMS Table 2 excluding parity, year, and parityXafter. We collapse down to parities 2+ versus 1 and effective year for columns (1)-(4) and 3+ versus 2 and effective year for columns (5)-(8). We then regress the first difference in the differences between parity 2+ versus 1 and 3+ versus 2 in low birth weight in each year on a single variable that equals 1 in 1994 and 0 otherwise (Donald and Lang, 2007). Results are inferentially no different from setting the treatment variable equal to 1 in each year from 1994-1996 to reflect changes in EITC between those years.