**The Cyclicality of Births and Babies' Health, Revisited: Evidence from Unemployment Insurance**

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This draft: December 19, 2022

We revisit the cyclical nature of birth rates and infant health and investigate to what extent the relationship between aggregate labor market conditions and birth outcomes is mitigated by the consumption smoothing income assistance delivered through unemployment insurance (UI). We introduce a novel empirical test of standard neoclassical models of fertility that directly tests the prediction of opposite-signed income and intertemporal substitution effects of business cycles by examining the interaction of the aggregate unemployment rate with a measure of potential income replacement from UI. Our results show that as UI benefit generosity reaches 100 percent income replacement, there is no effect of the unemployment rate on fertility rates. This implies that the well-documented cyclical nature of fertility rates is about access to liquidity. We also provide novel evidence that infant health is countercyclical based on timing of conception, but procyclical based on time *in utero*. The negative relationship between the *in utero* aggregate unemployment rate and infant health also disappears when potential UI replacement rates reach 100 percent. Our results indicate that the social insurance provided by UI has a pro-natalist effect and improves babies' health.

JEL Codes: J13, I1, J65, E32

*Acknowledgements and Disclaimer:* We thank Julia Hewitt and Eleanor Warren for excellent research assistance. We are grateful to Phil Levine, Elira Kuka, Will Peterman, David Simon, Aaron Sojourn, and seminar participants from the University of Nebraska, University of Connecticut, and the Florida Applied Micro Group for helpful comments. The analysis and conclusions set forth are those of the authors and do not indicate concurrence by other members of the research staff or the Board of Governors.

**Introduction**

Across advanced economies, falling birth rates have led to renewed concerns over declining population growth. In addition, the importance of health at birth to long-term outcomes is now well understood. As such, understanding how families' economic circumstances and government safety net programs affect childbearing decisions and babies' health have become increasingly critical questions for economists and policy makers. In this paper, we address these questions by revisiting one of the most well-established empirical facts in economics and demography: the pro-cyclical nature of births.[1] Specifically, we revisit how births and infant health respond to the business cycle and investigate to what extent these relationships are mitigated by consumption smoothing income assistance, namely, unemployment insurance (UI).

A re-visiting of the cyclical nature of births and infant health advances our understanding of the mechanisms through which families' economic circumstances affect childbearing decisions and infant health outcomes. We find that births and infant health are pro-cyclical, and that these relationships are fundamentally about cyclical changes in liquidity. This highlights the critical role of access to liquidity in family formation decisions, and points to the potential for policies which boost disposable incomes to affect birth outcomes. It also highlights a mechanism through which aggregate economic shocks are propagated to the next

[1] As early as the 1960s, procyclical fertility was described as "one of the most firmly based empirical findings in any of the social sciences" (Thomas, cited in Butz and Ward, 1979). Empirical studies on the topic include, for example, Galbraith and Thomas (1941); Silver (1965); Butz and Ward (1979); Ermisch (1988); Adsera (2005); Currie and Schwandt (2014); Schaller (2016); Dettling and Kearney (2014); Schaller, Fishback, and Marquardt (2020).

generation, that is, via changes in health at birth and cohort size. Our focus on the role of unemployment insurance additionally adds to our collective understanding of how social insurance mitigates potential effects of economic downturns on individuals. We find that when UI fully replaces lost incomes, births and infant health are no longer cyclical. This implies that UI is pro-natalist and improves the health and economic well-being of the next generation.

Since Becker (1960), economists have modelled couples' child-bearing decisions in a neo-classical decision-making framework, where potential parents decide on the optimal number of children that maximizes lifetime utility, subject to the budget constraint that they face.[2] Such a model generates testable predictions that have been borne out in empirical work: namely, that there is an increase in births associated with an increase in family income and with a decrease in the costs associated with having children.[3] The effect of exogenously determined changes in earnings on fertility rates is ambiguous, since increases in earnings increase family income, but also increase the opportunity cost of time devoted to raising children.[4]

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[2] In this paper we abstract away from the question about whether the decision to have a child is best modeled as the decision of an individual or a couple, as well as whether it is the outcome of one individual's dominant preferences, a consensus decision, or the result of interpersonal bargaining. Our reference to decisions made by a couple or by a woman when it comes to the decision to have a child is one of expositional convenience only.

[3] Studies documenting a causal link between shocks to income or liquidity and fertility rates include Lindo (2010), Black, Kolesnikova, Sanders, and Taylor (2013), Kearney and Wilson (2018); Cumming and Dettling (2020), and Goodman, Isen and Yannelis (2022). Studies showing a causal link between child-related prices and fertility rates include Milligan (2005); Cohen, Dehejia and Romanov (2013); and Dettling and Kearney (2014).

[4] Historically, economists have posited this off-setting substitution effect should apply to women's earnings (given historical gender patterns in time spent in childrearing), and they have treated increases in men's earnings as equivalent to increases in unearned income. However, as highlighted by Doepke et al, (2022), the historical negative correlation between women's earnings and childbearing under-pinning these types of assertations reversed in recent

This constrained optimization framework for the demand for having a child extends to decisions about the optimal timing of childbearing, including in response to the business cycle. The cyclical nature of fertility has been interpreted as reflecting optimal timing decisions in a life cycle model in which individuals respond to a temporary shock to income and wages. Assuming well-functioning capital markets, would-be parents could borrow and save to finance the cost of children and optimally choose when to have them over the life-cycle, unaffected by temporary changes in income (see, for instance, Hotz, Klerman, and Willis, 1997). But capital markets are imperfect, and the pro-cyclical nature of births is consistent with the proposition that liquidity or credit-constrained couples are more likely to choose to become parents when they have more disposable income available to pay for the associated costs of childbearing. Another common conjecture in the economics literature is that potential parents without access to savings or credit would opt into childbearing during an economic downturn when wages (and hence, the opportunity cost of time) are temporarily lower. This presumption is explicit in, for example, Dehejia and Lleras-Muney (2004) and Schaller (2016).[5]

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years and, as such, the economics of fertility has likely entered a "new era" in which distinctions based on traditional gender roles are increasingly less relevant.

[5] Dehejia and Lleras-Muney (2004) document improved infant health during economic downturns, which they attribute to both compositional shifts in new mothers and improved healthy behaviors among pregnant women. They write, "our evidence that the fertility response to temporary shocks in income differs substantially by socioeconomic status and by race is consistent with life-cycle models in which imperfect capital markets and skill depreciation during pregnancy play a role in the timing of fertility decisions (page 1093)." Schaller (2016) writes, "In the case of perfect certainty and perfect capital markets, transitory fluctuations in wages do not alter expected lifetime income and thus should not impact expected total fertility. They do, however, impact the timing of fertility if couples respond to transitory fluctuations in female wages by choosing to give birth when wages are low (page 4)."

This paper brings data to the question of whether and why fertility rates track the business cycle by combining testable predictions from economic models of fertility with high frequency data on cyclical movements in economic conditions and unemployment insurance benefit generosity. Our analysis begins by confirming that births are indeed pro-cyclical through recent cycles. We use administrative birth certificate data on the universe of live births in the United States from 2000-2019 and we date births to the likely month of conception. Guided by economic models of fertility and the empirical macroeconomics literature, we estimate OLS regressions relating monthly county-group birth outcomes to a broad array of cyclical economic variables which could plausibly influence child-bearing decisions. Specifically, our empirical analysis captures movements in employment opportunities with the county unemployment rate. We incorporate cyclical movements in unearned income with measures of asset prices –including county house prices and equity prices-- and allow for differential effects on groups with different ownership rates of such assets.[6] We also include measures of consumer sentiment and expectations. We both confirm and extend the existing literature, finding a separate role for movements in unemployment rates, asset prices, sentiment, and expectations. To the best of our knowledge, our paper is the first in the literature to take account of such a broad array of cyclical economic indicators.

[6] We do not include measures of output or consumer spending because economic models of fertility highlight a role for movements in employment opportunities, income, and liquidity in child-bearing decisions, but there is no mechanism by which consumer spending or GDP would be expected to independently drive child-bearing decisions. If anything, we would expect consumer spending or GDP to be reverse-casually related to births as future parents make additional purchases in preparation for a new child (e.g., furniture, clothes, etc.). Consistent with his notion, Buckles et al (2021) provide suggestive evidence that conceptions lead changes in GDP.

We then move on to analyses of *why* fertility is observed to be pro-cyclical. We introduce a novel empirical test of standard neoclassical models of fertility (e.g. Becker 1960, Hotz et al, 1996). Specifically, we propose an explicit test for the prediction of opposite-signed income and intertemporal substitution effects of business cycles by interacting the aggregate unemployment rate with a measure of potential income replacement from UI. If unemployment rates affect births because of an income effect associated with a transitory shock – as opposed to a perceived reduction in permanent lifetime income associated with a higher unemployment rate -- we would expect income replacement from UI to offset those effects. And when UI replaces 100 percent of income, the intertemporal substitution effect would be expected to lead to a positive relationship between unemployment rates and births. To test these predictions, we construct a refined measure of UI generosity that captures differences over time across states in UI income replacement rates that are due to changes in state-level policies, as well as how state system progressivity differentially affects different groups; importantly, our constructed measure is stripped of cyclical fluctuations in group-level incomes.

We find that as weekly UI benefits reach 100 percent income replacement, there is no negative effect of the unemployment rate on fertility rates. This implies that aggregate unemployment exerts a downward pressure on fertility rates entirely through a lost income channel. We find no evidence of a positive intertemporal substitution effect. We confirm this relationship using a second empirical test, again grounded in economic models of fertility. Specifically, we note that transitory changes in income should only affect child-bearing decisions if capital markets are imperfect and families cannot borrow and save. We construct a group-level measure of the incidence of liquidity or borrowing constraints and confirm that

unemployment rates exert no influence on fertility rates for groups that are not constrained, consistent with these models. In other words, the observed co-movement of births with unemployment rates is explained by movements in access to disposal income.[7]

Next, we revisit the question of whether infant health is pro-cyclical in the US and provide novel evidence that recessions are bad for infant health.[8] We confirm the finding of Dehejia and Lleras-Muney (2004) that infants who were *conceived* during periods of elevated unemployment in the US tend to be healthier, on average, but we show empirically that this finding is simply a downstream implication of the aforementioned change in the pool of mothers during downturns. When we narrow in on the effect of the aggregate unemployment rate *in utero*, we find that recessions lead to worse infant health -- specifically, a higher incidence of low birth weight and preterm births -- consistent with a causal negative effect of recessions on babies' health. We additionally find that these relationships are consistent with the unemployment effect being driven by income loss. Our analysis finds that when *in utero* UI income replacement rates reach 100 percent, there is no longer a negative effect of

[7] We include multiple robustness checks and show that this pattern holds across alternative model specifications. We also examine effects across birth parity and demographic groups.

[8] Dehejia and LLeras-Muney (2004) consider this question in the US context. Their analysis of national data exploiting state/year variation in unemployment rates finds that babies born during a year with a relatively higher unemployment rate tend to be of better health; their analysis of data from California using a sibling comparison design finds no evidence of a cyclical link. De Cao et al. (2022) use a sibling-comparison design to investigate the cyclicality of infant health in England. Their analysis implies that among babies born in the most affluent areas, infant health is counter-cyclical (meaning a sibling born during an economic downturn is healthier, on average, than their sibling born during a time of lower unemployment), but among babies born in lower income areas, infant health is pro-cyclical. Studies of the cyclicality of infant health in low to middle income countries tend to find evidence of pro-cyclical effects, as reported in De Cao et al. (2022).

unemployment on infant health. We explore the role of health inputs and find no evidence in favor or prenatal care as a mechanism, but there is evidence of a change in the incidence of maternal smoking.

Our paper makes important conceptual contributions to the economics of fertility literature by providing arguably one of the most direct tests of neoclassical models of fertility in the context of modern business cycles. We consider a broader set of measures of economic conditions than has typically been used, recognizing that business cycles are characterized by co-movements in labor market conditions, asset prices, and consumer sentiment. We also directly test the prediction of life-cycle fertility models that only liquidity or credit constrained families will respond to cyclical fluctuations in labor market conditions when timing their childbearing decisions, and that consumption smoothing social insurance could offset these effects. Our results suggest that *all* of the relationship between labor market conditions and fertility rates is driven by income and liquidity effects, not intertemporal substitution behavior.

Our paper also contributes to the literature on the economic effects of UI beyond job search. Existing papers document the positive effects of UI on consumption smoothing (e.g., Gruber, 1997, Browning and Crossley, 2001; Chetty, 2008, East and Kuka, 2015).[9] Our paper documents that UI smooths cyclical fluctuations in births, pointing to a new "liquidity effect" of UI – allowing potential parents to optimally time their childbearing decisions. [10] Cyclical

[9] There is also a large, tangential literature documenting the effects of UI on labor market outcomes (e.g., Kroft and Notowidigo, 2016, or Schmieder, vonWachter, and Bender 2012). A couple of recent papers also document UI effects on other outcomes, such as averting mortgage default and foreclosures (Hsu et al, 2018) and increasing access to health care (Kuka, 2020).
[10] Chetty (2008) identifies an effect of access to additional cash income from UI that allows otherwise liquidity constrained unemployed individuals to spend longer looking for a desirable

fluctuations in fertility rates matter for outcomes where cohort sizes matter, such as the allocation of educational resources.[11] Our results also have important long-run effects, since temporary postponements in births often lead to permanent reductions in the number of children a woman will have (see, e.g., Sommer, 2016). Consistent with this notion, there is empirical evidence that recessions reduce completed fertility (Currie and Schwandt, 2014). We find that UI mitigates the cyclical postponement of births in recessions, implying that UI is an incidental pro-natalist policy. Advanced economies across the globe are currently grappling with the implications of declining fertility rates for productivity and fiscal sustainability. Our paper provides new evidence that consumption-smoothing social insurance may help to counteract falling birth rates, and more generally, that policies which boost liquidity can increase birth rates even when those policies are not specifically targeted at doing so.

Furthermore, health at birth plays a key role in explaining lifetime health and human capital attainment (e.g., Almond and Currie, 2011), and there is growing causal evidence that infant health is affected by the pre-natal environment. We add to this literature by providing novel evidence that US babies who are *in utero* during periods of high unemployment have worse health outcomes.[12] Recognizing that mothers' economic well-being while pregnant could

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job and he labels this the UI liquidity effect. Chetty's naming of a liquidity effect of UI (which corrects for capital market imperfections) as distinct from a moral hazard effect of UI (which leads to suboptimal job search efforts) implies a higher optimal level of UI generosity.

[11] See, for example, Petrilli (2019) on how changing birth rates affect the education system.

[12] To the best of our knowledge, related papers do not look separately at the link between infant health outcomes and aggregate unemployment rates at the time of conception versus the time in utero. There are other differences between our paper and existing papers on this question. For instance, as noted in a footnote above, De Cao et al (2022) examine the link between aggregate unemployment and infant health outcomes in the UK by comparing outcomes for siblings born in years with different unemployment rates; Dehejia and Lleras-Muney (2004) run a similar

affect the lifetime health and well-being of the next generation, a growing body of work has examined how social policies indirectly affect infant health outcomes. For instance, Hoynes, Miller, and Simon (2015) demonstrate that the EITC has a beneficial effect on infant health outcomes. Our paper adds to this literature by studying the role of UI, which is the largest U.S. safety net program in recent recessions (Bitler and Hoynes, 2016). Our finding that UI ameliorates the negative effect of recessions on the health outcomes of babies who are affected in utero thus implies beneficial intergenerational effects of UI on economic well-being.

**Background: Recessions, Birth Outcomes, and the Potential Role of UI**

*a. Recessions and Fertility rates*

Economic models of fertility indicate that births may follow the business cycle owing to temporary changes in income or wages, which can alter the optimal timing of childbearing for couples who are liquidity or borrowing constrained (Butz and Ward, 1979). Pro-cyclical fertility is consistent with the proposition that couples without access to credit or savings will be less likely to choose to become parents when they have less disposable income available to pay for the associated costs of childbearing (e.g. during unemployment). Consumption-smoothing income assistance from unemployment insurance would be expected to offset these effects.

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siblings-comparison specification in their paper using data on infants born in California. (Both of these papers measure unemployment at the year level.) The implicit assumption in a sibling comparison is that maternal characteristics that lead to differential selection into childbearing during recessions (i.e., the presence of liquidity constraints) and differential infant health outcomes (i.e., income, wealth, socioeconomic status) are fixed for a given mother over time. This strikes us as a dubious proposition. Lindo (2010) studies a related, but different, question - the effect of spousal job loss on infant health - and finds some evidence suggesting spousal unemployment leads to lower average infant birth weight.

It is common in the cyclicality of fertility literature to encounter discussions that suggest that Becker's (1960) model implies that fertility could be counter-cyclical due to off-setting intertemporal substitution effects (e.g., Dehejia and Lleras-Muney, 2004; Schaller, 2016). Though such considerations likely apply to *permanent* changes in earned income, we propose they should not apply to short-run *transitory* changes in earned income. One obvious reason why is that such considerations are made with respect to the opportunity cost of time spent *rearing* children. But opting into pregnancy during an economic downturn likely means opting into having an infant or toddler during an economic recovery period. That typically would mean even greater child-related demands on parental time than during pregnancy. This raises practical doubt about the proposition often cited in previous papers that opting into pregnancy during a temporary economic downturn is potentially optimal.

Recessions and recoveries typically feature movements in more than just families' incomes and employment. Each recession and recovery is unique, and as such, there will be specific contextual factors that shape the relationship between aggregate economic conditions and the fertility response. Some recessions are accompanied by financial crises featuring dramatic swings in financial markets which reduce unearned income. Other recessions feature large declines in house prices, which have been shown to effect child-bearing decisions (Lovenheim and Mumford, 2011; Dettling and Kearney, 2014). Consumer sentiment and expectations also vary over the business cycle, which could potentially alter childbearing

plans.[13] The fiscal and monetary policy environment could also shape how fertility rates move during an economic downturn (e.g., Cumming and Dettling, 2020; Kearney and Levine, 2021).[14] Finally, more prolonged recessions and recoveries will mean that delayed or postponed births are more likely to lead to permanent reductions in completed fertility; this could result from biological fertility challenges associated with older age (Sommer, 2016) or from shifted preferences over time as one grows accustomed to their present family size.

*b. Recessions and Infant Health*

Aggregate economic conditions could affect infant health outcomes by shifting the demographic composition of mothers. Specifically, as noted above, economic models predict that liquidity or credit constrained mothers will opt out of childbearing during downturns. Because the incidence of liquidity and credit constraints is negatively correlated with socioeconomic status (e.g, Bhutta and Dettling, 2018), and lower socio-economic status mother's tend to have, on average, less healthy babies, this would potentially mean that the infants born during recessions would be of better relative health, on average. Indeed, Dehejia and Lleras-Muney (2004) show that babies conceived during periods of higher unemployment have a reduced incidence of low and very low birth weight, which they attribute in part to the shifted composition of mothers.

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[13] For example, in the context of the COVID-19 pandemic, Kearney and Levine (2022) document that the recession in the early spring of 2020 led to a large decline in conceptions, which was amplified by the extent of local-area COVID cases.

[14] Cumming and Dettling (2020) document a link between central bank policy rates and births. In particular, they find that monetary easing during the Great Recession increased births in the United Kingdom by increasing household liquidity via mortgage rate pass-through. Kearney and Levine (2021) find that during the COVID-19 pandemic, fertility rates increased when household spending recovered, which was bolstered by the fiscal policy response.

There are also direct ways in which an economic downturn that affected the income and employment of parents would lead to a change in infant health. First, job loss or household income loss during an economic downturn could, all else equal, reduce health inputs, leading to a reduction in infant health outcomes. This could be a result of people being income constrained to pay for adequate nutrition or prenatal care, or from unemployed parents losing access to health insurance that was attached to employment. [15] Second, recession-induced income losses may lead to either increases or decreases in either risky or healthy behaviors, depending on how time intensive such behaviors are or whether they are complements or substitutes with time in market work. Dehejia and Lleras-Muney (2004) document a decline in maternal smoking during recessions, which could reflect either a behavioral change related to unemployment or the compositional shift in who becomes a mother during an economic downturn. Finally, unemployment (both at an individual or aggregate level) may affect maternal stress, and maternal stress levels have been shown to be negatively related to infant health outcomes (Aizer, Stroud, Buka, 2016).[16] UI benefits – by providing liquidity during downturns – could mitigate any or all of these channels.

**Data and Summary Statistics**

[15] Kuka (2020) finds that UI benefit generosity is associated with great health insurance coverage, which raises the possibility that more generous UI might increase prenatal care access.

[16] The medical literature has also documented the effects of maternal stress -- specifically, elevated maternal cortisol levels-- on fetal development. See, for example, Hobel and Culhane (2003).

Our main source of data on births and infant health outcomes is Vital Statistics birth certificate data. Our data set includes all births in all years from 2000-2019. The data includes information on mother's race and Ethnicity, age, and marital status. Appendix A includes more information on the construction of this data. The data also include a variety of measures of infant health outcomes including birth weight and length of gestation, which we use to infer if the infant was born prematurely. The data also has information on maternal health behaviors, including prenatal care usage though-out the pregnancy (including the month of initiation of care and number of visits) and whether the mother smoked while she was pregnant.[17]

We applied for and obtained a confidential version of the Vital Statistics data which includes the mother's state and county of residence. To order to construct fertility rates, we collapse the birth data at the county-month-group-level, where demographic groups are defined by age category (18-34 and 35-49) and race/ethnicity (White, non-Hispanic, Black, non-Hispanic and Hispanic) and month refers to the month of conception, since that is the time relevant to the decision to get pregnant. We use the SEER data to estimate female population in each county-group-year in to use as the denominator in the construction of fertility rates. Table 1 reports summary statistics for these variables.

We link the birth certificate data to information on economic conditions and unemployment insurance benefits available in the mother's county of residence in the relevant month. Our main measure of local labor market conditions is the monthly unemployment rate in a mother's county of residence which we obtained from the Bureau of Labor Statistics Local

[17] Smoking and the month in which prenatal care usage began are missing in some states and year due to known issues with the Vital Stats data, as noted in Appendix A.

Area Unemployment Statistics (BLS LAU). Our models control for house prices at the level of county and month. Our house price measure comes from Zillow housing value index (ZHVI), whichrepresents the value of a typical home in the county-month. We put this measure in real dollars using the CPI-U less shelter series. We measure equity prices using the average value of the Wilshire 5000 Index in each month and put them in real dollars by deflating by the CPI-U. We use the Michigan Survey of Consumers measures of consumer sentiment and consumer expectations, which we construct at the level of month by group by region (the smallest geographic unit available).[18] The indices are defined as the percent of positive responses minus the percent of negative responses plus 100. Summary statistics for these variables are included in Table 1.

We also construct several demographic group-level measures from the 2001-2019 waves of the Survey of Consumer Finances (SCF). Specifically, we construct demographic group-level home and equity ownership rates, to allow for differential effects of movements in asset prices amongst owners and non-owners of those assets, following the intuition in Dettling and Kearney (2014). We also construct a group-level measure of the proportion of families who are both liquidity constrained, where families are identified as liquidity constrained if they are both credit constrained and have limited liquid savings. Credit constraints are defined as an affirmative response to either being denied credit or not applying for fear of denial, as is common in the literature (e.g., Bricker et al, 2020). Limited liquid savings is defined as having

[18] We average the monthly region-level and age-group level indices to construct a region-group level monthly index.

less than one month of one's own recurring expenses in liquid savings as in Bhutta and Dettling (2018).[19] Table 1 also includes summary statistics for these variables.

We construct a refined measure of UI generosity which we refer to as a "potential weekly UI income replacement rate." We use data collected from Department of Labor (DOL) publications to construct a UI benefit calculator which factors in two key sources of state-by-year policy variation: weekly income replacement formulas and maximum benefit levels.[20] We use all women age 18-49 in the 2000 Census (five percent sample) and feed their wage and salary incomes through the UI benefit calculator for their state of residence in each year in our sample to estimate potential weekly benefit levels and income replacement rates.[21] We then calculate median potential UI replacement rates by state, year, and group, where the groups are defined as in the birth data. The key to this measure is that it captures differences over time across states in UI generosity due to changes in state-level policies, as well as how state system progressivity differentially affects different groups; but, by using a fixed sample of women from the year 2000, it is stripped of cyclical fluctuations in group-level incomes and employment. [22] In extensions and robustness checks, we estimate our regression models using potential

[19]This measure is designed to mimic financial planner rule of thumb that families should maintain an emergency fund that is proportional to a certain number of months of their own expenses. Recurring expenses include housing costs, food, vehicle costs, health costs, debt payments, etc. but exclude durables spending.

[20] Appendix A includes more information on the UI benefit calculator and construction of potential replacement rates. For ease of construction, for the 2000-2009 period we use the maximum benefit levels from the Hsu et al (2018) replication, which were constructed from the same DOL data we use in the same manner.

[21] We use the CPI-U to inflate the 2000 incomes to 2001-2019 dollars, since the maximum benefit levels are in nominal terms.

[22] This is similar in spirit to the simulated instrument method initially proposed in Currie and Gruber (1996).

replacement rates for different subsamples (including men), and we also examine the sensitivity of results to different measures of UI generosity, such as maximum weekly benefit caps and maximum durations of UI eligibility.

Appendix Table 1 lists constructed UI replacement rates across states. As can be seen in the table, on average, weekly UI benefits are estimated to potentially replace 53 percent of weekly wage and salary income for the typical working woman ages 18 to 49. Across states, potential UI replacement rates range from below 50 percent in Alaska, Arizona, DC, Florida, DC, Illinois, and Mississippi to above 60 percent Hawaii, Oregon and Kentucky. Appendix Table 2 lists UI replacement rates constructed for different samples; later we will run robustness checks using these alternative constructions. Appendix Table 3 shows that there is considerable variation across years in our measure of potential UI replacement rate. Over time, potential replacement rates peak in the Great Recession at 58 percent and fall to 51 percent in 2019. Appendix Table 4 reports potential UI replacement rates across groups. Replacement rates are a bit higher for younger women than older women, reflecting typical life-cycle earnings profiles.

Our measure of UI benefit generosity has some key advantages over plausible alternatives like maximum durations of eligibility or maximum benefit caps for both conceptual and empirical reasons. Conceptually, our measure is designed to be interpretable as a percentage change in income, which is the parameter of interest in our conceptual framework. If we were to use, for example, maximum durations of benefits we would not be able to easily interpret those parameters within the economic models of fertility. Our measure also better captures the relevant features of state UI policy for the population we study than maximum benefit caps alone. Specifically, comparing our estimate of median weekly UI benefits for

women of child-bearing age (column 2) with state maximum benefit caps (column 4) in Appendix table 1 we see that across all 50 states, column 2 falls below column 4. This implies that the maximum benefit caps do not bind for a typical woman of child-bearing age, on average. Furthermore, we can see that our measure of potential UI replacement is only weakly correlated with maximum benefit levels.[23] For example, Louisiana has an average maximum benefit of $272 a week --the bottom 10 percent of all states in terms of maximums-- but at Louisiana's median women's income of $19,112, UI replaces 59 percent of income –placing Louisiana in the top 10 percent of all states in terms of potential UI replacement rates. In other words, states with more generous replacement formulas do not always have the highest maximum benefit caps, and vice versa. Thus, our measure –which combines both state replacement rate formulas and maximum benefit levels – is likely to better capture the relevant policy variation for the population we study than state maximums alone.[24]

**Empirical Strategy**

To examine the effects of recessions on fertility rates we estimate models of the following form:

$$\ln(fertrate_{c,g,t}) = \beta_1 UR_{c,t-3,t-1} + \mathbf{E}_{cgt-3,t-1}\boldsymbol{\beta_E} + \delta_c + \delta_t + \delta_g + \epsilon_{igct} \quad (1)$$

Where $fertrate_{c,g,t}$ is conceptions per thousand women in in county $c$ and age-group by race/ethnicity group $g$ in month $t$. $UR_{c,t-3,t}$ is the mean unemployment rate in the county from

[23] The correlation coefficient between column 2 and 4 is 0.16.

[24] Women of child-bearing age tend to have below average incomes. Studies focusing on higher income populations –e.g., homeowners (Hsu et al, 2018) – may not face these issues as the maximums will more often bind in higher income populations.

month *t-3* through month *t-1* (e.g., the quarter leading up to conception). $\mathbf{E}_{cgt-3,t-1}$ is a vector of time-varying county-group level economic conditions, including county-month house prices interacted with group home ownership rates, stock prices interacted with group equity ownership rates, and group-region-month Michigan survey measures (sentiment, expectations). $\delta_c$ are county fixed effects, $\delta_g$ are group fixed effects and $\delta_t$ are month fixed effects. We weight all regressions by the female population in each cell, and standard errors are adjusted for clustering at the county-level.[25]

In this paper we are interested in identifying a causal relationship between lagged economic conditions and conception rates, and our models are designed to overcome a number of different identification challenges that might be present in a cross-sectional approach. A key feature of our regression model is the inclusion of county fixed effects ($\delta_c$), which capture time-invariant differences across counties in fertility preferences. This would include any sorting patterns by preferences for children according to fixed labor market or economic characteristics of different cities or neighborhood types, as well the potential effects of state-level social policies which might affect birth outcomes but were mostly unchanged over the period we study (e.g., abortion policies or welfare generosity).[26] Another important aspect of the model is the inclusion of month fixed effects ($\delta_t$) which capture national trends in economic conditions and fertility rates, including how fertility decisions respond to national economic developments that

[25] We also estimated models adjusting for clustering at the state level and find similar results (not shown).

[26] In general, incremental changes in such policies have been found to have fairly small or negligible effects on aggregate fertility rates (e.g., Kearney, Levine, Pardue, 2022). Larger changes in abortion policy and Medicaid coverage happened outside the time period being studied here.

might co-vary with the cycle, such as the monetary or fiscal policy response. Put together, this means $\beta_1$ and $\boldsymbol{\beta_E}$ are identified off within-county changes in economic conditions over time.

In one specification we add an interaction term between the unemployment rate and a group-level measure of the incidence of credit and liquidity constrained ($Constrained_g$). Specifically, we augment equation (1) with the following additional term: $\beta_2 UR_{c,t-3,t-1}$* $Constrained_g$. In this model, $\beta_2$ describes the effects higher unemployment rates as the proportion of the group that is liquidity or credit constrained approaches 100 percent. Conditional on $\beta_2$, $\beta_1$ then describes the effects of unemployment on birth outcomes for women who are not liquidity or credit constrained. If the predictions of lifecycle models of fertility are borne out in the data, for equation (1) we would expect $\beta_1 = 0$ and $\beta_2 < 0$.

We then estimate a version of the model that includes an interaction term between the unemployment rate and the potential UI replacement rate. In this case we augment equation (1) with the following additional terms: $\beta_2 UR_{c,t-3,t-1}$* $UI_{s,g,t-3,t-1}$+ $\beta_3 UI_{s,g,t-3,t-1}$, where $UI_{s,g,t-3,t-1}$ is our measure of the potential UI replacement rate. In the augmented model, $\beta_2$ describes the effects higher unemployment rates as UI approaches 100 percent income replacement. Conditional on $\beta_2$, $\beta_1$ then describes the effects of unemployment on birth outcomes when there is no income replacement from UI. In this augmented equation, testing for income effects of unemployment on fertility rates is thus a test of whether $\beta_2 > 0$ and testing for positive intertemporal substitution effects on fertility rates is tantamount to testing whether or not $\beta_1 + \beta_2 > 0$. We formally implement the latter using 500 bootstrap replications.

To study infant health outcomes we estimate models of the following form:

$$Y_{c,g,t} = \beta_1 UR_{c,t,t+8} + \beta_2 UR_{c,t-3,t-1} + \mathbf{X}_{cgt}\boldsymbol{\beta_x} + \mathbf{E}_{st,t+8}\boldsymbol{\beta_E} + \delta_c + \delta_t + \epsilon_{isct} \qquad (2)$$

Where $Y_{c,g,t}$ is the proportion of babies in each county *c*, group *g*, conceived in month *t* born with various health attributes at birth or maternal health behaviors during pregnancy. The main infant health outcome measures include, alternately, the proportion of babies that are low birth weight (under 2500 grams at birth) and the proportion born prematurely (before 37 weeks).[27] In models that use a measure of maternal health behavior as the outcome variable of interest, we consider the proportion of mothers who initiated prenatal care in the first trimester, who obtained at least 5 prenatal care visits, and who smoked while pregnant. In these specifications, most models include two measures of the unemployment rate, $UR_{c,t,t+8}$, which is the unemployment rate in the mother's county during the nine months following conception (i.e., during pregnancy) and $UR_{c,t-3,t-1}$, the unemployment rate in the mother's county of residence in the three months leading up to conception. The former is designed to capture the causal effect of experiencing a recession during pregnancy on infant health or maternal behavior, while the latter is designed to capture health effects owing to any differential selection into childbearing during recessions.

These models explicitly control for various infant and maternal characteristics in vector $\mathbf{X}_{cgt}$, which includes the proportion of babies who are boys and from multiple births (which affect birthweight), as well as the proportion of mothers who are married, since married women's incomes are in part insured by their husbands' earnings and could react differently to

[27] Appendix table 8 has alternative health measures, including the fraction of babies born with a very low birth weight (under 1500 grams), the proportion born very prematurely (before 32 weeks) and birth weight.

the business cycle and government social insurance.[28] The remaining terms in equation (2) are defined as in equation (1).

We also estimate models where we introduce an interaction term between unemployment rates and UI benefit generosity. Specifically, we augment equation (2) with the following additional terms: $\beta_3 UR_{c,t,t+8}{}^{*}\, UI_{s,g,t,t+8} + \beta_4 UI_{sg,t,t+8}$ , where $UI_{s,t,t+8}$ are our measures of UI benefit generosity in state *s*, for group g, measured during pregnancy *(t,t+8)*. In these models, $\beta_2$ describes the effects of higher unemployment rates on infant health outcomes as UI approaches 100 percent income replacement. If experiencing a pregnancy during an economic downturn is, on net, bad for infant health because of various (indirect) effects of lost income, then a higher replacement rate should mitigate those negative effects, resulting in $\beta_2 < 0$ (since the infant health outcomes captured in $Y_{c,g,t}$ indicate worse health.)

**Results**

***i. Unemployment rates and Fertility rates***

Table 2 presents results from estimating equation (1). Column 1 starts with a sparse specification omitting $\mathbf{E}_{cgt-3,t-1}$ and columns 2-4 sequentially add the various economic indicators in $\mathbf{E}_{cgt-3,t-1}$. The relationship between unemployment rates and conceptions is stable and statistically significant (at the 1 percent level) in columns 1-4, indicating that each percentage point increase in local unemployment rates reduces conception rate by about 1

[28] We do not control for mother's education because of the well-known issue of birth certificates missing this information in nearly half of states during the period from 2009-2013. See Appendix A for more information on missing data.

percent. These magnitudes are extremely similar to those found in previous work (see, e.g., Kearney and Levine, 2021).

Column 5 of table 2 reports the results of estimating equation (1) with the interaction term between unemployment rates and the group level measure of the incidence of liquidity and credit constraints. As predicted by life-cycle models, the coefficient on the unemployment rate ($\beta_1$) is no longer statistically different from zero, but the coefficient on the interaction term between the proportion of the group that is liquidity or credit constrained and the unemployment rate ($\beta_2$) is negative and statistically significant at the 10 percent level. This implies that the childbearing decisions of groups who are not liquidity or credit constrained are unaffected by changes in unemployment rates.

Column 6 of table 2 reports the results of estimating equation (1) with the interaction term between unemployment rates and UI benefit generosity. Consistent with the notion that children are "normal goods" – that is, the demand for children increases in income -- $\beta_2$ is positive and statistically significant at the one percent level. The coefficient on $\beta_2$ implies that holding unemployment rates fixed, moving from zero to 100 percent income replacement increases fertility rates by 5.78 percent.[29] In column 6, $\beta_1$ is more negative than in column (4),

[29] This pattern of results is robust across various alternative specifications, as shown in Appendix Table 5. These specifications use UI replacement rates calculated over the earned income of different samples, including women with positive wage income, women who worked at least 26 weeks last year (i.e., typical UI eligibility work requirements), men with positive wage income, and men who worked at least 24 weeks last year. It is also robust across specifications that use the mean instead of the median replacement rate, define the outcome only for marital births and use the replacement rate over married men's earned income, and using the maximum benefit level instead of the potential replacement rate. In a specification that replaces the replacement rate with just the maximum weeks of eligibility, the coefficient on the interaction term goes to zero.

indicating that when UI does not replace any income, each 1 percentage point increase in the unemployment rate is associated with a 4.21 percent reduction in fertility rates. Recall that testing for an intemporal substitution effect is tantamount to testing if $\beta_1 + \beta_2 > 0$. We test this formally by running 500 bootstrap replications on our estimate of $\beta_1 + \beta_2$ , and find that we cannot reject that the substitution effect is zero at the ten percent level (the point estimate is 0.0157 with standard error of 0.0097).[30]

Column 6 of Table 2 also indicates that most of the variables included in $\mathbf{E}_{cgt-3,t-1}$ exert independent and statistically significant effects on births above and beyond changes in unemployment rates. Local house price increases exert a positive effect on homeowners, and negative effect on non-owners, confirming the results of Dettling and Kearney (2014). Increases in equity prices similarly exert a positive wealth effect on fertility rates for owners of equities. Improvements in consumer sentiment and consumer expectations both increase fertility rates.

A natural question arising from the results displayed in Table 2 is whether the estimated unemployment rate and UI effects on birth rates is driven by individuals who are unemployed and eligible to receive UI, or whether those effects capture other aspects of the recessionary environment. To probe this, we estimated a modified version of equation (1) using individual-

[30] In a contemporaneous working paper, Lindo et al (2022) uses SIPP data to examine the relationship between state maximum UI benefit caps and child-bearing decisions in the years following a man or woman experiencing a spell out of the labor force. They find that such spells reduce the probability a man will have a child and increase the probability a woman will have a child two to three years after the initiation of the spell. Although it is not their main specification, when they focus on spells in which the potential parent is actively looking for work (i.e., unemployed and eligible for UI), they find that higher maximum benefits levels increase men's and women's childbearing one and half to two years after the onset of those spells.

level data from the Current Population Survey Annual Social and Economic Supplement (CPS ASEC) on recent births and unemployment spells. (The Data Appendix gives details about the data construction and individual-level specification.) Appendix Table 6 displays the results. Column 1 indicates that women who report having been unemployed in the prior year are 0.6 percentage points less likely to have had a recent birth (about a 10 percent decline at the mean of the dependent variable). The specification reported in Column 2 adds in local unemployment rates; the results indicate that the unemployment rate does not exert a statistically significant effect on the probability of giving birth for individuals who are not unemployed. Finally, column 3 adds the interaction between own unemployment status and UI benefit generosity. Similar to the main results presented in Table 2, this specification shows that when UI benefits reach 100 percent income replacement, the negative effect of own unemployment on the probability of giving birth disappears. All told, Table A6 indicates that the results in Table 2 hold at the individual-level and suggests the effects are driven by women who experience an unemployment spell and receive UI benefits.

Table 3 displays the results of estimating equation (1) and the augmented models by birth parity. There is a negative effect of unemployment across all parity of births (columns 1, 4, 7), and these effects are concentrated on liquidity constrained women (columns 2, 5, 8). The interaction with UI has a similar sign across parities, however, the effects are quite a bit larger for third and higher births. Following the literature, which has proposed that third and higher

births may be more "marginal" (e.g., Dettling and Kearney, 2014), we interpret this as potentially consistent with the notion that UI increases completed fertility.[31]

### *ii. Unemployment rates and Infant Health*

Table 4 presents results from estimating equation (2) for two infant health outcomes. Columns 1-4 present results where the outcome is the percent of infants born with a low birth weight (under 2500 grams) and columns 5-8 present results where the outcome is the percent of infants born prematurely (fewer than 37 weeks). Columns 1 and 5 start with a specification including only the *in utero* unemployment rate ($UR_{ct,t+8}$). The estimated effect on percent low birth weight is not statistically different from zero (column 1), but the result reported in column 5 indicates that a 1 percentage point increase in the unemployment rate is associated with a 0.075 percentage point increase in the proportion of babies born prematurely, which is a 0.05 percent increase at the mean of the dependent variable.

In column 2 and 6 we omit the *in utero* unemployment rate and include only the unemployment rate during pregnancy with the unemployment rate prior to conception ($UR_{ct-3,t-1}$). In that specification, the data do not show a statistically significant relationship between labor market conditions prior to conception and birth outcomes. Then, in column 3 and 7 we add both the unemployment rate prior to conception and the *in utero* unemployment rate.

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[31] Appendix Table 7 reports results separately for White women, Black women, Hispanic women, women ages 18 to 34, and women ages 35 to 49. The largest negative effects of the unemployment rate and positive effects of the interaction term with the UI replacement rate are for Black women and for younger women. This could be indicative of Black women and younger women being especially liquidity constrained and sensitive to labor market conditions and income replacement rates.

In this case, the two unemployment rates exert opposite-signed effects on infant health, consistent with a positive selection effect of mothers into childbearing and a negative direct effect on *in utero* infant health development. Specifically, each one percentage point increase in the *in utero* unemployment rate increases the fraction of babies born prematurely by 0.1367 percentage points (statistically significant at the 1 percent level), which is a one percent increase at the mean of the dependent variable.

Although babies *in utero* during times of high unemployment suffer worse health outcomes, columns 3 and 7 indicate that babies *conceived* during times of high unemployment are a bit healthier. A 1 percentage point increase in the unemployment rate in the months leading up to conception reduces the fraction of babies born with low birth weight by 0.04 percentage point and the fraction of babies born prematurely by 0.07 percentage point. We interpret this as a downstream implication of the change in the pool of expectant mothers documented in table 1. In particular, the results reported in Table 4, column 6, indicate liquidity and credit constrained mothers –who tend to be lower socioeconomic status and would otherwise have had less healthy babies, on average--- opt out of childbearing during recessions.

Columns 4 and 8 report results from the model augmented with the interaction term between the unemployment rate and the potential UI replacement rate. The results suggest there may be a role for public policy to improve infant health outcomes by buffering income losses during recessions. The coefficients on $\beta_1$ implies that when there is no income replacement from UI, each 1 percentage point increase in the unemployment rate increases the proportion of infants that are low birth weight by 0.17 percentage point (about 2.5 percent at the dependent variable mean) and the proportion of births that are pre-term by 0.43 percentage

point (about 3.5 percent at the dependent variable mean). However, the coefficient on $\beta_3$ --the interaction between the in utero unemployment rate and potential UI replacement-- shows that, holding the unemployment rate fixed, higher UI replacement rates lead to better infant health outcomes --moving from zero to 100 percent UI replacement reduces the proportion of low birth weight births by 0.25 percentage point (3 percent) and preterm births by 0.54 percentage point (4.1 percent). In other words, at 100 percent UI replacement, the negative effect of the unemployment rate on infant health disappears.[32] Together, these estimates imply that UI would need to replace between 70-80 percent of incomes to offset the negative effects of higher aggregate unemployment rates on infant health – well above typical replacement rates in our sample.[33]

To put this estimate in a cost/benefit perspective, we note that the average cost of a birth hospitalization was $62,931 for commercially insured preterm infants and $43,858 for Medicaid-insured preterm infants compared to $2,401 and $1,894 for commercially insured and Medicaid-insured full-term infants, respectively (McLaurin et al, 2019). Our results imply that it would

[32] More formally, we run 500 bootstrap replications on $\beta_1 + \beta_3$ and find that when UI replaces 100 percent of income there is a small positive, but statistically insignificant (at the five percent level) effect of unemployment rates on infant health outcomes. For percent low birth weight, the coefficient is -0.076 with a standard error of 0.041. For percent born prematurely, the coefficient is 0.110 with a standard error of 0.065.

[33] Appendix Table 8 reports results across various specifications and definitions, analogous to Appendix Table 5 for fertility rates. The pattern of results is generally robust, as can be seen in the table. Appendix Table 9 displays results for alternate outcomes, including the proportion of births that are very low birth weight (less than 1500 grams), the proportion of births that are very preterm (less than 32 weeks), birth weight, gestational age, and the proportion of births that are male. The same patterns hold, indicating that UI has a protective effect both on average health at birth (birth weight) as well as more adverse outcomes (very low birth weight, very preterm). There is no effect on male share.

take an average of $383/week of UI benefits to erase the cyclical pattern of pre-term births, or an additional $129 per week on top of the current average UI benefit level of $254 per week. For the length of a full-time pregnancy (44 weeks), this would amount to nearly $17,000 in UI benefits. The implication is that the cost of replacing income loss among pregnant mothers has a positive social return based on the cost of a pre-term birth alone, not even accounting for the longer-term effects of infant health.

Columns 4 and 8 of Table 4 also point to some novel relationships between cyclical movements in asset prices and infant health which are also consistent with positive liquidity effect on infant health. Specifically, we find that higher house prices and equity prices during pregnancy lead to improvements in infant health among groups with higher ownership rates of those assets. This is consistent with a wealth effect on infant health, presumably also via an increase in liquidity.

Table 5 reports results separately by demographic group. For all groups there is evidence of positive selection effects from higher unemployment at the time of conception. And for all but Hispanic mothers, unemployment rates during pregnancy lead to worse infant health outcomes, and higher potential UI replacement rates mitigates those effects. Across groups, the largest magnitude effects are for Black women and for women 35-49 -- both groups which have a higher incidence of low birth weight and preterm births on average. For Black women, a 1 percentage point increase in unemployment rates lead to 0.64 percentage point (4.3 percent at the mean) increase in the proportion of low birth weights and 0.84 percentage point (4.5 percent) increase in the proportion of preterm births. For older women, a 1 percentage point increase in unemployment rates lead to 0.23 percentage point (2.5 percent at the mean) increase

in the proportion of low birth weights and 0.42 percentage point (2.9 percent) increase in the proportion of preterm births. Holding the unemployment rate fixed, moving from zero to 100 percent income replacement from UI reduces the proportion of low-birth weight births among Black women by 0.89 percentage points (6 percent) and preterm births by 1.2 percentage points (6.4 percent), and among older women by 0.36 percentage points (3.9 percent) and preterm births by 0.53 percentage points (3.7 percent).

Table 6 reports results separately by trimester. The estimated conditional relationship between the unemployment rate and infant health outcomes is largest in the first trimester, while the point estimate of the effect of the interaction of the unemployment rate and the UI replacement rate is generally stable across pregnancy trimesters. This would be consistent with the protective impact of UI on infant health being strongest during the first trimester. One interpretation of this pattern of results is that they are consistent with unemployment affecting infant health development through elevated maternal stress (with higher UI benefits reducing stress), as stress earlier in pregnancy has been shown to have more pronounced harmful effects on fetal development (e.g., Hobel and Culhane, 2003).

Table 7 probes other possible explanations for the procyclicality of infant health by reporting the results for maternal health behavior during pregnancy. Columns 1-4 display two measures of prenatal care usage: obtaining prenatal care in the first trimester (columns 1-2) and having at least 5 prenatal care visits (columns 3-4). The coefficient on the *in utero* unemployment rate is either positive or not statistically different from zero in columns 1-4, indicating that

prenatal care usage cannot explain the procyclicality of infant health.[34] This is consistent with the notion that prenatal care has only modest causal effects on babies' health (see Corman et al, 2019 for a review). Columns 5-6 displays the results for smoking during pregnancy. In column 5, there is no evidence that smoking is associated with aggregate unemployment either before or during pregnancy, however, in column 6, we find that higher *in utero* unemployment rates are associated with an increase in the proportion of mothers who smoked during pregnancy (marginally statistically significant at the 10 percent level), but that higher potential UI replacement rates mitigate those effects. We speculate that perhaps this pattern is related to maternal stress patterns, insofar as worse aggregate economic conditions exacerbates stress, which leads to increased rates of smoking, and income replacement mitigates that stress and associated smoking behavior.

**Conclusion**

This paper provides arguably one of the most direct tests of economic models of fertility (e.g., Becker, 1960; Hotz et al, 1996) in the context of modern business cycles. We provide direct evidence on the separate proposed mechanisms through which higher aggregate unemployment could affect fertility rates – an income loss effect (negative) and an intertemporal substitution effect (positive). We find that when UI provides 100 percent replacement, unemployment rates exert essentially no effect on births, consistent with the notion that children are "normal" and widespread and binding liquidity constraints causing

[34] The coefficient on unemployment at the time of conception in columns 1-2 indicates increased prenatal care usage, which also consistent with positive selection into childbearing during recessions.

families to delay childbearing when their incomes are temporarily low. We find no evidence to support intertemporal substitution effects. This finding strikes us as being consistent with the practical observation that the opportunity cost of time spent child-rearing would be expected to occur well after the typical temporary unemployment spell is over (specifically, about nine months through eighteen years later).

This paper was written in the aftermath of the COVID-19 pandemic recession, which featured a dramatic fiscal expansion -- for the first time, UI benefit generosity was expanded so that replacement rates exceeded 100 percent of income (Ganong, Noel and Vavra, 2020), and personal incomes ultimately *rose* during the recession (Bhutta et al, 2021). New research indicates that the pandemic also initially featured the typical child-bearing response; as unemployment rates sky-rocketed, there was a reduction in conceptions, but this pattern quickly reversed (Kearney and Levine, 2022). This reversal led some observers to speculate that perhaps the long-standing procyclical nature of fertility had broken down (Bailey, Currie and Schwandt, 2022). Our finding that procyclical fertility is ultimately about the cyclical nature of liquidity reconciles this apparent contradiction in the recent experience.

Our paper provides novel evidence that infant health is pro-cyclical in the US. We find that although differential selection into childbearing during recessions causes infants *conceived* during periods of high unemployment to be healthier, on average, babies who are *in utero* during periods of high unemployment experience worse health outcomes, namely, they are more likely to be low or very low birth weight and are more likely to be born prematurely. These effects are concentrated on liquidity constrained mothers, and when we allow for off-

setting income replacement from UI, the effects disappear. In other words, loss of income during pregnancy harms babies' health.

This paper adds to the literature on the effects of UI on consumption smoothing by documenting that UI acts to smooth cyclical fluctuations in birth outcomes, allowing would be parents to maintain their intended rates of fertility and protecting infant health. Because births that are delayed are averted altogether and health at birth plays a key role in explaining lifetime health and human capital attainment, our results suggest that UI is pro-natalist and leads to intergenerational improvements in economic well-being.

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**Table 1: Summary Statistics**

| | Mean | SD | Memo: Data Source |
|---|---|---|---|
| $fertrate_{c,t,g}$ | 4.42 | 3.30 | Vital Statistics and SEER |
| Percent low birth $weight_{c,t,g}$ | 8.65 | 8.53 | Vital Statistics |
| Percent $preterm_{c,t,g}$ | 13.05 | 10.42 | Vital Statistics |
| | | | |
| unemp $rate_{c,t-3,t-1}$ | 6.01 | 2.53 | BLS LAU |
| ui $replacement_{g,s,t-3,t-1}$ | 0.53 | 0.05 | DOL, Hsu et al (2018), and Census |
| liquidity constrained $rate_{g}$ | 0.19 | 0.07 | SCF & Bhutta and Dettling (2018) |
| house $prices_{c,t-3,t-1}$ ($10,000) | 27.04 | 17.13 | Zillow |
| equity $prices_{t-3,t-1}$ | 74.86 | 11.84 | Wilshire Index (FRED) |
| consumer $sentiment_{g,t-3,t-1}$ | 88.48 | 11.84 | Michigan SRC |
| consumer $expectations_{g,t-3,t-1}$ | 80.68 | 11.38 | Michigan SRC |

**Table 2. Unemployment, UI, and Fertility Rates**

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| unemp rate$_{c,t-3,t-1}$ | -0.0069 | -0.0114 | -0.0119 | -0.0111 | 0.0060 | -0.0421 |
| | (0.0014)*** | (0.0014)*** | (0.0012)*** | (0.0013)*** | (0.0096) | (0.0115)*** |
| unemp rate$_{c,t-3,t-1}$ * liquidity constrained rate$_{g}$ | | | | | -0.0864 | |
| | | | | | (0.0503)* | |
| unemp rate$_{c,t-3,t-1}$ * ui replacement$_{g,s,t-3,t-1}$ | | | | | | 0.0578 |
| | | | | | | (0.0212)*** |
| ui replacement$_{g,s,t-3,t-1}$ | | | | | | -0.1626 |
| | | | | | | (0.2764) |
| house prices$_{c,t-3,t-1}$ * home own rate$_{g}$ | | 0.0436 | 0.0440 | 0.0440 | 0.0447 | 0.0438 |
| | | (0.0044)*** | (0.0045)*** | (0.0045)*** | (0.0046)*** | (0.0045)*** |
| house prices$_{c,t-3,t-1}$ | | -0.0228 | -0.0222 | -0.0221 | -0.0226 | -0.0222 |
| | | (0.0020)*** | (0.0021)*** | (0.0021)*** | (0.0021)*** | (0.0021)*** |
| equity prices$_{t-3,t-1}$ * equities own rate$_{g}$ | | | 0.0146 | 0.0141 | 0.0153 | 0.0140 |
| | | | (0.0011)*** | (0.0011)*** | (0.0015)*** | (0.0011)*** |
| consumer sentiment$_{g,t-3,t-1}$ | | | | 0.0060 | 0.0070 | 0.0062 |
| | | | | (0.0016)*** | (0.0014)*** | (0.0015)*** |
| consumer expectations$_{g,t-3,t-1}$ | | | | 0.0028 | 0.0017 | 0.0026 |
| | | | | (0.0013)** | (0.0012) | (0.0013)** |
| N | 1,642,728 | 1,642,728 | 1,642,728 | 1,642,728 | 1,642,728 | 1,642,728 |

Notes: Estimated according to equation 1. Includes county, month of conception, and age-group-race/ethnicity group fixed effects. Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC, and SCF. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.05 **p<0.01 ***p<0.001

## Table 3: Unemployment, UI, and Fertility Rates, by Birth Parity

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| *Dependent Variable:* | | First Births | | | Second Births | | | Third+ Births | |
| unemp $rate_{c,t-3,t-1}$ | -0.0149 | 0.0126 | -0.0332 | -0.0115 | 0.0113 | -0.0310 | -0.0110 | 0.0108 | -0.0478 |
| | (0.0018)*** | (0.0105) | (0.0137)** | (0.0015)*** | (0.0102) | (0.0137)** | (0.0020)*** | (0.0101) | (0.0120)*** |
| unemp $rate_{c,t-3,t-1}$ * liquidity constrained $rate_g$ | | -0.1373 | | | -0.1146 | | | -0.1098 | |
| | | (0.0527)*** | | | (0.0508)** | | | (0.0514)** | |
| unemp $rate_{c,t-3,t-1}$ * ui $replacement_{g,s,t-3,t-1}$ | | | 0.0339 | | | 0.0361 | | | 0.0686 |
| | | | (0.0248) | | | (0.0251) | | | (0.0213)*** |
| N | 1,173,677 | 1,173,677 | 1,173,677 | 1,231,310 | 1,231,310 | 1,231,310 | 1,390,827 | 1,390,827 | 1,390,827 |

Notes: Estimated according to equation 1. Includes county, month of conception, and age-group-race/ethnicity group fixed effects; house prices, equity prices consumer sentiment and expectations. Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC, and SCF. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.05 **p<0.01 ***p<0.001

## Table 4. Unemployment, UI, and Infant Health

| | Percent Low Birth Weight (< 2500g) | | | | Percent Preterm (< 37 weeks) | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| unemp rate$_{c,t,t+8}$ | 0.0029 | | 0.0396 | 0.1747 | 0.0748 | | 0.1367 | 0.4252 |
| | (0.0209) | | (0.0245) | (0.0620)*** | (0.0284)*** | | (0.0361)*** | (0.0942)*** |
| unemp rate$_{c,t-3,t-1}$ | | -0.0113 | -0.0431 | -0.0440 | | 0.0281 | -0.0726 | -0.0737 |
| | | (0.0161) | (0.0148)*** | (0.0148)*** | | (0.0215) | (0.0221)*** | (0.0223)*** |
| unemp rate$_{c,t,t+8}$ * ui replacement$_{g,s,t,t+2}$ | | | | -0.2511 | | | | -0.5353 |
| | | | | (0.0928)*** | | | | (0.1400)*** |
| ui replacement$_{g,s,t,t+8}$ | | | | 1.3595 | | | | 5.6615 |
| | | | | (0.7230)* | | | | (1.2432)*** |
| house prices$_{c,t,t+8}$ * home own rate$_{g}$ | -0.0181 | -0.0185 | -0.0181 | -0.0177 | -0.0415 | -0.0425 | -0.0415 | -0.0423 |
| | (0.0042)*** | (0.0041)*** | (0.0042)*** | (0.0042)*** | (0.0066)*** | (0.0065)*** | (0.0066)*** | (0.0065)*** |
| house prices$_{c,t,t+8}$ | 0.0048 | 0.0030 | 0.0041 | 0.0044 | 0.0149 | 0.0123 | 0.0137 | 0.0146 |
| | (0.0074) | (0.0070) | (0.0074) | (0.0074) | (0.0103) | (0.0097) | (0.0103) | (0.0103) |
| equity prices$_{c,t,t+8}$ * equities own rate$_{g}$ | -0.0061 | -0.0047 | -0.0063 | -0.0058 | 0.0061 | 0.0053 | 0.0059 | 0.0072 |
| | (0.0037)* | (0.0033) | (0.0037)* | (0.0037) | (0.0049) | (0.0045) | (0.0049) | (0.0048) |
| consumer sentiment$_{g,t,t+8}$ | -0.0516 | -0.0580 | -0.0494 | -0.0493 | -0.1035 | -0.1131 | -0.0998 | -0.0921 |
| | (0.0345) | (0.0352)* | (0.0343) | (0.0340) | (0.0436)** | (0.0440)** | (0.0432)** | (0.0422)** |
| consumer expectations$_{g,t,t+8}$ | 0.0761 | 0.0799 | 0.0745 | 0.0749 | 0.1562 | 0.1639 | 0.1535 | 0.1491 |
| | (0.0339)** | (0.0342)** | (0.0337)** | (0.0334)** | (0.0411)*** | (0.0416)*** | (0.0408)*** | (0.0399)*** |
| N | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 |

Notes: Estimated according to equation 2. Includes county, quarter of conception, and age-group-race/ethnicity group fixed effects; the proportion of mothers who are married and the proportion of babies that are boys and twins . Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC, and SCF. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.05 **p<0.01 ***p<0.001

## Table 5: Unemployment, UI, and Infant Health, by Demographic Group

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | Low Birth Weight (<2500grams) | | | | | Preterm (<37 Weeks) | | | | |
| unemp rate$_{c,t,\ t+8}$ | 0.0832 | 0.6375 | -0.1072 | 0.1315 | 0.2247 | 0.2949 | 0.8428 | 0.3826 | 0.4291 | 0.4171 |
| | (0.0561) | (0.1937)*** | (0.2091) | (0.0718)* | (0.0816)*** | (0.0721)*** | (0.3033)*** | (0.2989) | (0.1204)*** | (0.1081)*** |
| unemp rate$_{c,t,\ t+8}$ * ui replacement$_{g,s,t,\ t+8}$ | -0.0827 | -0.8938 | 0.1731 | -0.1712 | -0.3597 | -0.3882 | -1.1838 | -0.3751 | -0.5612 | -0.5270 |
| | (0.0903) | (0.2897)*** | (0.3255) | (0.1044) | (0.1341)*** | (0.1085)*** | (0.4965)** | (0.5025) | (0.1809)*** | (0.1733)*** |
| unemp rate$_{c,t-3,t-1}$ | -0.0388 | -0.0751 | -0.0672 | -0.0077 | -0.0632 | -0.0514 | -0.0735 | -0.0879 | -0.0373 | -0.0948 |
| | (0.0163)** | (0.0451)* | (0.0307)** | (0.0093) | (0.0219)*** | (0.0241)** | (0.0564) | (0.0417)** | (0.0150)** | (0.0327)*** |
| Memo: Subsample | White | Black | Hispanic | 18-34 | 35-49 | White | Black | Hispanic | 18-34 | 35-49 |
| Dep. Var Mean | 7.8 | 14.6 | 7.2 | 7.7 | 9.3 | 12 | 18.8 | 12.5 | 11.4 | 14.2 |
| N | 784,787 | 403,463 | 391,102 | 954,167 | 625,291 | 784,787 | 403,463 | 391,102 | 954,167 | 625,291 |

Notes: Estimated according to equation 2. Includes county, quarter of conception, and age-group (columns 1-3 and 76-8) or race/ethnicity group (columns 4-5 and 9-10) fixed effects; the proportion of mothers who are married and the proportion of babies that are boys and twins; house prices, equity prices, consumer sentiment and expectations. Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC, and SCF. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.05 **p<0.01 ***p<0.001

## Table 6: Unemployment, UI, and Infant Health, by Trimester

| | Percent Low Birth Weight (< 2500g) | | | Percent Premature (< 37 weeks) | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| unemp rate$_{c,t+l,t+k}$ | 0.1796 | 0.1545 | 0.1284 | 0.3707 | 0.3621 | 0.3454 |
| | (0.0606)*** | (0.0589)*** | (0.0480)*** | (0.0821)*** | (0.0847)*** | (0.0861)*** |
| unemp rate$_{c,t+l,t+k}$ * ui replacement$_{g,s,t+l,t+k}$ | -0.2266 | -0.2354 | -0.2270 | -0.4475 | -0.4929 | -0.5176 |
| | (0.0906)** | (0.0885)*** | (0.0833)*** | (0.1264)*** | (0.1322)*** | (0.1373)*** |
| ui replacement$_{g,s,t+l,t+k}$ | 1.2969 | 1.2057 | 1.1739 | 5.1312 | 5.2522 | 5.4291 |
| | (0.7252)* | (0.6972)* | (0.6603)* | (1.1567)*** | (1.1868)*** | (1.2325)*** |
| unemp rate$_{c,t-3,t-1}$ | -0.0623 | -0.0342 | -0.0196 | -0.0759 | -0.0396 | -0.0257 |
| | (0.0157)*** | (0.0129)*** | (0.0158) | (0.0215)*** | (0.0183)** | (0.0200) |
| house prices$_{c,t+l,t+k}$ * home own rate$_{g}$ | -0.0179 | -0.0176 | -0.0180 | -0.0425 | -0.0422 | -0.0430 |
| | (0.0042)*** | (0.0041)*** | (0.0041)*** | (0.0067)*** | (0.0065)*** | (0.0065)*** |
| house prices$_{c,t+l,t+k}$ | 0.0033 | 0.0044 | 0.0048 | 0.0151 | 0.0143 | 0.0119 |
| | (0.0067) | (0.0074) | (0.0078) | (0.0092) | (0.0102) | (0.0109) |
| equity prices$_{c,t+l,t+k}$ * equities own rate$_{g}$ | -0.0053 | -0.0047 | -0.0047 | 0.0082 | 0.0095 | 0.0086 |
| | (0.0037) | (0.0035) | (0.0035) | (0.0048)* | (0.0047)** | (0.0047)* |
| consumer sentiment$_{g,t+l,t+k}$ | 0.0449 | -0.0479 | -0.0480 | 0.0428 | -0.0835 | -0.0504 |
| | (0.0230)* | (0.0234)** | (0.0305) | (0.0283) | (0.0289)*** | (0.0382) |
| consumer expectations$_{g,t+l,t+k}$ | -0.0062 | 0.0585 | 0.0590 | 0.0242 | 0.1102 | 0.0879 |
| | (0.0145) | (0.0223)*** | (0.0297)** | (0.0187) | (0.0272)*** | (0.0359)** |
| Memo: Trimester during which economic conditions and ui are measured | First l=0,k=2 | Second l=3,k=5 | Third l=6,k=8 | First l=0,k=2 | Second l=3,k=5 | Third l=6,k=8 |
| N | 1,637,912 | 1,608,595 | 1,579,463 | 1,637,912 | 1,608,595 | 1,579,463 |

Notes: Estimated according to equation 2. Includes county, quarter of conception, and age-group-race/ethnicity group fixed effects; the proportion of mothers who are married and the proportion of babies that are boys and twins; house prices, equity prices, consumer sentiment and expectations. Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC, and SCF. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.05 **p<0.01 ***p<0.001

**Table 7: Unemployment, UI, and Maternal Health Behaviors**

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Obtained prental care in 1st Trimester | | At least 5 prental care visits | | Smoked during pregnancy | |
| unemp rate$_{c,t,\ t+8}$ | 0.2626 | -0.2315 | 0.0188 | 0.0634 | -0.0807 | 0.5616 |
| | (0.0853)*** | (0.5722) | (0.0365) | (0.1017) | (0.0942) | (0.3067)* |
| unemp rate$_{c,t,\ t+8}$ * ui replacement$_{g,s,t,\ t}$ | | 0.9157 | | -0.0838 | | -1.1955 |
| | | (1.0577) | | (0.1780) | | (0.5102)** |
| unemp rate$_{c,t-3,t-1}$ | 0.1988 | 0.1984 | 0.0194 | 0.0179 | 0.0227 | 0.0184 |
| | (0.0556)*** | (0.0544)*** | (0.0221) | (0.0220) | (0.0370) | (0.0372) |
| | N=1,499,852; Dep. Var. Mean = 81.5 | | N=1568,648; Dep. Var. Mean = 95.6 | | N=1,436,963 ; Dep. Var. Mean = 7.9 | |

Notes: Estimated according to equation 2. Includes county, quarter of conception, and age-group-race/ethnicity group fixed effects; the proportion of mothers who are married; and the proportion of babies that are boys and twins; house prices, equity prices, consumer sentiment and expectations. Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.1 **p<0.05 ***p<0.001

**Appendix for "The Cyclicality of Births and Babies Health, Revisited: Evidence from Unemployment Insurance"**

## Appendix A: Data Construction

*Fertility rates and Infant Health Outcomes*

To construct county-group-month fertility rates and infant health outcomes, we applied for and obtained access to the confidential Vital Statistics birth certificate data from NCHS, which includes information on the mother's state and county of residence. We use SEER data on group-year female population counts to construct a fertility rate.

We assign births to the month of conception using the month and year of birth and the reported weeks of gestation, where we assume all births occur mid-month (because day or week of birth is missing from the data). When weeks of gestation are missing, we assume 40.

For the infant health outcomes, we construct the proportion of births (dated to conception) in each county-group-month with a low birth weight (defined as a birth weight under 2500 grams) and that are born prematurely (defined as a gestational period less than 37 weeks). In robustness checks, we also consider the following health outcomes: very low birth weight (below 1500 grams), very preterm (less than 32 weeks), and birth weight and the following behavioral outcomes: if the mother obtained prenatal care during the first trimester of pregnancy, if the mother had at least 5 prenatal care visits, and if the mother smoked during pregnancy. We also calculate the proportion of births to married women, and the proportion of babies that are twins and boys.

There are well known issues with missing data in the Vital Statistics data. In particular, the CDC updated birth certificate reporting requirements in 2003, but not every state altered their birth certificates to conform to the new reporting standards right away leading to missing data from 2009-2013. The states with missing information during at least some of the 2009-13 period are: AL, AK, AZ, AR, CT, HI, IL, LA, ME, MD, MA, MN, MS, MO, NJ, NC, RI, VA, WV, WI. This missing information affects a number of variables, including mother's educational attainment, smoking during pregnancy and the month prenatal care began. For the latter two, which are variables we use as robustness check outcomes, we drop the missing state-years.

*UI Benefit Generosity*

We construct potential weekly UI replacement rates by first constructing a state-by-year weekly UI benefit calculator. The UI benefit calculator is based on information we collected on weekly

benefit formulas found in Department of Labor publications from 2000-2019.[35] To construct the benefit calculator, we followed a similar procedure to Ganong, Noel and Vavra (2020) and recorded (a) the time period of wages the state uses (highest quarter, two highest quarters, three highest quarters, weekly, annual wages, etc.), (b) the percent of those period's wages used in the calculation of a weekly UI benefit, and (c) the cap on the maximum weekly benefit that can be received. When there were multiple formulas or maximums listed, we chose the higher formula or maximum.[36] When formulas include both a rate and intercept, we include both pieces of information in our calculation (in the table below, Pennsylvania is an example of such a formula). We do not use information on differences in benefits by the number of dependents.

The table below gives some illustrative examples of how we calculate the potential weekly UI replacement rate in various instances. Take the example of someone who earned $100,000 a year in Oregon. This person is eligible for the maximum benefit of $624 per week, since .0125 of their earnings exceeds that cap; the potential replacement rate is thus 624/(100,000/52) = .32. Now let's consider someone who earned $25,000 in the state of Florida. We assume evenly quarter earnings, so take ($25,000/4)*0.0385 = $240, which is below the cap of $275 and implies a potential replacement rate of 0.50.

| Annual Wage and Salary Earnings | State | Year | Wage Concept (w) | Formula | Maximum | Weekly UI Benefit | Potential Replacement Rate |
|---|---|---|---|---|---|---|---|
| $100,000 | OR | 2019 | Annual | 0.0125w | $624 | $624 | 0.32 |
| $50,000 | PA | 2017 | Highest Quarter | 0.0392w+1.96 | $569 | $492 | 0.51 |
| $25,000 | KY | 2007 | Annual | 0.0131w | $401 | $328 | 0.68 |
| $25,000 | FL | 2010 | Highest Quarter | 0.0385w | $275 | $240 | 0.50 |

[35] For 2006-2020, these can be found at URLs of the following format (modifying the year): https://oui.doleta.gov/unemploy/pdf/uilawcompar/2020/monetary.pdf. For years prior to 2006, the format for the URL is: https://oui.doleta.gov/unemploy/content/sigpros/1990-1999/July1990.pdf. For the maximum benefits, when possible, we used the Hsu, Matsa and Meltzer (2018) replication file data, which can be found at https://www.openicpsr.org/openicpsr/project/116160/version/V1/view.

[36] This differs from Ganong , Noel and Vavra (2020) who used the first formula listed. We chose the higher formula following Hsu, Matsa and Meltzer (2018), who chose the highest maximum benefit. The one exception to this rule is Alaska, where we used the lower formula and the highest maximum benefit. We did so because Alaska publishes tables with benefits amounts by income in each year, rather than using an explicit formula. When these tables are translated into a formula in the DOL publications it results in an unusually wide range, but only the very lowest income levels are eligible for the top of the range and the vast majority income levels are eligible for closer to the bottom of the range. Note that our results are robust to omitting Alaska entirely (not reported).

In our baseline we construct replacement rates using these benefit formulas for all women with non-zero wage and salary income between the age of 18-49 in the 2000 Census (based on the 5 percent sample, obtained from IPUMS-USA). We use a fixed sample of individuals to abstract away from cyclical fluctuations in population composition and realized incomes. We inflate this sample of women's wage and salary incomes using the CPI-U for each year of our sample (2000-2019), and then feed those annual real earned incomes through the UI benefit calculator based on each woman's state of residence for each year. Because Census records annual incomes, rather than quarterly incomes, we cannot exactly match many of the concepts used in the UI benefit calculator (e.g., highest quarter of the past four). Instead, we simply divide annual earned income by four to obtain quarterly income, two to two-quarter income, 52 to obtain weekly income, etc. We do not incorporate information on weeks or hours worked in these calculations. To account for federal expansions in benefit amounts, we add $25 a week to weekly benefits in 2009 and 2010, as provided by the Federal Additional Compensation program. Our constructed weekly UI replacement rate is the calculated weekly benefit amount divided by 1/52 times annual wage and salary income. Once we have the weekly benefit levels and replacement rates, we collapse by state, year, month, and group.

As noted above, in our baseline specification we construct replacement rates from earned income among women age 18 to 49 with non-zero wage and salary income. The collapsed median by state, year, and group thus reflects the replacement rate among this sample of women. In robustness checks, we use instead calculated means for this sample, as well as medians constructed from the following alternative samples: women who worked at least 26 weeks in the prior year (since many states have a 2-quarter work requirement for UI eligibility), men with non-zero wage and salary income, men who worked 26 weeks in the prior year, married men with non-zero wage and salary income, and unemployed men and women. Note that for the men's measures we use the exact same state-groups as we do for women, so the implicit assumption is that the economically dependent partner of a woman is in the same state-group as her.

We match the replacement rates to the birth data by calculating means of the each of these measures over (a) the three months prior to conception and (b) the nine months following (inclusive of) the month of conception. Although the data do not vary by month within in a state-year, many pregnancies scan multiple calendar years and by calculating means over these periods we allow for the possibility that state UI rules change prior to and within a pregnancy.

In other robustness checks, we also examine maximum durations of UI eligibility as an alternative measure of UI generosity. Regular and expansions to maximum durations were obtained from Farber, Rothstein and Valletta (2015) replication files. We use durations two ways, first, as a measure of generosity on its own and second, to scale our measure of *in utero* potential replacement rates to account for benefit exhaustion. That is, if maximum durations of

benefit eligibility are 13 weeks, we create a scaled potential replacement rate of 13/40 * replacement rate.

Appendix tables 1-4 display summary statistics on the various calculations of UI potential replacement rates.[37] Replacement rates are in general lower for men than women, and for individuals who worked more of the previous year, reflecting higher median incomes. Mean replacement rates are lower than medians, since replacement rates fall when benefit caps bind. Appendix Table 2 shows state benefit levels in 2019, and in column 7, for comparison, we add Ganong, Noel and Vavra (2020) median replacement rates for unemployed individuals in 2020 (obtained from Appendix table A1 of Ganong, Noel, and Vavra, 2020). Although their measure is based on a different sample (no age restriction, only individuals who were unemployed in 2020 and eligible for UI based on their work history), the resulting replacement rates are quite similar.

*Economic Variables*

We collected information on county-level unemployment rates from the Bureau of Labor Statistics Local Area Unemployment Statistics. We collected information on county-month house prices from Zillow. Specifically, we used the Zillow Home Value Index (ZHVI) which represents the value of a typical home. We put this value in real dollars using the CPI-U-less shelter series, and put the series in the $10,000 units. We captured equity prices using the average monthly value of the Wilshire-5000 index, which we downloaded from the St. Louis Federal Reserve FRED database. This series was adjusted to real dollars using the CPI-U.

We collected information on consumer sentiment and expectations using the University of Michigan Survey of Consumers, which we obtained from the Michigan Survey Research Center. We construct these measures by age-group and region of residence, which is the lowest level of geography available in those data. To construct these measures at this level, we average the region-level and age-group level indices in order to construct a region-group level index. The indices are defined as the percent of positive responses minus the percent of negative responses plus 100, implying that a higher value of the index can be interpreted as more positive sentiment/expectations.

We constructed group-level measures of asset ownership rates using the 2001-2019 Survey of Consumer Finances (SCF) using the SCF Bulletin measures of home ownership and direct or indirect equity ownership, where indirect equity ownership refers to ownership via retirement accounts, mutual funds, etc. We also used the SCF to construct a group-level

[37] Because this data begins in 2000, our analysis sample begins in 2000Q2 since we use the prior quarters' UI benefit information for our analyses.

measure of the incidence of liquidity constraints, which is the fraction of each group that report being credit constrained and having limited liquid savings. Credit constraints are defined as an affirmative response to either being denied credit or not applying for fear of denial, as is common in the literature (see, e.g., Bricker et al, 2020). Limited liquid savings is defined as having less than one month of one's own recurring expenses in liquid savings as in Bhutta and Dettling (2018), where recurring expenses include housing costs, food, vehicle costs, health costs, debt payments, etc. but exclude durables spending and liquid savings include all funds held in accounts or equities that can be accessed/sold without a fee or penalty, such as checking accounts, savings accounts, and directly held stocks and mutual funds.

*Individual-level Fertility Data*

We downloaded Current Population Survey (CPS) Annual Social and Economic Supplement (ASEC) data for 2001-2020 from IPUMS-CPS. The CPS ASEC data includes retrospective information on unemployment spells in the previous calendar year, the age of one's youngest child, various demographic information (age, race and ethnicity, marital status) and household information (home ownership), and state and county of residence (county is available for counties with over 100,000 residents). We define a birth as having occurred in the past year if the youngest own child in the home is under age 1 and we refer the prior calendar year as the "conception year". We limit the sample to women ages 18-49 as in our main analysis.

We merge our measure of unemployment insurance benefit generosity at the level of state and conception year. We also merge in information on the annual unemployment rates and median house prices at the level of conception year and county (when available) or state (when county is not available). For these measures, we define the state measure as the population weighted average of the county-level measures for all counties not reported in the CPS (that is, we omit counties for which we are able to merge the county-level data to the CPS from the state averages). Finally, we merge the measures of the sentiment and expectations by region and year and equity prices by year.

We use our state-group-year measure of UI benefit generosity to measure UI benefits rather than calculating individual-level UI benefits or using self-reported benefits for conceptual and practical reasons. Conceptually, as in our earlier analysis, we are interested in measuring changes in UI generosity owing to changes in policy rather than changes in the pool of unemployed individuals or changes in labor income owing to the economic environment. Practically, it would be difficult to construct such a measure because we only have information on income and hours worked for the portion of the prior calendar year that the individual was not unemployed, and furthermore, we do not know if that income and work occurred before or

after the unemployment spell. This means we cannot accurately estimate a benefit level or replacement rate. We do not use information on self-reported UI income because of known issues with severe underreporting of such benefits.[38]

[38] In our sample, only 20 percent of individuals who report being unemployed in the previous year report positive UI income. Similarly, only about 60 percent of individuals who report positive UI income report being unemployed last year.

## Appendix B: Additional Tables and Figures

Appendix Table 1: Calculated potential UI replacement rates by State

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | ui replacement$_{g,s}$ | ui benefit$_{g,s,t}$ | income$_{g,s,t}$ | max benefit$_{,s,t}$ |
| Alabama | 0.51 | 197 | 20,388 | 246 |
| Alaska | 0.47 | 255 | 28,047 | 396 |
| Arizona | 0.48 | 206 | 23,034 | 233 |
| Arkansas | 0.51 | 181 | 18,566 | 410 |
| California | 0.52 | 236 | 23,820 | 420 |
| Colorado | 0.53 | 254 | 25,079 | 472 |
| Connecticut | 0.50 | 284 | 29,283 | 591 |
| Delaware | 0.55 | 278 | 27,012 | 328 |
| District of Columbia | 0.48 | 282 | 32,112 | 361 |
| Florida | 0.49 | 219 | 23,375 | 275 |
| Georgia | 0.58 | 266 | 24,015 | 316 |
| Hawaii | 0.62 | 290 | 24,144 | 510 |
| Idaho | 0.51 | 176 | 17,987 | 368 |
| Illinois | 0.48 | 229 | 24,755 | 524 |
| Indiana | 0.53 | 229 | 22,725 | 386 |
| Iowa | 0.58 | 243 | 21,876 | 476 |
| Kansas | 0.56 | 231 | 21,491 | 430 |
| Kentucky | 0.66 | 248 | 19,657 | 400 |
| Louisiana | 0.59 | 214 | 19,112 | 272 |
| Maine | 0.60 | 239 | 20,797 | 514 |
| Maryland | 0.53 | 307 | 30,044 | 374 |
| Massachusetts | 0.51 | 284 | 29,044 | 928 |
| Michigan | 0.54 | 240 | 23,005 | 354 |
| Minnesota | 0.51 | 249 | 25,650 | 553 |
| Mississippi | 0.48 | 184 | 20,274 | 235 |
| Missouri | 0.52 | 227 | 22,473 | 301 |
| Montana | 0.53 | 179 | 17,486 | 460 |
| Nebraska | 0.51 | 204 | 20,827 | 335 |
| Nevada | 0.53 | 251 | 24,857 | 378 |
| New Hampshire | 0.53 | 256 | 25,321 | 396 |
| New Jersey | 0.60 | 340 | 29,211 | 569 |
| New Mexico | 0.54 | 210 | 20,112 | 439 |
| New York | 0.51 | 252 | 25,889 | 407 |
| North Carolina | 0.51 | 228 | 23,379 | 422 |
| North Dakota | 0.51 | 192 | 19,580 | 543 |
| Ohio | 0.51 | 219 | 22,523 | 504 |
| Oklahoma | 0.57 | 204 | 18,509 | 388 |
| Oregon | 0.66 | 268 | 21,248 | 489 |
| Pennsylvania | 0.56 | 247 | 22,798 | 542 |
| Rhode Island | 0.58 | 255 | 23,030 | 634 |
| South Carolina | 0.51 | 215 | 22,031 | 313 |
| South Dakota | 0.51 | 204 | 20,902 | 346 |
| Tennessee | 0.50 | 206 | 21,438 | 284 |
| Texas | 0.53 | 223 | 21,990 | 414 |
| Utah | 0.50 | 174 | 18,125 | 452 |
| Vermont | 0.58 | 234 | 20,804 | 401 |
| Virginia | 0.52 | 255 | 25,243 | 353 |
| Washington | 0.51 | 231 | 23,580 | 574 |
| West Virginia | 0.55 | 185 | 17,498 | 403 |
| Wisconsin | 0.53 | 239 | 23,605 | 356 |
| Wyoming | 0.53 | 179 | 17,663 | 449 |

Note: Displayed are median potential weekly UI replacement rates (Column 1), median potential UI weekly benefits (Column 2), and median annual wage and salary income (Column 3) for 2000-2019, based on a sample of women of child-bearing age with non-zero wage and salary incomes in the 2000 Census. UI benefits calculated as detailed in appendix. Median incomes inflated to 2000-2019 dollars using CPI-U in Column 3. Column 4 list the average state maximum benefit from 2000-2019. Source: Census, DOL, BLS.

Appendix Table 2: Median UI replacement Rates in 2019, Alternative groups and measures

| | Women, Wages>0 | Men, Wages>0 | Women, Worked >26 weeks | Men, Worked >26 weeks | Unemployed women and men | Women, Wages>0, Mean | Unemployed, Eligible for UI (Ganong and Noel, 2020) |
|---|---|---|---|---|---|---|---|
| Alabama | 0.48 | 0.47 | 0.39 | 0.37 | 0.50 | 0.43 | 0.47 |
| Alaska | 0.47 | 0.47 | 0.43 | 0.42 | 0.47 | 0.44 | 0.46 |
| Arizona | 0.47 | 0.42 | 0.37 | 0.34 | 0.51 | 0.41 | 0.34 |
| Arkansas | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.49 | 0.50 |
| California | 0.50 | 0.50 | 0.47 | 0.47 | 0.50 | 0.46 | 0.50 |
| Colorado | 0.50 | 0.50 | 0.50 | 0.49 | 0.50 | 0.48 | 0.60 |
| Connecticut | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.48 | 0.50 |
| Delaware | 0.52 | 0.50 | 0.44 | 0.40 | 0.55 | 0.46 | 0.57 |
| District of Columbia | 0.48 | 0.48 | 0.47 | 0.46 | 0.50 | 0.43 | NA |
| Florida | 0.47 | 0.47 | 0.41 | 0.39 | 0.50 | 0.42 | 0.47 |
| Georgia | 0.57 | 0.54 | 0.46 | 0.43 | 0.62 | 0.50 | 0.62 |
| Hawaii | 0.62 | 0.62 | 0.60 | 0.60 | 0.62 | 0.59 | 0.62 |
| Idaho | 0.50 | 0.50 | 0.47 | 0.47 | 0.50 | 0.48 | 0.50 |
| Illinois | 0.47 | 0.47 | 0.47 | 0.47 | 0.47 | 0.46 | 0.47 |
| Indiana | 0.47 | 0.47 | 0.43 | 0.43 | 0.47 | 0.44 | 0.47 |
| Iowa | 0.57 | 0.57 | 0.57 | 0.57 | 0.57 | 0.55 | 0.57 |
| Kansas | 0.55 | 0.55 | 0.52 | 0.51 | 0.55 | 0.52 | 0.55 |
| Kentucky | 0.62 | 0.62 | 0.57 | 0.57 | 0.62 | 0.59 | 0.62 |
| Louisiana | 0.65 | 0.60 | 0.47 | 0.42 | 0.68 | 0.57 | 0.39 |
| Maine | 0.59 | 0.59 | 0.59 | 0.59 | 0.59 | 0.58 | 0.59 |
| Maryland | 0.53 | 0.52 | 0.48 | 0.46 | 0.54 | 0.47 | 0.54 |
| Massachusetts | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 |
| Michigan | 0.53 | 0.52 | 0.43 | 0.40 | 0.53 | 0.47 | 0.53 |
| Minnesota | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.49 | 0.50 |
| Mississippi | 0.47 | 0.46 | 0.39 | 0.37 | 0.50 | 0.42 | 0.41 |
| Missouri | 0.52 | 0.50 | 0.45 | 0.43 | 0.52 | 0.46 | 0.51 |
| Montana | 0.52 | 0.52 | 0.52 | 0.52 | 0.52 | 0.51 | 0.52 |
| Nebraska | 0.50 | 0.50 | 0.48 | 0.48 | 0.50 | 0.48 | 0.50 |
| Nevada | 0.52 | 0.52 | 0.49 | 0.49 | 0.52 | 0.49 | 0.52 |
| New Hampshire | 0.52 | 0.52 | 0.45 | 0.45 | 0.52 | 0.48 | 0.48 |
| New Jersey | 0.60 | 0.60 | 0.56 | 0.56 | 0.60 | 0.57 | 0.60 |
| New Mexico | 0.54 | 0.54 | 0.52 | 0.52 | 0.54 | 0.52 | 0.53 |
| New York | 0.50 | 0.50 | 0.45 | 0.45 | 0.50 | 0.45 | 0.50 |
| North Carolina | 0.50 | 0.50 | 0.45 | 0.44 | 0.50 | 0.45 | 0.50 |
| North Dakota | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 |
| Ohio | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.49 | 0.50 |
| Oklahoma | 0.57 | 0.57 | 0.56 | 0.55 | 0.57 | 0.55 | 0.57 |
| Oregon | 0.65 | 0.65 | 0.63 | 0.62 | 0.65 | 0.62 | 0.51 |
| Pennsylvania | 0.51 | 0.51 | 0.51 | 0.51 | 0.52 | 0.51 | 0.51 |
| Rhode Island | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.49 | 0.50 |
| South Carolina | 0.50 | 0.49 | 0.45 | 0.43 | 0.50 | 0.45 | 0.49 |
| South Dakota | 0.50 | 0.50 | 0.48 | 0.48 | 0.50 | 0.48 | 0.50 |
| Tennessee | 0.48 | 0.47 | 0.40 | 0.38 | 0.50 | 0.43 | 0.44 |
| Texas | 0.52 | 0.52 | 0.50 | 0.50 | 0.52 | 0.50 | 0.52 |
| Utah | 0.48 | 0.49 | 0.48 | 0.48 | 0.47 | 0.45 | 0.49 |
| Vermont | 0.58 | 0.58 | 0.55 | 0.55 | 0.58 | 0.55 | 0.58 |
| Virginia | 0.52 | 0.51 | 0.46 | 0.45 | 0.52 | 0.46 | 0.52 |
| Washington | 0.50 | 0.50 | 0.50 | 0.50 | 0.50 | 0.49 | 0.50 |
| West Virginia | 0.55 | 0.55 | 0.52 | 0.52 | 0.55 | 0.52 | 0.55 |
| Wisconsin | 0.52 | 0.52 | 0.45 | 0.44 | 0.52 | 0.47 | 0.52 |
| Wyoming | 0.52 | 0.52 | 0.50 | 0.50 | 0.52 | 0.51 | 0.52 |

Note: Table displays replacement rates across various subsamples in 2019, except the final column which displays median replacement rates in 2020 from Ganong and Noel (2020), Appendix table A1.

Appendix Table 3: Replacement Rates and Unemployment Rates by Year

| Year | ui replacement$_{g,s,t}$ | unemp rate$_{c,t}$ |
|---|---|---|
| 2000 | 0.53 | 4.0 |
| 2001 | 0.52 | 4.4 |
| 2002 | 0.53 | 5.8 |
| 2003 | 0.53 | 6.0 |
| 2004 | 0.53 | 5.6 |
| 2005 | 0.53 | 5.1 |
| 2006 | 0.52 | 4.7 |
| 2007 | 0.52 | 4.6 |
| 2008 | 0.52 | 5.5 |
| 2009 | 0.57 | 8.9 |
| 2010 | 0.58 | 9.8 |
| 2011 | 0.53 | 9.2 |
| 2012 | 0.52 | 8.3 |
| 2013 | 0.52 | 7.6 |
| 2014 | 0.52 | 6.4 |
| 2015 | 0.52 | 5.4 |
| 2016 | 0.52 | 5.0 |
| 2017 | 0.52 | 4.5 |
| 2018 | 0.51 | 4.0 |
| 2019 | 0.51 | 3.7 |

Appendix Table 4: Replacement Rates and Unemployment Rates by Age and Race/Ethnicity

| | ui replacement$_{g,s,t}$ | unemp rate$_{c,t}$ |
|---|---|---|
| White (NH), 18-35 | 0.53 | 5.8 |
| Black (NH), 18-35 | 0.54 | 6.2 |
| Hispanic, 18-35 | 0.53 | 6.5 |
| White (NH), 35-49 | 0.52 | 5.9 |
| Black (NH), 35-49 | 0.53 | 6.3 |
| Hispanic, 35-49 | 0.52 | 6.5 |

Appendix Table 5: Different UI Replacement Rates, Fertility Rates

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| unemp rate$_{c,t-3,t-1}$ | -0.0421 | -0.0290 | -0.0410 | -0.0527 | -0.0493 | -0.0335 | -0.0533 | -0.0165 | -0.0118 |
| | (0.0115)*** | (0.0077)*** | (0.0137)*** | (0.0138)*** | (0.0133)*** | (0.0131)** | (0.0303)* | (0.0020)*** | (0.0018)*** |
| unemp rate$_{c,t-3,t-1}$ * ui replacement$_{g,s,t-3,t-1}$ | 0.0578 | 0.0322 | 0.0562 | 0.0853 | 0.0792 | 0.0484 | 0.1649 | 0.0014 | 0.0000 |
| | (0.0212)*** | (0.0139)** | (0.0257)** | (0.0290)*** | (0.0285)*** | (0.0296) | (0.0453)*** | (0.0004)*** | (0.0000) |
| Memo: Sample and measure used in ui replacement$_{g,s,t-3,t-1t}$ | Women with Wage income>0 | Women with Wage income>0, Mean instead of Median | Women, Worked 24 weeks last year | Men with Wage Income>0 | Men who worked 24 weeks last year | Married Men with Wage Income>0 | Married Men with Wage Income>0; Outcome is marital births | Maximum Benefit Level ($100s) | Maximum Weeks of Eligibility |
| N | 1,642,728 | 1,642,728 | 1,642,728 | 1,642,728 | 1,642,728 | 1,642,728 | 1,632,582 | 1,642,728 | 1,642,728 |

Notes: Estimated according to equation 1. Income measure refers to the group-level measure of income that is used to estimate UI replacement rates. Includes county, month of conception, and age-group-race/ethnicity group fixed effects; house prices, equity prices, consumer sentiment and expectations. Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC, and SCF. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.1 **p<0.05 ***p<0.001

Appendix Table 6: Individual-level unemployment, UI, and Probability of birth in the past year

| | (1) | (2) | (3) |
|---|---|---|---|
| $unemployed_{i,t-1}$ | -0.0060 | -0.0072 | -0.0232 |
| | (0.0011)*** | (0.0011)*** | (0.0091)** |
| $unemp\ rate_{c,t-1}$ | | -0.0003 | -0.0003 |
| | | (0.0003) | (0.0003) |
| $unemployed_{i,t-1}$ * $ui\ replacement_{g,s,t-1}$ | | | 0.0299 |
| | | | (0.0171)* |
| $ui\ replacement_{g,s,t-1}$ | | | -0.0432 |
| | | | (0.0287) |
| Controls for house prices, sentiment and expectations | No | Yes | Yes |
| Observations | 905,625 | 877,843 | 877,843 |

Notes: Estimated according to a modified version of equation 1 where the dependent variable is Pr(Baby Under Age 1) and the independent variable is an indicator for having been unemployed in the prior calendar year. Includes state, year, age, race/ethnicity group, and marital status fixed effects. Data sources is CPS ASEC, BLS LAU, Zillow, and Michigan SRC. Standard errors adjusted for clustering at the state-level in parentheses. *p<0.05 **p<0.01 ***p<0.001

Appendix Table 7: Unemployment, UI, and Fertility Rates, by Demographic Group

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| unemp $rate_{c,t-3,t-1}$ | -0.0045 | -0.0405 | 0.0201 | -0.0287 | -0.0194 |
| | (0.0111) | (0.0116)*** | (0.0189) | (0.0072)*** | (0.0156) |
| unemp $rate_{c,t-3,t-1}$ * ui $replacement_{g,s,t-3,t-1}$ | -0.0018 | 0.0657 | -0.0392 | 0.0324 | 0.0181 |
| | (0.0200) | (0.0199)*** | (0.0293) | (0.0122)*** | (0.0293) |
| Memo: Subsample | White | Black | Hispanic | 18-34 | 35-49 |
| N | 815,622 | 420,470 | 406,537 | 991,682 | 651,042 |

Notes: Estimated according to equation 1. Includes county, month of conception, and age-group-race/ethnicity group fixed effects; house prices, equity prices, consumer sentiment and expectations. Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC, and SCF. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.05 **p<0.01 ***p<0.001

Appendix Table 8: Different UI Replacement Rates, Infant Health

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| **Panel A: Percent Low Birth Weight** | | | | | | | | | |
| unemp $rate_{c,t,\ t+8}$ | 0.1747 | 0.1327 | 0.1653 | 0.1198 | 0.1157 | 0.1083 | 0.0558 | 0.0232 | 0.0215 |
| | (0.0620)*** | (0.0512)*** | (0.0613)*** | (0.0372)*** | (0.0371)*** | (0.0368)*** | (0.0387) | (0.0294) | (0.0415) |
| unemp $rate_{c,t,\ t+8}$ * ui $replacement_{g,s,t,\ t+8}$ | -0.2511 | -0.1665 | -0.2363 | -0.1676 | -0.1614 | -0.1522 | -0.0042 | 0.0003 | 0.0370 |
| | (0.0928)*** | (0.0607)*** | (0.0954)** | (0.0596)*** | (0.0565)*** | (0.0537)*** | (0.0049) | (0.0002) | (0.0589) |
| unemp $rate_{c,t-3,t-1}$ | -0.0440 | -0.0452 | -0.0440 | -0.0447 | -0.0447 | -0.0449 | -0.0434 | -0.0452 | -0.0430 |
| | (0.0148)*** | (0.0150)*** | (0.0148)*** | (0.0148)*** | (0.0149)*** | (0.0149)*** | (0.0150)*** | (0.0154)*** | (0.0149)*** |
| **Panel B: Percent Premature** | | | | | | | | | |
| unemp $rate_{c,t,\ t+8}$ | 0.4252 | 0.3202 | 0.3805 | 0.2494 | 0.2482 | 0.2271 | 0.2002 | 0.0995 | 0.1153 |
| | (0.0942)*** | (0.0657)*** | (0.0874)*** | (0.0499)*** | (0.0488)*** | (0.0471)*** | (0.0564)*** | (0.0412)** | (0.0602)* |
| unemp $rate_{c,t,\ t+8}$ * ui $replacement_{g,s,t,\ t+8}$ | -0.5353 | -0.3284 | -0.4574 | -0.2357 | -0.2366 | -0.1998 | -0.0158 | 0.0008 | 0.0419 |
| | (0.1400)*** | (0.0816)*** | (0.1309)*** | (0.0750)*** | (0.0712)*** | (0.0692)*** | (0.0084)* | (0.0003)** | (0.0952) |
| unemp $rate_{c,t-3,t-1}$ | -0.0737 | -0.0767 | -0.0740 | -0.0749 | -0.0750 | -0.0750 | -0.0745 | -0.0701 | -0.0728 |
| | (0.0223)*** | (0.0223)*** | (0.0222)*** | (0.0221)*** | (0.0221)*** | (0.0221)*** | (0.0224)*** | (0.0226)*** | (0.0222)*** |
| Memo: Sample used to construct ui $replacement_{g,s,t,\ t+8}$ | Women with Wage income>0 | Women with Wage income>0, Mean instead of Median | Women, Worked 24 weeks last year | Men with Wage Income>0 | Men who worked 24 weeks last year | Married Men with Wage Income>0 | Maximum Benefit Level ($100s) | Maximum Duration | Women with wage income>0, Add Durations |
| N | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 | 1,579,463 |

Notes: Estimated according to equation 2. Includes county, quarter of conception, and age-group-race/ethnicity group fixed effects; the proportion of mothers who are married; and the proportion of babies that are boys and twins; house prices, equity prices, consumer sentiment and expectations. Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.1 **p<0.05 ***p<0.001

Appendix Table 9: Unemployment, UI and Alternative Infant Health Outcomes

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | Very Low Birth Weight (< 1500g) | | Birth Weight | | Very Preterm (<32 weeks) | | Gestational Age | | Share Male | |
| unemp rate$_{c,t,\ t+8}$ | 0.0423 | 0.0881 | -1.2906 | -4.8899 | 0.0613 | 0.1296 | -0.1227 | -2.1401 | 0.0123 | 0.0161 |
| | (0.0178)** | (0.0392)** | (0.6424)** | (1.7212)*** | (0.0181)*** | (0.0403)*** | (0.3628)*** | (0.9404)** | (0.0261) | (0.0692) |
| unemp rate$_{c,t,\ t+8}$ * ui replacement$_{g,s,t,\ t-}$ | | -0.0850 | | 6.6955 | | -0.1267 | | 1.6943 | | -0.0072 |
| | | (0.0475)* | | (2.5709)*** | | (0.0527)** | | (1.4034) | | (0.1185) |
| unemp rate$_{c,t-3,t-1}$ | -0.0239 | -0.0241 | 0.8679 | 0.9032 | -0.0330 | -0.0333 | 0.5169 | 0.5211 | 0.0201 | 0.0201 |
| | (0.0080)*** | (0.0080)*** | (0.4352)** | (0.4328)** | (0.0095)*** | (0.0095)*** | (0.2150)** | (0.2140)** | (0.0226) | (0.0226) |
| | N=1,579,463 ; Dep. Var. Mean = 1.6 | | N=1,579,463; Dep. Var. Mean = 3291 | | N=1,579,463; Dep. Var. Mean = 2.1 | | N=1,579,463; Dep. Var. Mean = 38.5 | | N=1,579,463; Dep. Var. Mean = 51.1 | |

Notes: Estimated according to equation 2. Includes county, quarter of conception, and age-group-race/ethnicity group fixed effects; the proportion of mothers who are married; and the proportion of babies that are boys and twins; house prices, equity prices, consumer sentiment and expectations. Data sources are Vital Statistics, SEER, BLS LAU, Zillow, Michigan SRC. Standard errors adjusted for clustering at the county-level in parentheses. *p<0.1 **p<0.05 ***p<0.001