Fertility Response to the Tax Treatment of Children

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Abstract

One of the most commonly cited studies on the effect of child subsidies on fertility, Whittington, Alm, and Peters (1990), claimed a large positive effect of child tax benefits on fertility using time series methods. We revisit this question in light of recent increases in child tax benefits by replicating this earlier study and extending the analysis with an additional 20 years of data. We find that their results suffer from the spurious regression problem, and are not robust to differencing. We find evidence of a statistically significant fertility response to a change in the real value of child tax subsidies occurring with a one- to two-year lag, but a much smaller and statistically insignificant total effect after several years, suggesting that a change in the child tax subsidy most strongly affects the timing of births.

We would like to thank Brigitte Madrian for generously providing access to a letter from Leslie Whittington. We would also like to thank participants in the Stanford Macro Bag Lunch, Peter Hansen, and Mohitosh Kejriwal for helpful comments.

1 Introduction

Standard economic theory tells us that the demand for children is influenced by the cost of raising children. Holding other things constant, a decrease in the cost of raising children should lead to an increase in the demand for children. In recent years, the value of child tax benefits has increased substantially relative to estimates of the cost of raising children. As shown in Figure 1, the average value of the U.S. child tax subsidy adjusted for inflation has increased from under \$850 in 1980 to more than \$2,000 in 2005.[1] Using the U.S.D.A. estimates of expenditures on children in the U.S., the \$2,000 annual subsidy represents between 13 and 49 percent of the average U.S. expenditure on children, depending on how child expenditure is measured (Lino, 2006). Thus, the \$1,150 real increase in child tax benefits can be thought of as a 7 to 28 percent discount in the cost of raising children. How much of an effect (if any) did this reduction in the cost of raising children have on fertility? While a very important empirical question, the magnitude and timing of the fertility response to child tax benefits has received little attention.

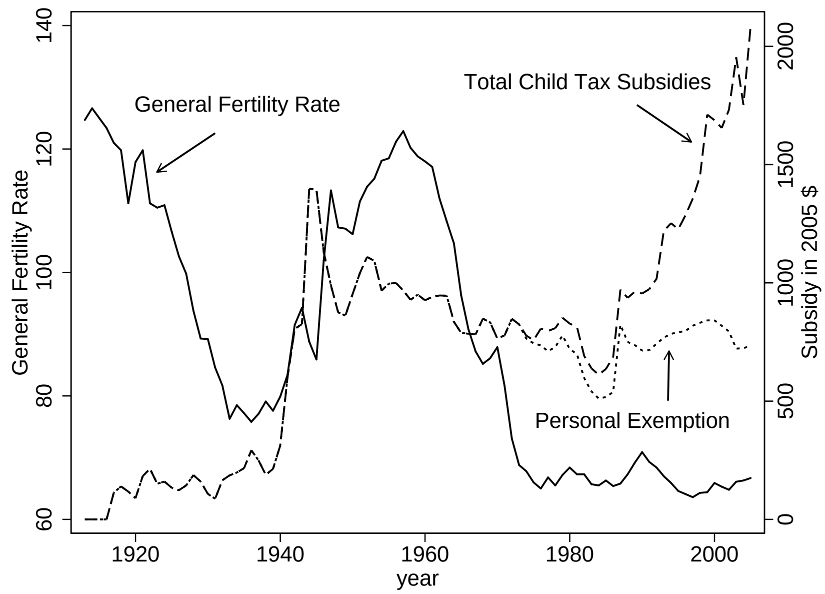
Whittington, Alm, and Peters (1990) was the first to seriously estimate the responsiveness of fertility to child tax benefit changes. Their analysis of time series data from 1913 to 1984 suggests that the U.S. fertility rate is very responsive to child tax benefits. They estimate that a \$100 increase (in 2005 dollars) in the tax value of the personal exemption would increase the general fertility rate by 2.1 to 4.2 births (a 3.2 to 6.5 percent increase).[2]

While the sign of the estimated effect is not unexpected, the magnitude of the Whittington et al. (1990) estimate is surprising. If a \$100 increase in annual child tax benefits could increase fertility by 6 percent, why are European countries, many with very generous and salient child subsidy programs that are not hidden in the complexities of the income tax, experiencing low and stagnant fertility rates? Or, to take the recent increase in U.S. child

[1] The details regarding the calculation of the average per-child tax subsidy are given in Section 4.

[2] Whittington, Alm, and Peters report their results in 1967 dollars. Their estimates of the effect of the value of the personal exemption in 1967 dollars on the general fertility rate range from 0.121 to 0.236. Converting the dollar amounts to 2005 dollars using the CPI-U, we find that their estimates range from 0.021 to 0.042.

Figure 1: General Fertility Rate and Real Average Per Child Tax Subsidy



tax benefits, should we have expected a 32 to 65 percent increase in the U.S. fertility rate in response to the $1,000 Child Tax Credit, holding all other factors constant?[3]

Since Whittington et al. (1990), there have only been a handful of empirical studies that estimate a fertility response from changes in child tax benefits or other child subsidies. The closest study in terms of methodology is Zhang et al. (1994). This paper employs the same specification and time series techniques, but uses Canadian data from 1921 to 1988. They find that the responsiveness of Canadian fertility to child tax benefits was about one half the magnitude reported by Whittington et al.[4] They interpret their results to be consistent with the Whittington et al. findings and explain the smaller magnitude as likely due to differences in the child benefit programs between Canada and the United States.

[3]From 1997 (the year the Child Tax Credit was passed) to 2005, the general fertility rate in the United States increased by 4.9 percent. Note however that eligibility restrictions and interactions in the tax code make the $1,000 Child Tax Credit worth much less than this amount on average. From 1997 to 2005, the average child subsidy increased by approximately $550 in real terms.

[4]Zhang and his coauthors use the total fertility rate instead of the general fertility rate as the measure of fertility. Adjusting the coefficient estimates to make them directly comparable to Whittington et al. (1990) we find that they range from 0.011 to 0.024.

Gauthier and Hatzius (1997) find evidence of similar magnitude to the results of Zhang et al. (1994), but only in some countries of their 22 country panel. Cohen, Dehejia and Romanov (2007) examine the role of financial incentives on fertility in Israel and find strong effects among low-income populations. Huang (2002), using time-series data from Taiwan, also finds evidence of a smaller fertility response than was reported by Whittington et al. (1990). Baughman and Dickert-Conlin (2003) examine the effect of the expansion of the Earned Income Tax Credit (EITC) in the 1990s on first-birth rates in the United States. They find no economically significant fertility response among unmarried women and only a small response for married women. By race, the largest estimated fertility response is for married non-white women, but even this estimate is less the half the magnitude reported in Whittington et al. Laroque and Salanie (2005) use individual data combined with detailed tax and child program information to estimate a structural model of fertility. While child subsides are quite generous in France, they find evidence of only a small effect on fertility.

Only Milligan (2005) reports estimates of the fertility response to child tax benefits of a similar magnitude as Whittington et al. (1990). Milligan finds a large effect on fertility from a 1988-1997 child subsidy program in Quebec. However, this large fertility effect is in part due to the temporary nature of the Quebec subsidy program; women may have had children earlier in order to claim the subsidy with no change in their completed fertility. Parent and Wang (2007) supports this hypothesis. Even so, Whittington et al. is cited by an increasing number of publications (many in non-economics journals) as evidence of a strong link between child tax benefits and fertility. Perhaps this is because the Whittington et al. paper came first and was published in the *American Economic Review*, while most of the other studies which indicate a smaller fertility response have been published in less prominent journals.

In this paper, we revisit and extend the analysis in Whittington et al. (1990) for several reasons. First, the availability of 21 additional years of data and the recent large increases in child tax benefits provide an opportunity to reassess the relationship between child tax benefits and fertility. The addition of more recent data provides estimates of this relationship that are perhaps more relevant in the current environment of large (and increasing) child tax subsidies. It also allows us to examine child tax subsidies other than those used by Whittington et al., who examined the response only to the real value of the personal exemption. In our extended analysis, we include the child tax benefits from the earned income tax credit, the child and dependent care tax credit, and the child tax credit.

Second, more recent time series methods allow us to address shortcomings in the original analysis and provide more nuanced interpretations of our results. We test for unit roots in the data series, and examine both the presence of a long-run equilibrium relationship between fertility and child tax subsidies and the short-run relationship between changes in tax benefits and changes in fertility. Lastly, more current published data sources for several data series are now available.

We start by replicating the original Whittington et al. (1990) study. We are able to reproduce Whittington et al.'s main findings, and we show that the results are likely due to the spurious regression problem and are not robust to correcting for this problem by differencing the data. We then extend the analysis using data through 2005 to examine what the additional data and more recent econometric techniques show about the responsiveness of fertility to the recent increases in the federal income tax subsidies for children. We find that changes in tax benefits two years prior are positively associated with changes in fertility rates in the current year. However, the overall effect of tax benefits on fertility is approximately half the magnitude found in Whittington et al. and not significantly different from zero. The results suggest that the timing of births may be influenced by tax subsidies more than the level of the general fertility rate. We do not find evidence of a long-run equilibrium relationship between child tax subsidies and the general fertility rate.

We employ a time-series approach and do not explore the other potential methods for estimating the fertility response to changes in the child subsidy level. Instead, we focus our efforts on the time-series methods employed and on the data used, but acknowledge that the identification of the fertility response to a change in the value of child tax benefits may be vulnerable to trends in unobserved variables. Milligan (2005) argues that time-series variation is not sufficient to identify fertility effects from child subsidy changes if unobservable characteristics important for child-bearing decisions change through time and are unobservable to the econometrician.

The time-series approach that we follow is vulnerable to this criticism. However, that the fertility rate is affected by unobservable factors is not enough to question the validity of the time-series results. One must argue in addition that at least one of these unobservable factors is correlated with an explanatory variable. For example, if U.S. households experience a growing preference for a larger family size (a pure change in tastes) over the same period in which there is an increase in the value of child tax benefits, the increase in the fertility rate due to the change in preferences may be incorrectly viewed as a response to the policy change. Removing trends or differencing the series would not correct the basic spurious correlation. With respect to the family size preference example, analysis of responses to questions about intended or ideal family size in the General Social Survey performed by Hagewen and Morgan (2005) reveals that there is "remarkable stability" in both the ideal family size and fertility intentions since 1972 when the survey began. In contrast, the real value of child tax benefits is far from stable; it declines slightly from the 1970's to the mid 1980's and then increases dramatically over the 1990's and into the current decade.

Section 2 describes our reconstruction of the dataset used in Whittington et al (1990). Section 3 describes the estimation methods used to replicate Whittington et al. and provides evidence for the spurious relationship between fertility and tax benefits reported in their earlier work. Section 4 describes refinements to the 1913-1984 data and extending the data through 2005. Section 5 reports the estimation results using the extended data and discusses the timing of the fertility response, and Section 6 concludes. The complete datasets and details of the data construction are given in the Appendix.

2 Data for Replication

Whittington et al. (1990) regressed the annual fertility rate from 1913 to 1984 on a set of explanatory variables that they argued would affect fertility: male and asset income, unemployment, infant mortality, immigration, female wage, and binary variables for World War II and the availability of the birth control pill. Some of the series that were not reported in the appendix of the published paper have been lost since the paper's publication. We have tried to reconstruct the missing series using the footnotes and references in Whittington et al. Here, we present only a brief summary of the data and how they were constructed. Remaining details as well as the full data series are found in Appendix A.

The dependent variable is the general fertility rate, the number of births per thousand women age 15-44. This measure is not as dependent on the age structure of the population as the simple birthrate which measures the number of births divided by the total population. However, the general fertility rate is affected by changes in the age structure within the 15-44 group.

The primary variable of interest for Whittington et al. (1990) is the real tax value of the personal exemption for dependents. Today, the personal exemption is only one of several child subsidy provisions in the federal tax code accounting for about one-third to one-half of the total child subsidy. However, for the 1913-1984 period considered in Whittington et al., the personal exemption was the primary source of the implicit child subsidy, never accounting for less than 85 percent of the total child subsidy. The statutory value of the personal exemption for dependents changed only nine times between 1913 and 1984; however, its real tax value fluctuates substantially due to changes in marginal tax rates and the price index.

The general fertility rate, value of the personal exemption, and the female wage series which was constructed by Whittington et al. to measure the real change in average female wages, were each reported in the paper's appendix. The introduction of the birth control pill and U.S. involvement in World War II are simple binary variables that equal one after

1963 for the birth control pill and between 1941-1945 for World War II. The male and asset income series is a measure of average family income less female earnings. While this series was not reported in the appendix of Whittington et al. (1990), it was recorded in a letter from Leslie Whittington.[5]

We were unable to obtain the original data for unemployment, infant mortality, and immigration which required reconstruction of these series by following the description in Whittington et al. (1990). The unemployment series is from Lebergott (1964) and the U.S. Census Bureau (2003).[6] The infant mortality series is from the U.S. Census Bureau (2003) and measures the number of children who die before reaching their first birthday (excluding fetal deaths), per thousand children born.[7] The immigration series measures the number of immigrants age 16-44 as a fraction of the resident population in the same age group. These data were obtained from various versions of the Historical Statistics of the United States as described in Appendix A.

It is clear from Table 1 that there are small differences between the reconstructed series and those used in Whittington et al. (1990). Differences in the data seem to be present even for some series that we copied directly from the Whittington et al. appendix. In fact, of those series for which we obtained original data (general fertility rate, personal exemption, male and asset income, and female wage), only the personal exemption series exactly matches the reported moments. The general fertility rate and female wage series reported in the appendix are either different than the series used to report the summary statistics or some error was made in computing the mean and standard deviation.[8] The male and asset income series

[5]Brigitte Madrian generously gave us access to a 1991 letter she received from Leslie Whittington in which the full male and asset income series used in Whittington et al. (1990) is reported.

[6]The U.S. Census Bureau unemployment data are from the Statistical Abstract of the United States: 2003, Mini-Historical Series HS-29 which covers the 1929-1984 period. The Lebergott data are from table A-3 and only the 1913-1928 data is used.

[7]The infant mortality data are from the Statistical Abstract of the United States: 2003, Mini-Historical Series HS-13.

[8]According to the letter received by Brigitte Madrian, the average female wage index values for 1972 and 1919 were typos. However, correcting these typos leads to greater discrepancies between both the reported moments and the replication results, so we use the series as reported in Whittington et al. in the replication analysis.

reported in previous correspondence from Leslie Whittington has the same problem. The unemployment, infant mortality, and immigration series that we constructed quite accurately match the reported moments.

Table 1: Summary Statistics, 1913–1984

		Replicated Data		Whittington et al.	
Variable	Obs.	Mean	Std. Dev.	Mean	Std. Dev.
General Fertility Rate	72	95.6	19.81	95.5	19.64
Personal Exemption	72	100.4	65.88	100.4	65.88
Male and Asset Income	72	7,467.38	2,926.06	7,466.37	2,982.78
Unemployment	72	0.071	0.054	0.071	0.053
Infant Mortality	72	43.02	26.84	43.02	26.84
Immigration	72	0.003	0.0036	0.003	0.0035
Female Wage	72	1.35	0.585	1.22	0.532
Pill	72	0.306	0.464	0.305	0.464
WW II	72	0.069	0.256	0.069	0.256
Time Trend	72	36.5	20.93	36.5	20.92

Variables expressed in constant 1967 dollars.

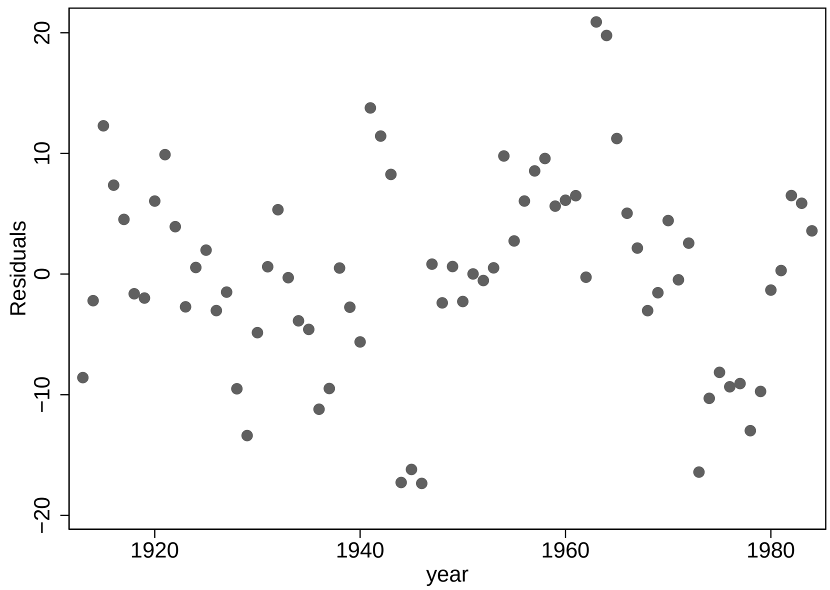
3 Replication and New Methods

Following Whittington et al. (1990) we estimate the following reduced form equation for the period 1913 to 1984:

$$\begin{aligned} \text{Fertility Rate}_t = \beta_0 &+ \beta_1 \,\text{Personal Exemption}_t + \beta_2 \,\text{Male and Asset Income}_t \\ &+ \beta_3 \,\text{Unemployment}_t + \beta_4 \,\text{Infant Mortality}_t + \beta_5 \,\text{Immigration}_t \\ &+ \beta_6 \,\text{Female Wage}_t + \beta_7 \,\text{Pill}_t + \beta_8 \,\text{WW2}_t + \beta_9 \,\text{Time Trend}_t + \epsilon_t. \end{aligned} \tag{1}$$

Whittington et al. (1990) give a general description of their estimation method: "Generalized least squares estimation is performed with a Yule-Walker first-order autocorrelation correction scheme," motivated by a stated concern about serial correlation. A Durbin-Watson test using the OLS residuals of our replicated data confirms the presence of serial correlation. The OLS residuals from a regression of Equation (1) are plotted in Figure 2. The residuals exhibit a clear cyclical pattern rather than white noise, indicating that the errors are indeed serially correlated.

Figure 2: OLS Residuals



We believe the somewhat vague terminology for the estimation procedure in Whittington et al. (1990) refers to FGLS estimation of an AR(1) model, such as a Cochrane-Orcutt or Prais-Winsten estimation procedure. Although not stated, some experimentation leads us to the conclusion that Whittington et al. uses a two-step procedure. In the first step, the OLS residuals are obtained in order to estimate ρ, where $u_t = \rho u_{t-1} + \epsilon_t$. The now-estimated $\hat{\rho}$ is used to quasi-difference all variables:

$$\tilde{x}_t = x_t - \hat{\rho}\, x_{t-1}. \tag{2}$$

The second step is to run an OLS regression using the quasi-differenced variables. This procedure gives the final results for the Cochrane-Orcutt estimation method. The Prais-Winsten method is similar, but uses, in addition, the first observation transformed by the factor $\sqrt{(1-\hat{\rho}^2)}$ in the OLS regression.

Although we believe that Whittington et al. (1990) uses a two-step estimation procedure similar to Cochrane-Orcutt or Prais-Winsten, the details of the procedure are not precisely specified. In addition, since we have slightly different data, it is unlikely that we would exactly replicate the Whittington et al. results. We report the original estimates of the primary specification as reported in Whittington et al. as Model (1) in Table 2. Next, we report the regular OLS estimates using the replicated data with Newey-West standard errors (robust to serial correlation) as Model (2) in Table 2. Finally, we report the results from two-step Prais-Winsten estimation using the replicated data with $\hat{\rho}$ based on adjusted autocorrelation as Model (3) in Table 2.

Model (3) closely replicates the original Model (1) results. The estimated coefficient on the tax value of the personal exemption is very close to the reported value in Whittington et al. (1990). In addition, the remaining coefficient estimates are also similar to Whittington et al.'s results. Slight differences in the data (including the series that were obtained from the paper itself) likely explain deviations from the original results.[9]

At first glance, Model (3) does not seem to fit the data as well as Model (1), as measured by the R^2. However, we believe the R^2 reported in Whittington et al. (1990) is inflated. Because the estimation method was not a pre-programmed procedure at the time Whittington et al. was written, it was likely performed manually. The manual implementation involves transforming each variable, and the regression performed in the second stage does not have a true constant term; therefore, the definition of the R^2 is ambiguous. Using the total sum of squares from the original OLS regression run in the first step of the Prais-Winston procedure (i.e. the non-transformed general fertility rate) and the sum of squared residuals from Model (3) yields an R^2 of 0.919. While this technique does not give an accurate description of the fit of Model (3), it represents a plausible method that Whittington et al. may have used to arrive at their reported R^2 of 0.916.

[9]Currently, there are several different methods possible to estimate ρ in the first step of the estimation procedure. Among the possible methods, estimating ρ by OLS gives the strongest fit to Whittington et al.'s results, and yields an estimate of $\hat{\rho} = 0.5963625$. This value of $\hat{\rho}$ was used to quasi-difference the variables, and transform the first observation. Estimating Equation 1 with $\hat{\rho} = 0.581$ yields an estimate of the key coefficient, β_1, as 0.121, the same as the reported estimate in Whittington et al. However, using this value for $\hat{\rho}$ does not eliminate the deviations in the other estimated coefficients.

Table 2: Comparison of Estimation Results

Variable	(1) Whittington et al.	(2) OLS	(3) Prais-Winsten
Personal Exemption	0.121	0.178	0.116
	(0.0446)**	(0.0977)	(0.0449)**
Male and Asset Income	-0.0004	0.0035	0.0007
	(0.0027)	(0.0031)	(0.0025)
Unemployment	-73.43	-68.12	-68.19
	(34.20)**	(25.818)*	(34.004)**
Infant Mortality	0.083	0.393	0.0351
	(0.255)	(0.321)	(0.251)
Immigration	774.24	964.13	760.71
	(311.31)**	(329.44)**	(304.98)**
Female Wage	5.647	15.427	5.629
	(15.686)	(5.286)**	(5.036)
Pill	-10.856	-25.383	-12.014
	(6.126)*	(11.961)*	(6.028)*
WW II	-17.223	-29.419	-17.863
	(4.989)**	(8.057)**	(4.854)**
Time Trend	-0.539	-0.843	-0.741
	(0.538)	(0.543)	(0.510)
Intercept	102.979	55.944	104.130
	(24.666)**	(25.831)*	(23.368)**
R^2	0.916	0.829	0.749

Standard errors in parentheses.
Model (2) reports Newey-West standard errors.
Variables expressed in constant 1967 dollars.
* significant at the 10% level
** significant at the 5% level

Zhang et al. (1994) mention that there is a concern that some series in both their study and the Whittington et al. (1990) study may be non-stationary. Therefore, both the Zhang et al. and the Whittington et al. results could suffer from what is called the spurious regression problem if some of the series are integrated of order one, or I(1). This is a well known criticism of Whittington et al. findings.[10] However, Zhang et al. dismiss these criticisms claiming that because the "time trend is insignificant" in their estimation, there is no concern that the results are being driven by a regression of "time against time." They argue further that when the time trend is dropped, "the result on the tax-transfer variables holds and [the] R squared is virtually unchanged."

However, an insignificant time trend coefficient estimate does not alleviate concerns that some of the series are I(1). Granger and Newbold (1974) in their seminal paper on spurious regressions argue that annual macro series, like those used in this study, are almost always I(1); thus, regressions involving the levels will be misleading, suggesting relationships when there may be none. Careful inspection of the data reveals that many of the Whittington et al. (1990) series do exhibit unit root behavior and should therefore be treated as I(1).[11] An initial approach to determine which series are I(1) is to regress each variable on its lagged value, and estimate the AR(1) model:

$$x_t = \alpha + \rho x_{t-1} + \epsilon_t \tag{3}$$

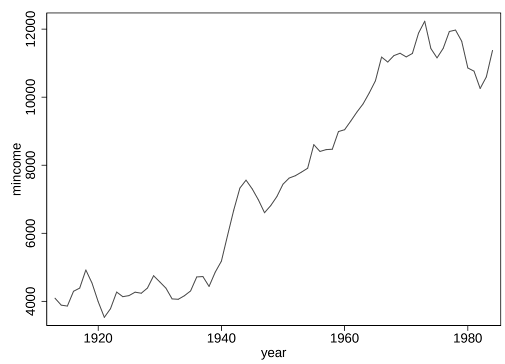
A series has a unit root if the true population parameter ρ equals one. We first determine if each series has a clear time trend, α. From inspection of the series in Figure 3, it appears that Male and Asset Income, Infant Mortality, and Female Wage have a time trend. If the series did not have a clear time trend, the constant was omitted from the AR(1) regression.

[10] Wooldridge (2006), a popular undergraduate econometrics text, uses the Whittington et al. data as an example of the spurious regression problem.

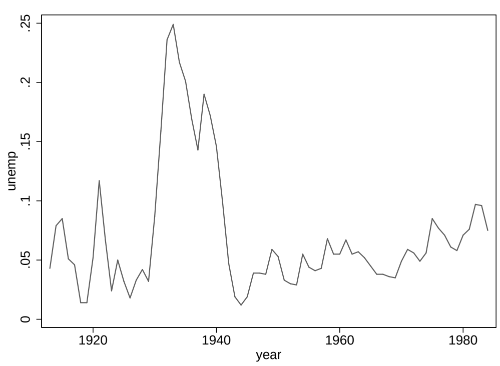
[11] In the analysis that follows, we correct the female wage series reported in Whittington et al. for typos documented in the letter from Leslie Whittington to Brigitte Madrian discussed in footnote 8. The corrected values are 0.548 in 1919 and 2.094 in 1972. The remaining data series are as reported in the Appendix. Results in the remainder of this section are robust to using the series as reported in Whittington et al.

Figure 3: Time Trend Inspection, 1913–1984

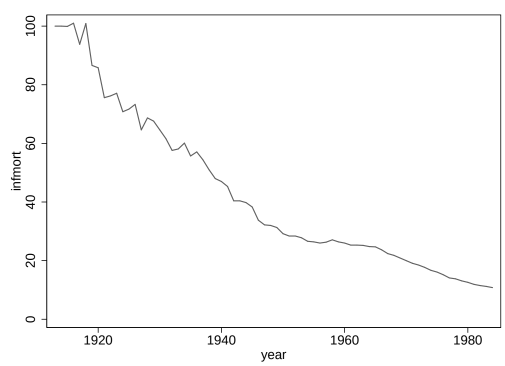
(a) Personal Exemption



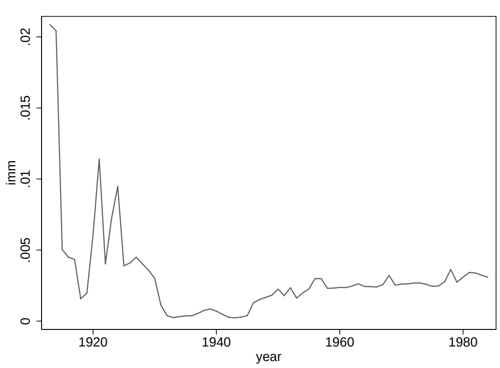
(b) Male and Asset Income



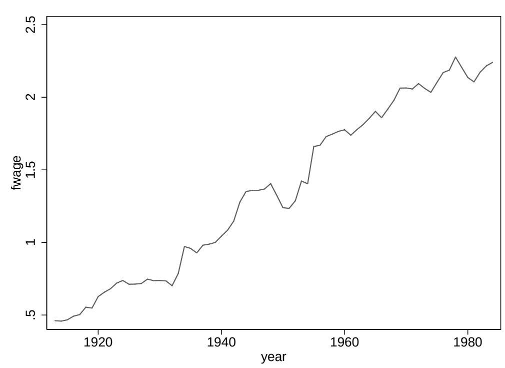
(c) Unemployment



(d) Infant Mortality



(e) Immigration



(f) Female Wage

The estimated values of ρ are reported in Table 3. To test if an estimated value of ρ is significantly different than one, we use the augmented Dicky-Fuller test.[12]

Table 3: Testing for a Unit Root, 1913–1984

Variable	Trend?	ρ	α	DF Approx. p-value	Unit Root?
General Fertility Rate	Yes	0.978	1.3049	0.9150	Yes
		(0.026)	(2.549)		
Personal Exemption		0.993		0.3717	Yes
		(0.018)			
Male & Asset Income	Yes	0.996	133.945	0.3762	Yes
		(0.015)	(117.305)		
Unemployment		0.961		0.1162	Yes
		(0.034)			
Infant Mortality	Yes	0.972	-0.024	0.7250	Yes
		(0.013)	(0.649)		
Immigration		0.740		0.0069	No
		(0.051)			
Female Wage	Yes	0.995	0.032	0.0698	Yes
		(0.011)	(0.017)		

Standard errors in parentheses
Dickey-Fuller test run with 2 lags

For each of these series, a value of ρ close to one is evidence that the series may be I(1). Using Dickey-Fuller tests, we are able to reject the null hypotheses that a series is I(1) for the immigration series only. Each of the other series show strong evidence of unit root behavior, as they have an estimated value of ρ close to one and the Dickey-Fuller test does not provide evidence to the contrary.[13]

Time series regressions with variables that are I(1) can give very misleading results, and our tests indeed indicate that estimation using the levels of our reconstructed data would suffer from the spurious regression problem. An approach to addressing this problem is to

[12]When ρ is close to 1, the sampling distributions of the estimator of $\hat{\rho}$ is extremely different. The augmented Dicky-Fuller test is similar to the AR(1) regression, but subtracts y_{t-1} from both sides. To perform the test, we estimate $\Delta y_t = \theta y_{t-1} + \epsilon_t$. Now $\theta = \rho - 1$, so we test the null hypothesis that $\theta = 0$ (i.e. the variable contains a unit root) against the alternative that $\theta < 1$. For variables with a time trend, the test is adjusted appropriately.

[13]We report the results of the Dickey-Fuller test assuming two lags. The results are generally robust to including other reasonable numbers of lags.

first-difference the variables that have a unit root. Each first-differenced series is weakly dependent, and thus usual OLS inference procedures are valid.[14] We run the specifications described in Whittington et al. using the first differences of the appropriate series. The results are reported in Table 4.

The first specification reported in Table 4 is comparable to those in Table 2 and includes no lagged values. Whittington et al. (1990) also run specifications with various lag structures to capture a potential delay in fertility response. They find that both the magnitude and the statistical significance of their results are robust to these lag structures. In our second and third specifications, we specify particular lag structures as in Whittington et al. The second specification has a three-year rectangular lag on the personal exemption, with equal weights on each of the three lags as shown below in equation (4). The coefficient on the rectangular lag can be interpreted as a measure of the total effect of the personal exemption on the general fertility rate.

$$\text{Rectangular Lag Structure} = \frac{1}{3}\left(\Delta PE_t + \Delta PE_{t-1} + \Delta PE_{t-2}\right) \tag{4}$$

The third specification gives personal exemption an inverted V-shaped pattern, with weights increasing until the second lag, and then decreasing. The rationale for this particular lag structure, as stated in Whittington et al., comes from the biological average of "24 to 31 months required to produce a birth."

$$\text{V-Shaped Lag Structure} = \frac{1}{9}\Delta PE_t + \frac{2}{9}\Delta PE_{t-1} + \frac{1}{3}\Delta PE_{t-2} + \frac{2}{9}\Delta PE_{t-3} + \frac{1}{9}\Delta PE_{t-4} \tag{5}$$

The results in Table 4 show that had Whittington et al. (1990) corrected for the spurious regression problem by first-differencing, they would not have found a strong positive relationship between child tax benefits and fertility in their specifications. The results are not

[14]The Dickey-Fuller test was run on the differenced series, and the null hypothesis of the presence of a unit root can be rejected in the differenced series.

Table 4: Impact of Personal Exemption on Fertility in the United States, 1913–1984

Variable	(1)	(2)	(3)
Δ Personal Exemption	-0.084		
	(0.042)*		
Δ Personal Exemption		-0.015	
(Rectangular Lag Structure)		(0.082)	
Δ Personal Exemption			0.116
(V-Shaped Lag Structure)			(0.152)
Δ Male and Asset Income	-0.003	-0.004	-0.003
	(0.002)*	(0.002)*	(0.002)
Δ Unemployment	-20.985	-23.184	-17.552
	(31.280)	(33.534)	(34.946)
Δ Infant Mortality	-0.042	-0.033	-0.038
	(0.315)	(0.351)	(0.380)
Immigration	68.878	-92.7	-48
	(119.073)	(325.409)	(295.193)
Δ Female Wage	7.472	8.368	4.802
	(5.792)	(4.896)*	(6.864)
Pill	-1.91	-1.71	-1.357
	(1.020)*	(0.995)*	(1.038)
WW II	5.138	2.324	-1.231
	(3.377)	(4.692)	(6.245)
Intercept	-0.618	-0.042	-0.229
	(0.954)	(1.330)	(1.085)
R-squared	0.2035	0.1258	0.1553

Robust standard errors in parentheses.
Variables expressed in constant 1967 dollars.
Model (1): no lags on independent variables.
Model (2): rectangular lag on the personal exemption only.
Model (3): five-year inverted V on personal exemption only.
* significant at the 10% level
** significant at the 5% level
*** significant at the 1% level

robust to the specification of the lag structure. None of the estimated coefficients on personal exemption are significant, and the signs of the point estimates in three of the specifications are negative rather than positive. These results suggest that the Whittington et al. results were due to the large degree of persistence in the data, and therefore should not be cited as evidence of a fertility response to child tax benefits.

4 Extending the Data

As explained in Section 3, there are several problems with the data used in Whittington et al. (1990). Rather than simply adding data for the additional 21 years (1985-2005) to the reconstructed 1913-1984 series, we examine each series to determine if better sources are available. The data construction is outlined in this section with additional details provided in Appendix B.

We found discrepancies between the general fertility rate series reported in Whittington et al. (1990) and general fertility rates available from more current published sources. We use fertility data from the National Vital Statistics Report (Martin et al. 2005) for those years in which it is available. For earlier years, we use the estimates from the Datapedia of the United States (Kurian 2001).

We follow the Whittington et al. (1990) methodology in calculating the value of the personal exemption, multiplying the statutory level of the dependent exemption by the average statutory marginal tax rate.[15] In the subsequent analysis, we use a measure of the total value of child tax benefits in the federal income tax, as recent tax changes have increased the importance of other child tax benefits in comparison to the personal exemption.[16] In addition to the tax value of the personal exemption, the total child subsidy series also includes the tax value of the child tax credit, the child and dependent care tax credit, and the earned income tax credit (EITC). The average tax value of these credits is calculated by dividing

[15] This measure was introduced by Barrow and Sahasakul (1983). The complete methodology is explained in the appendix.

[16] The results are not sensitive to the definition of child tax benefits that is used.

Table 5: Summary Statistics, 1913–2005

Variable	Obs.	Mean	Std. Dev.	Min	Max
General Fertility Rate	93	88.9	21.4	63.6	126.6
Child Tax Subsidy	93	760.9	492.7	0	2088.0
Male & Asset Income	93	31,287	11,681	17,043	50,169
Unemployment	93	0.0679	0.0476	0.0120	0.2490
Infant Mortality	93	35.15	27.77	6.7	101
Immigration	93	0.00351	0.00257	0.00028	0.01505
Female Wage	93	7.59	3.34	2.14	12.93
Pill	93	0.462	0.501	0	1
WW II	93	0.054	0.227	0	1

Variables expressed in constant 2005 dollars.

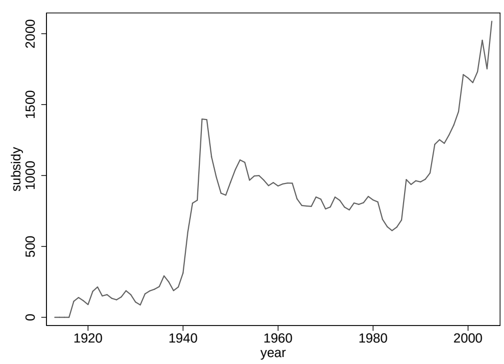
the total federal tax expenditure on these credits by the number of children in the United States.

We construct a revised male and asset income data series, using more recently available data. The male and asset income series is constructed from some of the same source material as Whittington et al. (1990), but also incorporates male income data reported by the U.S. Census Bureau (see the appendix). Similarly, we utilize more recent data for female wages and Whittington et al.'s historical series to create an extended female wage series.

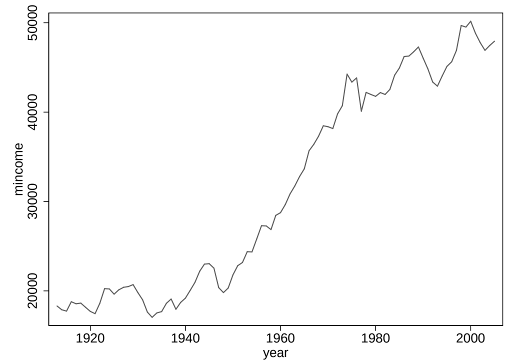
The series for unemployment, infant mortality, and immigration come from the same source material as the series used for replication, and are simply extended to 2005. The summary statistics for the extended data are reported in Table 5. We check each series for unit root behavior using the same procedures as explained in Section 3. Each series is inspected to determine if it has a time trend, as shown in Figure 4. The regression of each variable on its lagged value and the p-value from the augmented Dickey-Fuller test for a unit root are reported in Table 6.

The outcomes of the unit root tests are similar to those performed on the replicated series in Section 3. While the Dickey-Fuller test provides some evidence to support the hypothesis that unemployment does not have a unit root, we report results in Section 5 assuming this series does have a unit root due to the estimated ρ close to 1 and in order to maintain

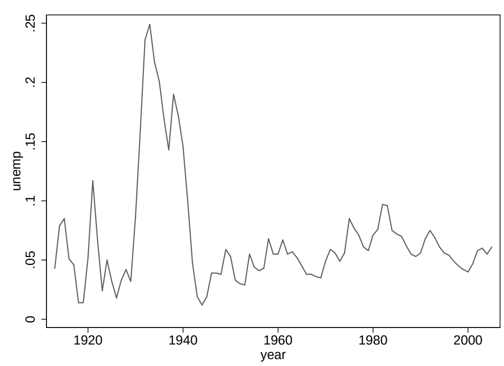
Figure 4: Time Trend Inspection, 1913–2005



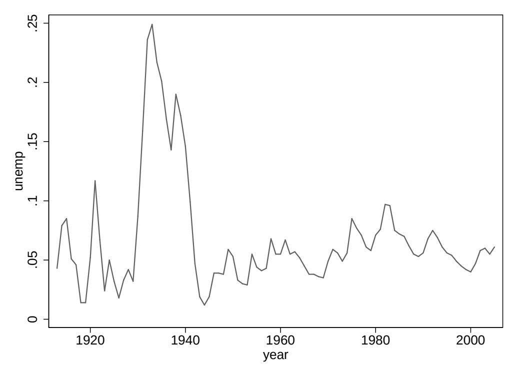
(a) Child Tax Subsidy



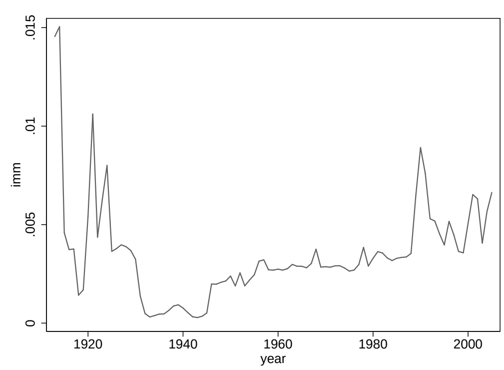
(b) Male and Asset Income



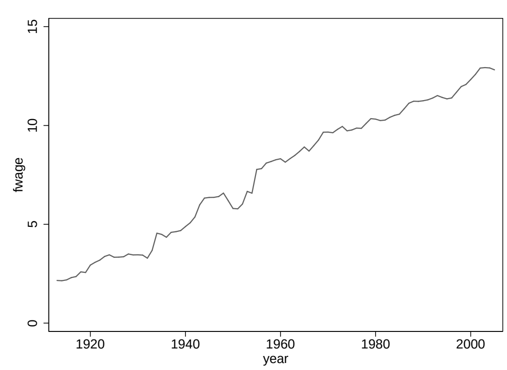
(c) Unemployment



(d) Infant Mortality



(e) Immigration



(f) Female Wage

Table 6: Testing for a Unit Root, 1913–2005

Variable	Trend?	ρ	α	DF Approx. p-value	Unit Root?
General Fertility Rate	Yes	0.975	1.633	0.8414	Yes
		(0.019)	(1.701)		
Child Tax Subsidy	Yes	1.002		0.8713	Yes
		(0.024)			
Male & Asset Income	Yes	1.000	310.31	0.3984	Yes
		(0.010)	(316.69)		
Unemployment		0.964		0.0501	Yes
		(0.029)			
Infant Mortality	Yes	0.971	0.010	0.5387	Yes
		(0.009)	(0.424)		
Immigration		0.871		0.0597	No
		(0.040)			
Female Wage	Yes	0.997	0.141	0.1297	Yes
		(0.007)	(0.059)		

Standard errors in parentheses

comparison with the prior analysis. However, we have performed the analysis under the alternative assumption and found that the results are robust.

5 Updated Results

In this section, we revisit the question of whether time series analysis shows an effect of child tax benefits on fertility using the data series from 1913-2005 described above. We first show that the main result from Table 4, namely that the strong results reported in Whittington et al. (1990) are not robust to first-differencing, follows through when we change the inflation base year, use more currently available data sources, change the time period analyzed, and examine additional features of the tax code that provide tax subsidies to families with children. Table 7 summarizes our findings. In Column (1), we report our replication of Whittington et al.'s main specification, copied from Column (3) in Table 2. These results are reported in constant 1967 dollars and are calculated using data series from the years 1913-1984. For Columns (2) and later, we make two changes: the value of the child tax subsidy, male income, and female wage are converted to constant 2005 dollars; and the typos in Whittington et al.'s series (discussed in footnote 8) are corrected. The effect of changing the base year can be seen clearly in the coefficient on the tax subsidy: whereas our replication of Whittington et al. in Column (1) showed that $100 in tax benefits (in 1967 dollars) are associated with an increase in the general fertility rate of 11.6, the results in Column (2) show that the comparable change in the general fertility rate for $100 in tax benefits (in 2007 dollars) is 1.7 births. This value provides a benchmark against which results from our subsequent analyses can be measured.

Moving from Column (2) to Column (3) illustrates the effect of first-differencing the series that contain a unit root. As was true in Table 4, the coefficient on the tax subsidy flips sign and decreases in magnitude. Columns (4) and (5) are similar to Columns (2) and (3), but summarize the analysis using our extended data series for 1913-2005. The results in Columns (4) and (5) show that using updated data sources and extending the data through 2005 do not substantively change the key coefficients estimated in Columns (2) and (3). Finally, Columns (6) and (7) repeat the analysis including other child tax benefits in the tax subsidy series. While the coefficients on the total child tax subsidy variable are of the same signs as in Columns (4) and (5), they are no longer significant and smaller in magnitude. Because of the increasing importance of tax subsidies for children other than the personal exemption (see Figure 1) and their more salient nature, the changes in the key coefficients that result from adding in these other tax benefits cast additional doubt regarding the true effect of tax subsidies on fertility. Overall, Table 7 shows that Whittington et al.'s result is sensitive to correcting for unit roots by first-differencing and adding the tax subsidies for children in other parts of the tax code, but not to changing the time period studied or using more currently available data sources.

For all following analysis, we use our extended data series for the 1913-2005 sample period, and use the total child tax subsidy (personal exemption + child tax benefits from the earned

Table 7: Comparison of Estimation Results

Variable	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Personal Exemption	0.116	0.017	-0.014	0.011	-0.013		
	(0.045)**	(0.008)**	(0.007)*	(0.006)*	(0.007)*		
Total Child Tax Subsidy						0.005	-0.007
						(0.004)	(0.006)
Male and Asset Income	0.001	-0.000	-0.001	-0.001	-0.001	-0.001	-0.001
	(0.002)	(0.000)	(0.000)*	(0.000)***	(0.000)	(0.000)**	(0.000)
Unemployment	-68.191	-68.019	-20.985	-86.711	-10.041	-84.979	-8.957
	(34.004)**	(33.684)**	(31.280)	(25.079)***	(26.985)	(24.274)***	(27.301)
Infant Mortality	0.035	-0.013	-0.042	0.057	-0.072	-0.088	-0.054
	(0.251)	(0.247)	(0.315)	(0.157)	(0.281)	(0.139)	(0.274)
Immigration	760.712	698.917	68.878	1,079.458	198.098	977.633	194.315
	(304.983)**	(299.761)**	(119.073)	(297.470)***	(135.746)	(289.513)***	(138.742)
Female Wage	5.629	2.829	1.278	4.137	2.127	4.303	1.924
	(5.036)	(2.416)	(0.990)	(2.349)*	(1.270)*	(2.240)*	(1.196)
Pill	-12.014	-10.937	-1.910	-6.080	-0.688	-5.427	-0.440
	(6.028)*	(5.902)*	(1.020)*	(4.697)	(0.866)	(4.643)	(0.841)
WW II	-17.863	-16.269	5.138	-13.736	4.703	-11.376	3.468
	(4.854)***	(4.772)***	(3.377)	(3.865)***	(2.758)*	(3.682)***	(2.572)
Time Trend	-0.741	-0.969		-0.527		-0.725	
	(0.510)	(0.590)		(0.348)		(0.368)*	
Constant	104.130	108.208	-0.618	119.724	-1.272	133.054	-1.174
	(23.368)***	(23.052)***	(0.954)	(15.527)***	(0.898)	(13.493)***	(0.943)
Levels or Differences?	Levels	Levels	Differences	Levels	Differences	Levels	Differences
Observations	72	72	71	93	92	93	92
R-squared	0.749	0.745	0.203	0.804	0.145	0.792	0.103

Robust standard errors in parentheses.
Model (1): Replication of Whittington et al. (1990) as shown in Column (3), Table 2.
Model (2): Variables expressed in constant 2005 dollars.
Model (3): Variables expressed in constant 2005 dollars, with estimation performed in first differences.
Model (4): Extended data series for sample period 1913-2005.
Model (5): Extended data series for sample period 1913-2005, with estimation performed in first differences.
Model (6): Model (4) with additional child tax benefits included.
Model (7): Model (5) with additional child tax benefits included.
* significant at the 10% level; ** significant at the 5% level; *** significant at the 1% level

income tax credit, the child and dependent care tax credit, and the child tax credit) as our measure of tax benefits that may affect fertility, except when otherwise indicated. We believe the revised data sources represent the most accurate available time series, and that ignoring the features of the tax code that subsidize children would misrepresent the responsiveness of fertility to tax benefits. We also perform the remaining analysis including first-differenced series for variables that were found to contain a unit root in Table 6.

As pointed out in Whittington et al. (1990), there are several reasons to believe that fertility response from changes in covariates may occur with a lag. The birth of a child will lag the decision to have a child by at least nine months and frequently longer, and therefore the relevant variable in analyzing fertility in year t may be the covariate's value in year $t-1$. Covariates in time t may have little influence on fertility in year t.[17]

There is a compelling reason to believe that the fertility response from changes in child tax benefits may be even more delayed. While a fertility response would not likely be observed until at least one year after a change to child tax benefits, it takes some time for taxpayers to learn that a tax change has taken place. Changes to the tax code are often made while the tax year is well underway. Individuals are not likely to learn about tax changes until they do their taxes (in April of the following year). While this may have an immediate effect on the decision to have a child, the actual birth is then realized with a delay. Therefore, while a single lag may be appropriate for the other regressors, the real value of child tax benefits should enter the fertility equation with at least two lags. That is, we posit that a tax policy change in year t may not affect the decision to have children until year $t+1$ and thus would not affect the total fertility rate until year $t+2$.

Table 8 reports the results from a regression of the differenced total fertility rate on the current and first four lags of the differenced real value of child tax benefits. The other controls (male and asset income, unemployment rate, infant mortality rate, immigration rate, female wage index, and indicators for the pill and World War II) are included in the

[17] Immigration by women of childbearing age is an exception since some women may be pregnant at the time of immigration.

Table 8: Child Tax Benefits Lagged Effect on Fertility

Variable	(1)	(2)	(3)	(4)	(5)
Δ Total Child Tax Subsidy	-0.006	-0.003	-0.005	-0.005	-0.005
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Δ Total Child Tax Subsidy$_{t-1}$	-0.001	0.000	-0.002	0.001	-0.001
	(0.005)	(0.005)	(0.005)	(0.004)	(0.005)
Δ Total Child Tax Subsidy$_{t-2}$	0.014	0.012	0.012	0.014	0.012
	(0.006)**	(0.007)*	(0.006)**	(0.006)**	(0.006)**
Δ Total Child Tax Subsidy$_{t-3}$	0.003	0.002	0.005	0.006	0.004
	(0.006)	(0.006)	(0.006)	(0.006)	(0.005)
Δ Total Child Tax Subsidy$_{t-4}$	-0.004	-0.003	-0.003	-0.003	-0.004
	(0.004)	(0.004)	(0.004)	(0.003)	(0.004)
Measure of Total Effect	0.006	0.008	0.008	0.013	0.007
	(0.012)	(0.012)	(0.011)	(0.010)	(0.010)
Current Covariates Included	Yes	Yes	No	No	No
Lagged Covariates Included	No	Yes	Yes	No	Yes
Error Correction Model	No	No	No	No	Yes
Observations	88	88	88	88	88
R-squared	0.252	0.358	0.288	0.214	0.320

Robust standard errors in parentheses.
* significant at the 10% level
** significant at the 5% level
*** significant at the 1% level
Note: only current values of Pill and WW2 included in Column (2)

estimations as indicated in the table although the estimated coefficients are not reported. Current values of the covariates are included in Columns (1) and (2), and lagged values are included in Columns (2) and (3). No other control variables are included in Column (4).

The results in Table 8 suggest that the estimated effect of the second lag of child tax benefits on fertility is smaller than the Whittington et al. (1990) estimate but still significantly different from zero across most of the specifications. This estimate ranges from 0.012 to 0.014 implying an increase of about 1.2 to 1.4 births per 1,000 women age 15-44 is associated with a $100 increase in the real value of child tax benefits in 2005 dollars. The second lag is the only lagged value that is significant across most specifications and provides evidence supporting our hypothesis that a tax policy change affects the fertility rate with a two-year lag.

The true total effect of tax benefits on the general fertility rate is not only the effect after two years, but rather the sum of the coefficients of all lagged values. Table 8 also reports the measure of this total effect with standard errors. The values across the four specifications are small and statistically insignificant, ranging from 0.006 to 0.013.[18] Excluding the estimate obtained in the specification reported in Column (4) which contains no additional controls, the estimates suggest that a $100 increase in the real value of child tax benefits in 2005 dollars is associated with an increase of approximately 0.6 to 0.8 births. The magnitude of this total effect is approximately half the magnitude of the Whittington et al. estimate of 1.7 births as calculated in Table 7, Column (2), and is statistically insignificant across all specifications. These results suggest that tax benefits may affect the timing of births and we find only weak evidence for an overall response of fertility to tax benefits. Our estimates of the total effect are consistently small and positive, but statistically insignificant.

Expressing the regression in levels allows us to evaluate the long-run equilibrium relationship between subsidies and the general fertility rate, while results of the first-difference models test the hypothesis that there is a short-term relationship between the subsidy and fertility. The results in Column (6) of Table 7 did not provide evidence of a long-run relationship between subsidies and the general fertility rate. In addition, in results not reported, we have estimated several different specifications of Column (6) using lagged values of both the subsidy level and covariates. In all of the specifications, the coefficients on the subsidy variable and its lagged counterparts are not statistically significant.[19]

[18]To calculate the total effect of a change in tax benefits on fertility and its standard error, the method outlined in Wooldridge (2006) is used as follows. Suppose the estimated coefficient on lag t of the differenced child tax subsidy is δ_t. We define θ as the total effect, where $\theta = \delta_0 + \delta_1 + \delta_2 + \delta_3 + \delta_4$. We then regress ΔGFR_t on $\Delta Subsidy_t$, $\Delta Subsidy_{t-1} - \Delta Subsidy_t$, $\Delta Subsidy_{t-2} - \Delta Subsidy_t$, $\Delta Subsidy_{t-3} - \Delta Subsidy_t$, $\Delta Subsidy_{t-4} - \Delta Subsidy_t$ and the other covariates. The coefficient and standard error on $\Delta Subsidy_t$ is a measure of the total effect of tax benefits on fertility.

[19]We also test for cointegration of the error terms in these specifications. We find there is evidence of cointegration, i.e. the residuals are stationary. This lends additional support to our negative finding of a long-run relationship between the subsidy and fertility.

However, the results in Table 8 suggest that there is a short-term relationship between the tax subsidy and general fertility rates and that this short-term relationship occurs with a lag. As a further robustness check on this hypothesis, we estimate an error correction model in which the lagged residuals from the regression in levels are included in the first difference model as the error correction term. This model estimates a short-run relationship while controlling for the long-run relationship between the variables. Given that we do not find evidence for a long-run relationship, it is unlikely that adding a control for the long-run relationship will change the estimated short-term relationship. However, we present the results of the error correction model in Column (5) of Table 8. The results provide additional support to the hypothesis that tax subsidies have a short-run influence on fertility that occurs with a lag.

There were large tax rate changes during World War II, significantly increasing the value of the personal exemption, and this period was followed by large increases in fertility that constituted the baby boom. We therefore examine the period 1960-2005 separately to determine whether the results in Table 8 are being driven by these two events. Performing the estimation on the 1960-2005 data decreases the sample size considerably, but allows us to estimate the modern level of responsiveness and eliminates the large fertility changes during the depression and World War II.

Table 9 reports the results from these regressions. The columns correspond to the columns in Table 8, the only difference being the time period examined. The results are similar with one important difference: effects of tax benefits on fertility are present after one lag rather than two. Table 9 suggests that the estimated effect of the first lag of child tax benefits on fertility is approximately 0.7 to 1 birth per $100 in child tax benefits, though the estimate is not as statistically strong, likely due to the decrease in the number of observations. The shift of the response to one year following a change in tax subsidies rather than two years may be explained by a decrease in the amount of time needed to realize that child tax benefits have increased. Increases in the child tax credit were heavily advertised including early payment

Table 9: Child Tax Benefits Lagged Effect on Fertility: 1960-2005

Variable	(1)	(2)	(3)	(4)	(5)
Δ Total Child Tax Subsidy	0.004	0.003	0.001	0.003	-0.000
	(0.004)	(0.005)	(0.004)	(0.003)	(0.004)
Δ Total Child Tax Subsidy$_{t-1}$	0.010	0.008	0.007	0.010	0.007
	(0.005)**	(0.007)	(0.006)	(0.004)**	(0.006)
Δ Total Child Tax Subsidy$_{t-2}$	0.005	0.003	0.006	0.008	0.007
	(0.005)	(0.005)	(0.005)	(0.004)*	(0.005)
Δ Total Child Tax Subsidy$_{t-3}$	-0.001	-0.001	0.003	0.004	0.004
	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)
Δ Total Child Tax Subsidy$_{t-4}$	-0.005	-0.004	-0.002	0.000	-0.002
	(0.005)	(0.005)	(0.004)	(0.002)	(0.004)
Measure of Total Effect	0.013	0.009	0.015	0.025	0.016
	(0.015)	(0.018)	(0.014)	(0.009)***	(0.014)
Current Covariates Included	Yes	Yes	No	No	No
Lagged Covariates Included	No	Yes	Yes	No	Yes
Error Correction Model	No	No	No	No	Yes
Observations	46	46	46	46	46
R-squared	0.323	0.394	0.302	0.223	0.330

Robust standard errors in parentheses.
* significant at the 10% level
** significant at the 5% level
*** significant at the 1% level
Note: only current values of Pill and WW2 included in Column (2)

of the credit to eligible families in one year. The estimated total effects are similarly positive, generally statistically insignificant, though larger in magnitude relative to the total effects shown in Table 8.

Overall, we find some suggestive evidence that tax subsidies for children have a modest short-term effect on fertility rates with a two-year lag. In more recent time periods, the effect is present one year after a change in tax subsidies. The overall effect is approximately half the magnitude found by Whittington et al. (1990) who reported an increase in the general fertility rate of roughly 2 births after an increase in child tax subsidies of $100 in 2005 dollars. However, the total effect of tax benefits on fertility is not statistically significant, and could be evidence in favor of the hypothesis that tax benefits affect the timing but not the overall level of fertility. Our revised results adjust for the persistence in the time series data that was driving the results reported in Whittington et al., include the total value of all child tax benefits rather than only the personal exemption for dependents (which has decreased in relative importance in recent years), and extend the analysis to include the most recent 20 years which saw large increases in subsidies for families with children.

6 Conclusion

The effect of tax policy on fertility rates is often neglected in the literature on federal tax policy, even though child tax benefits are large and have recently grown in importance. One of the most cited studies on this topic, Whittington et al. (1990), uses basic time-series methods and estimates a very large fertility rate response to the tax value of the dependent exemption. We attempt to replicate the Whittington et al. result, and show that the analysis suffers from the spurious regression problem and that the result is not robust to differencing.

We update the time-series data 21 years to 2005, and employ more recent time-series methods to revisit whether fertility responds to changes in the child tax benefits in the federal income tax. The results show that the general fertility rate is most responsive to the second lag of the value of child tax benefits. This is consistent with a view that families do not learn about changes in the tax benefits associated with having children until they do their taxes. However, estimation using data from only the post-war period show that fertility may begin to respond after only one year. The overall estimates of the fertility response to child tax benefits using the updated data and methods are about half the magnitude of those found in Whittington et al. (1990) and statistically insignificant.

We urge caution in any application of these results. While this study provides an updated look at methods employed by Whittington et al. (1990) and a more moderate estimate of the responsiveness of fertility to child tax benefits, there is concern that the identification strategy employed by using time-series analysis is flawed because of correlation between unobserved variables that influence fertility rates and the tax policy. Further work using an alternative identification strategy, such as using detailed micro data of fertility by income, would provide additional evidence regarding the magnitude of the estimated fertility response.

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Appendix: Data for Replication and Extensions

A Data for Replication

The general fertility rate, the real value of the personal exemption, and the female wage were obtained directly from Whittington et al. (1990).

Male and Asset Income

The male and asset income series was obtained from a letter in which Leslie Whittington lists this series. Whittington et al. derived these data for 1913-1948 from Historical Statistics Series D722-727 and D830-844 by calculating a male-to-average earnings ratio, and multiplying this by the average earnings. Years 1949-1955 were derived in the same manner, but used data from the CPS Series P-60 on median earnings. Years 1956-1984 are directly from CPS Series P-60. Nonwage income was obtained from the 1988 Economic Report of the President by subtracting Compensation from National Income, dividing by the population, and multiplying by average family size. The series is adjusted for inflation and is included as a measure of the income effect on fertility. The year to which the series is normalized is not reported.

Unemployment

Whittington et al. (1990) do not report their source for the annual national unemployment series. Unemployment rates for 1929 to 1984 is obtained from the Statistical Abstract of the United States Mini-Historical Series HS-29 (U.S. Census Bureau 2003). Unemployment rates from 1913 to 1928 is obtained from Lebergott (1964) Table A-3. While there is overlap of certain years between the two sources of unemployment data, we found that this specification gave us the best match of the mean and standard deviation reported in Whittington et al.

Infant Mortality

Infant mortatiliy data from 1915 to 1984 is obtained from the Statistical Abstract of the United States Mini-Historical Series HS-13 (U.S. Census Bureau 2003). However, no data appears to be available before 1915 and Whittington et al. do not record the source or give any indication of what values they used for 1913 and 1914. Some studies cite an estimated infant mortality rate of 200 in the early 1700s and then use a linear extrapolation for years between 1700 and 1915. Because the measured infant mortality rate for 1915 is 99.9, it is likely that Whittington et al. simply used values of 100 for both 1913 and 1914. Doing so closely matches their reported mean and standard deviation.

Immigration

The immigration series is listed as the immigration of the at-risk group as a fraction of the resident at-risk group. We assume that the at-risk group is the age group 16-44.[20] We use the original source material as provided in the previous correspondence from Leslie Whittington. For 1913-1970, immigration by age is obtained from the Historical Statistics of the United States: Colonial Times to 1970 Series C 138-142, and population totals by age come from Series A 29-42 of the same volume (U.S. Census Bureau 1975). The source of the remaining data for 1971-1984 is listed as various years of the Statistical Abstract; we use the Historical Statistics of the United States: Millenium Edition Online (U.S. Census Bureau 2006).[21]

[20] Defining the at-risk group as females aged 16-44 requires making an assumption that the percent of immigrants that are female is uncorrelated with the percent of immigrants that are aged 16-44, and yields a series that does not match the reported moments in Whittington et al. (1990).

[21] The ages for which data are available differ slightly over the years. The number of immigrants prior to 1918 was reported for 14-44 year olds. From 1940-1944, the reported age category was 16-45, and from 1971 onwards, 15-44 year-olds were reported. We do not attempt any correction for these differences.

Reconstructed 1913-1984 Data Series

Year	Fertility Rate	Personal Exemption	Male & Asset Income	Unemploy-ment	Infant Mortality	Age 16-44 Immigration	Female Wage
1913	124.7	0	4,090	0.043	100	0.02086	0.461
1914	126.6	0	3,887	0.079	100	0.02043	0.458
1915	125	0	3,860	0.085	99.9	0.00504	0.467
1916	123.4	0	4,294	0.051	101	0.00450	0.492
1917	121	19.27	4,388	0.046	93.8	0.00434	0.503
1918	119.8	23.94	4,920	0.014	100.9	0.00157	0.554
1919	111.2	20.07	4,536	0.014	86.6	0.00197	0.548
1920	117.9	15.33	3,990	0.052	85.8	0.00608	0.627
1921	119.8	34.32	3,529	0.117	75.6	0.01141	0.657
1922	111.2	36.65	3,782	0.067	76.2	0.00403	0.681
1923	110.5	25.83	4,271	0.024	77.1	0.00723	0.72
1924	110.9	27.34	4,136	0.05	70.8	0.00948	0.738
1925	106.6	22.85	4,167	0.032	71.7	0.00389	0.712
1926	102.6	21.13	4,268	0.018	73.3	0.00410	0.713
1927	99.8	24.61	4,237	0.033	64.6	0.00450	0.717
1928	93.8	31.96	4,390	0.042	68.7	0.00403	0.747
1929	89.2	27.29	4,751	0.032	67.6	0.00359	0.737
1930	89.2	18.4	4,570	0.087	64.6	0.00301	0.738
1931	84.6	14.91	4,386	0.159	61.6	0.00113	0.735
1932	81.7	28.36	4,070	0.236	57.6	0.00038	0.702
1933	76.3	31.95	4,059	0.249	58.1	0.00025	0.786
1934	78.5	33.91	4,164	0.217	60.1	0.00031	0.972
1935	77.2	36.98	4,304	0.201	55.7	0.00037	0.959
1936	75.8	50.12	4,716	0.169	57.1	0.00038	0.928
1937	77.1	42.79	4,727	0.143	54.4	0.00055	0.981
1938	79.1	32.22	4,437	0.19	51	0.00075	0.988
1939	77.6	36.53	4,857	0.172	48	0.00086	1
1940	79.9	53.33	5,179	0.146	47	0.00070	1.043
1941	83.4	102.49	5,936	0.099	45.3	0.00048	1.084
1942	91.5	137.7	6,678	0.047	40.4	0.00027	1.147
1943	94.3	141.2	7,327	0.019	40.4	0.00023	1.278
1944	88.4	243.83	7,561	0.012	39.8	0.00028	1.351
1945	85.9	238.4	7,304	0.019	38.3	0.00038	1.358
1946	101.9	193.16	6,983	0.039	33.8	0.00129	1.359
1947	113.3	168.9	6,604	0.039	32.2	0.00152	1.368
1948	107.3	149.79	6,811	0.038	32	0.00167	1.405
1949	107.1	147.05	7,076	0.059	31.3	0.00183	1.323
1950	106.2	163.1	7,442	0.053	29.2	0.00225	1.239
1951	111.5	178.14	7,622	0.033	28.4	0.00179	1.235
1952	113.9	189.43	7,691	0.03	28.4	0.00235	1.287
1953	115.2	186.51	7,797	0.029	27.8	0.00162	1.423
1954	118.1	165.46	7,910	0.055	26.6	0.00198	1.404
1955	118.5	170.57	8,603	0.044	26.4	0.00227	1.661
1956	121.2	171	8,404	0.041	26	0.00299	1.669
1957	122.9	165.12	8,458	0.043	26.3	0.00299	1.729

Year	Fertility Rate	Personal Exemption	Male & Asset Income	Unemploy-ment	Infant Mortality	Age 16-44 Immigration	Female Wage
1958	120.2	158.66	8,470	0.068	27.1	0.00231	1.746
1959	118.8	162.19	8,989	0.055	26.4	0.00232	1.765
1960	118	158.28	9,043	0.055	26	0.00237	1.776
1961	117.2	160.71	9,298	0.067	25.3	0.00236	1.739
1962	112.2	161.58	9,563	0.055	25.3	0.00247	1.777
1963	108.5	161.61	9,802	0.057	25.2	0.00263	1.812
1964	105	142.73	10,125	0.052	24.8	0.00244	1.855
1965	96.6	134.6	10,481	0.045	24.7	0.00243	1.903
1966	91.3	133.94	11,178	0.038	23.7	0.00240	1.859
1967	87.6	133.8	11,032	0.038	22.4	0.00258	1.918
1968	85.7	145.1	11,221	0.036	21.8	0.00321	1.979
1969	86.5	142.62	11,290	0.035	20.9	0.00253	2.063
1970	87.9	130.58	11,183	0.049	20	0.00261	2.064
1971	81.8	132.99	11,284	0.059	19.1	0.00262	2.057
1972	73.4	144.85	11,882	0.056	18.5	0.00268	2.094
1973	69.2	140.87	12,231	0.049	17.7	0.00269	2.061
1974	68.4	130.49	11,429	0.056	16.7	0.00259	2.034
1975	66	122.36	11,154	0.085	16.1	0.00245	2.103
1976	65.8	120.08	11,434	0.077	15.2	0.00247	2.17
1977	66.8	116.11	11,930	0.071	14.1	0.00277	2.187
1978	65.5	118.98	11,972	0.061	13.8	0.00363	2.277
1979	67.2	132.93	11,646	0.058	13.1	0.00274	2.206
1980	68.4	123.17	10,857	0.071	12.6	0.00310	2.136
1981	67.4	119.31	10,765	0.076	11.9	0.00342	2.106
1982	67.3	102.04	10,255	0.097	11.5	0.00339	2.173
1983	65.8	92.49	10,595	0.096	11.2	0.00324	2.216
1984	65.4	83.9	11,370	0.075	10.8	0.00309	2.24

B Extending the Data

General Fertility Rate

For our extended data series, we use the general fertility rate in years 1913-1959 from the Datapedia of the United States (Kurian 2001) and years 1960-2005 from the National Vital Statistics Report (Martin, Hamilton, Sutton, Ventura, Menacker and Munson 2005). The general fertility rates reported in the Datapedia match those reported in Whittington et al. in all but two years; however, the National Vital Statistics Report's general fertility rates differ slightly in several years. Since we believe the National Vital Statistics Report to have the most current and reliable fertility data, we use these data for all years they are available.

Child Tax Benefits

The value of the personal exemption for a parent claiming a child as a dependent is calculated by multiplying the statutory amount of the personal exemption by the marginal tax rate.[22] From 1913 to 1916, there was no personal exemption for dependents. Starting in 1917, a personal exemption for dependents was introduced and set at $200, one fifth of the personal exemption for a individual. In 1944, the separate category for dependents was removed; the personal exemption for a dependent was equal to the personal exemption for the taxpayer or a spouse.

Because the value of the personal exemption depends on the marginal tax rate, an average marginal tax rate for each year is needed. Whittington et al. use an arithmetic average marginal statutory income tax rate weighted by adjusted gross income that was first introduced by Barrow and Sahasakul (1983) and then updated to include all years from 1916 to 1983[23] in (Barrow and Sahasakul 1986). Stephenson (1998) updated the series to 1994.[24] We use the Barrow and Sahasakul methodology to extend the average marginal tax rate series to 2005 using data from the IRS Statistics of Income (see http://www.irs.gov/pub/irs-soi/04in01tr.xls). The IRS tables report the number of taxpayers and the amount of income at each marginal tax rate. Using this data, we take the arithmetic average weighted by AGI to update the Barrow-Sahasakul statutory marginal tax rate series. Some of the AGI cells in the IRS data are negative and are dropped from the calculation.

The value of the personal exemption is not the only tax benefit for a parent claiming a child as a dependent. To calculate the total benefit, we add the tax value of the Earned Income Tax Credit (EITC), Child and Dependent Care Credit (CDCC), and the Child Tax Credit (CTC) to the value of the personal exemption. Unlike the additional personal exemption that can be claimed by nearly every taxpayer with a dependent child, the EITC can only be claimed by taxpayers in

[22]The personal exemption level series is commonly available. We used the series provided by the tax policy center, online at http://www.taxpolicycenter.org.

[23]As noted by Whittinton et al. Barrow and Sahasakul calculate the average marginal tax rate starting in 1916 because this is when the IRS statistics of income data become available. However, since from 1913 to 1916 the personal exemption for dependents was zero, no values for the value of the personal exemption series are missing.

[24]Stephenson notes that the average marginal tax rates reported by Barrow and Sahasakul (1986) for 1981 and 1983 are slightly different than the values that he calculates. Stephenson attributes the difference to Barrow and Sahasakul's use of preliminary statistics of income data.

a specific income range and the CDCC can only be claimed by taxpayers where there is no stay-at-home parent. Thus, rather than calculate the tax value of these benefits for a taxpayer in the particular situation, we take the real value of all benefits from these three tax provisions and divide by the number of children to produce an average benefit level. The value of the personal exemption and the total value of benefits are the same until the mid 1970's when these tax provisions are introduced. The tax expenditure on the EITC, CDCC, and CTC were gathered from the OMB *Analytical Perspectives, Budget of the United States Government* Tables 5-1 and 19-1 from various years.

Male and Asset Income

Male and asset income data were constructed from a variety of sources. From 1947-2005, male income data were obtained from the Historical Income Table P-53 constructed by the U.S. Census Bureau.[25] Male income data before 1947 were constructed by estimating the equation

$$MaleIncome_t = \alpha_0 + \beta_0 MedianIncome_t + \epsilon_t \tag{A-1}$$

for years 1947-2005, and using these estimated coefficients to impute male income from median income prior to 1947[26]. The series that Whittington et al. used includes asset income, which was obtained from two additional sources: the Statistics of Income for years 1916-1936, and the National Income and Product Accounts for years 1929-2005. Finally, the series was adjusted to 2005 dollars.

Other Series

As in the unemployment series for replication, unemployment data after 1929 is obtained from the Bureau of Labor Statistics. The infant mortality series is also extended to 2005 using the same source as the replication data, the U.S. Census Bureau.

For years 1986 - 2005, the Department of Homeland Security publishes the number of immigrants

[25]see http://www.census.gov/hhes/www/income/histinc/incpertoc.html.

[26]Median income from 1913-1960 is from Lebergott (1964). Using the overlapping years 1947-1960, a scaling factor was estimated and applied to the imputed male income series to make the transition between the two series smooth

by age and gender in the Yearbook of Immigration Statistics. These reports are available on the Department of Homeland Security's website.[27] These data were appended to the immigration data used for replication.

While the constructed female wage series was used for replication purposes, for our later analysis, we obtain female wages for 1973-2005 from the Economic Policy Institute and estimate a scaling factor which is applied to Whittington et al.'s series to fill in the values from 1913-1972.

Complete 1913-2005 Data

Year	Fertility Rate	Child Tax Benefits	Male Income	Unemploy-ment	Infant Mortality	Age 16-44 Immigration	Female Wage
1913	124.7	0.00	18,309.34	0.043	100	0.01455	2.159
1914	126.6	0.00	17,886.99	0.079	100	0.01505	2.145
1915	125	0.00	17,737.60	0.085	99.9	0.00459	2.187
1916	123.4	0.00	18,786.45	0.051	101	0.00374	2.304
1917	121	112.91	18,559.87	0.046	93.8	0.00376	2.356
1918	119.8	139.68	18,632.94	0.014	100.9	0.00142	2.594
1919	111.2	117.41	18,160.82	0.014	86.6	0.00169	2.562
1920	117.9	89.84	17,704.93	0.052	85.8	0.00543	2.936
1921	119.8	183.30	17,451.41	0.117	75.6	0.01062	3.077
1922	111.2	213.90	18,617.67	0.067	76.2	0.00436	3.189
1923	110.5	150.76	20,249.26	0.024	77.1	0.00626	3.372
1924	110.9	159.89	20,218.60	0.05	70.8	0.00800	3.456
1925	106.6	133.92	19,634.65	0.032	71.7	0.00364	3.334
1926	102.6	123.58	20,122.23	0.018	73.3	0.00379	3.339
1927	99.8	143.67	20,401.31	0.033	64.6	0.00398	3.358
1928	93.8	187.31	20,484.62	0.042	68.7	0.00388	3.498
1929	89.3	159.89	20,699.47	0.032	67.6	0.00369	3.451
1930	89.2	107.59	19,807.67	0.087	64.6	0.00324	3.456
1931	84.6	87.37	18,999.14	0.159	61.6	0.00137	3.442
1932	81.7	165.36	17,625.68	0.236	57.6	0.00049	3.287
1933	76.3	186.29	17,043.05	0.249	58.1	0.00031	3.681
1934	78.5	198.21	17,536.04	0.217	60.1	0.00038	4.552
1935	77.2	216.68	17,682.97	0.201	55.7	0.00046	4.491
1936	75.8	292.25	18,606.59	0.169	57.1	0.00047	4.346
1937	77.1	249.55	19,093.16	0.143	54.4	0.00064	4.594
1938	79.1	188.37	17,941.25	0.19	51	0.00087	4.627
1939	77.6	213.57	18,711.81	0.172	48	0.00093	4.683
1940	79.9	312.48	19,190.55	0.146	47	0.00077	4.884
1941	83.4	600.51	20,055.10	0.099	45.3	0.00054	5.076
1942	91.5	805.16	20,946.84	0.047	40.4	0.00032	5.371
1943	94.3	825.79	22,196.92	0.019	40.4	0.00028	5.985

[27] see http://www.dhs.gov/ximgtn/statistics/publications/yearbook.shtm.

Year	Fertility Rate	Child Tax Benefits	Male Income	Unemploy-ment	Infant Mortality	Age 16-44 Immigration	Female Wage
1944	88.8	1,398.17	22,995.08	0.012	39.8	0.00035	6.327
1945	85.9	1,394.23	23,045.00	0.019	38.3	0.00051	6.359
1946	101.9	1,131.74	22,541.69	0.039	33.8	0.00199	6.364
1947	113.3	989.64	20,363.85	0.039	32.2	0.00198	6.406
1948	107.3	875.20	19,809.01	0.038	32	0.00207	6.58
1949	107.1	861.62	20,323.56	0.059	31.3	0.00214	6.196
1950	106.2	953.00	21,795.75	0.053	29.2	0.00239	5.802
1951	111.5	1,041.10	22,819.57	0.033	28.4	0.00189	5.783
1952	113.9	1,109.89	23,177.59	0.03	28.4	0.00255	6.027
1953	115.2	1,092.80	24,385.18	0.029	27.8	0.00189	6.664
1954	118.1	967.06	24,359.67	0.055	26.6	0.00219	6.575
1955	118.5	996.90	25,817.02	0.044	26.4	0.00245	7.778
1956	121.2	999.48	27,291.53	0.041	26	0.00315	7.816
1957	122.9	967.46	27,266.51	0.043	26.3	0.00321	8.097
1958	120.2	928.52	26,854.98	0.068	27.1	0.00271	8.176
1959	118.8	950.33	28,446.68	0.055	26.4	0.00269	8.265
1960	118	926.36	28,753.13	0.055	26	0.00274	8.317
1961	117.1	940.58	29,653.57	0.067	25.3	0.00269	8.144
1962	112	946.75	30,843.61	0.055	25.3	0.00277	8.322
1963	108.3	945.86	31,734.30	0.057	25.2	0.00298	8.485
1964	104.7	835.38	32,786.44	0.052	24.8	0.00289	8.687
1965	96.3	788.64	33,657.01	0.045	24.7	0.00289	8.912
1966	90.8	784.82	35,673.43	0.038	23.7	0.00281	8.706
1967	87.2	782.37	36,400.13	0.038	22.4	0.00303	8.982
1968	85.2	848.54	37,302.76	0.036	21.8	0.00375	9.268
1969	86.1	833.35	38,471.56	0.035	20.9	0.00284	9.661
1970	87.9	764.46	38,369.11	0.049	20	0.00286	9.666
1971	81.6	777.95	38,162.67	0.059	19.1	0.00284	9.633
1972	73.1	848.01	39,802.10	0.056	18.5	0.00291	9.806
1973	68.8	824.75	40,713.68	0.049	17.7	0.00291	9.951
1974	67.8	776.97	44,256.83	0.056	16.7	0.00280	9.73
1975	66	757.65	43,358.02	0.085	16.1	0.00264	9.773
1976	65	806.34	43,829.38	0.077	15.2	0.00269	9.869
1977	66.8	796.56	40,102.89	0.071	14.1	0.00297	9.856
1978	65.5	808.47	42,210.55	0.061	13.8	0.00384	10.103
1979	67.2	852.46	41,978.00	0.058	13.1	0.00290	10.346
1980	68.4	827.71	41,766.75	0.071	12.6	0.00329	10.322
1981	67.3	813.83	42,185.98	0.076	11.9	0.00363	10.248
1982	67.3	690.79	41,977.79	0.097	11.5	0.00356	10.275
1983	65.7	638.60	42,543.73	0.096	11.2	0.00331	10.414
1984	65.5	611.46	44,132.05	0.075	10.8	0.00318	10.514
1985	66.3	636.05	44,941.38	0.072	10.6	0.00329	10.573
1986	65.4	686.11	46,223.44	0.07	10.4	0.00334	10.844
1987	65.8	971.47	46,272.96	0.062	10.1	0.00336	11.126
1988	67.3	936.77	46,748.23	0.055	10	0.00354	11.229
1989	69.2	962.84	47,289.59	0.053	9.8	0.00646	11.22
1990	70.9	955.05	46,044.02	0.056	9.2	0.00891	11.251
1991	69.3	973.56	44,831.14	0.068	8.9	0.00760	11.299

Year	Fertility Rate	Child Tax Benefits	Male Income	Unemploy-ment	Infant Mortality	Age 16-44 Immigration	Female Wage
1992	68.4	1,017.63	43,371.91	0.075	8.5	0.00530	11.389
1993	67	1,219.79	42,906.69	0.069	8.4	0.00518	11.514
1994	65.9	1,252.32	44,043.06	0.061	8	0.00453	11.42
1995	64.6	1,227.08	45,117.02	0.056	7.6	0.00397	11.347
1996	64.1	1,286.12	45,635.49	0.054	7.3	0.00517	11.394
1997	63.6	1,356.05	46,908.29	0.049	7.2	0.00449	11.682
1998	64.3	1,450.81	49,689.69	0.045	7.2	0.00364	11.969
1999	64.4	1,711.55	49,516.12	0.042	7.1	0.00357	12.076
2000	65.9	1,687.82	50,168.74	0.04	6.9	0.00505	12.325
2001	65.3	1,655.20	48,822.69	0.047	6.8	0.00653	12.589
2002	64.8	1,732.16	47,774.14	0.058	7	0.00631	12.906
2003	66.1	1,953.81	46,914.04	0.06	6.9	0.00406	12.929
2004	66.3	1,752.48	47,459.20	0.055	6.8	0.00568	12.912
2005	66.7	2,087.97	47,932.25	0.061	6.7	0.00663	12.816