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ABSTRACT

The Impact of Relative Cohort Size on U.S. Fertility, 1913–2001*

This paper tests for the long-term and short-term relationships between fertility and relative cohort size for the United States using the annual time series data between 1913 and 2001. An error correction model, imbedded with the cointegration theory, is coupled with the general impulse response function. Empirical evidence on relationships is found lending support to the Easterlin hypothesis in that the change in relative cohort size is an important explanatory variable to include in studies of human fertility both in the short run and in the long-run for the United States. In addition, our results support the catching-up hypothesis and that the child tax deduction has been an important policy variable influencing births.

JEL Classification: J13, D13

Keywords: Easterlin hypothesis, relative cohort size, age structure, catching-up, child tax deduction

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1. Introduction

Empirical studies of fertility have long faced the difficulty that socioeconomic variables might influence the tempo of fertility and the long-run fertility level. Furthermore, the short and long-term relationships of a variable with fertility might be in opposite directions. For example, parents might delay the next birth during an economic downturn and then catch up by spacing subsequent births closer together. Indeed, much of the early demographic literature on the baby boom was dominated by this 'catching up' hypothesis.[1] An alternative and perhaps the best known explanation of the baby boom and subsequent baby bust is due to Richard Easterlin (1961, 1968, and 1987) who posits that swings in relative cohort size influence fertility in part through their impact on relative income.[2]

In this paper, advanced time series models are used to estimate the long-term and the short-term determinants of U.S. fertility. Cointegration analysis is used to test for long-term relationships, and a vector error correction mechanism will be modeled to find short-term relationships for a model that incorporates relative cohort size as one of its possible explanatory variables. The generalized impulse response functions, adopted for avoiding variable ordering problems, indicate both spot response and accumulated response for the fertility rate to the innovations of other explanatory variables.

The time series model extends the empirical model used in Whittington, Alm, and Peters (1990, henceforth WAP) in that we introduce relative cohort size and changes in relative cohort size as possible variables reflecting the Easterlin hypothesis. We also extend the annual data of WAP's period of 1913 to 1984 to the updated period of 1913 to 2001. Detailed information on the update is discussed.

This paper is organized as follows. Section 2 reviews the Easterlin hypothesis in the framework of a household production model. In section 3, we

[1] See Whelpton (1954), Freedman, Whelpton and Campbell (1959), Goldberg, Sharp and Freedman (1959), and Whelpton, Campbell and Patterson (1966).

[2] See Pampel and Peters (1995) and Macunovich (1996b, 1998ab, 2000) for a review of the many facets of the Easterlin hypothesis.

briefly discuss the updated data set and the time series model to be examined. In the penultimate section, the empirical results are presented and interpreted. The final section provides a few concluding remarks.

2. The Easterlin Hypothesis

The Easterlin hypothesis places an emphasis on the appropriate income variables to be used in studies of fertility. The appropriate specification of the income-fertility relationship has long been a dominant theme in population economics. This theme concerns the sign of the relationship between income and fertility. Indeed, much of the early work on the economics of fertility was motivated by the desire to explain the sometimes negative relationship between income and fertility found in many studies.[3] Both positive and negative income-fertility relationships have been found. In part, these sign differences might be characterized in comparative static terms and in part in terms of whether the income-fertility relationship being observed is a long-term or a short-term relationship.

In a comparative static framework, parents choose the number of children, expenditures per child, expenditures on the parents, and the labor-leisure of the parents. The basic income-fertility relationship is often expressed in terms of household production theory. The family maximizes

(1) $U = u(q, s, n)$

subject to

(2) $I\ (= H + wT) = B_q qn + B_s s.$

The arguments in the utility function are the quality of children q, other commodities s and the number of children n. Note that q and s are produced from inputs of goods purchased on the market and the time of the household. The shadow prices of commodities, B_q and B_s, are derived on the assumptions of linear homogeneity, no

[3] For further discussion, see Becker (1960), Freedman and Thornton (1982), and Grabowski and Shields (1996).

joint production and employment of the wife.[4] The observed income of the family is

(3) $$Y = H + w(T - t) = H + wT - wt,$$

where $(T - t)$ is the wife's time in market employment. This division of income emphasizes the impact of the labor market choices of women on fertility. Observed income, Y, is partly determined by the wife's wage rate, w, and other income, H, and also to some degree determined by the labor-leisure choice of the wife, t.[5]

A difficulty immediately arises if the wife is not employed. In this case, there is, by assumption, no substitution effect because only the husband earns income.[6] Furthermore, the budget constraint becomes nonlinear resulting in the shadow price of a commodity depending upon the commodity bundle chosen. However, a difficulty in using micro data occurs when some women will be fully employed and others will not be employed. If the price of a child is thought to be determined by the opportunity cost of a woman's time, there is no immediate proxy such as the wage rate for her opportunity cost.

The household production model of fertility just developed can be modified to incorporate the Easterlin hypothesis that relative cohort size has a negative impact on fertility (Easterlin 1968, 1969). In the model above, the quality of children is an endogenous variable chosen by parents. Easterlin has suggested the quality of children might be exogenous as well. Easterlin assumes that both q and s are

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[4] Employment of the wife means that $T - t > 0$, where T is the total time available to the wife for employment and household production, and t is the wife's time devoted to household production. For these assumptions, B_q and B_s are independent of H, but dependant on w. Full income, $I = H + wT$, is divided into two components, the husband's income, H, and the wife's potential income, wT, where w is the wife's wage rate.

[5] The importance of having data on male plus asset income and the female wage rate has long been stressed in fertility theory and has received attention in empirical work that uses microeconomic data See Easterlin, (1969); Becker and Lewis, (1973); Willis, (1973); Butz and Ward (1979); and Borg (1989).

[6] When the wife is employed, the wage rate is the opportunity cost of her time. When she chooses not to be employed, the opportunity cost of her time exceeds the wage rate. In this case, the opportunity cost of the wife's time is no longer an exogenous variable to be thought of as determining the demand for children. An advantage of macroeconomic data on fertility is that there is no problem with unobserved wage rates. The wage rate for women is simply assumed to be the observed, market wage rate. A representative woman chooses her market employment based on the wage rate, which is the opportunity cost of their labor on the assumption that $t<T$. Recent articles, such as WAP (1990) and Macunovich (1995), have looked directly at this relationship for U.S. data.

functions of relative income. The notion of relative income comes from Duesenberry's (1949) relative income hypothesis. Consumption is argued to be a function of income relative to past income and to the average income in the community. Easterlin emphasizes past income arguing that the tastes and aspirations of today's adults are greatly influenced by the income of their parents in earlier years. A decline in their income relative to their parents' earlier income will decrease fertility and a rise will increase fertility. During the baby boom, this relative income was high and during the subsequent baby bust it was low.

Easterlin applies this socioeconomic view to an explanation of fertility cycles, called the Easterlin hypothesis.[7] He contends that twists in the age structure influence fertility for many reasons, most notably through their influence on relative income and on entry-level job opportunities. The proportion of young adults to the total work force is thought to influence the relative income of young adults. A low proportion, as was the case after World War II, leads to high relative income and, hence, higher fertility. A high proportion, as was the case when the Baby Boom generation moved into the work force, leads to low relative income and low fertility.

The Easterlin hypothesis that twists in the age structure can influence births could hold for several reasons. First, the age structure could affect fertility because it changes the income of young adults relative to the income of their parents' generation. Second, the age structure could affect fertility because it changes the income of young adults relative to the income of older workers today. Third, and more narrowly, the age structure combined with the average values of the husband's income and wife's potential income would provide a better measure of the income of young adults than would their average values alone. The proportion of the labor force consisting of young adults is negatively related to the income of young adults relative to average income. Hence, even if tastes and aspirations do not depend

[7] Easterlin's view can be introduced into the utility function by rewriting the utility function in equation (1) in Stone-Geary form as $U = u(q - q^*, s - s^*, n)$, where q^* and s^* are interpreted as subsistence levels of q and s and are assumed to be positive functions of the consumption levels younger adults became accustomed to when they were teenagers. Of course, q^* and s^* might be determined by the current income of older adults or by other socioeconomic variables.

upon the age structure, age structure could still influence fertility in a macroeconomic model as long as income influences fertility.

Relative cohort size can affect relative income for two basic reasons. First, younger and largely inexperienced workers may be competing for different, entry-level jobs than older and more experienced workers. A higher proportion of younger workers would then imply lower wage rates in comparison to older workers. With time, however, employers might adjust their hiring practices to take advantage of the lower wage rates for younger workers. Second, an increase in the proportion of younger versus older workers might suppress relative income regardless of the long-term impact of the level of this proportion on relative income. Previous studies of relative cohort size have focused on the first explanation with the possible exception of Macunovich (2000), who includes both an age structure variable and changes in age structure in her model.[8] We will examine time series properties of the age structure in the United States.

3. Empirical Time Series Model

Time plays a role in an empirical model that it does not play in the essentially timeless comparative static model discussed in the last section. With micro data, questions arise concerning the spacing versus the total number of births.[9] A decrease in income might result in a postponement of a child without reducing the total number of children. At the macro level, the distinction is between the tempo of fertility and the quantity of fertility. Changes in the tempo of fertility might influence annual fertility measures such as the crude birth rate, the general fertility rate, or the total fertility rate while leaving the underlying level of the quantity of fertility for each cohort over their lifetimes largely unchanged.[10] For example, fertility might be

[8] The explanation in Macunovich (2000) differs somewhat from our explanation in that the rate of change in relative cohort size is introduced to account for what is viewed as asymmetry in the relationship of age structure with fertility.
[9] See for example Cigno and Ermisch (1989), and Grabowski and Shields (1996).
[10] For a more detailed discussion of the relationship between period (annual) fertility rates and cohort rates see Ryder (1964), Shields and Tracy (1983), and Foster (1990).

positively related to income in the short run, over the business cycle, but negatively related in the long run. In this case, a catching-up phenomenon might exist. This catching-up phenomenon was thought by many to be the major factor explaining the baby boom in the U.S. This explanation of the baby boom has been challenged by other interpretations of the empirical evidence including that provided by Easterlin.

The Easterlin hypothesis has been incorporated into fertility models at the macroeconomic level by Wachter (1975), Butz and Ward (1979), Shields and Tracy (1986), and Macunovich (1996a and 1998c) for U.S. data. The results generally support the Easterlin hypothesis for the U.S., Canada, Australia, England, and Wales while studies for other countries have found little support for this hypothesis.[11] More recently, the Easterlin hypothesis has been examined using panel data by Pampel (2001), Gauthier and Hatzius (1997), and Jeon and Shields (2005). These results find a smaller but substantial impact of relative cohort size on fertility. The differences may depend more on differences in economic structure, particularly as they affect the labor market, than on cultural differences.[12]

The fertility variable chosen to analyze is the total fertility rate, *TFR.* TFR is chosen as the aggregate measure of fertility for two basic reasons. First, it controls for the age structure of the population. It may be important in this model to use *TFR* or some other age-specific fertility rate because many of the explanatory variables also depend on age structure. This dependence may occur with the unemployment rate, the tax exemption variable, the immigration variable and the income variables. Second, *TFR* has the advantage that it can be interpreted as the number of births per woman during her lifetime that would occur if the age-specific rates remain at their current levels.[13] The TFR is preferable to age-specific measures for narrower ranges of ages because it captures both the current impact of a variable through

[11] See Easterlin and Condran (1976), Ermisch (1979), and Wright (1989).

[12] See Pampel (1993). Carlson (1992), for example, suggests that there may be an inverted relationship between relative cohort size and fertility in Eastern European countries due to differences in economic structure.

[13] A difficulty in using *TFR* is that the series in *Historical Statistics* only goes back to 1940. Bogue (1969), however, has *TFR* back through 1917. In this paper, the *TFR* series was extended from 1917 to 1913 on the assumption that *TFR* and *GFR* move by the same percentage rate.

changing the spacing of births and the long-term impact on the completed number of births on the assumption that age-specific birth rates remain the same.

There are two types of variables that have been used to represent the Easterlin hypothesis. First, there are variables that measure relative income or employment prospects. For example, Wachter (1975) looks at *w/w**, where *w* is the current wage rate and *w** is a lagged function of past wage rates capturing the desired standard of living. Second, there are age-structure variables designed to capture relative cohort size.[14] Following Shields and Tracy (1986), relative cohort size, *AGE,* is the ratio of the population aged 18-24 to those 25-64. Ages 18 to 24 are critical years for accumulating human capital and entering the labor force. It is for these years that supply and demand in the labor market are most crucial for determining lifetime income.[15] It is readily apparent, in looking at Figure 1, that there could be some negative relationship between the age-structure variable and the total fertility rate (per 1000 women), which would appear to be most evident during the years of the baby boom. Although WAP examined U.S. fertility for the period 1913-1984 and emphasized the income tax exemption variable, they did not include a variable for relative cohort size that is introduced in this paper.

The other variables included in this paper are the husband's wage income plus asset income, the wife's wage rate, the size of the tax exemption per child, the infant mortality rate, the unemployment rate, and the immigration rate. As mentioned, the husband's income represents the income effect, which is usually thought to be positive. In contrast, the female wage rate is thought to reflect the price of children and, hence, has both an income and a substitution effect. Thus, unless children are to be thought of as being inferior goods, male plus asset income will be positively related to fertility, while women's wage rate could be either positively or

[14] For example, Shields and Tracy (1986) use the size of the population aged 18 to 24 divided by the size of the population 25 to 64, Wright (1989) uses the inverse of the size of the number of men aged 15-29 to those aged 30-64 and Ermisch (1988) uses the size of the cohort born in a year relative to the size of the cohort born in 1948.

[15] Age-specific male and female wage rates are available in the current population survey to construct relative wage income for young versus older adults. However, these data are only available from the 1960s and, consequently, cannot capture the important baby boom years.

negatively related to fertility. Also, the tax exemption variable reduces the costs of children and as such would have a positive impact on fertility.

The infant mortality rate is thought to be positively related to fertility for many reasons. In demographic transition theory, a long-term lag exists between declines in infant mortality and fertility, which is caused by the eventual adjustment of social norms and mores to the reduced need for higher births.[16] Microeconomic choice theory suggests the possibility of shorter lags, as assumed in this model. Since *n*, in the utility function, is the demand for surviving children, a decline in infant mortality will reduce the number of births needed to achieve a given number of surviving children.[17]

The unemployment rate is introduced as a business cycle variable. Unemployment or the threat of unemployment can influence fertility in two basic ways. First, to the extent that unemployment represents a transitory loss of income, it will increase the spacing of the next birth.[18] Second, unemployment can dramatically alter expectations of future income and feelings of economic security. As a consequence, families might save more money have fewer children and work more hours. In addition, a loss or threatened loss of the husband's income might lead the wife to enter the labor force. Her labor force participation is usually thought to lead to lower fertility. In either event, fertility is expected to be negatively related to unemployment.[19]

Immigration might also be positively related to fertility. Both Easterlin and WAP expect a positive effect of immigration on fertility. WAP expect a positive effect because immigrants are thought to have higher fertility due to pronatalist cultures in the original countries. However, this effect cannot fully be captured

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[16] Shields and Tracy (1986) found long lags between infant mortality and fertility.
[17] See Schultz (1981) for a more detailed discussion. Also Chowdhury (1988) found fairly short lags between infant mortality and fertility for a number of countries. He also found mutual causality between fertility and infant mortality in many cases.
[18] Becker (1960) assumes that the consumer durable analogy of childbearing holds. Furthermore, see Cigno and Ermisch (1989) for a theoretical framework for analyzing the timing of births.
[19] WAP speculate that it is possible to have a positive sign if unemployment of the wife reduces the opportunity cost of her time. This possibility seems unlikely.

because immigration is the flow of recent immigrants and it is not necessarily the flow that would influence births. Another reason for a positive relationship, suggested by Easterlin, is that immigration may be a response to excess demand for younger workers. This demand would be negatively related to *AGE* because immigration could be the response to the demand for entry-level workers. Since *AGE* is negatively related to fertility, immigration might also be positively correlated with the fertility of non-immigrants. However, the data do not support this possibility because immigration and *AGE* are positively correlated. Immigration to the U.S. may be more related to legal barriers and supply conditions in other countries than to demand condition in the U.S.

The early year data (1913-1984) are provided by WAP (1990), and updated with three exceptions; the age-structure data, the TFR data and the immigration data. WAP used the GFR and not the TFR as the fertility variable and did not consider age structure. Their immigration variable was replaced by a more complete historical series from *Immigration Statistics 2004*. Early years of TFR were taken from Shields and Tracy (1986) and updated using overlapping versions of *Statistical Abstracts of the US.* The same was done for the age-structure variable.

Updating most of the data for most of the series is straightforward. There are numerous sources for the unemployment rate. The infant mortality rate can be found in overlapping editions of *Vital Statistics* in Statistical Abstracts of the U.S., and the female wage rate is available from the *Current Population Survey.* Variables requiring calculation are the average marginal tax rate and the husband's income. First, the average marginal tax rate variable was updated until 1994 by Stephenson (1998) and further updated using data from annual issues of *Statistics of Income* data provided by the Internal Revenue Service. Data are provided for 31 groups of tax payers according to their adjusted gross income (AGI), from low to high. For each of those groups (i=1,2,…,31), we have the proportion of taxpayers in the group (W_i), average AGI of the group (X_i), and the average taxes they pay ($TaxX_i$). The average marginal income tax rate (AMITR) is then

$$AMITR = \sum_i W_i * (TaxX_i - TaxX_{i-1}) / (X_i - X_{i-1}) \ .$$

Second, the male income data are extended by calculating average asset income per male as total asset income divided by the size of the male population and then adding this to median wage income of males. The series is adjusted using 1984 values so that our series was the same in 1984 as the WAP series. Each subsequent year was multiplied by this adjustment factor.

The use of all these variables in a time series as explaining fertility can be justified on the grounds that they are nonstationary individually but cointegrated. McNown and Rajbhandary (2003) tested the cointegrability of fertility with relative cohort size as a possible explanatory variable. The results are supportive of the Easterlin hypothesis. The data, however, are for a relatively short time period, 1948 through 1998. The power of tests of long-run relations increases with the length of the time period. We will test these hypotheses in the next section.

4. Error Correction Model and its Generalized Impulse Response

Having specified a vector of variables in which we are interested, we now turn to testing for the existence of both long-term and short-term relationships among the variables. The basic model for determining whether there is a long-run relationship among a vector of variables comes from cointegration theory. A long-term relationship is thought to exist between these variables if they are cointegrated. Therefore, short-run and long-run relationships are modeled in an error correction mechanism whereby the relationships among cointegrated variables converge to their long-term association. We will test for cointegration and then estimate the vector error correction model with the focus on the impulse response function.[20]

Prior to examining the long-term cointegrating relationship among variables, a test for the nonstationarity of the individual series will be conducted to avoid spurious results (Granger and Newbold 1986). Unit root tests evaluate whether or not a

[20] See Ermisch (1988) and Cigno and Rosati (1996) for a similar approach.

variable is nonstationary. This paper employs two widely used unit root tests: the augmented Dickey-Fuller (ADF) test and the Phillips-Perron (PP) test.[21] When the test results conflict between the two tests, the GLS transformed Dickey-Fuller (DF-GLS) test is used to determine the final statistical decision (Elliot, Rothenberg, and Stock 1996). For the time series except age structure, both unit root tests for the first differences reject the null hypothesis of nonstationarity at the 1% significance level, while those for level cannot; as a consequence, those series are integrated of order 1, I(1). The ADF unit root test indicates quite unusual statistical decision on the age structure variable; The ADF for the level can reject the null hypothesis (that implies stationarity) but the ADF test for the first difference of age structure variable cannot reject the null hypothesis of nonstationarity (that is nonstationarity), while the tests for its second differences reject the null hypothesis (that is stationarity), as shown in Table 1. In contrast, the PP test indicates nonstationarity for level but stationarity for the first and second differences. This indeterminacy makes us rely on the DF-GLS test, which confirms the test results by the ADF unit root tests. We suspect that the statioinarity of level data is due to the insufficient time-modeling in both ADF and DF-GLS test, which assume a linear trend. Rather the quadratic trend is preferred. Also the AGE variable is very smooth with large swings, which is slightly different from the fertility rate variable. Therefore, the age structure variable is concluded to be integrated of order 2 (that is, I(2)) for our sample period.[22]

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[21] Stationarity of a time series variable implies that the variable's stochastic properties are invariant with respect to time; thus, for example, the mean and covariance with other variables do not depend on time. The ADF approach (Dickey and Fuller (1979) and Said and Dickey (1984)) controls for higher-order correlation by adding lagged difference terms of the dependent variable x_t to the right-hand side of the regression: $\Delta x_t = \mu + \gamma_1 x_{t-1} + \delta_1 \Delta x_{t-1} + \delta_2 \Delta x_{t-2} + ... + \delta_q \Delta x_{t-q} + \varepsilon_t$. The corresponding null and alternative hypotheses are: $H_0 : \gamma_1 = 0$ vs $H_1 : \gamma_1 < 0$. The PP test, proposed by Phillips and Perron (1988), makes a nonparametric correction from the AR(1) regression of the error term to account for the serial correlation in the error term. The correction is made by using an estimate of the spectrum of the error term at frequency zero, so the PP test is robust to any unknown form of heteroskedasticity and autocorrelation. MacKinnon (1991) implements a much larger set of simulations and estimates the response surface using the simulation results, permitting the calculation of Dickey-Fuller critical values for any sample size and for any number of right-hand variables.

[22] This result is supportive of McNown (2003) which finds that a relative income variable is I(2).

The cointegration test aims to check whether a linear combination among nonstationary processes can be stationary, which indicates possible long term relationships. In our test, the first differences of AGE are used due to the data property of I(2) and Johansen's method is applied for a cointegrated relationship among nonstationary variables used in this study, in order to test the restrictions imposed by cointegration on the unrestricted vector autoregression involving the series.[23] Table 2 shows that the cointegration trace test yields two cointegrated relationships but the maximum eigenvalue test indicates one cointegration at 5% significance level among eight time series. Due to this discrepancy on the statistical decision, we draw the cointegrated relationships under two difference statistical decisions. The first cointegrated relationship under two cointegration assumptions (by the trace test) is very similar to the cointegration under only one cointegration (by the maximum eigenvalue test). The other cointegrated relation under two cointegration assumptions is very flat, close to 0 over the sample period. Thus, it should not be harmful to assume only one cointegration, which is used for our further analysis in estimating an error correction model and also for a generalized impulse response analysis. The top panel of Table 3 indicates our cointegrated relationship, where the coefficient of fertility is normalized to unity separately. In the long run, the total fertility rate moves in the same direction as the tax exemption, wife's wage rate,

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[23] We consider the *p*-dimensional autoregressive process X_t defined by the equations:
$X_t = \Pi_1 X_{t-1} + \Pi_2 X_{t-2} + ... + \Pi_k X_{t-k} + \Phi D_t + \varepsilon_t$, $t = 1,2,...,T$ for fixed values of $X_{-k+1},..., X_0$, and independent identically distributed errors ε_t that are $N_p(0,\Omega)$. The deterministic terms D_t can contain a constant, a linear term, seasonal dummies, intervention dummies, or other regressors that we consider fixed and non-stochastic. This equation can be rewritten in error correction form:

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \varepsilon_t, t = 1,2,...,T \text{ where } \Pi = \sum_{i=1}^{k} \Pi_i - I \text{, and } \Gamma_i = -\sum_{j=i+1}^{k} \Pi_j .$$

Granger's representation theorem asserts that if the coefficient matrix has reduced rank $r < k$, then there exist $k \times r$ matrices α and β each with rank *r* such that $\Pi = \alpha\beta'$ and $\beta' X_t$ is stationary. *r* is the number of cointegrating relations (the *cointegrating rank*) and each column of β is the cointegrating vector. Johansen's method is to estimate the Π matrix in an unrestricted form, and then test whether the restrictions implied by the reduced rank of Π can be rejected or not. To determine the number of cointegrating relations *r*, we can proceed sequentially from *r* = 0 to *r* = *k*-1 until we fail to reject. See Johansen and Juselius (1990), Hamilton (1994), Banerjee *et al* (1993) and Johansen (1995).

infant mortality, immigration and the change of age structure. The fertility rate moves in the opposite direction of the husband's income, but the unemployment rate does not have any long-run impact on the fertility rate.

The lower panel of Table 3 also presents the estimated vector error correction model, representing the short-run as well as long-run relationships. The lag terms of fertility are included in order to control for the serial correlation in the time series model. The tax exemption and age structure have an impact on the fertility rate not only through the short-run path (that is, through differencing operators) but also through the long-run path (that is, through cointegrated relationship). In contrast, the unemployment rate has an impact on the fertility rate only in the short term. However, the wife's wage rate, infant mortality, husband's income and immigration only show the long-run impact through the cointegrated relationships, but no short-run influence. That infant mortality has a long-run but not a short-run impact on fertility tends to support the theory of the demographic transition over static choice theory as being the dominant explanation of the decline in fertility resulting from a decline in infant mortality (see Chowdhury, 1988). The lack of a short-run relationship of income variables for both husband and wife with fertility is consistent with the notion that lifetime income determines fertility. An exogenous shock to income should only impact income through its impact on long-run lifetime earnings.

Now, each variable's impact on fertility is measured. A shock to each variable to fertility is transmitted to all of the other endogenous variables within the dynamic structure of the error correction model. That is, an impulse response function traces the effect of a one-time shock to each variable's innovations on current and future values of the total fertility rate. To orthogonalize the impulses, the Cholesky decomposition is popularly used to calculate the inverse of the Cholesky factor of the residual covariance matrix. However, these impulse responses may change dramatically when the ordering of the variables are switched in an error correction models. Consequently, an alternative method that is suggested by Pesaran and Shin

(1998), called the generalized impulse method, is used in the paper.[24] Figure 3 shows the impulse response function tracing the effect of a one-time positive shock of each variable to current and future values of the fertility variable, while Figure 4 accumulates the responses.

Figure 3 and Figure 4 indicates that the age structure, Tax exemption and the income-related variables, the wife's wage rate and the husband's income, not only have significant short-run impact on fertility (that is within five years), but also have a significant positive influence over thirty years. Hence, the Easterlin hypothesis in the U.S. and the arguments of WAP are confirmed. Our results also support the catching-up hypothesis and policy implication of tax exemption. However, the infant mortality rate, immigration and unemployment do not have significant impact on fertility both in the short-run and the long-run.

5. Concluding Remarks

Relative cohort size appears to be an important consideration for macroeconomic models of US fertility and offers support for the Easterlin hypothesis over the long run, when the Easterlin hypothesis is interpreted in terms of changes in age structure affecting the level of fertility. That relative cohort size is I(2), which means that the variable becomes stationary after taking the differences twice, during this time period casts suspicion on studies using the level of relative cohort size as an explanatory variable having a long-term relationship with fertility.[25] Studies looking at different time periods or different countries might well consider whether relative cohort size is I(2) and not I(1). Also, our results support the catching-up hypothesis. There is a significant short run relationship between the wife's wage rate or male plus asset income and fertility, and in the long run this relationship is negative. In addition, the child tax deduction appears to be an important policy

[24] A statistical program *Eviews 5.1* is used to this calculation. See Levtchenkova, Pagan and Robertson (1998) and Gonzalo and Ng (2001) for identifying the shocks in the co-integrated systems.
[25] A similar result for a different time period using different definitions of RCS is found by McNown (2003).

variable influencing births.

It should be cautioned that the importance of relative cohort size implies little about the reasons why age structure is important. An explanation concerns the aspirations of young adults. Their aspirations might be determined by the living levels of older adults or of their parents a decade or so ago. By influencing relative income, age-structure influences tastes and, hence, fertility. Another explanation is simply that age structure influences the income of young adults relative to average income. No taste argument is implied. Further research is needed to determine which of these best explains the impact of age structure on US fertility. However, the results do offer support for the Easterlin hypothesis. However, the baby boom is not entirely explained by the Easterlin hypothesis because the results also provide evidence for a catching-up phenomenon.

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Table 1: Unit Root Tests

(A) ADF test and PP test

	Integration	ADF test		PP test	H_0 of I(1)	Comments
	of data	statistics	lags	statistics		
Total Fertility Rate*	Level	-3.022	5	-1.626	cannot reject	Nonstationary
	Difference	-6.178	0	-6.227	reject	Stationary
Tax Exemption	Level	-2.208	1	-1.582	cannot reject	Nonstationary
	Difference	-6.063	0	-5.855	reject	Stationary
Wife's wage rate *	Level	-2.738	0	-2.906	cannot reject	Nonstationary
	Difference	-8.519	0	-8.485	reject	Stationary
Infant mortality *	Level	-2.595	1	-2.010	cannot reject	Nonstationary
	Difference	-14.366	0	-15.165	reject	Stationary
Immigration	Level	-2.859	0	-2.909	cannot reject	NonStationary
	Difference	-9.232	0	-9.279	reject	Stationary
Husband's income *	Level	-0.040	1	0.186	cannot reject	Nonstationary
	Difference	-6.543	0	-6.543	reject	stationary
AGE*	Level	-4.313	5	-1.747	undetermined	undetermined
	1st diff	-2.675	6	-4.793	undetermined	undetermined
	2nd diff	-16.046	0	-20.532	reject	stationary
Unemployment	Level	-3.345	1	-2.647	cannot reject	nonstationary
	Difference	-6.684	0	-6.625	reject	stationary

Note that ADF test means Augumented Dickey-Fuller Test, and PP test means Phillips Perron Unit Root Test. Wife's wage rate, infant mortality and husband's income are modeled with trend and constant, while others are only with a constant term. MacKinnon critical values for rejection of hypothesis of a unit root for considering trend and constant are **-4.066 for 1% Critical Value,** -3.461 for 5% CV, and -3.157 for 10% CV. Those of only considering constant are **-3.509 for 1% CV**, -2.896 for 5% CV, and -2.585 for 10% CV. The optimal lags for ADF are selected by the Bayesian information criteria.

(B) DF-GLS test for AGE variable

	Integration	DF-GLS test	Test critical values			H_0 of I(1)	Comments
	of data	statistics	1%	5%	10%		
AGE*	Level	-4.389	-3.645	-3.084	-2.791	undetermined	nonstationary
	1st diff	-2.730	**-3.652**	-3.091	-2.797	cannot reject	nonstationary
	2nd diff	-16.202	-3.633	-3.075	-2.782	Reject	stationary

Table 2 - Long-run Transmission Mechanism (Cointegration Test)

Hypothesized		Trace			Max-Eigen		
No. of CE(s)	Eigenvalue	Statistic	Critical Value*	Prob.**	Statistic	Critical Value*	Prob.**
None *	0.478	187.025	159.530	0.001	55.284	52.363	0.024
At most 1 *	0.404	131.740	125.615	0.020	43.927	46.231	0.087
At most 2	0.316	87.813	95.754	0.155	32.232	40.078	0.291
At most 3	0.218	55.581	69.819	0.395	20.920	33.877	0.690
At most 4	0.151	34.661	47.856	0.466	13.940	27.584	0.827
At most 5	0.141	20.721	29.797	0.375	12.937	21.132	0.458
At most 6	0.057	7.785	15.495	0.489	5.011	14.265	0.741
At most 7	0.032	2.774	3.841	0.096	2.774	3.841	0.096

Note that "Trace test" indicates 2 cointegrating equations at the 0.05 level while "Maximum eigenvalue test" indicates 1 cointegrating eqn(s) at the 0.05 level.
* denotes rejection of the hypothesis at the 0.05 level.
**MacKinnon-Haug-Michelis (1999) p-values.

Table 3 - Estimated Cointegrated Relation and Error Correction Model

(A) Estimated Cointegrating Equation

$$Ci_{t-1} = 5876.068 + TFR_{t-1} - \underset{[-3.78]}{6.478} * EXEMPT_{t-1} - \underset{[-2.93]}{3828.152} * FEMALEW_{t-1} - \underset{[-5.16]}{107.457} * InfantMortality_{t-1}$$

$$- \underset{[-2.46]}{0.0015} * Immigration_{t-1} + \underset{[2.98]}{0.572} * MALE\_INC_{t-1} - \underset{[-5.49]}{164671.535} * \Delta(PAGE_{t-1}) + \underset{[0.07]}{2.427} * U_{t-1}$$

where the values in the parentheses are t-statistics.

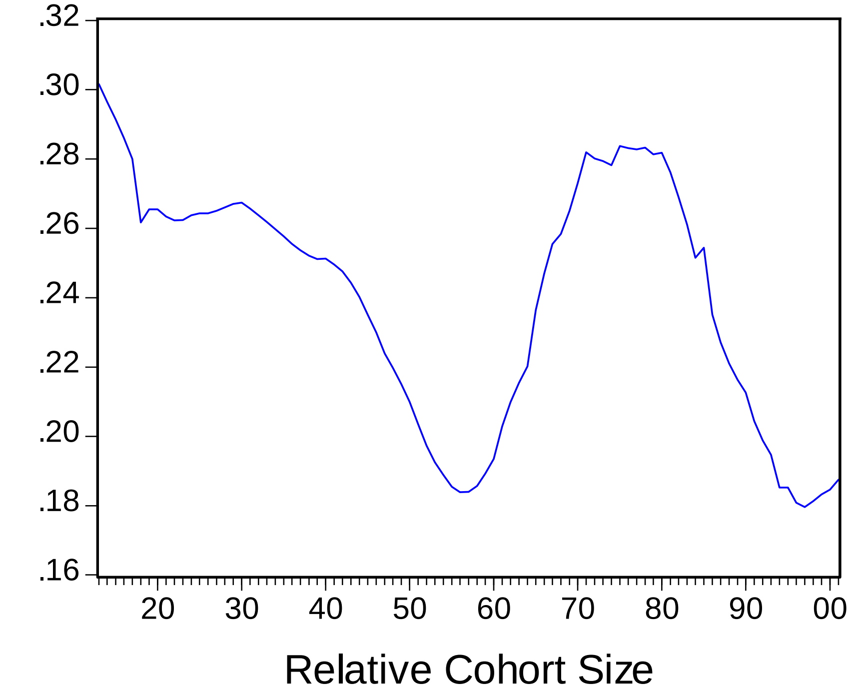
(B) Estimated Vector Error Correction Model

The dependent variable is ΔTFR_t.

	Coefficient	t-statistics
CointEq1	0.035	2.54
$\Delta(TFR_{t-1})$	0.376	3.22
$\Delta(TFR_{t-2})$	-0.199	-1.50
$\Delta(EXEMPT_{t-1})$	-0.443	-1.32
$\Delta(EXEMPT_{t-2})$	1.094	3.61
$\Delta(FEMALEW_{t-1})$	85.406	0.38
$\Delta(FEMALEW_{t-2})$	19.113	0.08
$\Delta(INFANTM_{t-1})$	-1.977	-0.41
$\Delta(INFANTM_{t-2})$	-6.668	-1.48
$\Delta(Immigration_{t-1})$	-0.0001	1.27
$\Delta(Immigration_{t-2})$	0.00004	0.69
$\Delta(MALEW_{t-1})$	-0.063	-0.92
$\Delta(MALEW_{t-2})$	0.001	0.02
$\Delta^2(PAGE_{t-1})$	7182.713	2.44
$\Delta^2(PAGE_{t-2})$	277.623	0.10
$\Delta(U_{t-1})$	-15.940	-2.71
$\Delta(U_{t-2})$	4.468	0.69
Constant	-22.500	-1.14

Note that the full version of ECM estimations is available upon request.

Figure 1: Age Structure Variable and Total Fertility in the U.S.



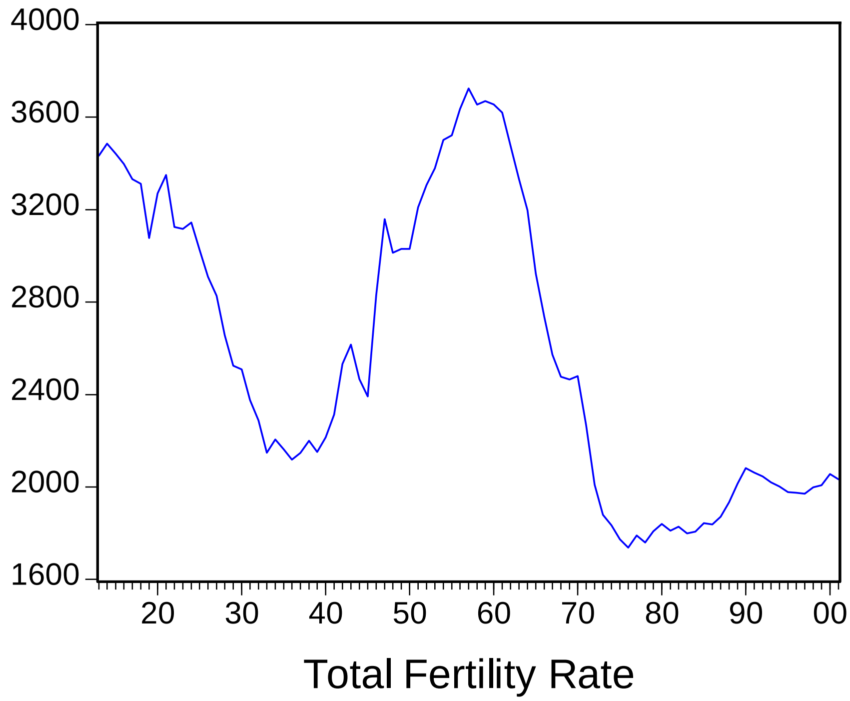
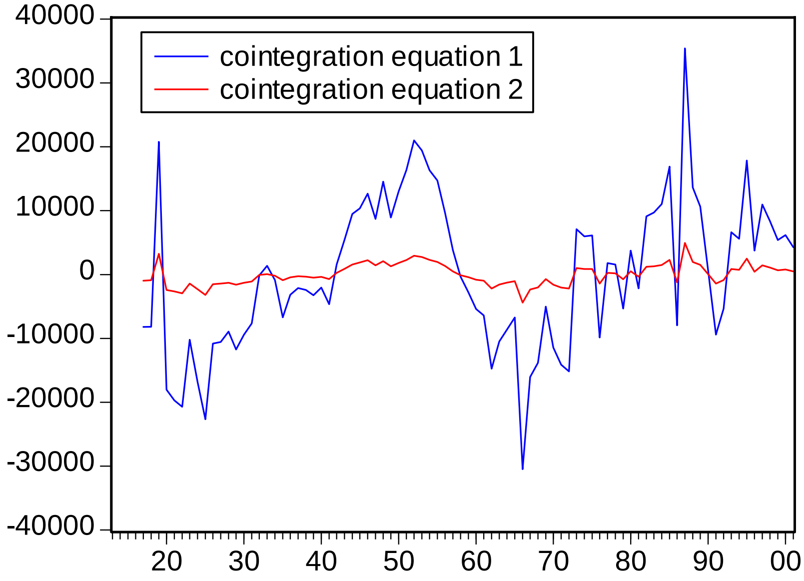
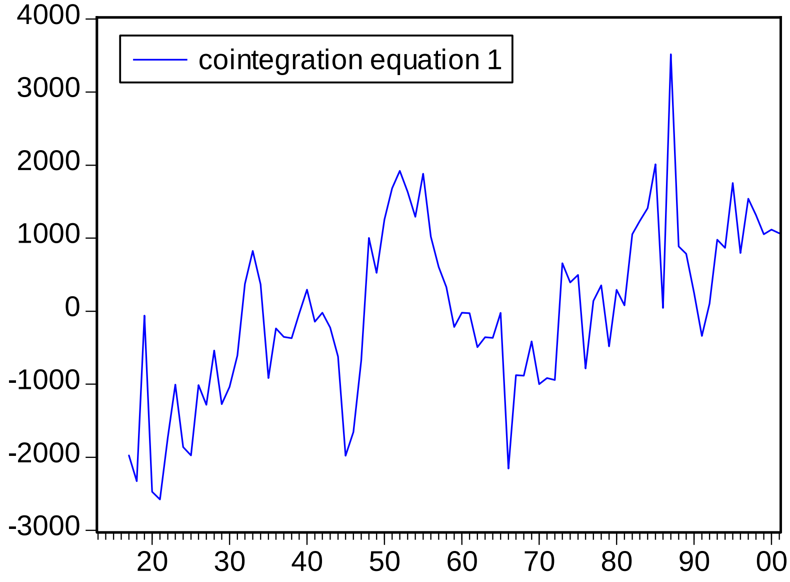


Figure 2: Long-run relationships under two different modeling strategies



(A) Under two cointegration assumption



(B) Under one cointegration assumption

Figure 3: Impulse response function of the fertility rate to generalized one S.D. innovations

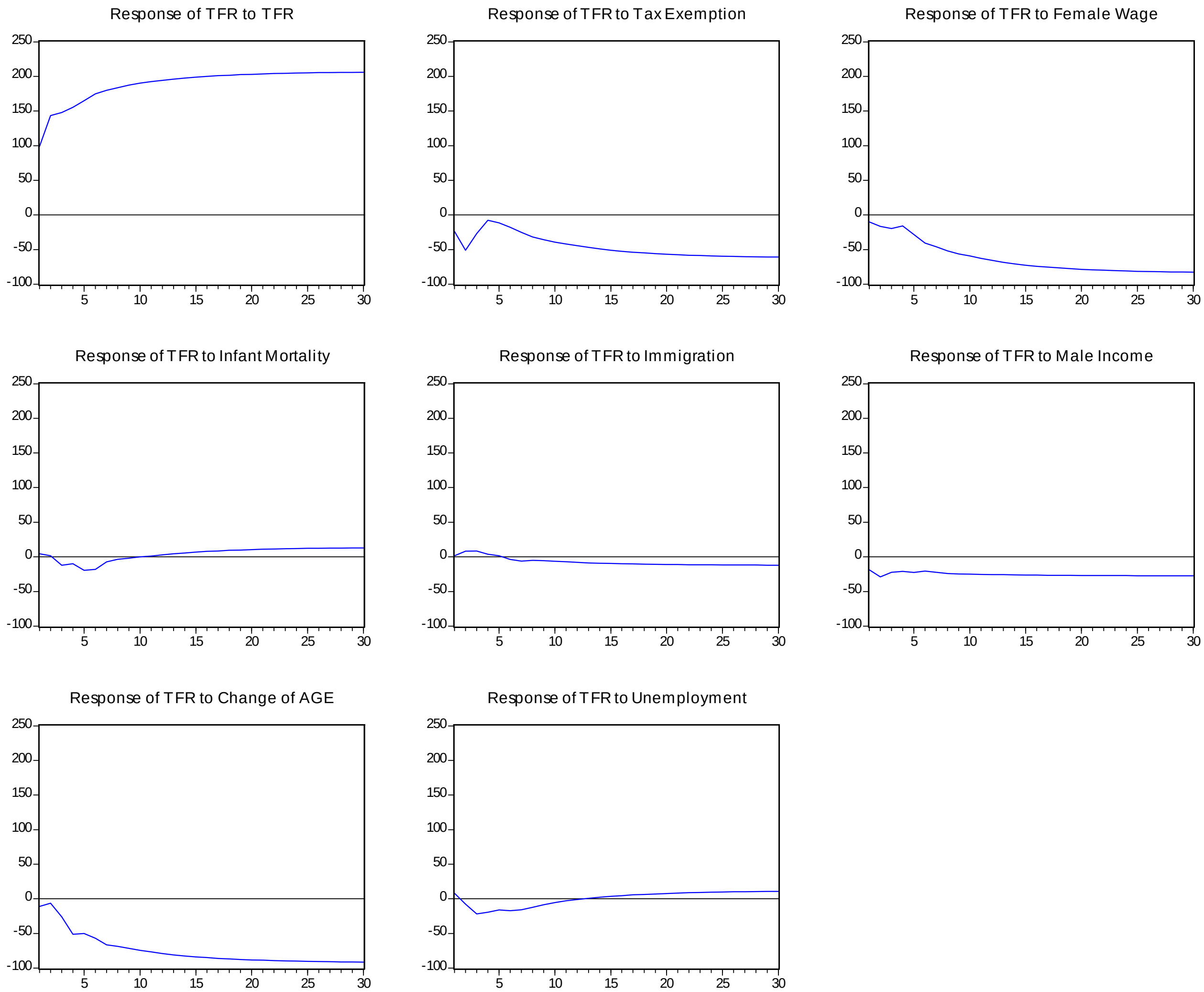


Figure 4. **Accumulated** response for the fertility rate
to one standard deviation innovations of other explanatory variables