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The Effect of Tax-Benefit Policies on Fertility and Female Labor Force Participation in the United States

Cristóbal Ridao-Cano
Department of Economics, University of Colorado at Boulder
Boulder, Colorado

Robert McNown
Department of Economics, University of Colorado at Boulder
Boulder, Colorado

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Center for Economic Analysis
Department of Economics

University of Colorado at Boulder
Boulder, Colorado 80309

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Cristóbal Ridao-Cano and Robert McNown

Department of Economics and Population Program at IBS

University of Colorado at Boulder

Campus Box 484

Boulder, Colorado 80309

E-mail: Robert.Mcnown@colorado.edu

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Abstract

This paper presents an investigation of the effects of the tax exemption for dependents and the child care tax credit on age-specific fertility rates and female labor supply for the U.S. 1948-1997. Effects of alternative tax and benefit programs on fertility and investments in child quality are analyzed in a theoretical model. Implications of the model are tested within a cointegration framework. Statistical evidence supports the existence of two cointegrating relations for each age group. One relation is identified as a fertility equation, with theoretically appropriate signs and significance of all coefficients, and the second is identified as a female labor supply equation. The two tax variables show positive effects on fertility, with the tax exemption accounting for a large portion of the short run variation in fertility for 20-24 year old women.

Keywords: Fertility, tax policies, time series models, cointegration

1. Introduction

Fertility rates in the US have experienced a substantial downward trend over the twentieth century, only temporarily interrupted by the post-war baby boom. These changes in fertility rates reflect fluctuations in both the *tempo* and the level of completed fertility. These changes in fertility have been accompanied by a precipitous rise over time in female labor force participation, particularly among mothers with young children.

Contrary to some European countries, where tax laws and benefit programs have been introduced with the aim of increasing fertility (Gauthier 2000), the United States has not adopted explicit pronatalist policies. However, the United States has had a long tradition of tax and welfare policies that affect the cost of bearing and raising children and may therefore have influenced fertility rates. Tax provisions include the federal tax exemption for dependents (introduced in 1917), the Child Care Tax Credit (introduced in 1976), and the Child Tax Credit (introduced in 1998). These programs provide cash transfer to families with children regardless of their income, provided this income is high enough to be subject to federal income taxation.[1] Welfare programs include the Aid to Families with

[1] The value of the tax exemption and the Child Care Tax Credit (CCTC) depend, however, on income, and the latter is contingent upon the purchase of child care.

Dependent Children (introduced in 1935 and replaced in 1997 by the Temporary Assistance to Needy Families), and the refundable Earned Income Tax Credit (introduced in 1975), which provide cash transfers to poor families with children.[2] By affecting the cost of children these programs may affect female labor force participation indirectly through changes in fertility. Furthermore, some of these programs may provide direct work incentives (e.g. CCTC and EITC) while others may provide work disincentives (e.g. TANF).

There is a long tradition of studies looking at fertility and female labor supply, from both a theoretical and an empirical perspective.[3] However, the interest of economists in the effects of the tax-benefit system on fertility and female labor supply, and the implications for the optimal fiscal treatment of the family, is relatively more recent. Theoretical work has combined both static and dynamic frameworks.[4] Much of the empirical work is based on microdata,[5] although there are some studies using aggregate time series data.[6]

[2] Starting in 1991, the Earned Income Tax Credit (EITC) was made available to childless families, although the maximum credit for those is much smaller than that for families with children. In addition the EITC, as well as the TANF, depend on the number of children in the family.

[3] See, for example, Moffitt (1984), Lehrer and Nerlove (1986), and Hotz and Miller (1988).

[4] Theoretical work includes Cigno (1986, 1991, 1996), Batina (1986), and Walker (1995).

[5] See, for example, Cigno (1990), Whittington (1992), Averett et al. (1997), Schultz (1997), Moffit (1998), and Hotz (2001).

[6] Ermish (1988), Whittington, Alm and Peters (1990), Georgellis and Wall (1992), Zhang, Quan, and Van Meerberger (1994), Cigno et al. (2001).

The objective of this paper is to reexamine the effect of family policies and other economic variables on age-specific fertility and female labor market behavior, applying contemporary time series methods to aggregate data from the United States over the post-World War II period. The analysis of tax-benefit policies is mainly focused on the tax exemption for dependents since this is the only program (along with AFDC) that has been in place for the entire time period considered here. Furthermore, in contrast with other programs the tax exemption for dependents applies to most families with children, it is not contingent on specific behavior, and it is received over the entire dependency period of the child. Thus the tax exemption for dependents is more likely to affect fertility rates over the long run than other policies.

Although household survey data offer the possibility of testing behavioral hypotheses, estimation of aggregate relations may be more relevant for policy analysis. Some variables, when measured at the household level reflect each household's relative position within society, and such effects may be absent at the aggregate level. Conversely, social contagion may induce aggregate responses to changes that are not reflected in individual differences. Therefore, estimates of the aggregate response of fertility rates or labor supply to policy changes require time series analysis. The potential aggregation bias is reduced in this study by considering the fertility and labor market behavior of different age cohorts separately, which also allows us to consider issues related to the timing of fertility.

In time series studies, the tax exemption for dependents has been found to have a positive effect on fertility in the US (Whittington et al. 1990; Georgellis and Wall 1992) and Canada (Zhang et al. 1994). The paradox of these studies is that while they find significant effects of the tax exemption on fertility, presumably operating through economic incentives, the standard variables of the economic model of fertility (i.e. female wages and household or male income) have insignificant effects.

Methodological issues may account for this anomaly in empirical investigations of tax policy effects on fertility. In particular, regressions with aggregate time series present considerable challenges. Variables common to models of fertility and labor market behavior are most likely nonstationary time series that trend or drift away from their initial values. Such nonstationarity may undermine traditional estimation procedures, leading to spurious inferences about the relations among the variables. In addition, models of fertility and female labor supply are plagued by problems of endogeneity. Economic models of fertility generally include measures of female labor force participation, women's wages, female education levels, and relative male income as explanatory variables. However, all of these variables are quite plausibly endogenous, reflecting the outcome of household decisions concerning education, work, and child rearing by young adult men and women. The presence of endogenous regressors may undermine the identifiability of the model, so that the relations are not estimable. Even if the relations are identified, the endogeneity of regressors leads to inconsistent least squares estimators of the model parameters. Traditional treatment of this problem with instrumental variables procedures requires additional exogenous instruments, which raises the same concerns of endogeneity of these variables.

In view of the above considerations, this paper makes use of contemporary time series methods. In particular, estimation and testing is performed using the cointegration model of Johansen (1995) that is appropriate for analyzing relations among nonstationary time series. This methodology, described in section 3, is applied to the variables defined in section 4. Results presented in section 5 resolve the paradox described above, and policy implications of the estimates are discussed in the concluding section 6. A theoretical model presented in the following section provides foundation for the empirical analysis.

2. A Simple One-Period Lifetime Model of Family Choice

The following model provides a simple framework within which to study the effects of different tax-benefit policies, particularly the tax exemption for dependents. The model is developed in the spirit of Cigno (1991). The static nature of the model precludes the consideration of the timing of fertility and intertemporal labor supply. However, a dynamic model, while certainly richer and more realistic, requires further simplifications and yields fewer unambiguous predictions.[7] Finally, although the quality of children is not modelled in the empirical section, a model of family choice with tax-benefit policies would lack completeness without the consideration of the relationship between quantity and quality of children.

For ease of exposition, the model considers a prototypical household consisting, at the outset, of a woman and her spouse, who are assumed to act in unison to make fertility and time and resource allocation decisions. The couple derives utility from their own consumption, x, and from child services they receive during their lifetime, c, according to the utility function,

$$U = U(x, c) \tag{2.1}$$

[7] See Hotz et al. (1997), and Arroyo and Zhang (1997) for a review of the literature on dynamic models of fertility.

where $U(.)$ is twice differentiable and quasi-concave. Child services are governed by the following production process

$$c = nq$$

where n represents the number of children and q denotes child quality.[8] Child quality represents parental perception of a child's lifetime utility. The latter is assumed to be a function of the mother's child care time and child related market goods in excess of the minimum amount of the mother's time, t_c^0, and child-related goods, e^0, necessary to raise a child of whatever quality (i.e. subsistence levels),[9]

$$q \quad V(t_c, I)$$

where V is an indirect utility function with properties similar to $U($

$t_c \quad t_c' - t_c^0$ and $I \quad e - e^0$. Another simplification will be to assume that, above t_c^0, maternal time is perfectly substitutable with the father's time or nonparental time in childcare.[10] Assuming, further, that the wage rate of hired helpers is

[8]This formulation rules out the possibility that, with n variable, q could be complementary for x or n. It also assumes that all siblings are treated equally. The model further assumes that parents can control n perfectly (or, equivalently, that n is an expectation and parents are risk neutral) and the cost of fertility control is either negligable or independent of n.

[9]Child-related goods include, for example, consumption, health, and education. If parents give their children just the subsistence levels of these goods, q is zero.

[10]In this simplified structure, the father's allocation of time is given, and thus his income is

always lower than those of the parents, t_c^0 is all the maternal time that a child will get. Purchased child care is then treated as a child-related good. Thus we can rewrite (2.3) as

$$q = V(I)$$

Setting t_c^0 equal to unity, by measuring time in terms of the number of children, the mother's labor supply is given by

$$h = T - n$$

An implicit assumption in the above formulation is that leisure is exogenous (T is constant). Adding leisure as a choice variable does not alter the main results of the model. Furthermore, the mother's labor supply is still endogenous through the endogeneity of n

Assume the mother's human capital is given by the following linear homogeneous production function [11]

exogenous.

[11] The assumption of constant returns to scale is only an approximation to what is probably a function with increasing returns initially, and subsequent dimishing returns.

$$k = k^0 + \beta h \ , \ \beta > 0$$

where k^0 is the initial endowment (i.e. a measure of her innate talent, education and work experience before marriage) and β is the rate at which human capital increases with work experience. Her wage rate is endogenously determined by

$$w = \omega k$$

where ω is the market rate of return to k. Suppose maternal paid leave is available and takes the form $\chi = \eta w$, where η is the after-tax earnings replacement rate.[12] Under these assumptions, the budget constraint may be written as

$$x + [(1-s)p(e^0 + I) - \tau\theta]n = \varepsilon + A + (1-\tau)wh \equiv Y$$

where A denotes the husband's income, p is the price of child-related goods, s is a subsidy on p,[13] ε is a lump-sum subsidy, and θ is the statutory value of

[12] Note that withdrawal from the labor force for childbearing of an additional child is confined to time t_c^0, which is normalized to 1, and thus the relevant replacement rate per unit of maternal time is η. Whereas most industrialized countries offer paid maternity leave, the US law provides only 12 weeks of unpaid leave.

[13] This subsidy can be interpreted, for example, as an education subsidy or even as a child

the tax exemption per dependent. Lump-sum subsidies, child allowances, and maternity leave payments are financed by a proportional tax on earnings at a rate τ.[14] The value of the exemption to the family is equal to its statutory value times the tax rate τ, since the exemption reduces taxable income. The tax exemption per dependent is designed to partly compensate parents for the 'survival' cost of raising a child.[15] The basic results do not change if the child allowance θ takes the form of a public transfer per child, where the value of the transfer to the household is θ, if the transfer is not taxed, and $(1-\tau)\theta$ if taxed, or a child tax credit, where θn would be substracted from the tax liability τwh. Finally, note that Y stands for actual income and not full income, since it does not include the potential earnings foregone from childbearing.

The household decision problem is

$$\underset{\{x,n,I\}}{Max}\ U(x,c) \tag{2.9}$$

$$s.t.\ x+[(1-s)p(e^0+I)-\tau\theta]n=\varepsilon+A+(1-\tau)wh$$

care tax credit.

[14]Linearity of the tax system can be taken as a local approximation to a non-linear tax schedule.

[15]The statutory value of the exemption is roughly based on the income needed to maintain an adequate diet (Pechman 1983).

$$w = \omega k \quad \omega(k^0 + \beta h) \ , h \quad T \quad n$$

$$0 \leq n \leq n_0$$

The latter is the biological constraint, where n_0 is the mother's maximum reproductive capacity (defined as the difference between fecundity and infant mortality). Henceforth, I shall focus just on the case where n is endogenous, that is, the biological constraint is not binding.

From the first order conditions for (2.9), parents will equate the marginal utility of income spent on themselves to that of income spent on each child

$$U_x \quad U_c V'(I) \tag{2.10}$$

Likewise, parents also equate the marginal rate of substitution of quantity for quality to their relative price

$$\frac{V(I)}{V'(I)} = \frac{\pi}{(1-s)p} \tag{2.11}$$

where π is the shadow price of a child

$$\pi = \pi_0 + (1 \quad s)pI + (1-\tau)h\omega\beta \tag{2.12}$$

where $\pi_0 = (1-s)pe^0 + (1-\eta)(1-\tau)w - \tau\theta$ is the subsistence cost of a child. The shadow price of a child has three components. The first, $(1-\eta)(1-\tau)w$, is the direct opportunity cost of foregone earnings from childbearing. The second, $(1-\tau)h\omega\beta$, represents the potential wage increases missed through foregone human capital accumulation.[16] The third term, $(1-s)p(e^0 - I) - \tau\theta$, represents net lifetime expenditures incurred on a child.

2.1. The Effects of Changes in the Tax-Benefit System

The effect of an increase in θ on n and I are given by

$$\frac{\partial n}{\partial \theta} = -\tau s_{nn} + \tau n z_n > 0 \quad (2.13)$$

$$\frac{\partial I}{\partial \theta} = -\tau s_{in} + \tau n z_i \gtrless 0 \quad (2.14)$$

where s_{ij} is the substitution effect of the marginal cost of i on the demand for j, and z_i is the income effect. Suppose that n and I are normal goods, so that z_n and z_i are positive. Then since $s_{nn} < 0$ and $s_{in} > 0$, an increase in the statutory value of the tax exemption per dependent will induce parents to have

[16] We can think of this cost as the appreciation in the woman's stock of human capital that will not be realized if she stays at home for one unit of time.

more children (and the effect is increasing in τ), but has an ambiguous effect on child-quality investment. The ambiguity of $\partial I/\partial\theta$ arises because θ reduces the price of n, but not of I (and q), and thus makes the latter relatively more expensive. The effect on the mother's labor supply is of the same magnitude as that on n but of opposite sign. In contrast with θ, the effects of lump-sum subsidies, ε, are pure income effects

$$\frac{\partial n}{\partial \varepsilon} = \frac{\partial n}{\partial A} = z_n > 0 \tag{2.15}$$

$$\frac{\partial I}{\partial \varepsilon} = \frac{\partial I}{\partial A} = z_i > 0 \tag{2.16}$$

Now consider the effects of a change in the subsidy to child-related goods

$$\frac{\partial n}{\partial s} = p(enz_n - es_{nn} - s_{ni}) \gtrless 0 \tag{2.17}$$

$$\frac{\partial I}{\partial s} = p(enz_i - es_{in} - s_{ii}) \gtrless 0 \tag{2.18}$$

Thus s has ambiguous effects on both n and I. However, under the plausible assumption that the sum of income and own price effects dominates the cross price effect, an increase in the subsidy to child-related goods induces parents to have more children and invest more in child-related goods.[17] Next consider changes in the exogenous variables affecting the opportunity cost component of the shadow price of a child. The effects of those on child-quality investments are unambiguous

$$\frac{\partial I}{\partial \eta} = -ws_{in} < 0 \tag{2.19}$$

$$\frac{\partial I}{\partial \tau} = -[(wh - \theta n)z_i + (\theta + w + h\omega\beta)s_{in}] < 0 \tag{2.20}$$

$$\frac{\partial I}{\partial k^0} = (1-\tau)[\omega s_{in} + \omega h z_i] > 0 \tag{2.21}$$

$$\frac{\partial I}{\partial \beta} = (1-\tau)[2\omega h s_{in} + \omega h^2 z_i] > 0 \tag{2.22}$$

$$\frac{\partial I}{\partial \omega} = (1-\tau)[(k + h\beta)s_{in} + khz_i] > 0 \tag{2.23}$$

but their effects on the desired number of children (and then the mother's labor supply) are, except for the maternity leave benefit, ambiguous

[17] If we interpret this subsidy as a child care tax credit, the model would then predict a negative effect of this credit on female labor supply through an increase in fertility. In this simplified model all child care time in excess of the minimum required for the survival of the child (t_c^0) is purchased in the market, precluding any effect of this subsidy on the mother's labor supply through changes in the mother's child care time. The latter could partly compensate the negative effect of the subsidy on labor supply through increased fertility.

$$\frac{\partial n}{\partial \eta} = -ws_{nn} > 0 \tag{2.24}$$

$$\frac{\partial n}{\partial \tau} = -[(wh - \theta n)z_n + (\theta + w + h\omega\beta)s_{nn}] \gtrless 0$$

$$\frac{\partial n}{\partial k^0} = (1-\tau)[\omega s_{nn} + \omega h z_n] \gtrless 0 \tag{2.26}$$

$$\frac{\partial n}{\partial \beta} = (1-\tau)[2\omega h s_{nn} + \omega h^2 z_n] \gtrless 0$$

$$\frac{\partial n}{\partial \omega} = (1-\tau)[(k + h\beta)s_{nn} + khz_n] \gtrless 0 \tag{2.28}$$

However, making the standard assumption that own substitution effects dominate income effects, a rise in the rate of income tax is likely to increase the number of children, whereas increases in k^0, ω and β are likely to induce parents to have fewer children. Female education could also be introduced as a preference shifter in the utility function, whereby more educated women would prefer smaller or larger families. This "preference" effect would then be added to the above income and substitution effects for female education. In any case, the number of children born to a woman is likely to move in the opposite direction to her wage rate and labor supply.

In short, in the context of this model, the tax exemption for dependents provides an incentive to procreation, but is not a reliable policy instrument for improving the well-being of children. By contrast, a subsidy to the price of child-related goods appears to be an effective policy for improving the quality of children, but this policy is also likely to induce parents to have more children. An increase in maternal benefits (through an increase in the replacement rate) reduces the foregone earnings component of the opportunity cost from raising a child, thus inducing parents to have more children but of lower quality. Finally, increases in the variables affecting the woman's gross wage rate, or a more lenient fiscal treatment of her earnings, are likely to reduce the number of children (and hence increase the mother's labor supply) and improve the quality of their lives. For instance, the model suggests that the latter could be achieved by an increase in the amount of education received by women (i.e. higher k^0), increases in labor productivity (through technical progress), and the mitigation or reduction of gender discrimination.[18]

2.2. Implications for the Optimal Fiscal Treatment of the Family

In most countries, one of the goals of government policies is to 'help families with dependent children' through direct subsidies and/or favorable tax treatment. However, it remains to be explained what possible arguments can be used, in the

[18] These last two changes would raise the rate of return to human capital (corresponding to increases in β and ω).

context of the present model, to justify pronatalist policies and/or child-welfare policies.

One possible argument for helping families with dependent children through, say, tax exemptions per dependent is to compensate parents for having a child, in the sense of leaving them with the same utility level as they would have with no children. To see this, suppose fertility is exogenous and define the compensating variation in income for an exogenous change in the number of children from 0 to n

$$C(\varphi, n, u) = E(\varphi, n, u) - E(\varphi, 0, u) \quad (2.29)$$

where $E(.)$ is an extended version of the expenditure function indicating what income a family with n children must have, given the state of the economic environment described by vector φ (prices, wages, taxes, benefits), in order to achieve the utility level u. However, in the present model n is endogenous, and hence the function $C(.)$ does not exist (Cigno 1996), which would invalidate the standard compensation argument.

More to the point, parents decide to have children up to the point where the marginal benefit of a new born equals marginal cost. To this extent, in the absence of any population externality,[19] a tax exemption per dependent would, by reducing the private cost of a child, encourage parents to have more children than is socially desirable, and hence there would not be any argument for state intervention. If, on the other hand, the contribution of children (the quantity) to social welfare is not fully internalized by parents (who bear the full cost), then there may indeed be an argument for public intervention. In any case, the tax exemption per dependent is typically justified as a mean of equating ability to pay by households of different sizes, and not to serve as a pronatalist policy (Pechman 1983). The same efficiency argument applies to policies that reduce the cost of child-quality investment, such as subsidies to child-related goods, except that in this case we would talk about externalities associated with the well-being of children. In contrast to the exemption, the externality argument in the case of subsidies to child-related goods is consistent with the intended effect of these policies, which is to improve the quality of children.

Another possible argument for intervention is to achieve horizontal equity. Couples with different numbers of children, but otherwise identical, should enjoy the same after-tax utility. However, in this model identical individuals behave

[19] A population externality exists if the utility of each member of the society contributes to social welfare directly, as well as indirectly via its contribution to the utility of his or her parents.

identically, and will thus desire the same number of children.[20]

Finally, there is the proposition that channelling resources towards families with dependent children (and tied to their number) is an effective way to combat poverty.[21] This is based on the presumed negative correlation between family income and fertility, and in the recognition of the practical difficulty of adequately identifying needy families. The present model is likely to generate this negative correlation if the compensated own price elasticity of fertility is sufficiently greater than the income elasticity, and the main source of inequality between poor and non-poor families is earning ability disparity. Therefore, poverty relief may potentially be used as a valid argument for raising child benefits. However, the present model predicts that a rise in such benefits would induce an increase in births, particularly in low-income families, where child benefits account for a larger share of the budget. Hence, since this effect on fertility is uncompensated, this policy would paradoxically result in more poor children in poorer households, since female labor earnings would be lower.

Two general conclusions emerge from the above analysis. First, if the govern-

[20]The introduction of a taste parameter in the utility function to account for heterogeneity would produce different desired numbers of children for otherwise identical couples. However, this feature would not justify public intervention on horizontal equity grounds.

[21]The tax exemption for dependent would certainly not qualify as an anti-poverty policy since many low-income families do not pay taxes so they do not benefit from the exemption.

ment chooses its tax-benefit policy based on the belief that fertility behavior is exogenous, then a suboptimal policy will be chosen (Batina 1986). Second, if the policy maker has at heart the well-being of children and parents in needy families, a preferred policy, in the context of this model, would be to reduce the price of child-related goods by means of subsidies on these goods.

3. Empirical Methodology

Traditional regressions with time series data are grounded in the implicit assumption that the variables in the model are stationary. Heuristically, a stationary time series returns quickly and frequently to its mean value (or to a deterministic trend line), a proposition that does not appear to hold for common measures of fertility (see Figure 1) or its determinants. A time series that must be differenced d times to become stationary is integrated of order d, or $I(d)$, or equivalently it has d unit roots. A series' order of integration may be tested with a sequence of Dickey-Fuller (1979) tests, as suggested by Dickey and Pantula (1987). The initial hypothesis of two unit roots is tested from the significance of ρ in equation (3.1) using the critical values tabulated by Fuller (1996).

$$\Delta^2 x_t = \alpha + \rho \Delta x_{t-1} + \sum_{j=1}^{p-2} \gamma_j \Delta^2 x_{t-j} + \varepsilon_t$$

If the null hypothesis of two unit roots is rejected, the null of a single unit root is tested with the standard Dickey-Fuller regression (3.2), allowing a deterministic linear trend, if appropriate, under the alternative hypothesis:

$$\Delta x_t = \alpha + \beta t + \rho x_{t-1} + \sum_{j=1}^{p-1} \gamma_j \Delta x_{t-j} + \varepsilon_t$$

A set of integrated series may be linked by one or more stationary linear combinations called a cointegrating equation. A cointegrating equation represents a steady state or long run equilibrium relation that should reflect predictions of the underlying theory. If a vector of time series $\mathbf{x}_t$ is cointegrated, it can be given a dynamic representation as an error correction model,

$$\Delta \mathbf{x}_t = \boldsymbol{\alpha} + \sum_{j=1}^{q} \gamma_j \Delta \mathbf{x}_{t-j} + \delta \beta' \mathbf{x}_{t-1} + \mathbf{e}_t$$

where δ and β are nxr matrices, r is the number of cointegrating equations, and $\mathbf{e}_t$ is a vector of nonautocorrelated errors. Each variable responds to the lagged cointegrating relations, $\beta' \mathbf{x}_{t-1}$, according to the magnitudes of the adjustment parameters in δ.

Johansen (1995) presents likelihood ratio tests for the number of cointegrating equations (r) and the maximum likelihood estimators of δ and β. Subject to valid identifying restrictions, these estimators are consistent even in the presence of endogenous explanatory variables. Furthermore, these estimators are governed by asymptotic normal distributions, permitting valid statistical inference with conventional test statistics. This study has employed the Cointegration and Time Series (CATS) module (Hansen and Juselius 1995) of the Regression and Time Series (RATS) program (Doan 1996) for cointegration testing and estimation.

Analysis of the dynamic relations among the variables is provided by an innovation analysis, which traces the effects of an exogenous shock, or innovation, on each variable in the model. Confidence bands around the impulse response functions can be computed from asymptotic approximations to the standard errors of the impulse response coefficients (Lutkepohl 1990). The magnitudes of these responses are also described through a decomposition of a variable's forecast error variance into relative contributions from each variable's innovation.

Since the error terms in the system of dynamic relations are most likely correlated, they cannot be uniquely identified as innovations to one specific variable. A common strategy is to associate any contemporaneous correlation among the original errors with the variable that appears nearest the first position in the VAR. The ordering of the variables therefore requires identifying assumptions concerning the direction of contemporaneous linkages between the variables. Sensitivity of the analysis to these identifying assumptions can be checked by reversal of the ordering.

4. Variable Definitions and Characteristics

In modeling age-specific fertility rates explanatory variables have been defined to correspond with the ages of the women giving birth. Fertility rates were chosen to span the ages of highest childbearing, with age divisions matching the data available on explanatory variables. Consistency across variables was achieved with age categories of 20-24 and 25-34. Data are collected spanning the years 1948 through 1997.

Fertility rates are collected for women aged 20-24, 25-29, and 30-34 from *Historical Statistics of the US: Colonial Times to 1970* (U.S. Bureau of the Census, 1975), and from Table 4 of the *National Vital Statistics Report* (National Center for Health Statistics, 2000). The last two rates are aggregated using the relative populations of women in these two age groups tabulated from Current Population Reports, P-25 (U.S. Bureau of the Census, 1954-95) and U.S. Bureau of the Census internet site, "Resident Population of the United States: Estimates by Age and Sex."

Labor force participation rates for women aged 20-24 and 25-34 are collected from the *Handbook of Labor Statistics* (U.S. Bureau of Labor Statistics 1989), *Employment and Earnings* (U.S. Bureau of Labor Statistics 1990-1991, 1997-1998), and the *Statistical Abstract of the United States* (U.S. Bureau of the Census 1991-1998)

Male incomes are defined as median incomes of all males aged 20-24 and 25-34, matching the age categories of the fertility series. These data come from the U.S. Bureau of the Census Internet Site "Table P-7. Age-People by Median Income and Gender: 1947 to 1997."

The female wage series is constructed from the income in 1997 dollars of female workers, aged 20-24 and 25-34, reported in the U.S. Bureau of the Census internet site "Table P-7. Age-People by Median Income and Gender: 1947 to 1997." Assuming young women are not likely to receive large portions of unearned income, these income data are reasonably accurate measures of female labor income. Dividing by 1750 hours of full time work per year (50 weeks at 35 hours per week of full time work) yields estimates of an hourly wage figure. These constructed wage series closely track those constructed by Macunovich (1995) from the Current Population Survey (CPS). There is one outlier in 1973 for the CPS data for younger women's wages, which does not appear in the income-based data. When this observation is removed, the correlations between the CPS and income based wage series are 0.97 for the 20 to 24 year olds and 0.99 for the older group.

Based on Wolfe's (1980) finding that only education at the higher levels significantly affects fertility, educational attainment is measured as the percentage of women aged 25 to 34 who have completed at least one year of college. A broader measure of education, such as median years of schooling completed, shows little variation over the sample, and therefore would not be related to any of the trends in fertility or labor market variables over this period. Data on years of schooling completed, females 25-34, comes from the internet site of the U.S. Census Bureau, www.census.gov/population/socdemo/education/tablea-01.txt, with data for missing years calculated by linear interpolation.

The key policy variable, the value of the tax exemption per dependent, is computed as the product of the statutory value of the exemption times the average marginal tax rate, in 1997 dollars. The exemption series is from *Facts and Figures on Government Finance* (Tax Foundation, 1948-1989) and tax forms (Internal Revenue Services, 1990-1997), and the average marginal tax rate is taken from Stephenson (1998) for the period 1948-1994 and computed from *Statistics* *of Income* (Internal Revenue Services, 1995-1998). A time plot of this series is displayed in Figure 2.

The value of the tax exemption is substantial and shows considerable variation over the period since 1948. After a peak in 1952 at $839 (in 1997 dollars), its value trended downwards until the late seventies as inflation eroded the real value of the exemption. The reduction in marginal tax rates during the early 1980's led to a further sharp drop in the value of the personal exemption, followed by a subsequent rise in its nominal value and indexation to the rate of inflation in 1986. Tax policy changes in the late seventies and eighties resulted in a decline and then increase of over $200 in the real value of the exemption.

Availability of the birth control pill is represented by a dummy variable that is zero before 1963 and one from 1963 onwards. The child care tax credit is captured by another dummy that equals zero prior to 1976 and one thereafter. Although the U.S. tax code has provided some tax relief for childcare expenses since 1954, until 1976 deductibility of such expenses was limited to lower income groups. In 1976 childcare expenses became deductible for all income groups at a flat rate equal to 20 percent of such costs. In 1981 this was modified to a rate that declines with income. Eligibility is limited to families with both parents, or the single parent, working or in school, and those with children under thirteen years of age. The sliding rates and eligibility requirements make it difficult to determine an average dollar value of the childcare tax credit, so in this study its impact is captured by a dummy variable marking the most important change in the program.

Table 1 presents the results of the unit root tests applied to all stochastic variables. Following Dickey and Pantula (1987) tests for the highest expected number of unit roots (in this case two) are implemented first. The hypothesis of two unit roots is rejected soundly for all series, and the tests of a single unit root indicate that all are integrated of order one. Only female wages for the older group is close to being stationary around a trend, with rejection of the unit root hypothesis for this series at the 10% level but not at the 5% level. Concluding that all variables are I(1), traditional regression methods that assume stationarity are precluded. There is, however, the possibility of cointegration among these variables that would allow further investigation of long run relations between fertility, female labor force participation, and their covariates.

5. Cointegration Models of Fertility and Female Labor Supply

One or more stationary long run relations may connect the six nonstationary variables described in the previous section. In fact, a finding of cointegration among these variables is a minimal condition for support of the underlying theory and an essential prerequisite for subsequent analysis of the long run relations. The absence of cointegration means that there is no stationary functional relation that links these series together, and regressions among such nonstationary variables are therefore spurious.

In the specification of the models for each age group, the tax exemption variable is assumed to be exogenous, while the remaining five nonstationary series are allowed to be endogenous. The lack of conscious pronatalist policy in U.S. tax laws justifies the exogeneity assumption for the tax exemption. Imposing this condition reduces the number of parameters to be estimated, which is still considerable in an error correction model containing five endogenous variables. The models include, in addition, the dummy variables for the availability of the birth control pill and for the adoption of the child care tax credit. Furthermore, the baseline models include a time trend in the cointegrating relations to allow for the differing trend characteristics of the variables in the models. Sensitivity to this particular specification is examined by deletion of the time trend and also by consideration of the model without educational attainment. All stochastic variables enter into the model in logarithmic form; thus their coefficients are interpreted as elasticities.

Each cointegrating relation is tested and estimated within an error correction model that contains sufficient lagged differenced terms to produce nonautocorrelated and normally distributed residuals.[22] For the older group a single lagged difference on all nonstationary variables yields an error correction model that meets the diagnostic criteria. For the younger group it is only necessary to include a single lagged difference on male incomes to meet these criteria. The tests for cointegration and diagnostic statistics are presented in Table 2. For the 20-24 year olds neither the normality assumption nor the absence of autocorrelation can be rejected at the 10 percent level; for the 25-34 age group the test statistics are not significant at the eight- percent level.

[22] Autocorrelation is tested with the multivariate Ljung-Box statistic for lag one autocorrelation and cross correlations in the residuals from every equation in the error correction model. Under the null hypothesis of no serial correlation, this statistic has the chi-squared distribution with n^2 degrees of freedom, where n is the number of equations in the system. Normality is tested with the Jarque-Bera statistic that combines measures of skewness and kurtosis computed for each time series of residuals. For each equation the Jarque-Bera statistic has the chi-squared distribution with 2 degrees of freedom under the null hypothesis of normality. When combined across all n equations, the statistic has 2n degrees of freedom.

Although Johansen (1995) presents critical values for likelihood ratio tests of the number of cointegrating relations, these values are appropriate only for the standard models that contain no exogenous nonstationary variables or deterministic dummy variables. Consequently, the determination of the number of cointegrating relations involves some judgement based on the formal tests, the roots of the complete dynamic system,[23] and the plausibility of the estimated parameters in each relation.

All pieces of evidence consistently affirm that the series in these models are cointegrated. For the 20-24 year olds Johansen's trace tests strongly reject the null hypothesis of no cointegration at the one-percent level using the critical values appropriate for the standard models. At this same stringent significance level, the trace test indicates exactly two cointegrating relations. On the other hand, only two roots of the complete dynamic system are close to one, suggesting three stationary relations. As discussed below, the two cointegrating relations that show the greatest degree of stationarity can reasonably be interpreted as a fertility equation and a female labor supply equation, respectively. Based on this collection of evidence, the analysis proceeds under the conclusion that there are

[23] A system of n endogenous I(1) variables has a maximum of n unit roots, attaining this maximum if there is no cointegration. Each cointegrating relation reduces the number of unit roots in the system by one. Therefore, the existence of n-r roots close to one is evidence of r cointegrating relations.

two stationary relations among the six stochastic variables in the system for this age group.

For the 25-34 year olds the evidence points to a similar conclusion. Comparing the trace statistic to the one-percent critical values appropriate for the standard models, exactly two cointegrating relations emerge. Here three roots of the dynamic system are reasonably close to one, with all other substantially lower, indicating again the existence of two cointegrating relations. Finally, these two relations can be sensibly interpreted as a fertility relation and a female labor supply equation, and the subsequent analysis is based on this interpretation.

Maximum likelihood estimates of the parameters of the cointegrating relations and their standard errors are presented in Table 3.1 for the 20-24 year olds and Table 3.2 for the 25-34 year olds. Each cointegrating equation excludes one variable to exactly identify the system (Johansen and Juselius 1994), and the coefficients of each equation are normalized by a linear transformation that sets one coefficient to unity. Based on the theoretical model, the identifying restrictions are that female labor force participation is excluded from the fertility equation, and fertility is excluded from the labor supply equation.

For the 20-24 year olds all coefficients in the first relation have the signs anticipated for a fertility equation, and all are statistically significant at the five-percent level or less. In particular, fertility is positively related to male incomes with an elasticity of 1.6 and negatively associated with female wages. This elasticity, at -4.5 is very large, especially since educational attainment is controlled in this model. This large elasticity reflects the possibilities of postponement of fertility for these younger women, as well as a decision by some to have fewer children, in response to favorable wages. Educational attainment also has a strong negative effect on fertility, independent of women's wages. With wages controlled in the model, this negative effect suggests that more educated women prefer smaller families.

Of particular interest here is the effect of the tax exemption on fertility, which is statistically significant with a positive elasticity of 0.74. The sign and significance of this parameter holds across several alternative specifications, with and without the trend term and with or without educational attainment. When the education variable is omitted, the tax exemption elasticity rises to 2.0; removing trend from the cointegrating relation has little effect on this point estimate. Computed at mean values, the 0.74 elasticity estimate implies that a $100 increase in the tax exemption would produce nineteen additional babies for every 1000 women in this age group. This may be compared with the range of estimates in Whittington, et al. (1990) of between twelve and twenty-four additional babies for every 1000 women between 15 and 44 year of age.

The labor supply equation estimates are also generally consistent with economic theory. Female wages and educational attainment have positive and significant effects on female labor force participation, with elasticities of 0.98 and 0.32, respectively. Male income has no significant effect on female labor supply, and none of the other specifications considered produced the significant negative effect predicted by the theory. The tax exemption has a small negative effect on female labor supply, reflecting the withdrawal from the labor force by some childbearing women.

For the older group the coefficients in the cointegrating relations have signs consistent with the theory and all are significant at the five- percent level or less. Elasticities in the fertility equation are smaller in absolute value than those for the younger age group. These differences are plausible since opportunities for deferment of childbearing in response to changing economic conditions are reduced for this older group. A rise in women's wages, for example, has a lower proportionate effect on fertility among 25-34 year olds because postponement of childbearing becomes increasingly difficult for older women. Likewise, the tax exemption elasticity at -0.2 is substantially smaller in absolute value than that for the younger adults (-0.74). Based on this estimate a one hundred-dollar increase in the value of the tax exemption would add 3.6 births per thousand women aged 25-34. Educational attainment at the college level still exerts a strong negative effect on fertility in this age group, which is consistent with the hypothesis that higher education shifts preferences towards smaller families.

In the labor supply equation of 25-34 year olds, male income now has a significant negative effect on female labor force participation. All other elasticities, although statistically significant, are smaller in absolute value than those estimated for the younger group. Women's labor force participation responds negatively to the dependent tax exemption, a reflection of the increased fertility response to this policy variable.

In addition to the long run elasticities of the cointegrating relations, the error correction model embodies information on dynamic responses of the endogenous variables to exogenous changes. Table 3.1 and Table 3.2 display estimates of the coefficients on the dummy variables for introduction of the birth control pill (PILL) and for the adoption of the child care tax credit (CCTC). The contraceptive pill has had a significant negative effect on fertility for both younger and older women, with a slightly stronger effect for the 20-24 year olds. The child care tax credit has increased fertility in both age groups, but has a statistically significant (negative) effect on labor supply only for the younger group. As with the other tax variable, this negative response is a reflection of labor supply responses to increased childbearing.

A more complete picture of the dynamic relations among the variables is presented by the impulse response functions in Figures 3 and 4. In each chart the impulse reponse function, represented by the solid line, is surrounded by 95 percent confidence bounds, indicated by dashed lines. Statistically significant responses correspond to confidence bounds that do not contain the horizontal axis. This information is supplemented by verbal summaries from the variance decompositions, with full tables of these results available from the authors. In this analysis the assignment of contemporaneous residual correlations follows the ordering: tax exemption, education, male incomes, female wages, female labor supply, and fertility. This ordering attributes contemporaneous residual correlations to those variables that are earliest in the ordering, an assumption that maximizes the impact of all other variables on fertility. The verbal summary below describes any qualitative differences from reversing the ordering.

For the younger women only the tax exemption variable and educational attainment have statistically significant effects on fertility. The tax exemption has significant positive effects on fertility for three to seven years after an exogenous policy change, and by year seven these shocks account for 26 percent of fertility variations. Negative effects of educational attainment innovations on fertility are significant in years one through six, when these account for an additional twenty-percent of fertility variations. Under the reverse ordering, educational attainment no longer significantly affects fertility, while the tax exemption effects remain significant and powerful.

Female labor supply responds significantly to innovations in every variable in the model in this dynamic analysis. According to the variance decompositions, the greatest proportionate response is to shocks in the tax exemption, which have significant negative effects for two through eight years after the policy change. By year eight these policy shocks account for 35 percent of the variation in female labor supply. The impulse response functions show significant effects of fertility innovations on female labor supply, but the reverse effect is not significant. With the reverse ordering of variables, labor supply still responds significantly to exogenous changes in fertility and the tax exemption, but the other effects lose significance.

The impulse response functions for fertility behavior of the older group show significant negative responses to exogenous shocks in educational attainment and female wages. Both effects are greatest in the third year after the shock, with wages accounting for 27 percent and education for sixteen percent of the variation in fertility at that point. Tax exemption effects on fertility are not statistically significant for this age group in this short run analysis.

For female labor force participation there are marginally significant and positive responses to exogenous changes in female education and women's wages. The most prominent source of short run variation in female labor supply, however, is from fertility shocks, exhibiting the usual negative association for as long as twelve years after the shock. Fertility shocks account for up to forty-one percent of the variation in female labor supply, but as with the younger group there is no evidence of feedback from labor force participation to fertility. Finally, tax exemption effects in the short run analysis of female labor supply are not significant.

Several notable features emerge from this dynamic analysis. First, the strong negative association between fertility and female labor force participation is confirmed for both age groups, with the evidence showing a single direction of causality from fertility variations to labor supply. This result contributes to the debate over the direction of this well-documented linkage surveyed by Lehrer and Nerlove (1986). Second, the importance of changes in the dependent tax exemption for fertility behavior is confirmed for the younger group. The lack of significance for this variable in the short run fertility behavior of the older group is somewhat consistent with the small elasticity on this variable in the long run relation. Third, exogenous changes in women's educational attainment have strong negative effects on fertility in both age groups, as result consistent with the estimates of the long run relations.

6. Summary and Conclusion

This paper has presented an economic model of fertility and labor force participation that incorporates the effects of provisions of the U.S. tax code on the decision to have children. Both the tax exemption for dependents and the child care tax credit are expected to have positive effects on fertility rates, with reverse effects on female labor supply. The impact of these tax policies on fertility and female labor force participation rates is investigated with time series data for the U.S. spanning the years 1948 to 1997. Two-equation cointegration models are constructed for age-specific fertility rates and for women's labor force participation rates of women 20-24 and 25-34 year of age. Based on the theoretical model, these equations include women's wages, male incomes, female educational attainment, the value of the tax exemption per dependent, and dummy variables for the adoption of the child care tax credit and the launch of the birth control pill.

Statistical evidence indicates the existence of two cointegrating equations for each age group, and these equations are identified as a fertility equation and a labor supply equation, respectively. The estimated coefficients in these cointegrating equations have the signs predicted by the underlying theory, and are statistically significant in every case except one. By capturing the long run relations among these variables, the cointegration framework resolves an anomaly in previous investigations of tax policy effects on fertility. While prior studies by Whittington et al. (1990), Whittington (1992), and Georgellis and Wall (1992) found statistically significant effects of U.S. tax policies on fertility, coefficients on the other key economic variables such as female wages and household incomes

insignificant. The paradox in these results is that tax incentives are expected to work through changes in household incomes or the opportunity cost of children, but the direct measures of female wages and household incomes show no significant impact on fertility.

In this investigation the dependent tax exemption shows its strongest long run impact on the fertility behavior of the younger age group. For this group the estimated elasticity of fertility with respect to the value of the tax exemption is

When evaluated at the mean values of these two variables, this estimate implies that a one hundred-dollar increase in the value of the tax exemption per dependent would cause an increase of nineteen births per thousand women aged 20-24. This variable also contributes substantially to the short run fluctuations in fertility of 20-24 year olds. The dummy variable indexing the adoption of the childcare tax credit also has significant positive effects on fertility rates.

Although these provisions of the United States tax code have not been motivated by pronatalist policies, these tax benefits do show consistent and significant effects on fertility across a range of studies for the U.S. and Canada. The magnitude of the tax exemption effect is small, in part because the value of the tax exemption per child is small relative to typical annual costs of children. However, the finding of a significant effect confirms that households do respond to financial incentives in the decision to have children, so that policies that substantially alter the effective costs of children can have important consequences for aggregate fertility rates.

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Table 1. Dickey-Fuller Tests for Unit Roots

Variables	Null Hypothesis I[2]	Null Hypothesis I[1]
Real wage of women aged 20-24	-7.12 (0)	-2.23 (0,t)
Real wage of women aged 25-34	-3.23 (1,4)	-3.41 (1-4,t)
Real income of males aged 20-24	-5.98 (1)	-1.69 (1,2)
Real income of males aged 25-24	-3.61 (2)	-1.93 (1,t)
Labor force participation of women aged 20-24	-4.50 (0)	-1.33 (1,t)
Labor force participation of women aged 25-34	-3.25 (0)	-2.20 (1,2,t)
Fertility rate of women aged 20-24	-3.02 (0)	-0.98 (1)
Fertility rate of women aged 25-34	-3.17 (0)	-1.19 (1)
Female education	-4.25 (0)	-2.17 (1,t)
Real value of the tax exemption	-6.26 (0)	-1.95 (0)

Notes: Number of lagged first and second differences for the I[1] and I[2] model, respectively, are shown in parentheses, with t if deterministic trend is included in the test equation. Five and ten percent critical values are –2.93 and –2.60 with no trend, and –3.50 and –3.18 when the trend term is included.

Table 2. Tests of cointegration among fertility and labor market variables

Model for the 20-24 age group

H_0: r=	λ-max	1% c.v.	Trace	1% c.v.
0	46.76	25.16	135.64	95.38
1	40.36	21.63	88.88	70.22
2	25.45	17.94	48.52	48.59
3	16.64	14.26	23.07	30.65
4	6.43	16.39	6.43	16.39

Residual diagnostics

Autocorrelation: $\chi^2_{[25]}$ (p-value)	29.62 (0.24)				
Error correction equation for:	Fertility	Labor supply	Female wage	Male income	Female edu.
Normality: $\chi^2_{[2]}$ (p-value)	5.07 (0.08)	3.93 (0.14)	0.72 (0.70)	2.17 (0.34)	0.73 (0.69)
ARCH: $\chi^2_{[3]}$ (p-value)	9.81 (0.02)	0.06 (0.99)	0.40 (0.94)	0.94 (0.82)	0.41 (0.94)

Roots of the autoregressive system

0.8753 0.8753 0.6188 0.6188 0.1950

Model for the 25-34 age group

H_0: r=	λ-max	1% c.v.	Trace	1% c.v.
0	49.31	25.16	130.88	95.38
	36.28	21.63	81.58	70.22
2	28.07	17.94	45.30	48.59
3	15.07	14.26	17.22	30.65
4	2.15	16.39	2.15	16.39

Residual diagnostics

Autocorrelation: $\chi^2_{[25]}$ (p-value)	35.46 (0.08)				
Error correction equation for:	Fertility	Labor supply	Female wage	Male income	Female edu.
Normality: $\chi^2_{[2]}$ (p-value)	0.15 (0.93)	8.52 (0.01)	0.99 (0.61)	1.48 (0.48)	3.95 (0.14)
ARCH: $\chi^2_{[3]}$ (p-value)	1.81 (0.61)	5.65 (0.13)	1.62 (0.65)	0.76 (0.86)	1.51 (0.68)

Roots of the autoregressive system

0.9471 0.9471 0.9069 0.7358 0.7358 0.5916 0.5916 0.3214 0.3214 0.1855

Notes: The model for the 20-24 age group is based on a two-lag structure, whereas the model for the 25-34 age group is based on a one-lag structure. ARCH tests for conditional autoregressive heteroscedasticity.

Table 3.1. Maximum likelihood estimates of the long-run and short-run coefficients of the Error Correction Model (20-24 age group)

	Fertility	Labor supply	Female wage	Male income	Female edu.	Tax exemp.	Trend
Fertility equation	-1.000		-4.454	1.599	-1.819	0.740	0.039
			(0.519)	(0.368)	(0.436)	(0.215)	(0.011)
Adjustment to disequilibria	0.079	-0.004	-0.166	0.177	-0.183		
in the fertility equation	(0.054)	(0.021)	(0.061)	(0.086)	(0.033)		
Labor supply equation			0.982	-0.005	0.319	-0.372	-0.001
			(0.164)	(0.117)	(0.138)	(0.068)	(0.004)
Adjustment to disequilibria	0.268	-0.176	-0.036	0.394	-0.662		
in the labor supply equation	(0.165)	(0.062)	(0.187)	(0.261)	(0.100)		
Coefficients on Birth	-0.058	-0.010	0.083	0.017	-0.038		
Control Pill	(0.019)	(0.007)	(0.022)	(0.031)	(0.012)		
Coefficients on Childcare	0.046	-0.020	0.028	-0.065	0.022		
Tax Credit	(0.016)	(0.006)	(0.018)	(0.029)	(0.009)		

Notes: Within each of the first two panels the top two lines present the long-run parameter estimates and their standard errors (in parentheses) respectively, and the second lines present the estimated adjustment coefficients and their standard errors. Short-run coefficients on the two dummy variables and their standard errors are reported in the third panel.

Table 3.2. Maximum likelihood estimates of the long-run and short-run coefficients of the Error Correction Model (25-34 age group)

	Fertility	Labor supply	Female wage	Male income	Female edu.	Tax exemp.	Trend
Fertility equation	-1.000		-0.793	0.099	-0.793	0.202	0.027
			(0.062)	(0.041)	(0.083)	(0.041)	(0.002)
Adjustment to disequilibria	-0.389	-0.228	-0.303	0.296	-0.395		
in the fertility equation	(0.103)	(0.068)	(0.180)	(0.103)	(0.108)		
Labor supply equation		-1.000	0.315	-0.336	0.265	-0.282	0.005
			(0.074)	(0.048)	(0.098)	(0.049)	(0.002)
Adjustment to disequilibria	-0.135	-0.118	-0.423	-0.486	-0.038		
in the labor supply equation	(0.099)	(0.065)	(0.173)	(0.098)	(0.104)		
Coefficients on Birth	-0.048	0.001	0.040	0.003	-0.024		
Control Pill	(0.010)	(0.011)	(0.018)	(0.011)	(0.011)		
Coefficients on Childcare	0.051	0.008	0.003	0.019	0.007		
Tax Credit	(0.010)	(0.007)	(0.018)	(0.011)	(0.011)		

Notes: Within each of the first two panels the top two lines present the long-run parameter estimates and their standard errors (in parentheses) respectively, and the second lines present the estimated adjustment coefficients and their standard errors. Short-run coefficients on the two dummy variables and their standard errors are reported in the third panel.

Figure 1. Fertility rates for women 20-24 and 25-34: 1948-1997

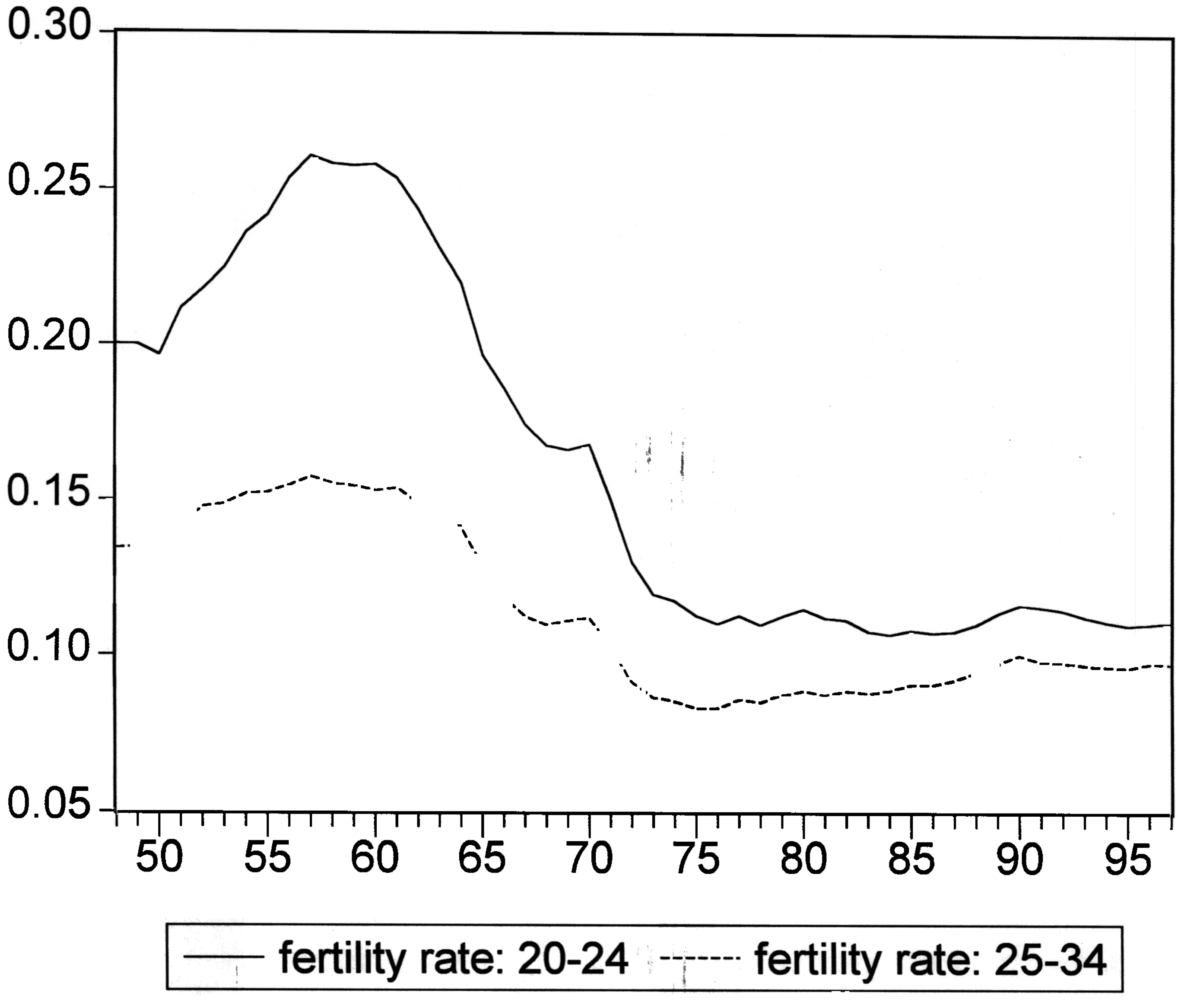


Figure 2. Value of tax exemption per dependent.

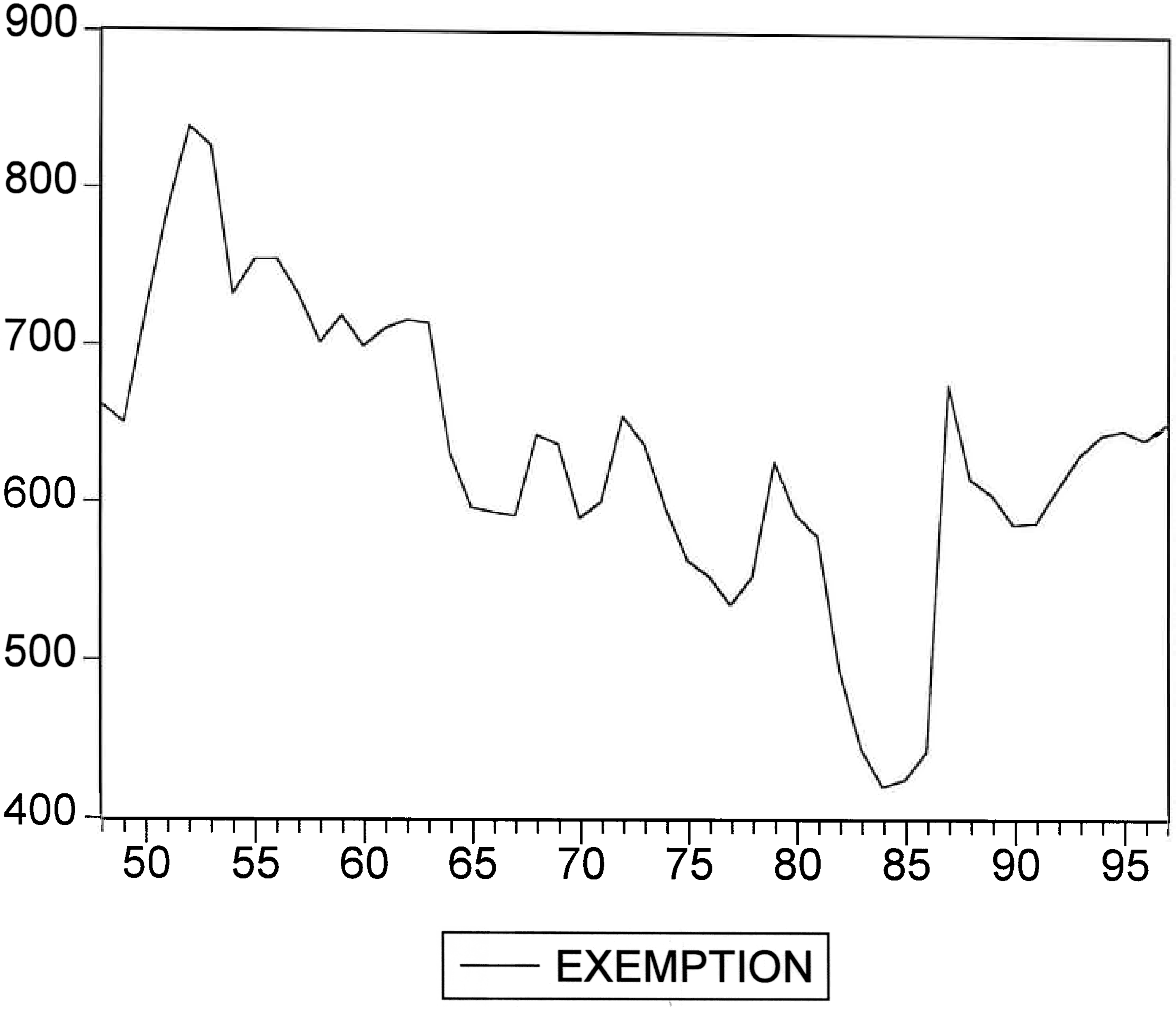
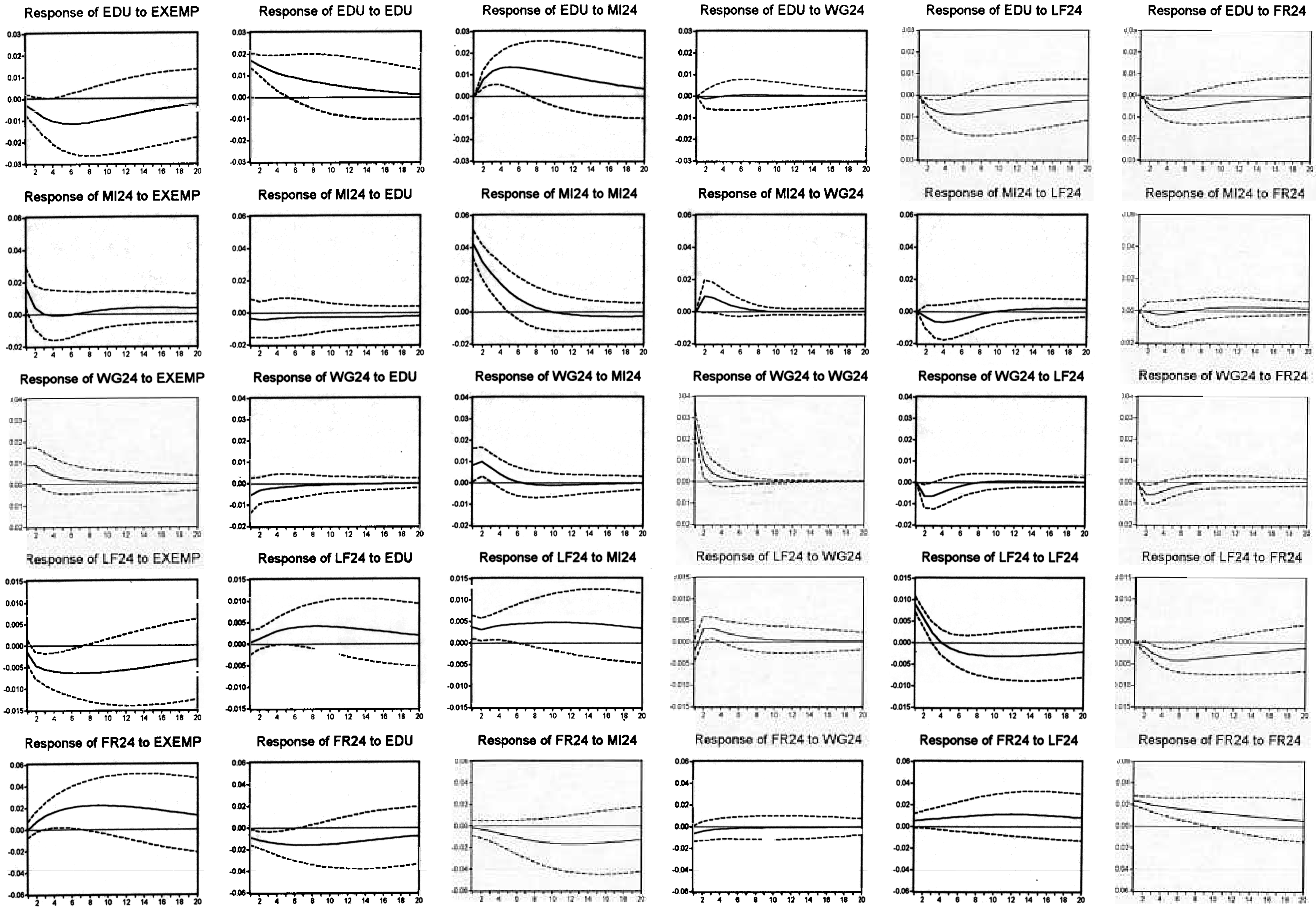
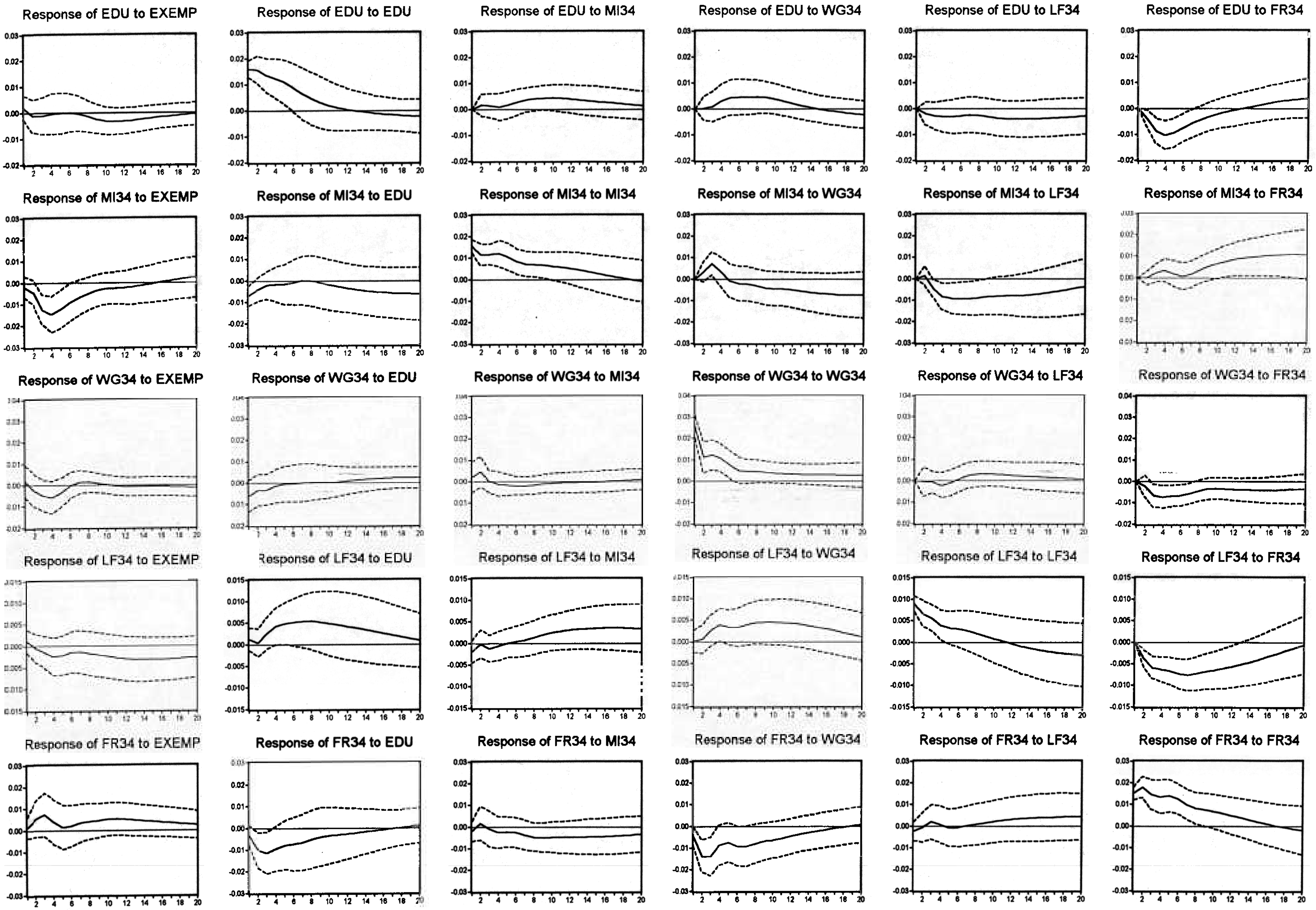


Figure 3. Impulse response functions: 20-24 age group.



EXEMP = tax exemption, EDU = education, MI = male income, WG = female wage, LF = female labor force participation, and FR = fertility rate. Measures specific to this age group are indicated by 24.

Figure 4. Impulse response functions: 25-34 age group.



EXEMP = tax exemption, EDU = education, MI = male income, WG = female wage, LF = female labor force participation, and FR = fertility rate. Measures specific to this age group are indicated by 34.