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Three essays in applied time series analysis

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THREE ESSAYS IN APPLIED TIME SERIES ANALYSIS

by

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A dissertation submitted to the Graduate Faculty in Economics in partial fulfillment of the requirements for the degree of Doctor of Philosophy, The City University of New York

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Abstract

THREE ESSAYS IN APPLIED TIME SERIES ANALYSIS

by
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Adviser: Professor Michael Grossman

This thesis consists of three essays which use time series techniques to analyze the relationships between macroeconomic, and demographic variables. The first essay, using Vector-Autoregressions, demonstrates that an increase in the proportion of young marriages does not increase the divorce rate if the model includes labor force participation and fertility. On the other hand, in a model, which omits labor force participation and fertility, a shock to the proportion of young marriages increases the divorce rate, implying that the often reported inverse relationship between age at marriage and divorce is not causal. Rather, it is an artifact of underlying relationships between labor force participation and fertility.

The second paper sheds light onto the conflicting findings of previous research on real wage cyclicality. Using Vector-Autoregressions between output, money supply, unemployment, real wages, relative price of oil, and productivity, it is shown that the shocks in aggregate

supply produce procyclical real wages, whether the nominal wage is deflated by the wholesale price index or the consumer price index. On the other hand, shocks in aggregate demand generate countercyclical behavior of real wages, when the real wage is measured as the ratio of nominal wage rate to the wholesale price index. However, if the nominal wage is deflated by the consumer price index, aggregate demand shocks yield movement in the same direction of output and real wages.

The third essay, using an Interrupted Time Series Analysis, estimates the change in adolescent childbearing that followed the liberalization of the New York State abortion law in 1970. It is found that the level of births to black adolescents fell 18.7 percent between 1970 and 1971, or approximately 142 fewer births per month. The level of white births fell 14.1 percent, or approximately 111 fewer births per month. The forecasts yield that if abortion were banned January 1, 1989, there would have been 2143 Black births, and 1067 White births to NYC adolescents in 1988 and 1989 above what would have been expected had the law remained unchanged. The total marginal cost supporting the births to the mothers with AFDC eligibility would be 11.5 million dollars.

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Last, but not least, I have to thank Murat Sertel, who had first encouraged me to pursue the doctoral studies in the United States.

To the memory of my father, and to my mother.

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INTRODUCTION

A Time Series is a collection of observations that are assumed to be generated sequentially by a stochastic process through time. The Analysis of a Time Series involves choosing a model for that process, and estimating it using statistical tools. Time series models, which utilize the past values of a particular variable only, [typically Autoregressive Integrated Moving Average (ARIMA) models], or the ones which involve more than one variable [e.g. Vector-Autoregressions (VAR)] have proved successful, and in many cases provided forecasts that are more accurate to projections from large-scale econometric models.

Time Series Analysis traditionally has been applied to macroeconomic variables. Along the same lines, the second paper of this dissertation investigates the relationships between production, real wages and unemployment. In particular, using a VAR, the paper demonstrates that the cyclicality of real wages depends on whether there is a demand or supply disturbance in the system, and whether one uses consumer or producer prices as the deflator for nominal wages. This outcome explains the conflicting findings of previous research, which employed different measures of the real wage rate, and used different sample periods which may have been dominated by different types of shocks.

The first and third papers of the thesis are the applications of the VAR and ARIMA techniques to the population and health economics fields. The first paper analyzes the dynamic interactions between age at marriage, divorce, fertility, and labor force participation of women using a VAR. It brings a comprehensive perspective to bear on a topic which had been exhausted by conventional analyses. Previous studies, based on cross-sections of micro data, uniformly concluded that the probability of divorce and age at marriage were negatively related. This paper demonstrates that this apparent relationship is not causal. Instead, it is an artifact of underlying relationships between labor force participation and fertility.

There is great speculation in the popular press that the U.S. Supreme Court will overturn the 1973 decision in Roe versus Wade, the case which legalized abortion across the United States. The objective of the third paper is to examine the probable changes in teenage chidbearing among New York City residents following a ban on legalized abortion. To this end, we fit an ARIMA model with the Intervention Component that estimates the change in the number of births to adolescents following the 1970 New

York State law which liberalized abortion. It is found that the level of births to black adolescents fell approximately 142 fewer births per month between 1970 and 1971. The level of white births fell approximately 111 fewer births per month. Projections based on the fitted model suggest that a ban on legalized abortion today would have a major impact on adolescent childbearing in New York City as well as other parts of the country, although the magnitude of the change would vary according to local conditions.

AGE AT MARRIAGE, DIVORCE, FERTILITY, AND LABOR FORCE PARTICIPATION OF WOMEN: A TIME SERIES PERSPECTIVE

I. INTRODUCTION

This paper aims to synthesize two close relationships which have previously been investigated separately; namely, the relation between age at marriage and probability of divorce on the one hand, and the relation between labor force participation and fertility of women on the other.

The mutual determination of fertility and labor force participation has long been emphasized. Traditionally, a rise in the labor force participation rate, say, due to an increase in the real wage rates, is expected to increase the "price" of a child, hence decrease fertility. Similarly, an increase in fertility is expected to discourage labor force participation of women. However, the findings of empirical research have not been conclusive, and in some cases have not fulfilled the theoretical expectations. Cain and Dooley (1976) and Link and Settle (1981) find that fertility depends on labor supply but not vice versa; whereas Hoffman (1985) reports the opposite relationship. Shultz (1978) demonstrates the negative effect of fertility on the labor force participation of married women, while Fleisher and Rhodes

(1980) find no association between the number of children and the mother's life-time labor supply. Gregory (1982), Dooley (1982), Joerding (1982), and Michael (1985) report a positive coefficient of the labor force participation variable in the fertility equation.

Becker (1973,1974) developed the theory on love, marriage, and dissolution which predicts that the probability of dissolution is lowered by an increase in the earnings of the men, an increase in the age at marriage, the number of young children, or by an increase in the duration of marriage. Becker et al. (1977) tested the predictions of the theory by considering the effects on the probability of divorce for the age at marriage, schooling, income, and children, and found an inverse relation between age at marriage and the probability of divorce. However, since they include age at marriage among the set of exogenous variables rather than treating it endogenously, a problem of potential bias arises in their estimation.[1] (There has been theoretical and empirical

[1] One can also find numerous studies in the sociology literature consistently pointing out an inverse relation between age at first marriage and the probability of divorce, where age at marriage treated exogenously [Burchinal and Chancellor (1963), Burchinal (1965), Bauman (1967), Parke and Glick (1967), Carter and Glick (1970), Glick and Norton (1971), Bumpass and Sweet (1972), Weed (1974), Shoen (1975), Lee (1977), Booth and Edwards (1985)].

work pertaining to the determinants of age at marriage. See for instance Keeley (1977, 1979), where household production theory and search theory are incorporated to explain the incentives to marry and the determinants of age at first marriage.)

Age at marriage, divorce, labor force participation and fertility constitute a general equilibrium system, where none of the variables can be considered exogenous. Isolating bivariate relationships between age at marriage and divorce, or fertility and labor force participation from this system clearly generates a simultaneous equation bias. Moreover, in a static model one assumes that all interaction among variables occur simultaneously. There is reason, however, to believe that the variables of this system can have lagged effects on each other. Therefore an accurate representation of the system should be the one that takes into account the dynamic nature of the interdependence of the variables. Consequently, an estimation technique which does not put a priori restriction on the model, and captures the dynamics of the variables is needed. In these regards, the vector-autoregression (VAR) technique which is employed in this study is superior to all other estimation methodologies.

II. THE FRAMEWORK

This paper is the first to attempt to investigate the dynamic interrelations between age at marriage, divorce, fertility and labor force participation of women in a multivariate context using vector-autoregressions. Due to the nature of the technique employed, time series data with a large number of observations was needed. Consequently, this study employed monthly U.S. data which covered the time period 1963-1982.[a]

The application of a VAR enables us to consider labor force participation, fertility, divorce and age at marriage as endogenous variables, and to observe the dynamic influence of each of them on the others. A VAR can be interpreted as the reduced form relationship that arises from a dynamic stochastic structural model, the underlying structural parameters of which are based on the utility functions and the constraints. In this case, the use of a VAR does not suggest that the dynamic theory can not identify the underlying structural system, and therefore should not be considered a substitute for the estimation of the structural model. Rather, through the application of a VAR, we will gain insight into the

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[a] The data are obtained from various issues of National Center for Health Statistics Monthly Vital Statistics Report, Vital Statistics of the U.S., Department of Health, Education and Welfare, Statistical Abstract of the United States and the Citibase.

relevance of many different hypothesized relationships between the variables, and hence it is a complement to building structural models.

In a VAR system, each variable is regressed on its own lagged values, lagged values of the other endogenous variables as well as lagged values of the relevant exogenous variables. Because the right-hand side variables in each equation consist of past values, they are all predetermined. Thus, the system can be consistently estimated using OLS, without being concerned about the existence of simultaneous equation bias. Furthermore, estimating each equation separately using OLS produces asymptotically efficient estimates, because the right-hand side variables are the same in every equation (Hakkio and Morris 1984).

IIA. ESTIMATION

Fertility (FERT) is measured as the number of births per 1000 women between the ages of 16 and 44. The variable LFPR stands for the labor force participation rate for women who are 16 years of age and over. YOUNGMAR measures the proportion of brides between ages 16 and 24 in their first marriages to the number of women in the same age group and shows the tendency to marry early. The divorce rate (DIVORCE) is measured as the number of divorces per

1000 married women who are 18 years of age and over. The ratio of males to females in population (SEXRATIO) is a proxy for the available alternatives to women. This fifth and the only exogenous variable is expected to have an influence on the timing of marriage as well as on divorce, and therefore on the whole system.

To ensure stationarity, all variables are expressed in logarithms and linear and quadratic trend variables are included in each equation. Since the variables are not seasonally adjusted, each equation contains eleven dummy variables to account for monthly variation.
The estimated system has the following form:

$$(1) \qquad Y_{i,t} = C_i + \sum_{k=1}^{13} \alpha_{i,k} D_{i,k,t} + \sum_{j=1}^{5} \sum_{s=1}^{12} \beta_{i,j,s} Y_{j,t-s} + \epsilon_{i,t}$$

where C_i represents the constant, and $D_{(.)}$ represents eleven seasonal dummies and the two trend terms in each equation. $Y_{(.)}$ stands for the endogenous variables, $\alpha_{(.)}$ and $\beta_{(.)}$ are the coefficients. Since twelve lags of SEXRATIO are included in each equation as exogenous variable, j goes from 1 to 5.

Table 1 reports the coefficients for vector-autoregressions, except for the constant, trends, and seasonal dummy variables. The individual coefficients are relatively uninformative due to the multicollinearity in such a heavily parameterized model (Corman and Joyce 1988). Consequently, we focus on the F-statistics for the set of lags in each equation. F_x stands for the F-statistics for the hypothesis that all lags of the variable x have zero coefficients in the corresponding equation.

According to Table 1, no variable appears to be close to a random walk. The variable which is affected mostly by the others is the proportion of young marriages. While all other variables are influenced by their own lags, the structure of the autoregressive behavior in the fertility equation is not well determined. As is evident, the F-statistics of FERT, YOUNGMAR and DIVORCE in LFPR equation are not significant, implying that the lagged values of these variables have no predictive power for the current labor force participation rate. In other words, the labor force participation rate is not caused by other variables in the Granger-causality sense.*

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* The Granger-causality is defined as follows: A stationary stochastic time series X causes another stationary stochastic time series Y within a set of information in the universe, if the current value of Y is more accurately predicted by using the information that includes at least the own-past series of Y and the past series of X than by using the information that excludes the past series of X (Granger 1969).

TABLE 1

ESTIMATED VECTOR AUTOREGRESSIONS

INDEPENDENT VARIABLE	LAG	LFPR	FERT	YOUNGMAR	DIVORCE
LFPR	1	0.625[a]	0.158	-1.071[b]	-0.395
LFPR	2	0.212[a]	0.121	-0.194	0.434
LFPR	3	0.028	-0.176	1.441[b]	-0.298
LFPR	4	0.072	0.114	0.568	0.570
LFPR	5	-0.211[a]	-0.129	0.767	-1.194[b]
LFPR	6	0.156[b]	0.234	-0.017	-0.644
LFPR	7	0.062	-0.292[b]	-1.891[b]	1.000[b]
LFPR	8	-0.177[b]	0.417[b]	1.551[b]	0.187
LFPR	9	0.079	0.161	0.147	-0.000
LFPR	10	0.134[b]	-0.360[b]	-0.417	-1.257[b]
LFPR	11	-0.017	-0.025	-0.265	0.678
LFPR	12	-0.071	0.233[b]	0.194	0.264
FERT	1	-0.001	0.799[a]	0.527[b]	0.458[b]
FERT	2	-0.018	0.028	0.377	-0.158
FERT	3	0.073[b]	0.067	-0.731[b]	0.183
FERT	4	-0.090[a]	-0.016	0.262	0.258
FERT	5	0.008	0.016	-1.179[a]	-0.261
FERT	6	-0.006	-0.029	0.647[b]	-0.677[b]
FERT	7	0.016	0.045	0.467[b]	0.225
FERT	8	0.041[b]	0.051	-0.348	-0.162
FERT	9	-0.013	-0.011	-0.002	0.420[b]
FERT	10	-0.057[b]	-0.023	0.560[b]	-0.155
FERT	11	0.029	-0.075	-0.921[a]	0.290
FERT	12	0.006	0.066	1.105[a]	-0.140
YOUNGMAR	1	-0.003	0.006	-0.004	0.070[b]
YOUNGMAR	2	0.003	-0.001	0.000	0.071[b]
YOUNGMAR	3	0.005	-0.012	0.033	-0.045[b]
YOUNGMAR	4	-0.005[b]	0.028[a]	-0.133[a]	-0.043
YOUNGMAR	5	0.000	-0.012	0.082[b]	0.063[b]
YOUNGMAR	6	-0.003	-0.006	-0.143[a]	-0.030
YOUNGMAR	7	-0.006[b]	-0.007	-0.036	-0.002
YOUNGMAR	8	0.002	0.042[a]	-0.071[b]	-0.123[a]
YOUNGMAR	9	0.002	-0.020[b]	0.079[b]	-0.038
YOUNGMAR	10	0.001	0.023[b]	-0.196[a]	0.052[b]
YOUNGMAR	11	-0.005[b]	-0.014[b]	0.180[a]	0.021
YOUNGMAR	12	0.005	-0.013	0.579[a]	-0.039
DIVORCE	1	0.014[b]	-0.058[a]	0.162[b]	0.136[b]
DIVORCE	2	-0.014[b]	0.003	0.094	0.274[a]
DIVORCE	3	-0.000	0.011	0.035	0.280[a]
DIVORCE	4	0.010[b]	0.027[b]	-0.115[b]	0.099[b]
DIVORCE	5	0.002	0.009	0.094	-0.073

TABLE 1 (concluded)

DIVORCE	6	0.005	-0.000	-0.205[b]	0.091[b]
DIVORCE	7	-0.017[b]	-0.027[b]	0.112[b]	-0.049
DIVORCE	8	-0.012[b]	0.022	-0.062	0.032
DIVORCE	9	0.010[b]	0.018	-0.022	-0.063
DIVORCE	10	-0.001	0.019	-0.219[a]	-0.094[b]
DIVORCE	11	0.004	0.013	-0.118[b]	0.108[b]
DIVORCE	12	-0.010[b]	-0.043[b]	0.147[b]	0.135[b]
SEXRATIO	1	0.861[a]	1.177[b]	-7.720[b]	0.058
SEXRATIO	2	-0.178	-0.989	9.384[b]	-3.695
SEXRATIO	3	-0.638[b]	1.477[b]	-2.589	8.921[b]
SEXRATIO	4	0.010	1.357	-1.128	-5.939[b]
SEXRATIO	5	0.565[b]	-2.628[b]	0.924	-4.947[b]
SEXRATIO	6	-0.328	1.543[b]	-12.428[a]	7.967[b]
SEXRATIO	7	-0.846[b]	-2.050[b]	7.189[b]	-4.425
SEXRATIO	8	0.586[b]	-0.106	-3.715	5.575[b]
SEXRATIO	9	-0.357	1.155	9.814[b]	-4.773[b]
SEXRATIO	10	0.457	-0.204	-14.900[a]	-0.192
SEXRATIO	11	0.065	0.528	-3.058	7.611[b]
SEXRATIO	12	0.074	-0.914	5.952[b]	-7.100[b]
R^2		.998	.996	.981	.994
F_{LFPR}		28.01	1.33	.97	.75
		(.000)	(.207)	(.479)	(.702)
F_{FERT}		.96	44.75	3.74	1.49
		(.492)	(.000)	(.000)	(.135)
F_{YM}		.63	1.92	16.59	1.98
		(.810)	(.036)	(.000)	(.029)
F_{DIV}		1.15	1.15	1.88	16.81
		(.324)	(.326)	(.040)	(.000)

The numbers in the parentheses are the marginal significance levels. a: $|t| \geq 2$ b: $1 < |t| < 2$

The labor force participation rate is exogenous with respect to fertility, the divorce rate, and the proportion of young marriages.

This outcome, although surprising, is consistent with the finding of Michael (1985), who reports that "...the short-run shocks in labor force participation rate appear to be temporally prior to the shocks in most of the other demographic and economic time series investigated." Along the same lines, one observes that there is strong interaction among fertility, the proportion of young marriages and the divorce rate. These variables are influenced by the past values of each other.

IIB. SPECIFICATION TESTS

To check the specification of the system, several test were performed. First, to see if the lag specification adequately captures the dynamics of the model, the estimated system with 12 lags of each variable was tested as a restriction on systems with 15 and 17 lags. The null hypotheses of no difference was accepted in both cases. The Chi-square was 81.97 with 60 degrees of freedom in the former, and 128.3 with 100 degrees of freedom in the latter. [This is the modified Chi-square proposed by Sims (1980, footnote 18)]. Secondly, a Sims exogeneity test (Sims 1972), coupled with the a priori knowledge as to the exogeneity of SEXRATIO, enabled us to perform a specification test of the model. For illustrative purposes suppose the following simple model represents the correct specification:

$$X_t=\alpha_1X_{t-1}+\alpha_2X_{t-2}+\alpha_3Y_{t-1}+\alpha_4Y_{t-2}+\alpha_5E_{t-1}+\alpha_6E_{t-2}+\epsilon_t$$

$$Y_t=\beta_1X_{t-1}+\beta_2X_{t-2}+\beta_3Y_{t-1}+\beta_4Y_{t-2}+\beta_5E_{t-1}+\beta_6E_{t-2}+u_t$$

$$E_{t-1}=\tau E_{t-2}+\mu_t$$

where E is the exogenous variable. If one estimates the model

$$X_t=\delta_1X_{t-1}+\delta_2X_{t-2}+\delta_3E_{t-1}+\delta_4E_{t-2}+\delta_5E_{t+1}+v_t,$$

omitting the lagged values of Y, but including the future value of the exogenous variable, the chain of correlation between E_{t+1} and Y_t may give rise to a significant estimate of $\delta_5\neq 0$, which represents the specification error of the model, since the future values of the exogenous variable is not expected to influence the current value of the endogenous variable.

The first row of Table 2 reports the F-statistics for the 12 leads of SEXRATIO in each of the four equations of the system. They are not significantly different from zero, implying the rejection of the hypothesis that there is a specification error of the model.

Lastly, a Lagrange Multiplier (LM) test was applied to the residuals. The LM statistic provides a general test for autocorrelation of errors, and is valid when the set of regressors includes lagged dependent variables (Breush 1978, Godfrey 1978). The test were based on additional regressions, in which each dependent variable is regressed on the same set of variables plus the set of lagged residuals. One rejects the null hypothesis of no autocorrelation (white noise errors) if the set of lagged residuals are different from zero. As Breush and Godfrey (1981) show, the reported F-statistic is asymptotically equivalent to the usual LM statistic.

The last five rows of Table 2 show the F-statistics for the coefficient of lagged errors in each equation for 1, 2, 4, 6 and 12 lags of the errors. As is evident, the F values are small, leading us to the acceptance of the hypothesis that the errors of the estimated system are white noise, hence the system is well specified.

TABLE 2

SPECIFICATION TEST

Test Statistic	Equation LFPR	FERT	YOUNGMAR	DIVORCE
F-stat for twelve leads of SEXRATIO	1.323 (.213)	.688 (.760)	.962 (.488)	1.778 (.581)
F-stat for 1 error lag	3.303 (.071)	3.284 (.072)	.442 (.507)	1.692 (.195)
F-stat for 2 error lags	1.300 (.276)	2.058 (.131)	.311 (.733)	1.950 (.146)
F-stat for 4 error lags	1.264 (.289)	1.572 (.185)	.690 (.600)	1.679 (.158)
F-stat for 6 error lags	1.303 (.259)	1.105 (.363)	.575 (.750)	1.461 (.196)
F-stat for 12 error lags	1.132 (.340)	1.791 (.056)	1.559 (.112)	.766 (.684)

The numbers in the parentheses are the marginal significance levels.

TABLE 3

DECOMPOSITION OF VARIANCE, 60 MONTHS AHEAD

Innovations in	Shock to LFPR	FERT	YOUNGMAR	DIVORCE
LFPR	79	42	14	11
FERT	15	49	15	23
YOUNGMAR	1	1	61	6
DIVORCE	6	8	10	61

Note: The column entries may not add up to 100 due to rounding.

IIC. IMPULSE RESPONSE FUNCTIONS

A very useful way to characterize the dynamic relationships among the variables is to map out their responses to unanticipated shocks in one of the variables. To that end, after estimating the model in (1), it is converted into a linear combination of the innovations. The moving average representation is of the following form:

$$Y_{i,t} = \sum_{j=1}^{4} \sum_{s=0}^{\infty} \theta_{i,j,s}\, u_{j,t-s}; \qquad i=1,\ldots,4,$$

where $u_{(.)}$ stands for the error term and $\theta_{(.)}$ represents the moving average coefficients.

As long as the errors of the variables are contemporaneously correlated (the off-diagonal elements of the variance-covariance matrix is nonzero), changes in errors occur simultaneously, hence a change in a variable can not be attributed to the innovation in that variable alone. Therefore an orthogonalization of the errors is needed, which is obtained by triangularization of the variance-covariance matrix of residuals, to create a block recursive system among errors.[4] The transformation imposes a causal ordering among the contemporaneous errors. The impulse response functions presented below is based upon

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[4] For details, see Gordon and King(1982), Sims(1980), and Litterman(1979).

the following causal ordering: LFPR, FERT, YOUNGMAR, DIVORCE.[a]

Table 3 reports the decomposition of variances which is calculated along with the impulse responses. Each entry stands for the proportion of the forecast error of a variable, generated by innovations in other variables 60 months after the shock. The variance decomposition is a way of gauging the interaction between the variables [Sims (1980,1981)]. If the variance of a variable is attributable mostly to its own innovations, then that variable appears to be relatively exogenous in the system. In particular, a strictly exogenous variable would have a 100 in its diagonal cell, and zeroes in the other cells of its column.

According to Table 3, fertility is the most endogenous variable, with only 49 percent of its forecast

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[a] The imposition of a causal ordering among the contemporaneous innovations only matters when their correlations are substantial. Consider Fisher's z-transformation. $z=\ln[(1+r)/(1-r)]$, where r is the sample correlation. The statistic $n^{1/2}z$ is distributed approximately N(0,1) in large samples, where n is the degrees of freedom for the least squares residuals. In our case with n=190, a correlation must exceed .07 in absolute value to be significantly different from zero. This was only the case with the correlations between YOUNGMAR and FERT, and YOUNGMAR and DIVORCE, where the correlations were -.17 and -.09 respectively. The alternative causal orderings for these variables did not change the results in any significant way; neither quantitatively, nor qualitatively. In particular, no ordering did generate different results from the one without a causal ordering.

error variance explained by its own innovations. Labor force participation innovations have almost the same explanatory power on the forecast error variance of fertility as fertility innovations; LFPR errors explain 42 percent of the variance in FERT. The labor force participation rate accounts for 79 percent of its own forecast error variance, and once again appears to be relatively exogenous in the system. One also observes that the innovations in the proportion of young marriages have no predictive power of explaining the forecast errors of other variables. In particular, the innovations in the proportion of young marriages account only for one percent of the forecast error variance of the labor force participation rate and fertility, and six percent of the divorce rate.

The relation between Table 1 and Table 3 is of interest. Table 3 indicates that the innovations in the proportion of young marriages explain 6 percent of the variance in the divorce rate, and divorce explains 10 percent of the variance in the proportion of young marriages. Similarly, in Table 1 the F-statistic of YOUNGMAR in DIVORCE equation is very close to the F-statistic of DIVORCE in YOUNGMAR equation. This is the evidence of bidirectional causalities between age at marriage and divorce.* In Table 3 one can see that the labor force participation rate explains only 11 percent of the variance of the divorce rate, and 14 percent of the proportion of young marriages; however it accounts for 42 percent of the forecast error variance of fertility. Table 1 too, demonstrates that the F-statistic for LFPR is .75 in DIVORCE equation, and .97 in YOUNGMAR equation, whereas the F-statistic in FERT equation is 1.33.

To make the impulse responses comparable, the response of each variable is expressed as a proportion of its standard deviation. The first characteristic of the system is its stability. The figures indicate that the responses to shocks in the system tend to dampen, with variables converging to the initial values within five years.

Figures 1 and 2 present responses to a shock in the labor force participation rate. In Figure 1, the shock in

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* Although the mutual causalities between any two variables in this system are theoretically well determined, not much has been said on the reverse causation from the divorce rate to age at marriage. A simple search model, however, demonstrates that optimal stopping for marriage (age at marriage) depends on the expected duration of marriage. A decrease in the expected duration of marriage lowers the potential benefits of searching longer for a mate, hence decreases age at marriage. If one assumes that the actual divorce rate shapes the expectations of the individuals on the duration of their potential marriages, one obtains causality running from the divorce rate to age at marriage (Mocan 1987).

the labor force participation rate triggers a jump in fertility, equal to 40 percent of its standard deviation before returning to its pre-shock level in approximately fifty months. Michael (1985), who encounters the same outcome in his bivariate autoregressions between fertility and labor force participation rate, argues that this might be due to the fact that labor force participation rate is a measure of participation by women with young children. An exogenous decrease in say day care costs can increase both labor force participation and fertility. Note that the effect of an exogenous variable, which is not a part of the system, but correlated with both fertility and the labor force participation rate is embodied in the error terms in the following way:

$$\epsilon_{Ft}=\emptyset_t+\mu_{Ft}, \quad \epsilon_{Lt}=\emptyset_t+\mu_{Lt},$$

where μ_{Ft} and μ_{Lt} are the white noise error terms of fertility and the labor force participation rate respectively which are independently distributed of each other. $\emptyset_t$ represents the influence of the common external factor, e.g. the decrease in day care costs. Under this structure, a one standard deviation shock to the labor force participation rate will affect the error of fertility instantaneously due to the contemporaneous correlation (θ_t), and it might not be possible to trace out the true response of fertility to the named shock in the labor force participation rate. To circumvent this inaccuracy, the impulse responses reported in this paper are based on the orthogonalization transformation, hence the absence of the contemporaneous correlation among errors. Nevertheless, one observes the increase in fertility as a response to an innovation in the labor force participation rate. This picture supports the findings of previous research which points to a positive relationship between economic activity and fertility behavior [e.g. Kirk (1960), Thomas and Galbraith (1956), Silver (1965), Ben-Porath (1973)].

In Figure 2, the proportion of young marriages and the divorce rate exhibit similar responses to the shock in the labor force participation rate. After some oscillations during the first 16 periods following the shock, they increase above their initial levels and stay there for the next 40 months. Note that, according to this picture, during the first three years after the shock, as the proportion of young marriages increases so does the divorce rate. After the third year,the decline in the proportion of young marriages brings about the decline in the divorce rate.

Figure 3 presents the reactions of the proportion of young marriages and the divorce rate to an unexpected increase in fertility. Note again that the proportion of young marriages and the divorce rate exhibit a co-movement similar to the one in Figure 2. Figures 2 and 3 demonstrate the strong interaction between the proportion of young marriages and the divorce rate, which is consistent with a priori expectations and the reports of past research. According to Figures 2 and 3, the increase in the proportion of young marriages couples with an increase in the divorce rate and vice versa. It should be noted, however, that this co-movement of the proportion of young marriages and the divorce rate is due to the unexpected increases in the labor force participation rate and fertility. Figure 4 on the other hand clearly demonstrates that the divorce rate is not influenced in any significant way by an increase in the proportion of young marriages. The shock to the proportion of young marriages generates oscillations in the divorce rate, which lose their magnitudes after the twenty-first month and approach the pre-shock level. According to this picture, an increase in the proportion of early marriages does not cause an increase in the divorce rate contrary to the evidence of cross-sectional analyses [e.g. Becker et al. (1977), Booth and Edwards (1985), Lee (1977), Schoen

(1975), Weed (1974), Bumpass and Sweet (1972)]. This implies that the proportion of young marriages and the divorce rate covary due to changes in labor force participation and fertility, but there is no causality from age at marriage to divorce.

III. SOME EXTENSIONS

A. ROBUSTNESS

To check the robustness of the results, different versions of the model are estimated. First, the model is estimated using a different measure of economic activity. Unemployment rate for women who are 16 years of age and over substituted for the labor force participation rate. The impulse responses of this system are reported in Figures 5-8. Figures 6 and 7 demonstrate the joint movement of the proportion of young marriages and the divorce rate when there are innovations in the female unemployment rate and fertility. Figure 8 shows the behavior of the divorce rate after a shock in the proportion of young marriages. These pictures are very similar to the ones reported earlier (Figures 2, 3 and 4). In Figure 5 one observes the reaction of fertility to an increase in the unemployment rate of women. The increase in the female unemployment rate, which is a proxy for an economic downturn yields a decline in fertility. An increase in economic activity, exemplified by an increase in the labor force participation rate of women, generated an increase in fertility (Figure 1). Figure 5 is the counterpart of Figure 1 and both simulations underscore the procyclical movement of fertility. When the unemployment rate of men, who are 16 and over is employed as the proxy of the economic activity, the same results are obtained (not shown).

To increase the degrees of freedom, SEXRATIO is dropped from the system. The system with LFPR, FERT, YOUNGMAR, and DIVORCE generated the same results as the initial system containing SEXRATIO as an additional variable. Similarly, the exclusion of SEXRATIO from the system where unemployment variables were substituted for LFPR did not alter the results.

B. A MISSPECIFIED MODEL

A VAR consisting of YOUNGMAR, DIVORCE, SEXRATIO, the constant, seasonal dummies, and trend variables is estimated. This is a misspecified model since it omits fertility and the labor force participation rate. In that sense it is the VAR representation of the methodology employed in previous studies. After estimating the model, a one standard deviation shock is applied to the proportion of young marriages.The behavior of the divorce rate is presented in Figure 9. In this model, as in previous studies, an increase in the proportion of young marriages generates an increase in divorce rate. However,as is shown in the previous section, adding fertility and labor force participation to the system makes the observed causality from age at marriage to divorce disappear (See Figure 4). It is not the causality from age at marriage to divorce that creates the picture of joint movement between the two; rather, the intervening effects of labor force participation and fertility generate that co-movement.

IV. SYNOPSIS

Given the inconclusive empirical evidence in the literature as to the relation between fertility and labor force participation of women on the one hand, and the expected interdependence between age at marriage, divorce, labor force participation and fertility on the other, a VAR technique is considered to be relevant to describe the dynamic interrelations among these variables. The estimated VAR and variance decompositions underscored the interdependence of the variables of the model, and allowed us to observe the short-run reactions of the variables to the unexpected perturbations in the system.

Contrary to the evidence presented by cross-sectional analyses, no significant influence of young marriages on the divorce rate is found: A positive shock to the proportion of young marriages did not generate an increase in the divorce rate. On the other hand, young marriages and the divorce rate simultaneously increased or decreased when there were innovations in the fertility or labor force participation rate. This outcome implied that the inverse association between age at marriage and divorce is not a causal relationship; rather divorces and marriages covary due to changes in labor force participation and fertility.

Another interesting outcome pertained to the exogeneity of the labor force participation rate. The estimated VAR and the variance decompositions highlighted that the labor force participation rate of women was not influenced by the fertility, age at marriage, or the divorce rate. Labor force participation decision was causally prior to other variables.

Also concluded that an unexpected increase in labor force participation generated an increase in fertility. This procyclical behavior of fertility was shown to be persistent even when different measures of economic activity like the female or male unemployment rates were employed.

FIGURE 1
SHOCK TO THE LABOR FORCE PARTICIP. RATE

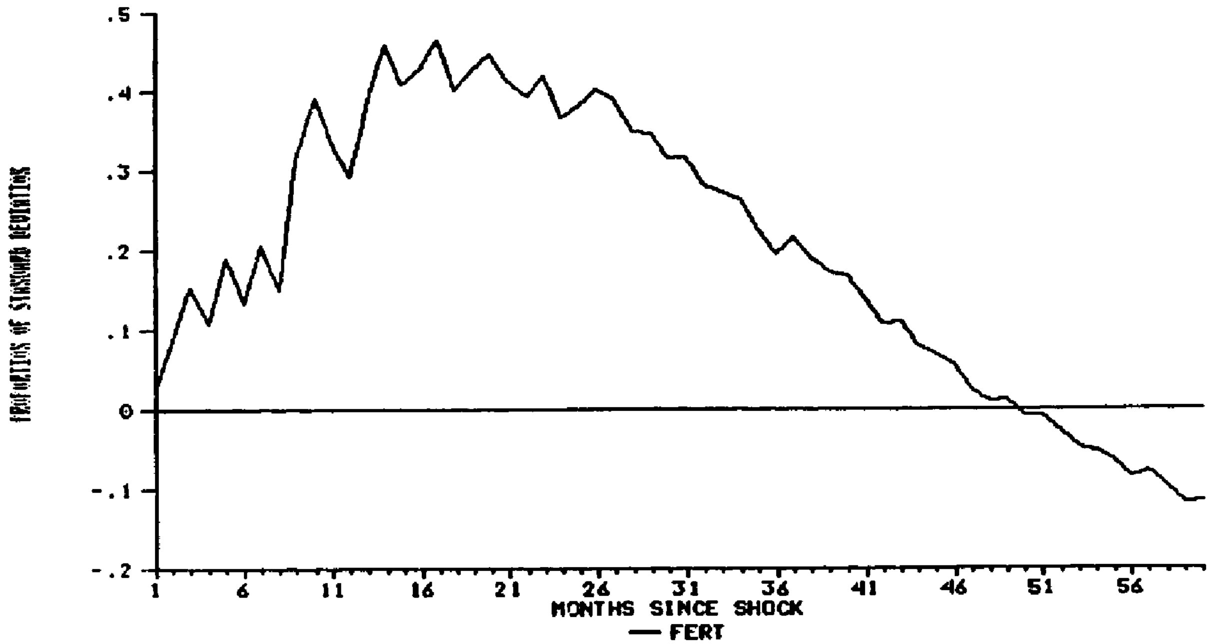


FIGURE 2
SHOCK TO THE LABOR FORCE PARTICIP. RATE

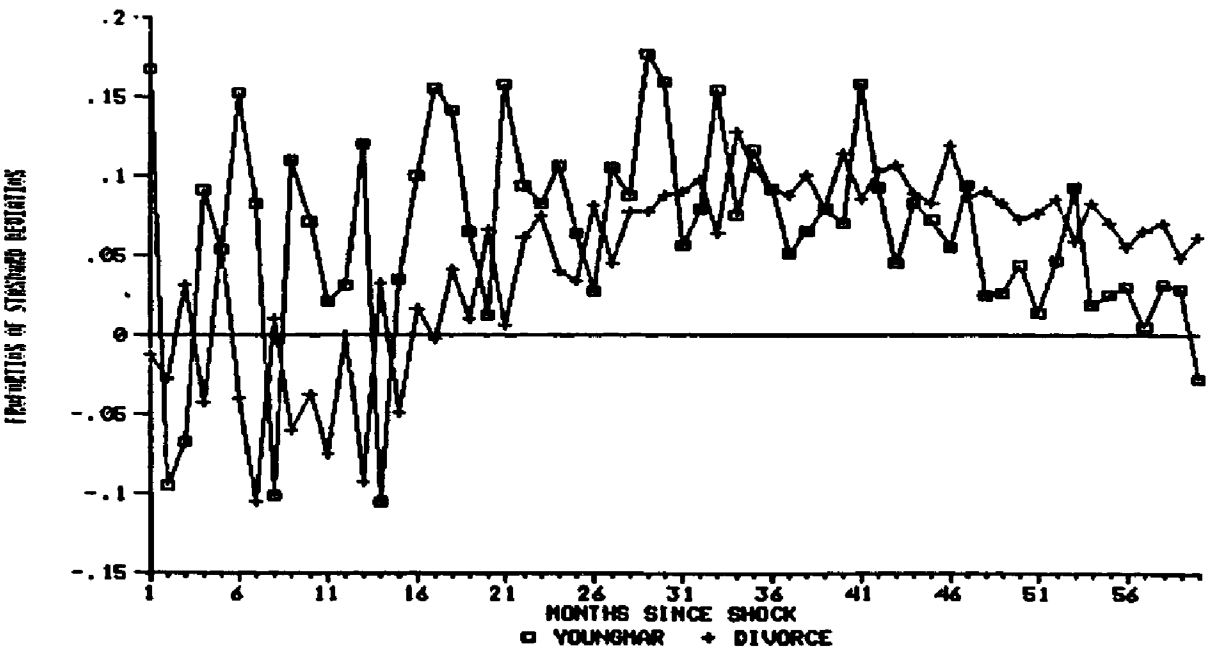


FIGURE 3
SHOCK TO FERTILITY

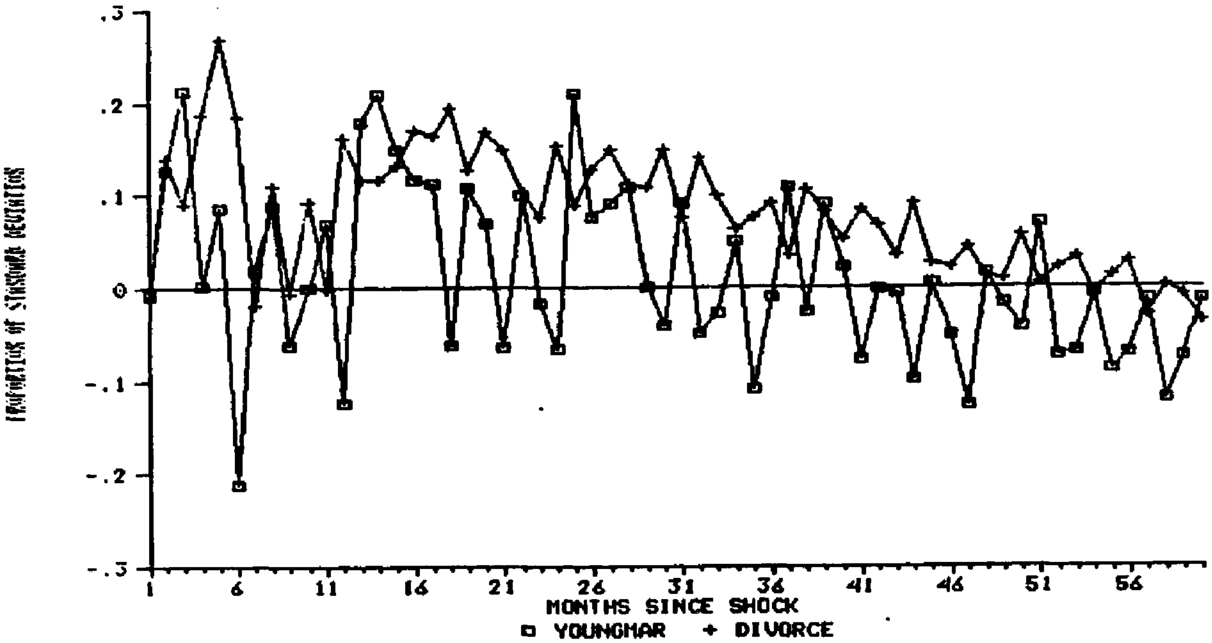


FIGURE 4
SHOCK TO THE PROPOR. OF YOUNG MARRIAGES

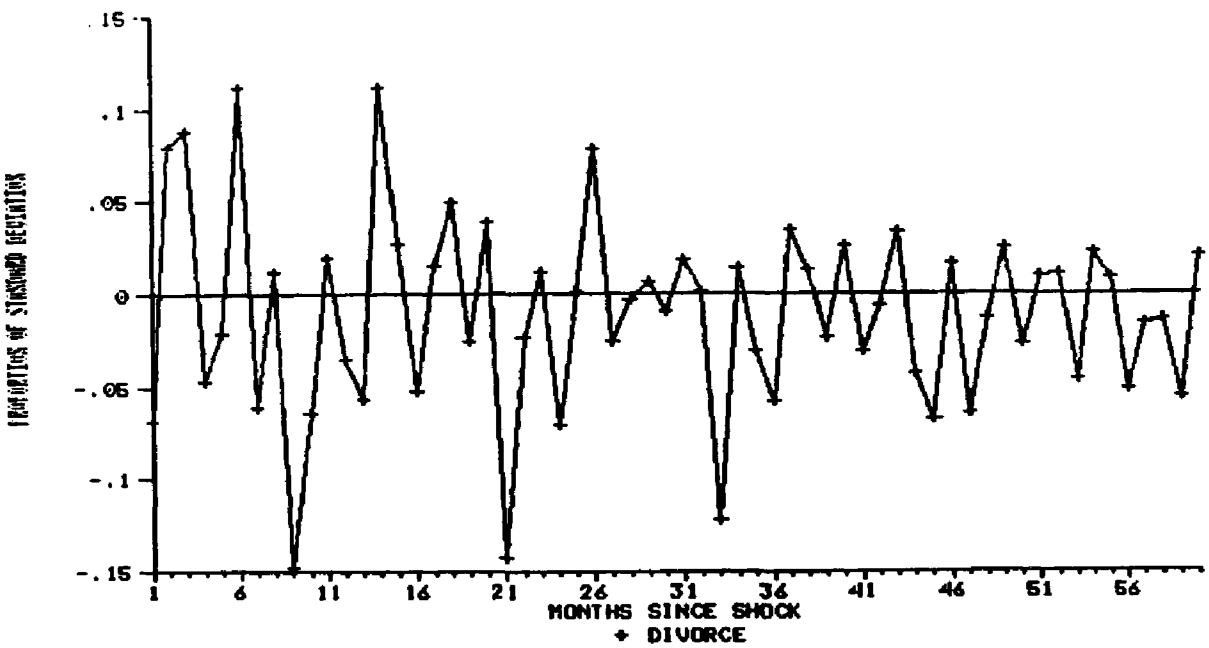


FIGURE 5
SHOCK TO THE FEMALE UNEMPLOYMENT RATE

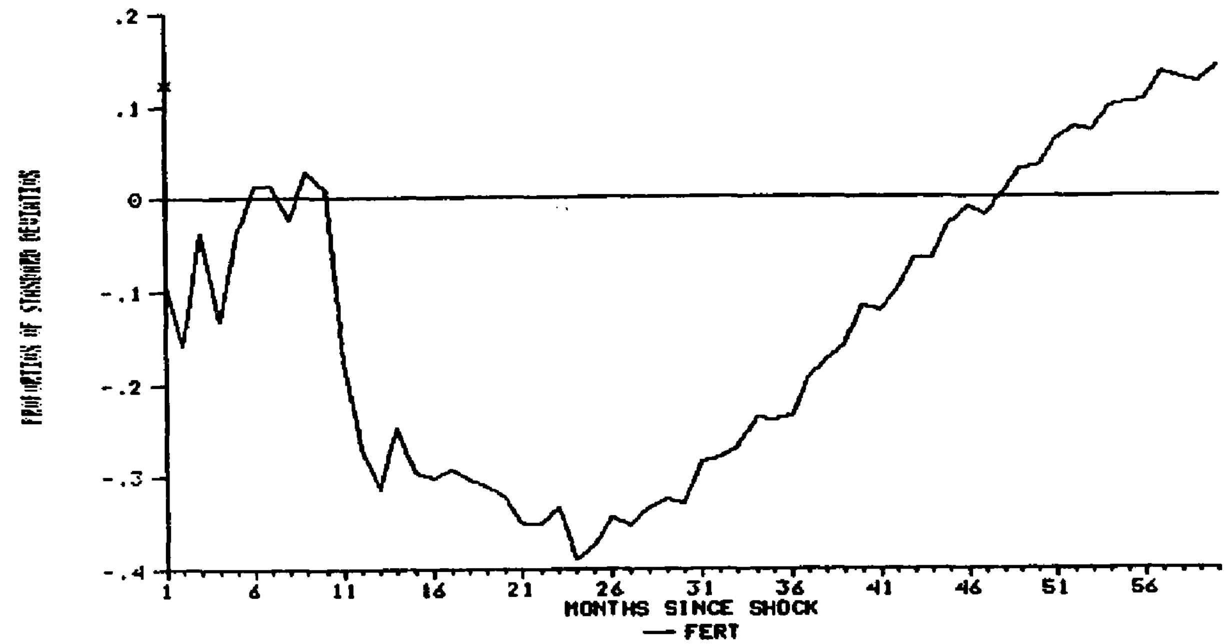


FIGURE 6
SHOCK TO THE FEMALE UNEMPLOYMENT RATE

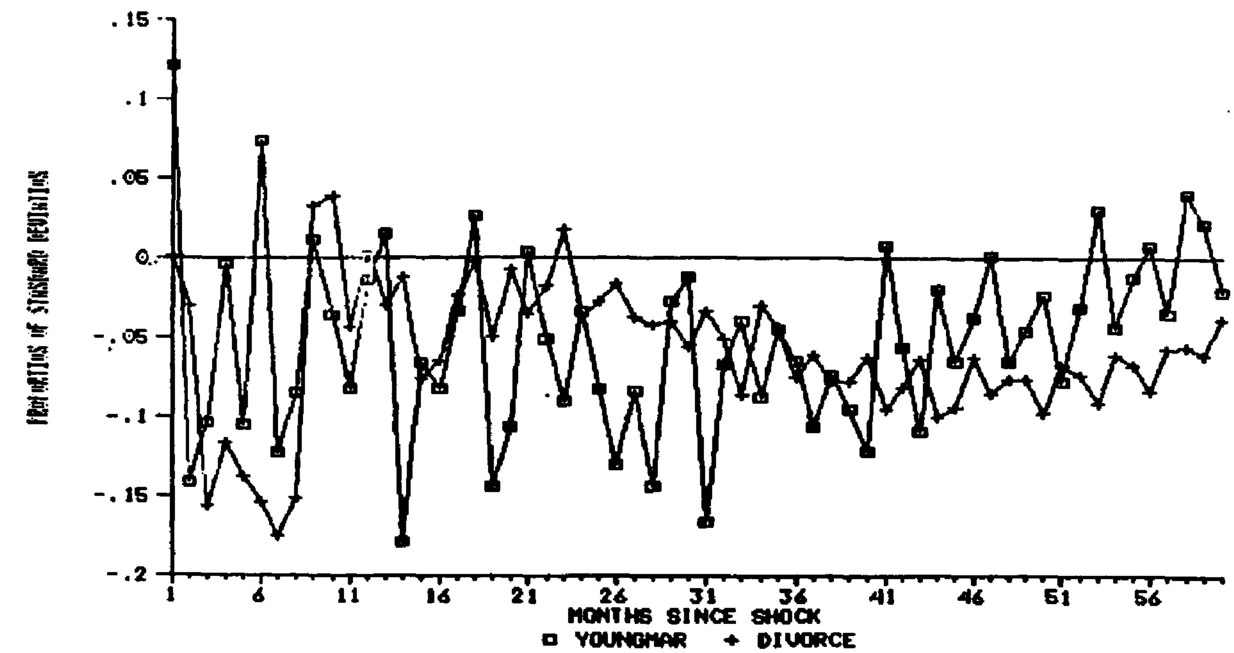


FIGURE 7
SHOCK TO FERTILITY

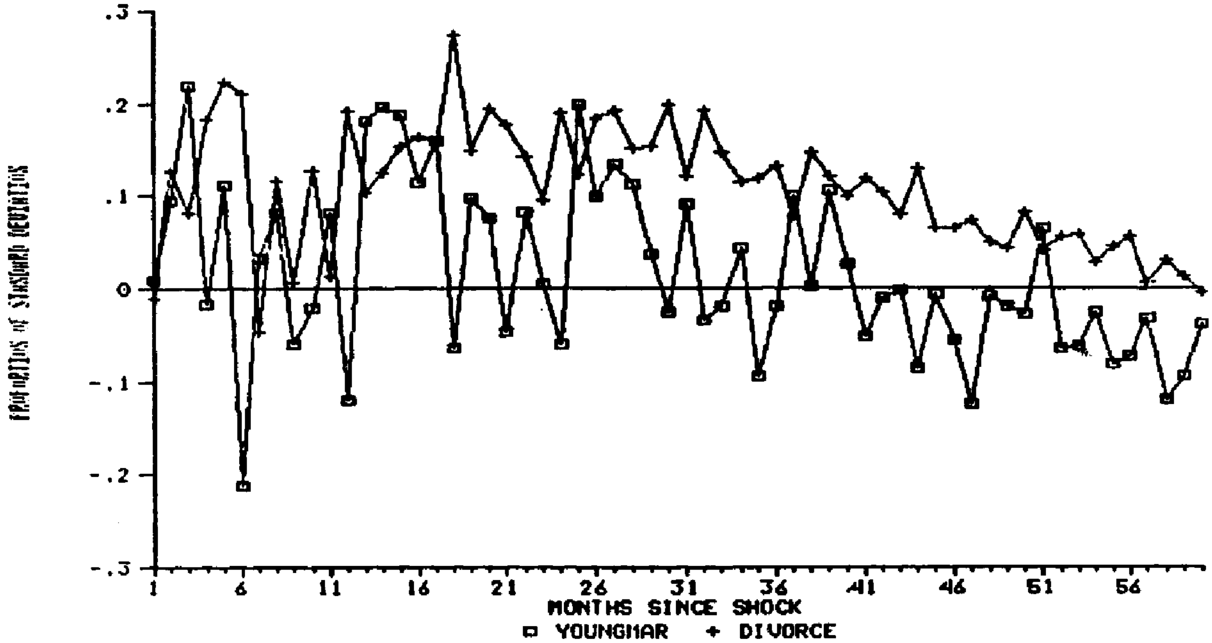


FIGURE 8
SHOCK TO THE PROPOR. OF YOUNG MARRIAGES

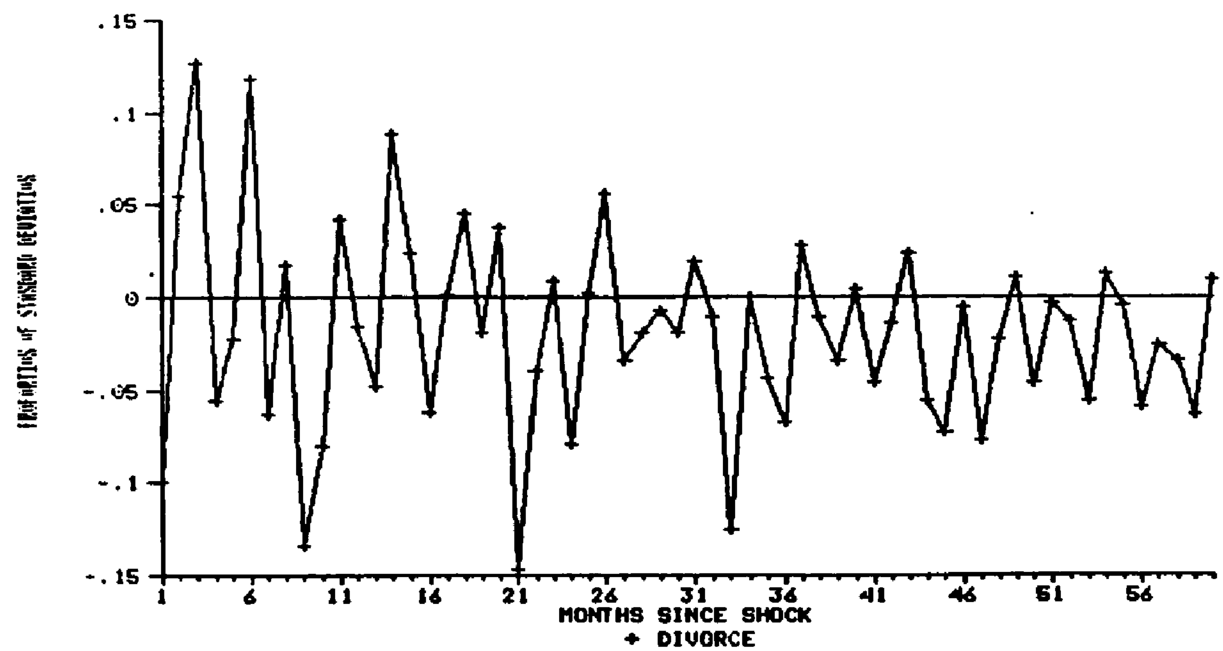
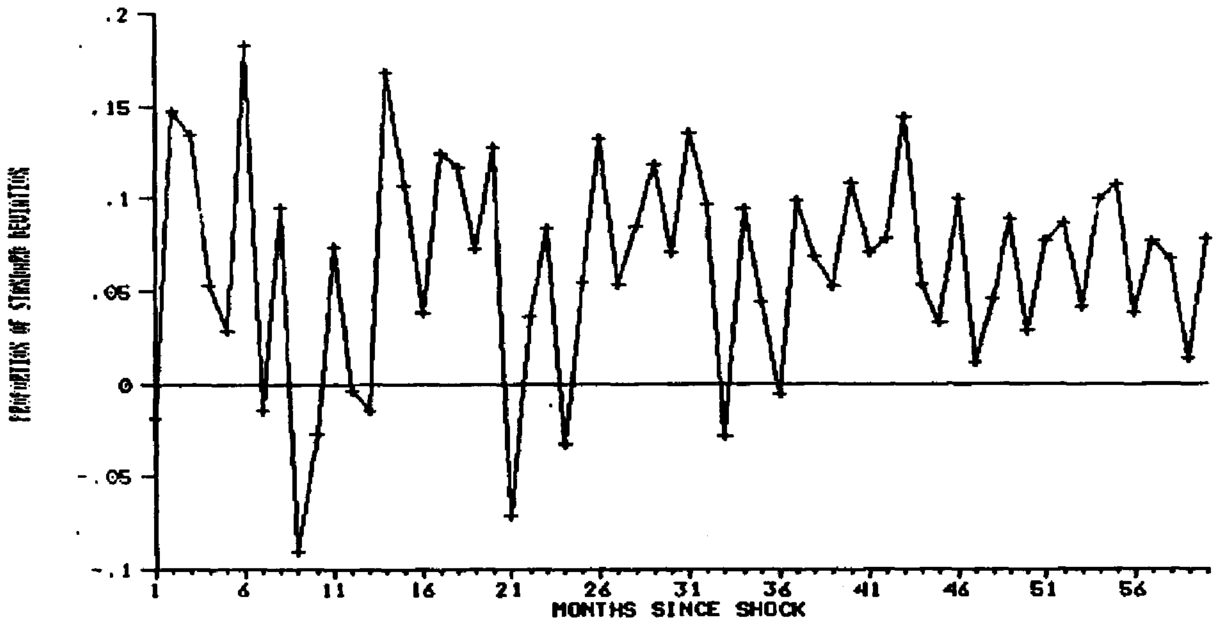


FIGURE 9
SHOCK TO THE PROPOR. OF YOUNG MARRIAGES



DEMAND SHOCKS, SUPPLY SHOCKS, AND THE CYCLICALITY OF REAL WAGES

I. INTRODUCTION

The cyclical behavior of real wages has received increasing attention in recent years (e.g. Bodkin 1969, Neftci 1978, Sargent 1978, Geary and Kennan 1982, Raisian 1983, Bils 1985, Keane et al.1988). These studies have tried to determine the sign of the relationship between real wages and either unemployment or employment. Some have found countercyclical behavior of real wages, whereas others have reported procyclicality. The cyclicality of real wages may be useful in determining the plausibility of competing business cycle theories. There exist two dominant approaches to business cycles: The real business cycle (RBC) models and non-RBC (mostly Keynesian disequilibrium) models. The RBC models attribute fluctuations in real quantities like output primarily to shocks in aggregate supply, stressing the role of technology and agents' preferences. The usual policy implication is that monetary policy plays only an insignificant role, if any. In a stronger language, money is neutral. The Keynesian models are based on wage and/or price rigidities, thus emphasizing imperfect market competition or frictions in contractual arrangements.

Business cycles are attributed to disturbances in aggregate demand, and the role of stabilization policies is acknowledged.[7]

Most textbooks and econometric models view the aggregate economy as consisting of aggregate supply and aggregate demand. In this standard model, in the short-run, disturbances in aggregate demand lead to movements in output and prices in the same direction, while supply shocks move them in opposite directions. Hence, disequilibrium business cycle models assuming sticky nominal wages predict countercyclical real wages when shocks are caused by aggregate demand disturbances, and procyclical real wages when shocks occur in aggregate supply. Indeed, the disequilibrium Keynesian models with long-term contracts (Fischer 1977) or staggered contracts (Taylor 1980) would generate countercyclical real wages when unanticipated shocks in money supply or velocity take place. In general, an unanticipated increase, say in money supply, raises output and prices and reduces real wages as firms remain on their labor demand schedules. In fact, both classical economists and Keynes shared this view. The countercyclicality of real wages was, in the words of Bils, "based on the premise of competition and a given

7 For a detailed description of these views, see Long and Plosser (1983), King and Plosser (1984), Eichenbaum and Singleton (1986), and Shapiro (1987).

short-run capital stock. Increases in employment then correspond to more intensive use of capital and, through diminishing product, lower real wages." However, since some current Keynesian versions of disequilibrium models stress aggregate demand shocks as the primary source of business cycles, they try to explain procyclicality of real wages--if it is in fact observed-- not by supply shocks, but by "excess-capacity" or "labor-hoarding" theories (Shapiro 1987). According to these theories, firms do not adjust their labor input in light of short-run fluctuations in demand because it is costly to do so.

The RBC models, on the other hand, by attributing business cycles to shifts in technology, as Shapiro argues, "explain the joint movement of output and measured productivity virtually by definition." Therefore, it is not difficult to reconcile the RBC theories with procyclical real wages.

We do not intend to argue for one approach against the other --whether the real or monetary shocks are the most important source of variation in output over the business cycle or whether there is room for activist demand-management policies. Ideally, the relative importance of real and monetary shocks could be assessed in identified structural models where the variation in a variable of interest, such as output, could be decomposed into variation attributable to the two types of shocks in question. This, however, is beyond the scope of this study, as our purpose is to determine whether output and real wages move in the same direction following demand and supply disturbances.

In this paper we use unrestricted vector autoregressions (VAR) to gain insights into the dynamic interrelations between output, unemployment, and real wages. Our goal is to uncover statistical evidence about various hypotheses, and since the technique we employ is equivalent to estimating reduced form equations, it does not allow us to make structural interpretations. Nevertheless, this study should be viewed as a complementary way of shedding light to important correlations in the data. Hence, it can increase our understanding of business cycles and the cyclicality of real wages, supplementing the literature in this field.

In particular, we analyze the responses of output, unemployment and real wages when shocks occur in nominal money (aggregate demand) and in the relative price of oil or productivity (aggregate supply). We find that supply shocks generate procyclicai real wages, while a demand shock yields countercyclicality when the product wage (nominal wage deflated by the WPI) is used as a measure of real wage rate. However, when the consumption wage (nominal wage deflated by the CPI) is employed, both types of shocks produce procyclicality, although the magnitude of the response of the consumption wage to a demand shock is much smaller than that of the product wage.

Previous studies on the real wage cyclicality have employed different measures of the real wage rate. They also used different sample periods, which may have been dominated by different types of shocks. The conflicting findings of those studies, therefore, are not due to the different methodologies that they used, but due to the different underlying disturbances in the time intervals they focused on, and also due to the different measures of real wages they employed.

II. ESTIMATION

A. VARIABLES AND THE DATA

Our basic system consists of real output (Y), the rate of unemployment (U), nominal money (M), the relative price of oil (OIL), and the product wage (W_p).* Versions of this system containing a measure of productivity (PROD) and a different measure of real wages--the consumption wage (W_c)--are also estimated.

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* This system is similar to the one employed by Sims (1980) which includes import price in addition to output, unemployment, price level, nominal wage, and nominal money.

The variables are defined as follows: Y is the real GNP (1982=100); U is the rate of unemployment of the civilian population; M is nominal M1; OIL is generated by taking the ratio of the price of crude petroleum to the GNP deflator; W_p is given by the ratio of nominal average hourly earnings in manufacturing to the Wholesale Price Index, industrial goods; W_c is the nominal average hourly earnings in manufacturing deflated by the Consumer Price Index; and PROD is the ratio of real GNP to total employment. The data are obtained from Citibank Economic Database and cover the period 1948:1-1986:4.

B. SPECIFICATION

Recent developments in time-series econometrics have underscored the importance of differentiating between difference-stationary processes (DSP) and trend-stationary processes (TSP) (Nelson and Plosser 1982). Failure to distinguish between the two can generate seriously misleading results (Nelson and Kang 1984, Stock and Watson 1988). Series belonging to DSP type should be detrended by differencing, whereas inclusion of time variables is most appropriate for TSP type series. A test developed by Dickey and Fuller (1981) can be used to test the hypothesis that a particular series belongs to DSP type against the alternative that it belongs to TSP type.

To implement the Dickey-Fuller test we estimate

$$(1)\quad \Delta z_t = \alpha_0+\beta_0 t+\beta_1 z_{t-1}+\sum_{i=1}^{k}\delta_i\Delta z_{t-i}+\epsilon_t,$$

where z is in natural logs.

The variable z belongs to DSP type if $\beta_0=\beta_1=0$. Following Nelson and Plosser the choice of k in equation (1) is based on the autocorrelations of first-differences and the partial autocorrelations of the deviations from the trend. Table 4 summarizes the results of Dickey-Fuller tests. The hypotheses that the money supply and the unemployment rate have unit roots are rejected. Consequently, in our VAR systems all variables are in first differences, except for the unemployment rate and the money supply.

We examine the responses of real wages, unemployment and output to unanticipated one standard deviation shocks in nominal money, the relative price of oil or productivity within the following five-variable unrestricted VAR systems:

1. OIL, M, Y, U, W_p
2. OIL, M, Y, U, W_c
3. PROD, M, Y, U, W_p
4. PROD, M, Y, U, W_c.

TABLE 4

DICKEY-FULLER TESTS FOR UNIT ROOTS

z	k	F-stat
Real GNP (Y)	2	4.84
Nominal M1 (M1)	2	14.03
Price of Petroleum (OIL)	6	2.37
GNP-Employment Ratio (PROD)	3	5.08
Production Wage (W_p)	2	2.33
Consumption Wage (W_c)	6	5.38
Unemployment Rate (U)	2	7.72

All variables are in logs. F-stat represents the F-statistics under the null hypothesis $\beta_0=\beta_1=0$ in the regression

$$\Delta z_t = \alpha_0+\beta_0 t+\beta_1 z_{t-1}+ \sum_{i=1}^{k}\delta_i \Delta z_{t-i}+\epsilon_t,$$

where t is a linear trend term. The F-ratio tabulated by Dickey and Fuller is 6.49 at the 5% level when the sample size is 100, and 6.34 when it is 250. Our critical F-ratios lie between these two since our sample sizes vary between 149 and 153.

Each of the five equations in all systems contains a set of three seasonal dummies as well as five lags of each variable. Also, a constant and linear and quadratic trend terms are included. All variables are expressed in natural logs.

To check the specification of the model two tests were performed on the basic system (1). First, to see if the lag specification adequately captures the dynamics of the model, the system with five lags was tested as a restriction on the same system with eight lags. The test was based on the difference in the log-determinants of the system-wide variance-covariance matrix of the restricted versus unrestricted equations.* The null hypothesis of no difference was accepted. The chi-square was 73.88 with 75 degrees of freedom.

Secondly, a Lagrange Multiplier (LM) test was applied to the residuals. The LM statistic provides a general test for autocorrelation of errors, and is valid when the set of regressors includes lagged dependent variables (Breush 1978, and Godfrey 1978). The tests are based on additional regressions in which each dependent variable is regressed on the same set of variables plus a set of lagged residuals. One rejects the null hypothesis of no

* This is the modified chi-square test proposed by Sims(1980).

autocorrelation (white noise errors) if the set of lagged residuals is different from zero. As Breush and Godfrey (1981) show, the reported F-statistic is asymptotically equivalent to the usual LM statistic. Table 5 shows the F-statistics for the coefficients of lagged errors in each equation for 1, 2, 3 and 4 lags of the errors. As is evident, none of these coefficients are significant at the 5% level. Thus we accept the hypothesis that the errors of the system are white noise, and the system is well specified.

Recently, Engle and Yoo (1987) have suggested that time series which are stationary after differencing may have linear combinations which are stationary without differencing--that is, they may be cointegrated. If so, the system may be "overdifferenced," and an error correction procedure must then be adopted (Engle and Granger, 1987).

The cointegrating regression is:

(2) $X_{1t}=\alpha_t+\beta_2X_{2t}+\beta_3X_{kt}+\epsilon_t$,

where X_1 represents one of the three DSP type series in each system. The residuals obtained from (2) are then used in $\Delta\epsilon_t=\tau\epsilon_{t-1}+u_t$,

where the hypothesis $\tau=0$ is tested. As Table 6 demonstrates, none of the variables were found to be cointegrated in any of our four systems.

TABLE 5

LAGRANGE MULTIPLIER TESTS

Lags of Error	Variable OIL	M	Y	U	W_p
1	.88	1.96	.0001	.65	.16
	(.35)	(.16)	(.99)	(.42)	(.64)
2	2.16	1.32	.16	.40	1.73
	(.12)	(.27)	(.85)	(.67)	(.18)
3	1.53	1.64	.10	.52	2.47
	(.21)	(.18)	(.96)	(.67)	(.06)
4	1.19	1.51	.09	2.17	2.30
	(.32)	(.21)	(.98)	(.08)	(.06)

The entries are the F-statistics.
The numbers in parenthesis are the marginal significance levels.

TABLE 6

COINTEGRATION TESTS

System	t-value for τ	System	t-value for τ
Y	.395	Y	1.503
OIL	.236	PROD	.120
W_p	.812	W_p	2.012
Y	1.790	Y	1.359
OIL	.221	PROD	.325
W_c	.580	W_c	1.725

With 3 variables and a sample size of 100, the critical t-value by Engle and Yoo is 3.93 at the 5% level, and 3.78 when the sample size is 200. With our sample size of 154, our critical t-values lie between the two.

III. IMPULSE RESPONSE FUNCTIONS

In this section we describe the impulse response functions generated by unanticipated shocks in aggregate demand and aggregate supply. These functions enable us to characterize the dynamic interactions among variables, and observe the speed of adjustment of variables in the system. One criticism of such simulations is that, if the contemporaneous correlations among innovations are substantial, then the assumption of a unique error structure, on which the interpretation of the responses depends, is invalid. This problem can be avoided by triangularizing the variance-covariance matrix of residuals, which transforms the unrestricted VAR systems to block-recursive systems (Sims, 1980). This procedure requires imposition of some causal ordering of the variables. The ordering matters especially when the variables exhibit strong correlations.[10] In the systems we estimate the variables are ordered as OIL (or PROD), M, Y, U, and W_p (or W_c).[11]

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[10] Although in most studies employing VAR methodology the ordering remains arbitrary and controversial, the researchers try several orderings, placing the variables which are known to respond most strongly to contemporaneous events at the bottom of the ordering list (Gordon and King, 1982).

[11] In all systems several orderings were tried and yielded very similar impulse responses.

The impulse responses of output, unemployment and the product wage to an unanticipated one standard deviation shock in the relative price of oil in the basic system 1 are given in figures 10-A to 10-C. Output starts declining two quarters after the shock very sharply, falling as much as 0.7% below its initial level in quarter 6. It then starts rising and returns to its steady-state level. Thus, output is adversely affected by an unfavorable supply shock for a relatively longer period. The response of unemployment reveals the negative nature of the output-unemployment correlation. Unemployment initially rises sharply, thereafter declining and eventually also returning to its steady-state level. Finally, the shock in the relative price of oil permanently reduces the real wage rate. Specifically, it leads to a decrease in the product wage of about 0.4% in the first quarter and slightly more than 1% in the first year. After this abrupt decline, it stabilizes and remains more than 1% below its initial level in the long-run. These results clearly indicate the procyclicality of real wages--that is, negative correlation between unemployment and real wages.

The impulse responses of output, unemployment and the product wage to a one standard deviation shock in money in the same system are presented in figures 10-D to 10-F. A positive demand shock leads to an increase in output for about 10 periods and a decrease in unemployment for 10 periods. However, afterwards both variables return to their mean levels, indicating that the monetary shock is effective for a relatively shorter time period. On the other hand, real wages decline about 0.3% by the third quarter following the initial shock before it begins rising and approaching its steady-state value in 35 quarters. Hence, a demand shock, exemplified by an unexpected increase in money supply, leads to countercyclical real wages--positive correlation between unemployment and real wages.

Figures 11-A to 11-C show the impulse responses to a shock in the relative price of oil in our second system where the consumption wage is used in place of the product wage. A comparison with the previous findings suggests that the responses of output and unemployment to an unfavorable supply shock are virtually the same as before (see Figures 10-A and 10-B). The consumption wage is reduced permanently also, remaining about 0.15% below its initial value in the long-run. The absolute magnitude of the initial sharp decline in the product wage is smaller than the one in the consumption wage. However, qualitatively, the results are not affected: the consumption wage is also procyclical. As figures 11-D to 11-E demonstrate, a monetary shock yields virtually the same responses in output and unemployment as before (See Figures 10-D and 10-E). However, the real wage rate rises immediately after the shock by 1.2%, and although it declines after the fifth quarter, it remains slightly above its pre-shock value in the long-run. Therefore, the consumption wage is procyclical whether shocks originate in aggregate demand or aggregate supply. Note, however, that the magnitude of the change in the consumption wage following the demand shock (Figure 11-F) is considerably smaller than all other real wage responses to both types of shocks in all systems.

Our approach does not allow us to arrive at any causal explanation of this persistent procyclical behavior of real wages when the CPI replaces the WPI as the deflator. However, these findings may suggest that relative to the economy as a whole, the industrial sector is highly cyclical during periods dominated by aggregate demand shocks. If so, an increase in money supply could reduce the ratio of nominal wages to the WPI, given that the increase in the WPI is greater than the increase in the relatively rigid nominal wages. Yet, our results also imply the possibility that the CPI is much less sensitive to the changes in money supply and increases in the CPI may be dominated by the nominal wage increases.

The dynamic effects of a favorable supply shock are characterized in figures 12-A to 12-C. The system now includes productivity in place of relative price of oil, along with money, output, unemployment and the product wage. A one standard deviation shock in productivity raises output by 0.7% in the first quarter and as much as 0.95% in the third quarter following the shock. It then declines and approaches its initial level. Unemployment again moves in opposite direction. Finally, the real wage rate rises permanently; after rising in the first six quarters, it remains at that level. Thus, a favorable supply shock, like an unfavorable one, leads to procyclicality and has a long-term effect on real wages. The impulse responses in this system generated by a monetary shock are summarized in figures 12-D to 12-F. A comparison with figures 10-D to 10-F shows clearly that a demand shock produces similar responses in the present system and yields countercyclical real wages.

Finally, in our fourth system we replace the product wage by the consumption wage. The impulse response functions to a productivity shock are presented in figures 13-A to 13-C. As expected, unemployment falls and remains below its initial level for about 20 quarters before the effect of the shock dies out. Output rises as much as 1% in the first year and eventually returns to its initial level. The real wage rate is procyclical, as it initially rises sharply, then falling and remaining slightly more than 0.1% above its initial level. As figures 13-D to 13-F indicate, once again the demand shock results in procyclical real wages if the nominal wages are deflated by the CPI.

To sum, supply shocks produce procyclical real wages, while a demand shock produces countercyclicality in systems where the product wage is used as a measure of the real wage rate. On the other hand, both supply shocks and a shock in money supply produce procyclical real wages when the consumption wage is used as the real wage rate variable.

IV. CONCLUSION

The question of whether real wages are procyclical or countercyclical has perplexed economists since Keynes (who predicted countercyclicality in the General Theory). Most recent studies, which found conflicting evidence have only added to the uncertainty pertaining to this question. In this paper using vector-autoregressions we demonstrate that the cyclicality of real wages depends on the source of the disturbances. In particular, we show that the shocks in aggregate supply produce procyclical real wages whether the nominal wage is deflated by the wholesale price index of industrial goods or the consumer price index. On the other hand, shocks in aggregate demand generate countercyclical behavior of real wages when the real wage is measured as the ratio of nominal wage rate to the wholesale price index. However, if the nominal wage is deflated by the consumer price index, aggregate demand shocks yield movement in the same direction of output and real wages.

In other words, real wages can be procyclical or countercyclical depending upon the type of shock which generates the cycle, and the measurement of the real wage rate. Since the previous studies on the real wage cyclicality have employed different measures of the real wage rate and, more importantly, different sample periods which may have been dominated by different types of shocks, it is not surprising that they report conflicting findings.

FIGURE 10-A
SHOCK TO OIL - SYSTEM: OIL,M,Y,U,WP

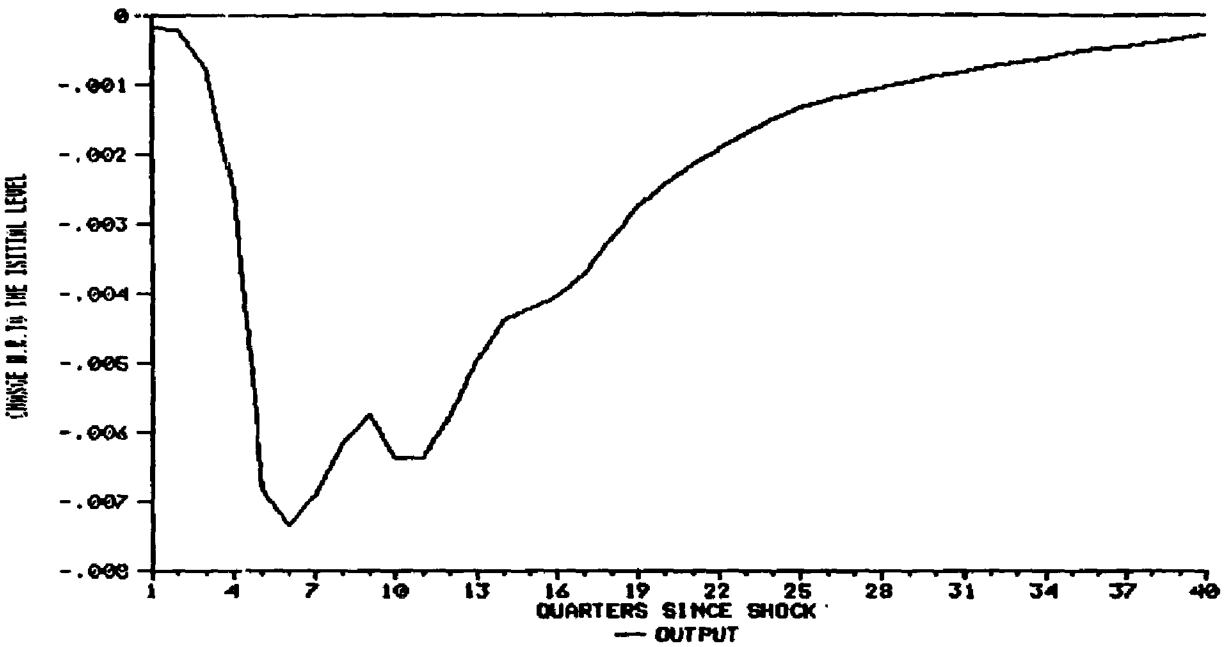


FIGURE 10-B
SHOCK TO OIL - SYSTEM: OIL,M,Y,U,WP

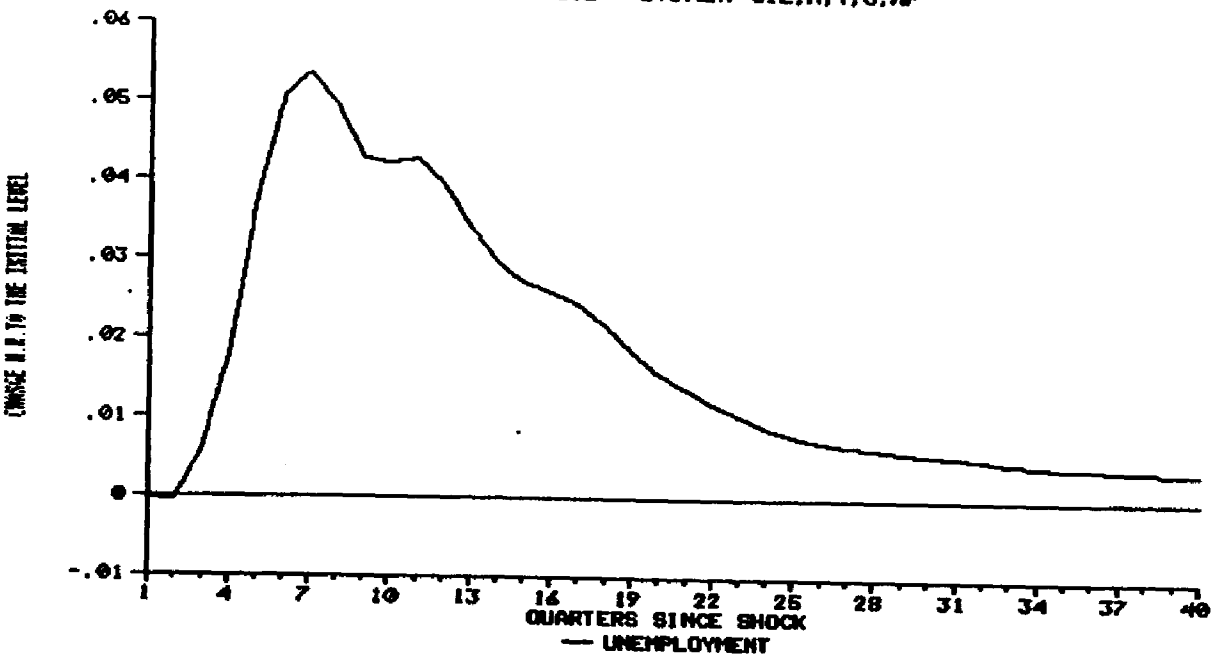


FIGURE 10-C
SHOCK TO OIL - SYSTEM: OIL,M,Y,U,WP

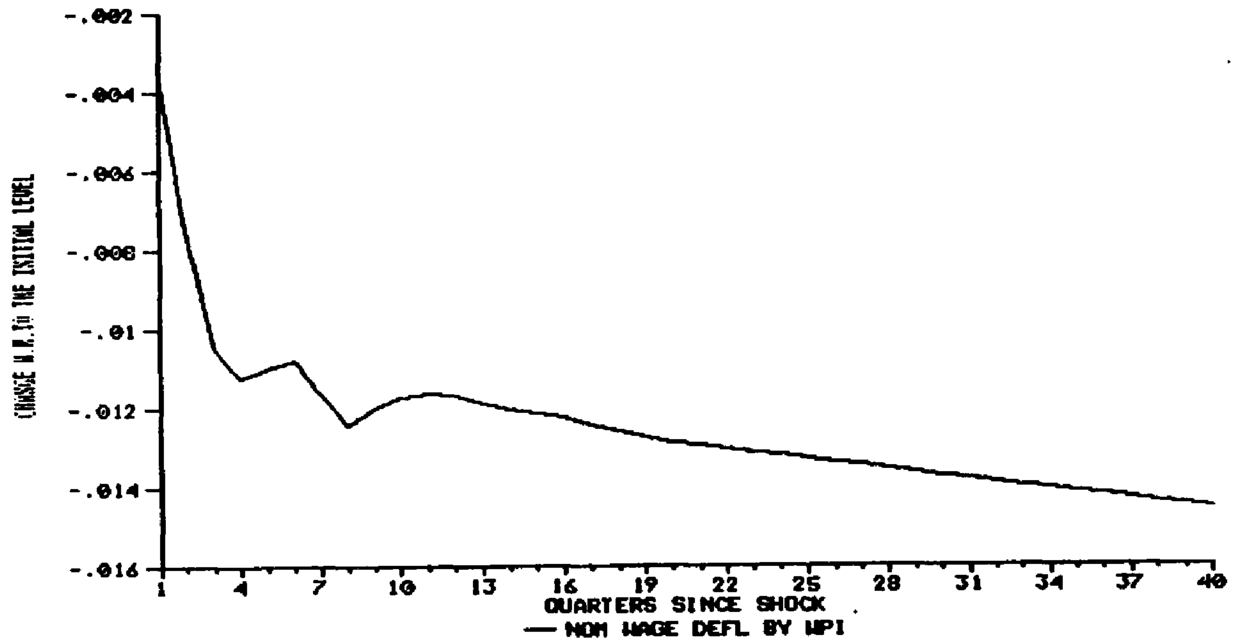


FIGURE 10-D
SHOCK TO MONEY - SYSTEM: OIL,M,Y,U,WP

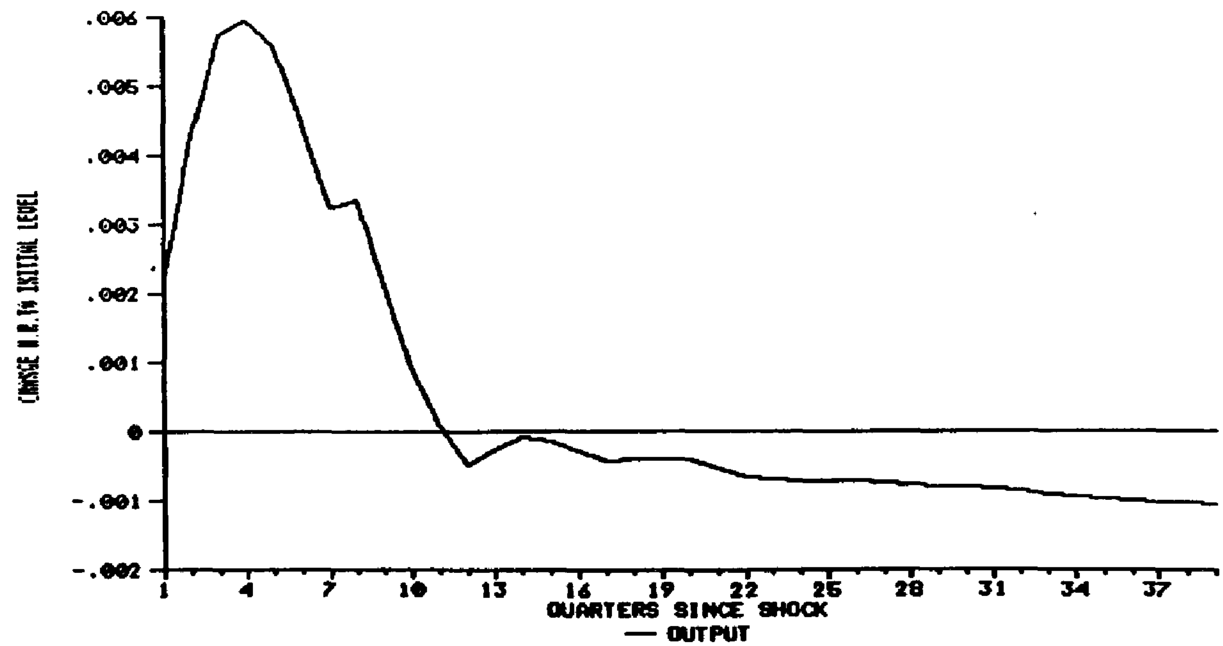


FIGURE 10-E
SHOCK TO MONEY - SYSTEM: OIL,M,Y,U,WP

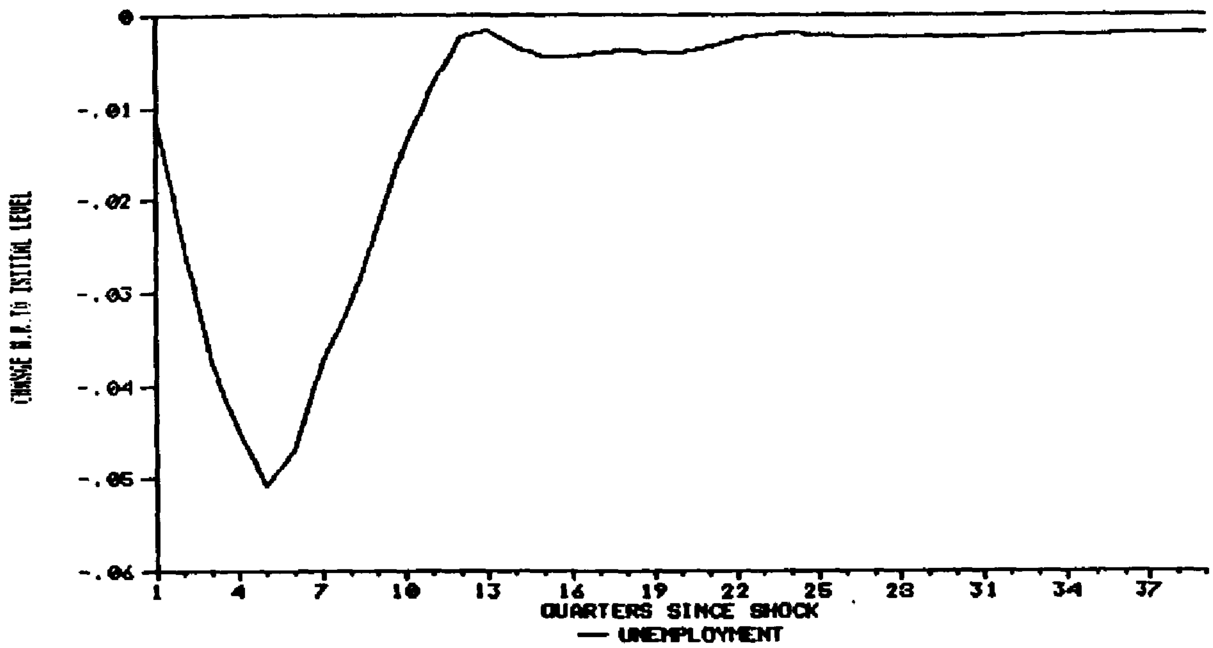


FIGURE 10-F
SHOCK TO MONEY - SYSTEM: OIL,M,Y,U,WP

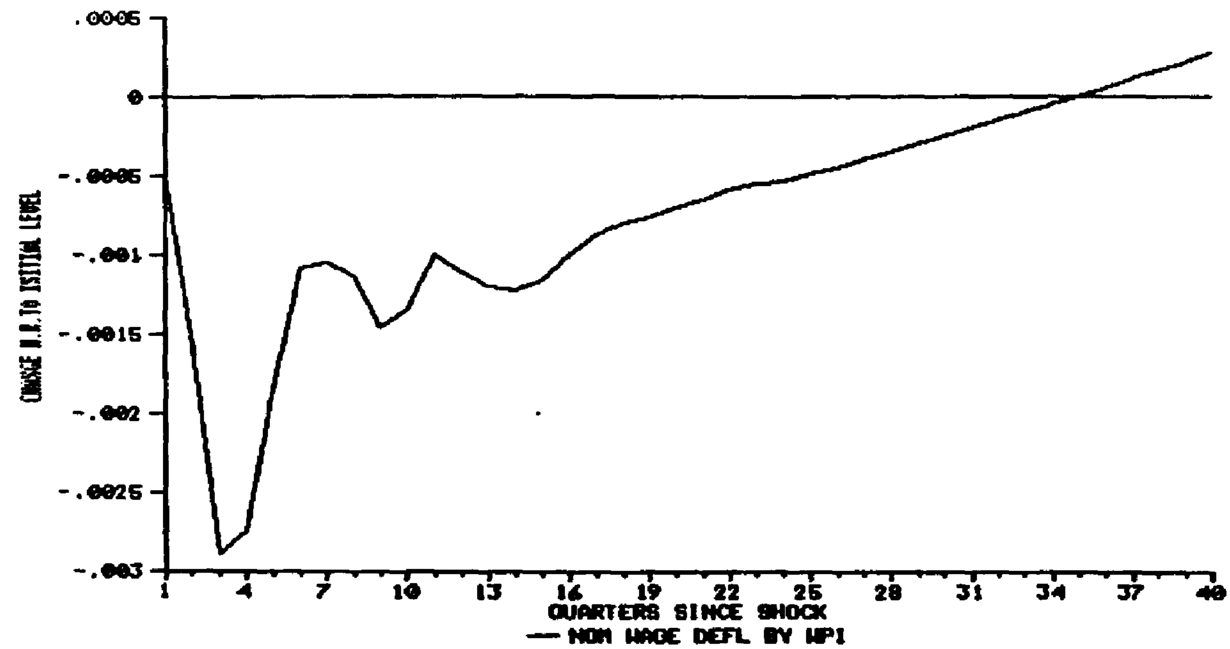


FIGURE 11-A
SHOCK TO OIL - SYSTEM: OIL,M,Y,U,WC

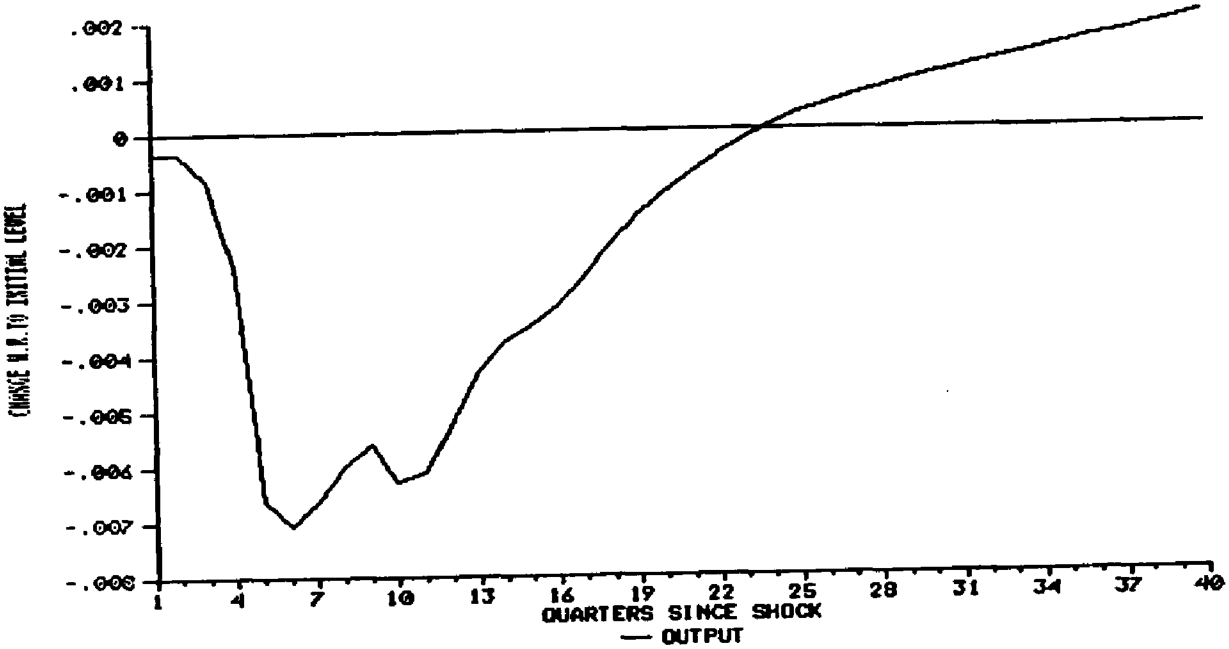


FIGURE 11-B
SHOCK TO OIL - SYSTEM: OIL,M,Y,U,WC

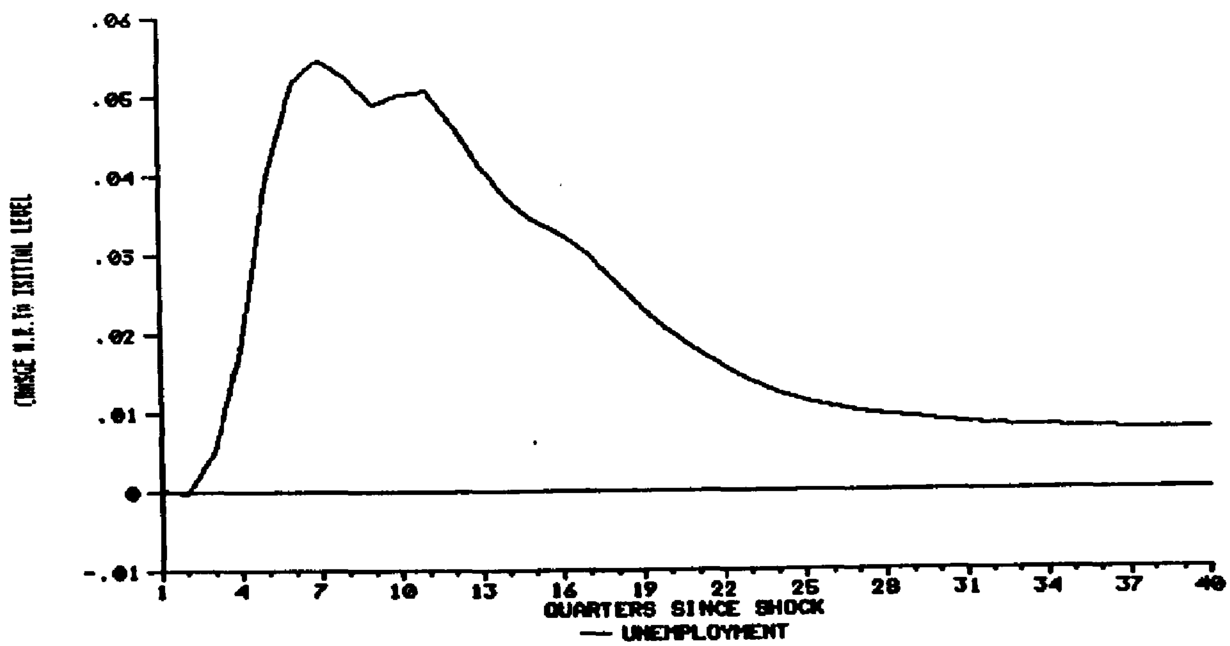
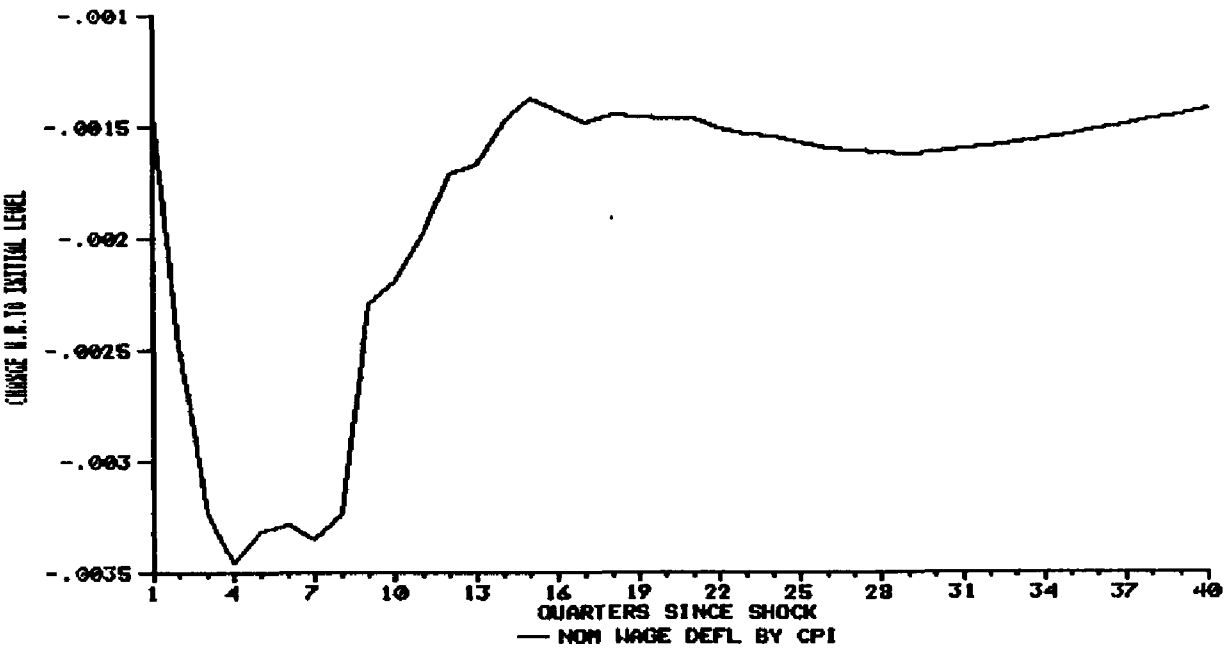
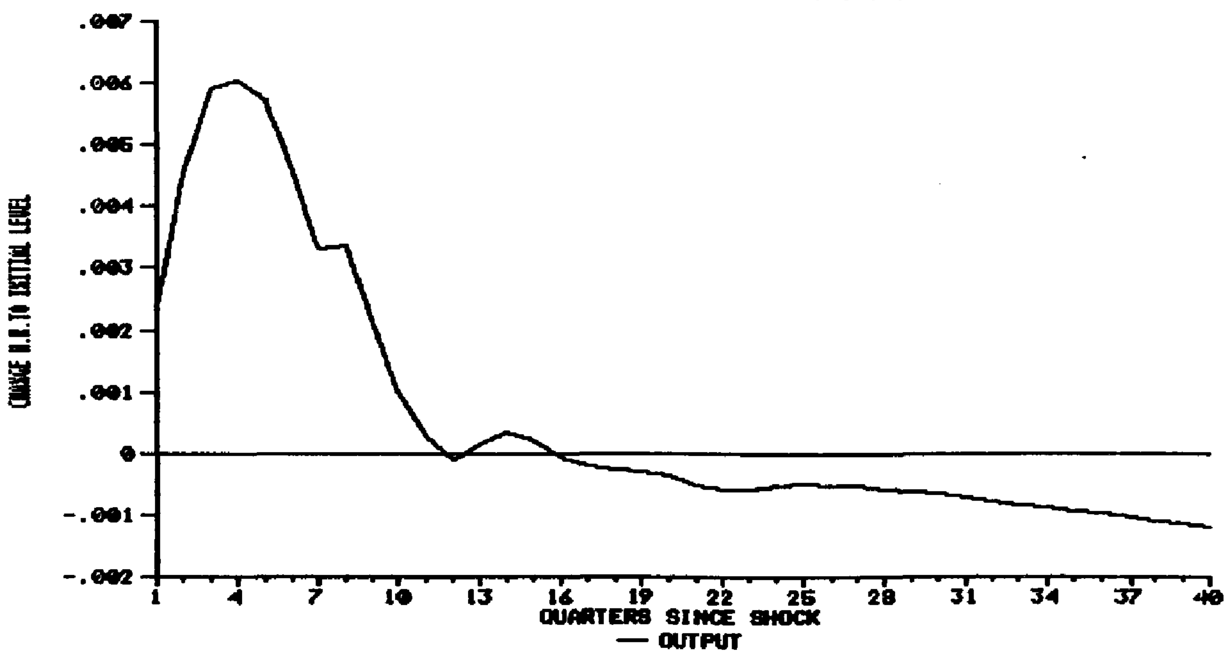
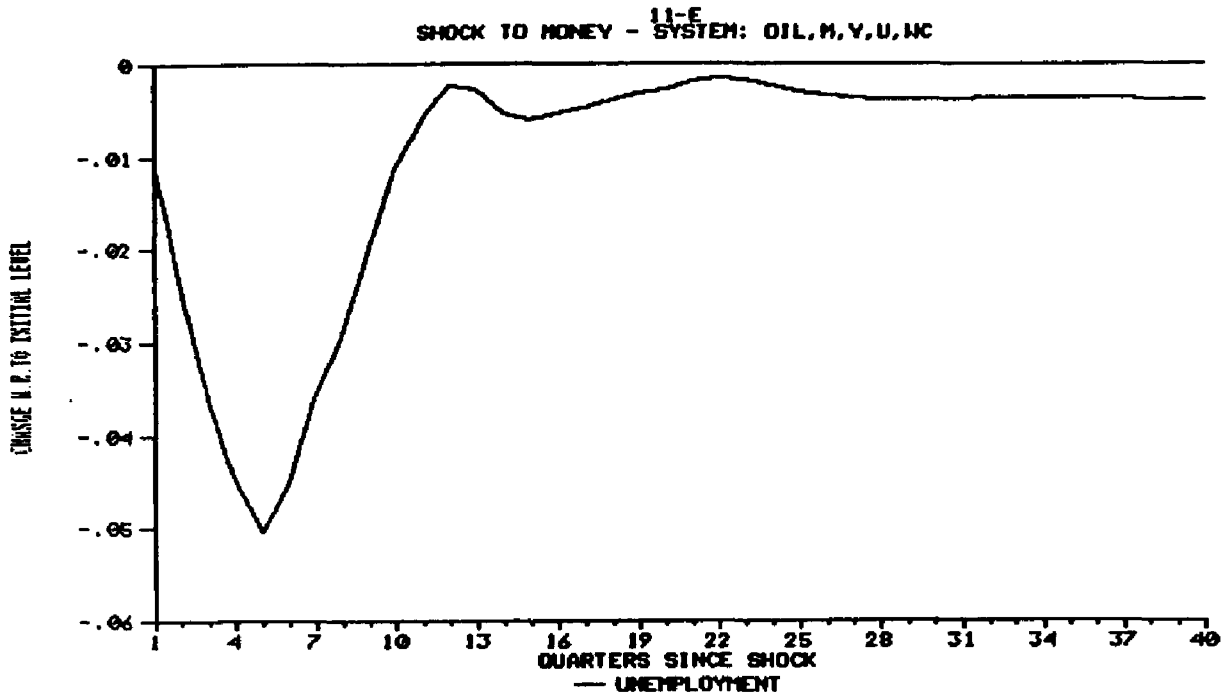


FIGURE 11-C
SHOCK TO OIL - SYSTEM: OIL,M,Y,U,WC



11-D
SHOCK TO MONEY - SYSTEM: OIL,M,Y,U,WC




11-E
SHOCK TO MONEY - SYSTEM: OIL,M,Y,U,WC
CHANGE W.R.TO INITIAL LEVEL
0
-.01
-.02
-.03
-.04
-.05
-.06
1
4
7
10
13
16
19
22
25
28
31
34
37
40
QUARTERS SINCE SHOCK
— UNEMPLOYMENT

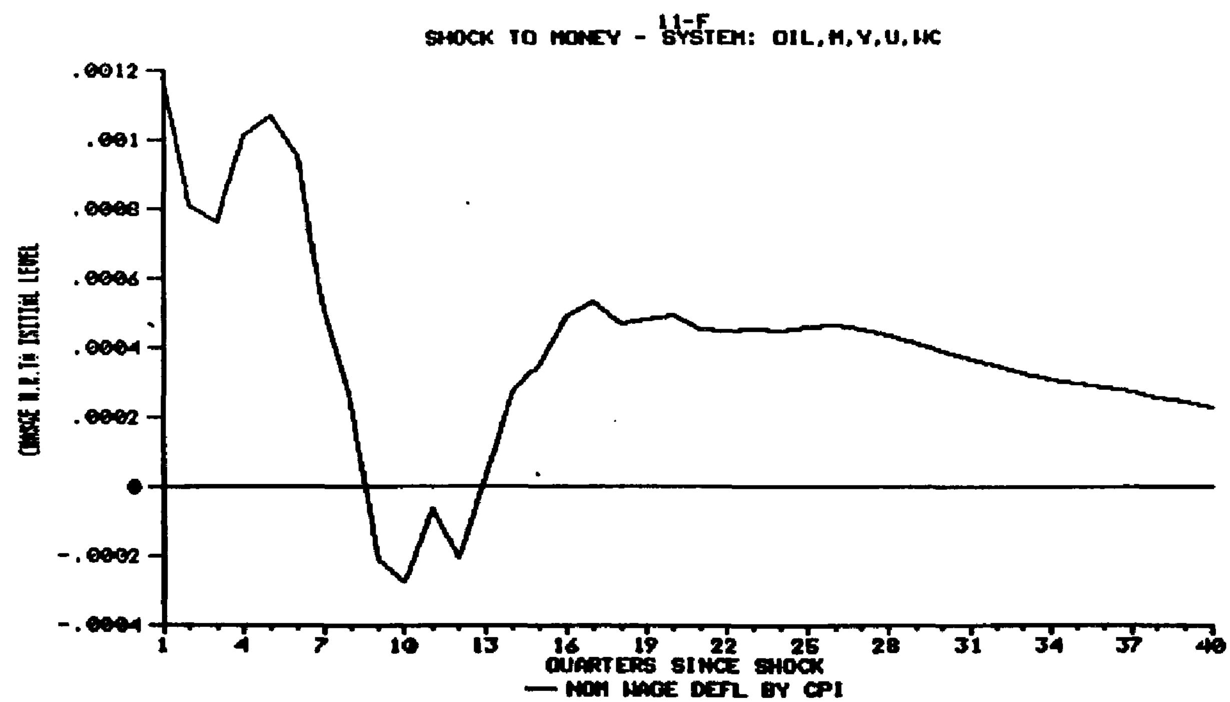

11-F
SHOCK TO MONEY - SYSTEM: OIL,M,Y,U,WC
CHANGE W.R.TO INITIAL LEVEL
.0012
.001
.0008
.0006
.0004
.0002
0
-.0002
-.0004
1
4
7
10
13
16
19
22
25
28
31
34
37
40
QUARTERS SINCE SHOCK
— NOM WAGE DEFL BY CPI

FIGURE 12-A
SHOCK TO PROD - SYSTEM: PROD,M,Y,U,WP

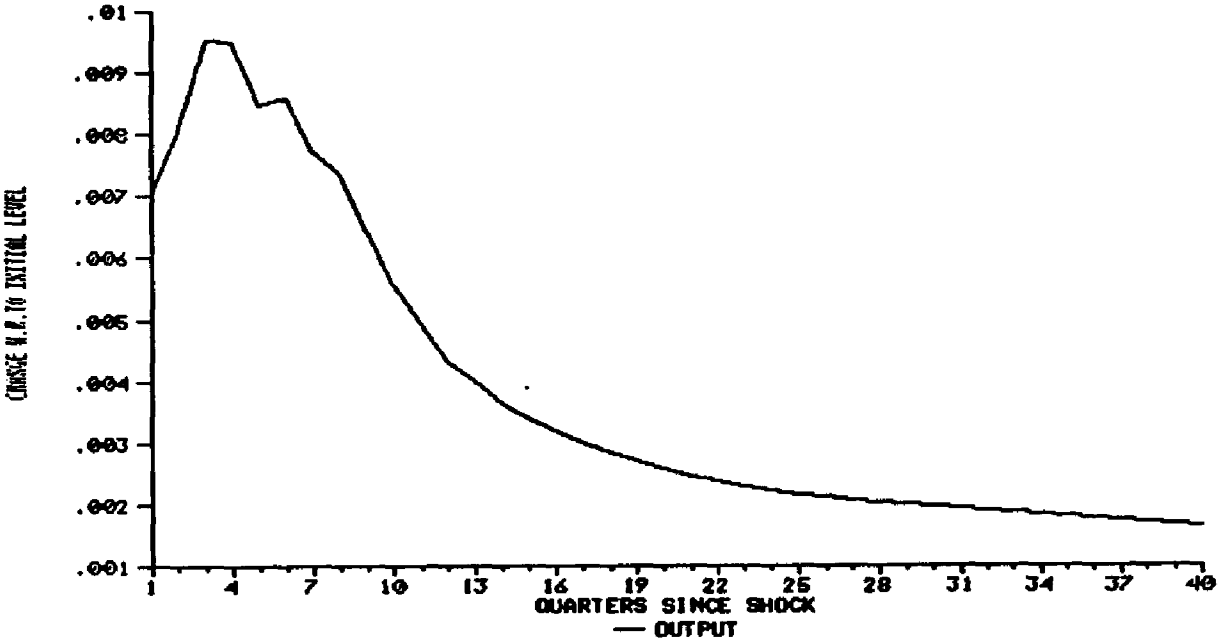


FIGURE 12-B
SHOCK TO PROD - SYSTEM: PROD,M,Y,U,WP

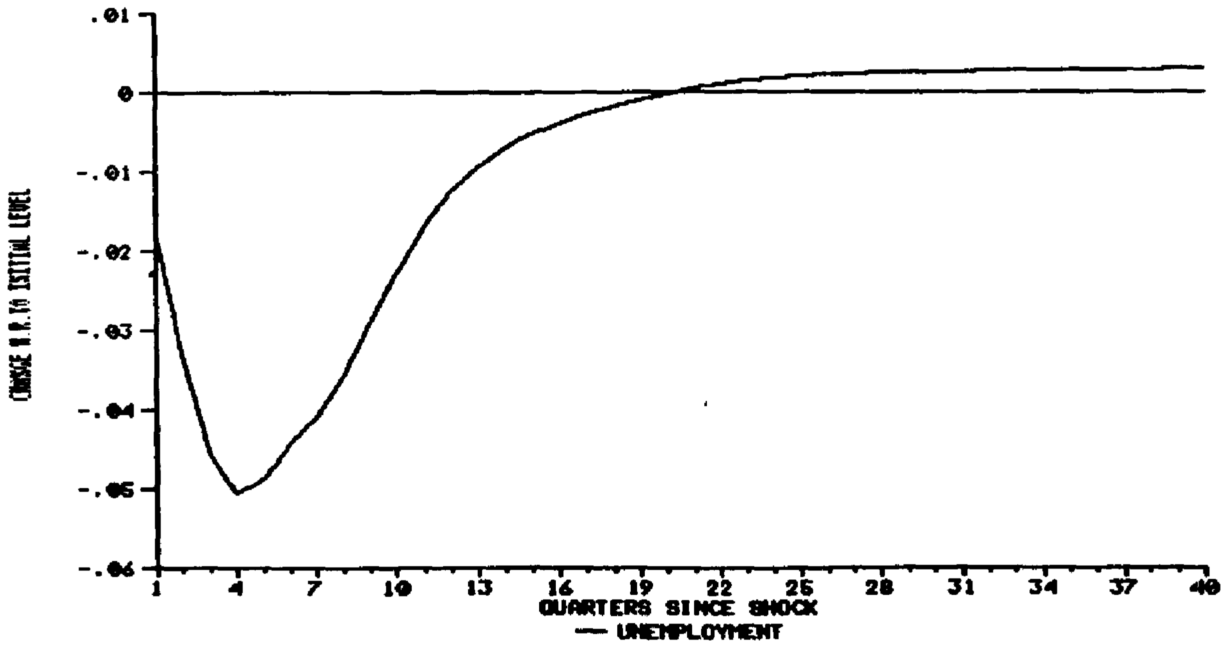


FIGURE 12-C
SHOCK TO PROD - SYSTEM: PROD,M,Y,U,WP

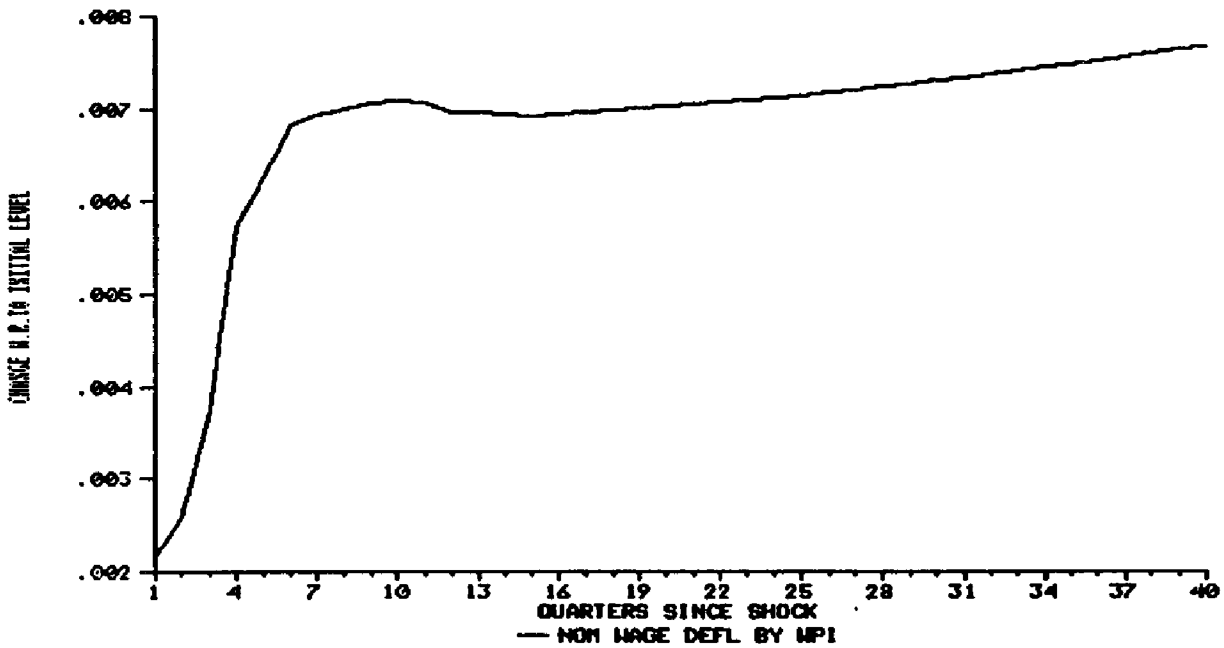


FIGURE 12-D
SHOCK TO MONEY - SYSTEM: PROD,M,Y,U,WP

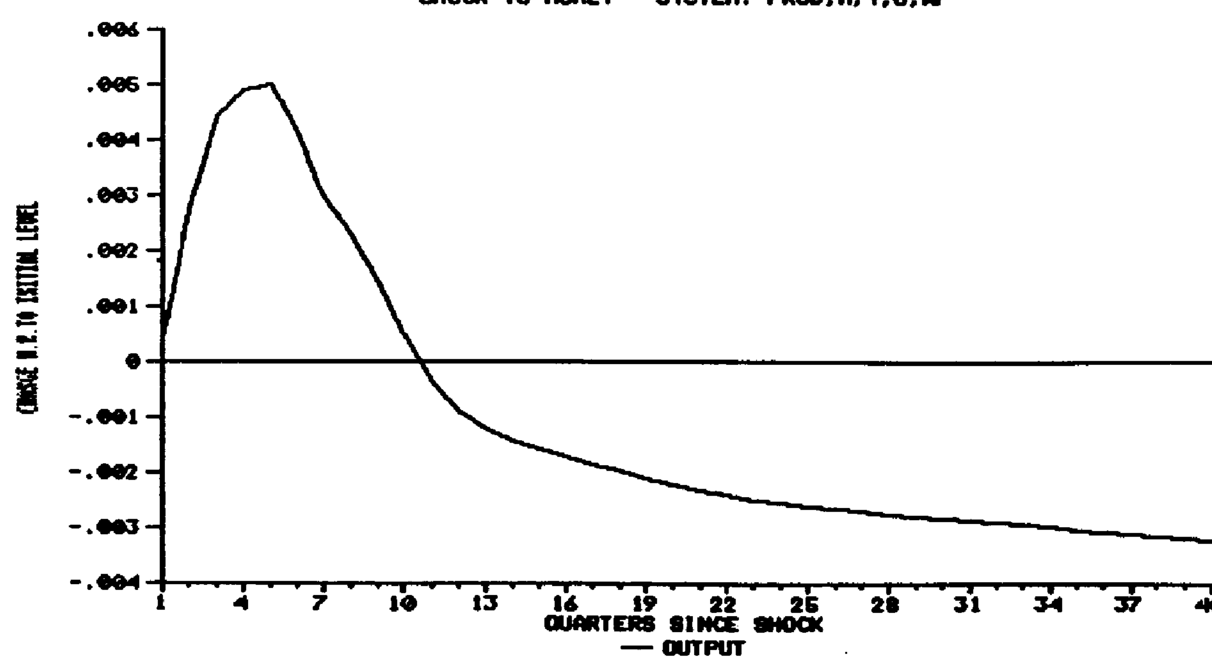


FIGURE 12-E
SHOCK TO MONEY - SYSTEM: PROD,M,Y,U,WP

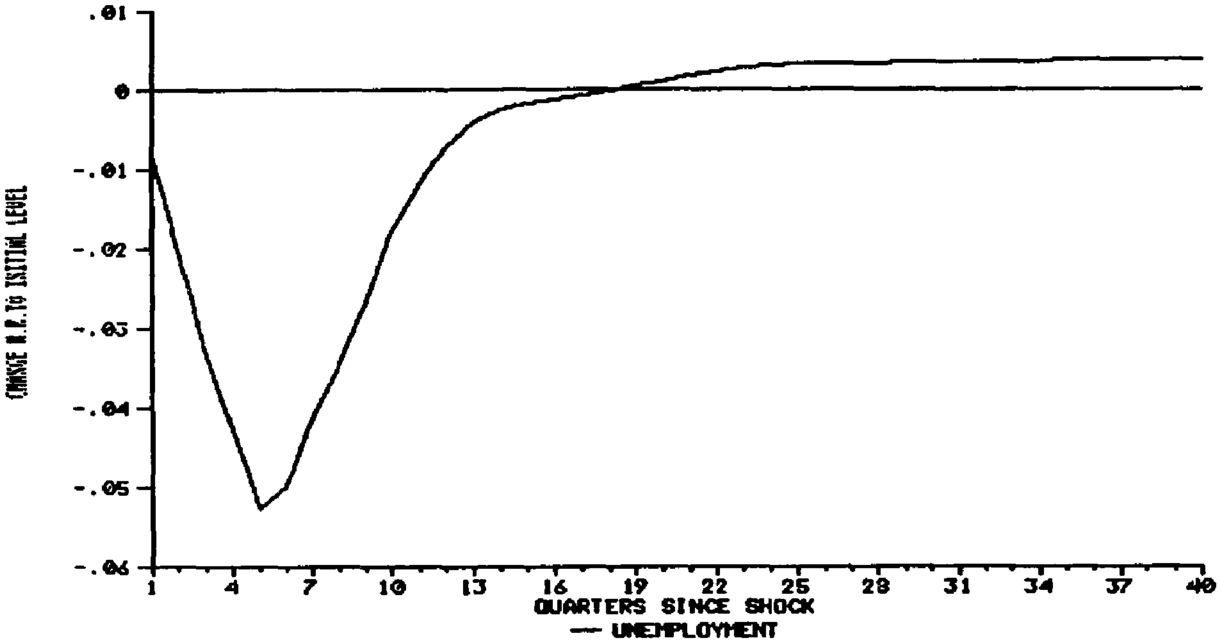


FIGURE 12-F
SHOCK TO MONEY - SYSTEM: PROD,M,Y,U,WP

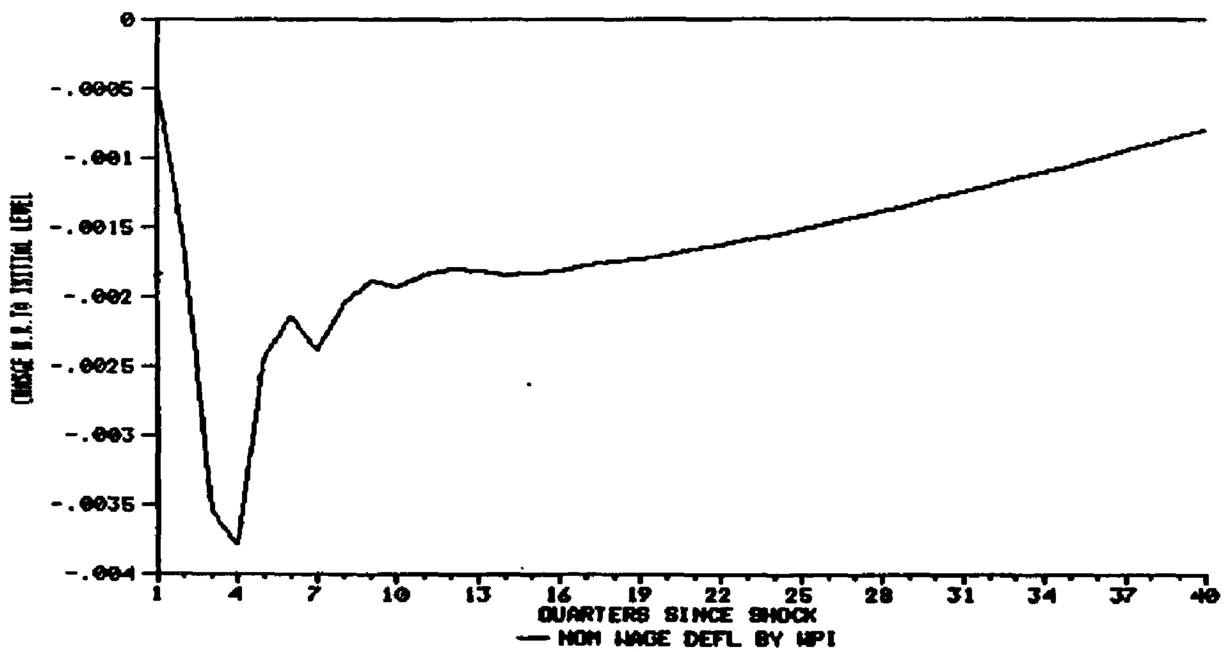


FIGURE 13-A
SHOCK TO PROD - SYSTEM: PROD,M,Y,U,WC

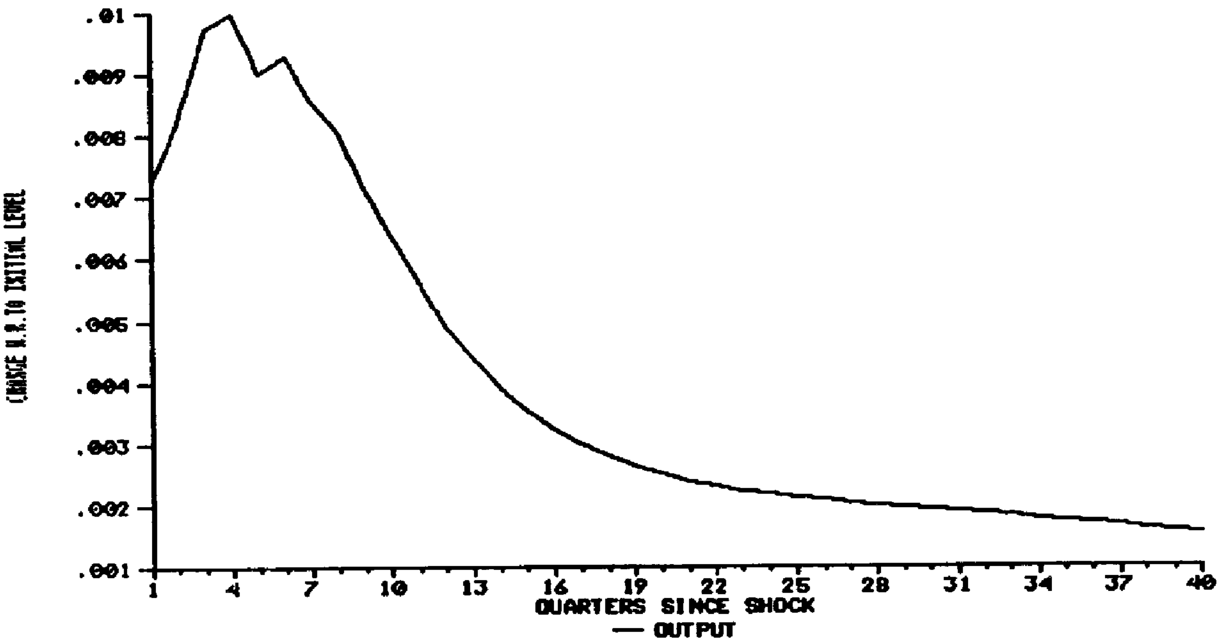
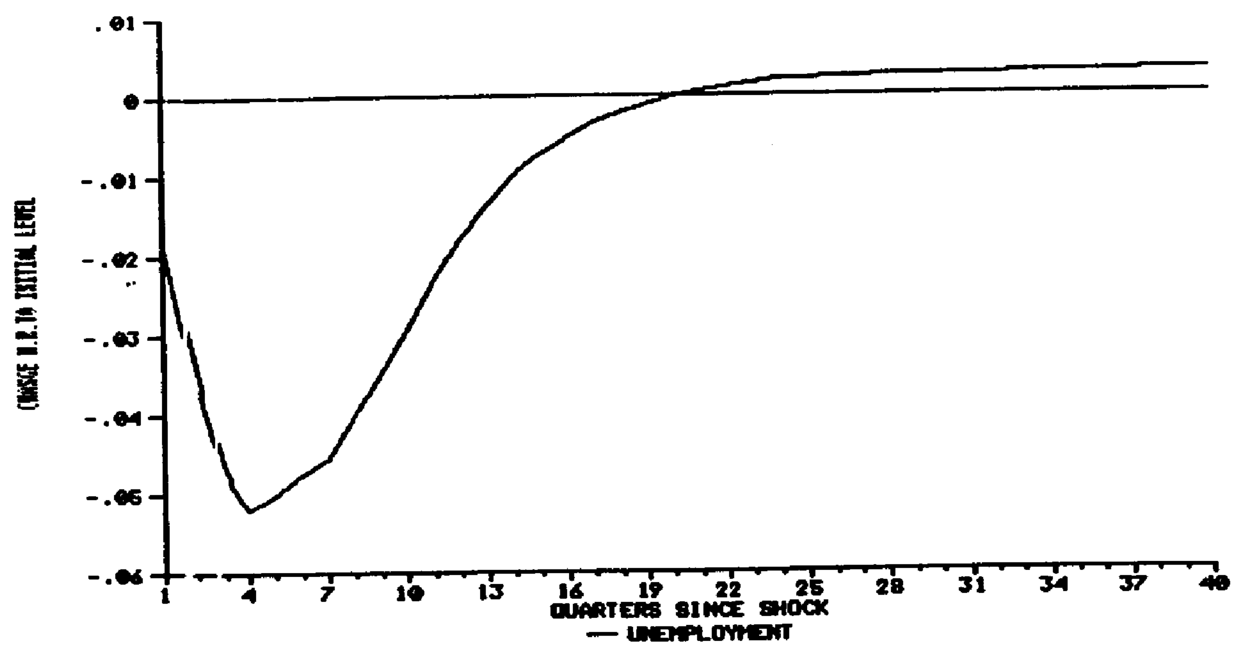


FIGURE 13-B
SHOCK TO PROD - SYSTEM: PROD,M,Y,U,WC



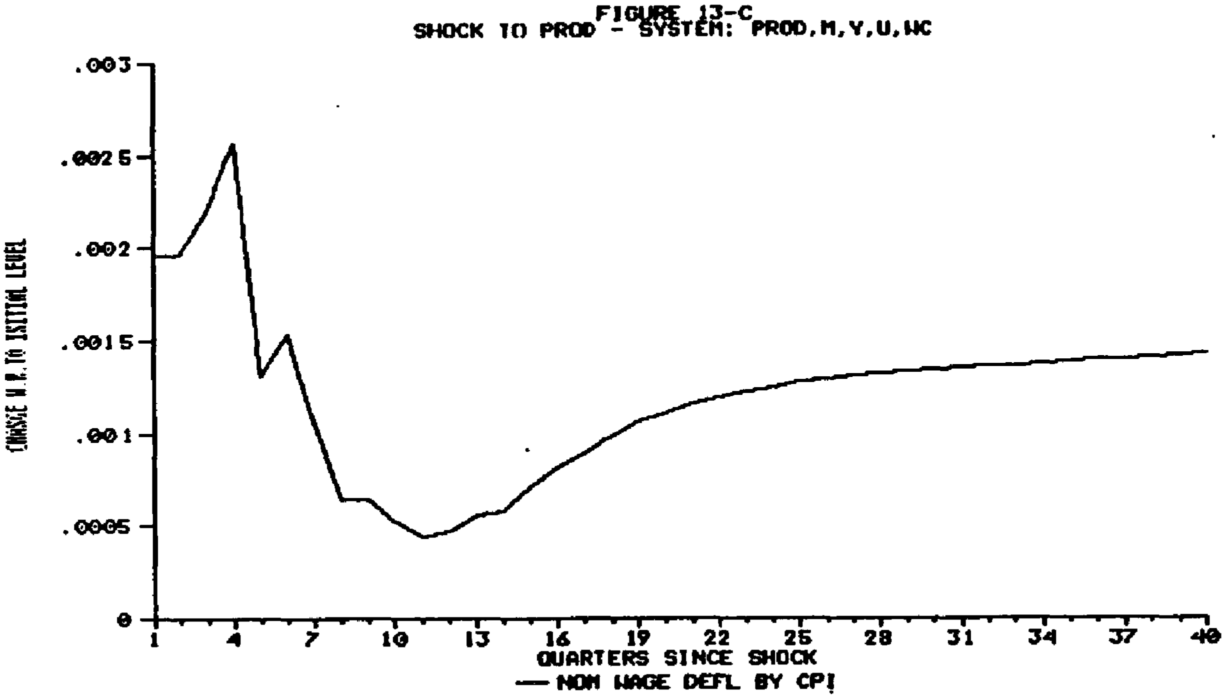


FIGURE 13-C
SHOCK TO PROD - SYSTEM: PROD,M,Y,U,WC
CHANGE W.R.TO INITIAL LEVEL
.003
.0025
.002
.0015
.001
.0005
0
1
4
7
10
13
16
19
22
25
28
31
34
37
40
QUARTERS SINCE SHOCK
— NOM WAGE DEFL BY CPI

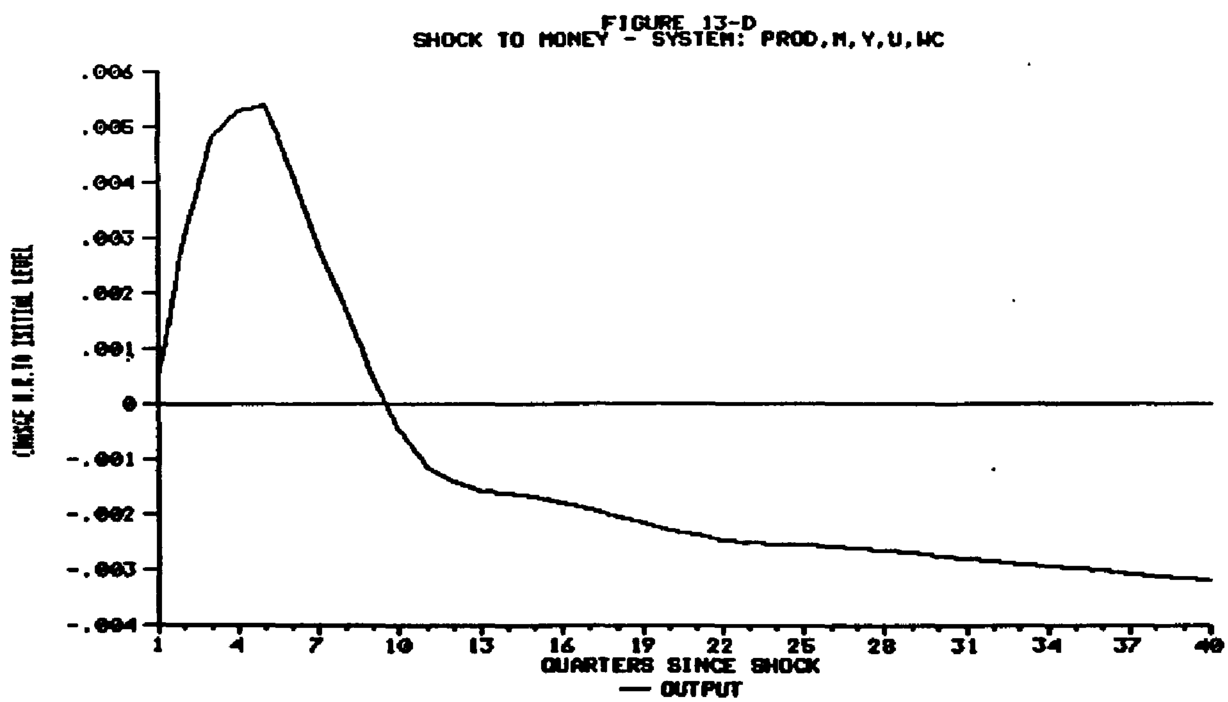


FIGURE 13-D
SHOCK TO MONEY - SYSTEM: PROD,M,Y,U,WC
CHANGE W.R.TO INITIAL LEVEL
.006
.005
.004
.003
.002
.001
0
-.001
-.002
-.003
-.004
1
4
7
10
13
16
19
22
25
28
31
34
37
40
QUARTERS SINCE SHOCK
— OUTPUT

FIGURE 13-E
SHOCK TO MONEY - SYSTEM: PROD,M,Y,U,WC

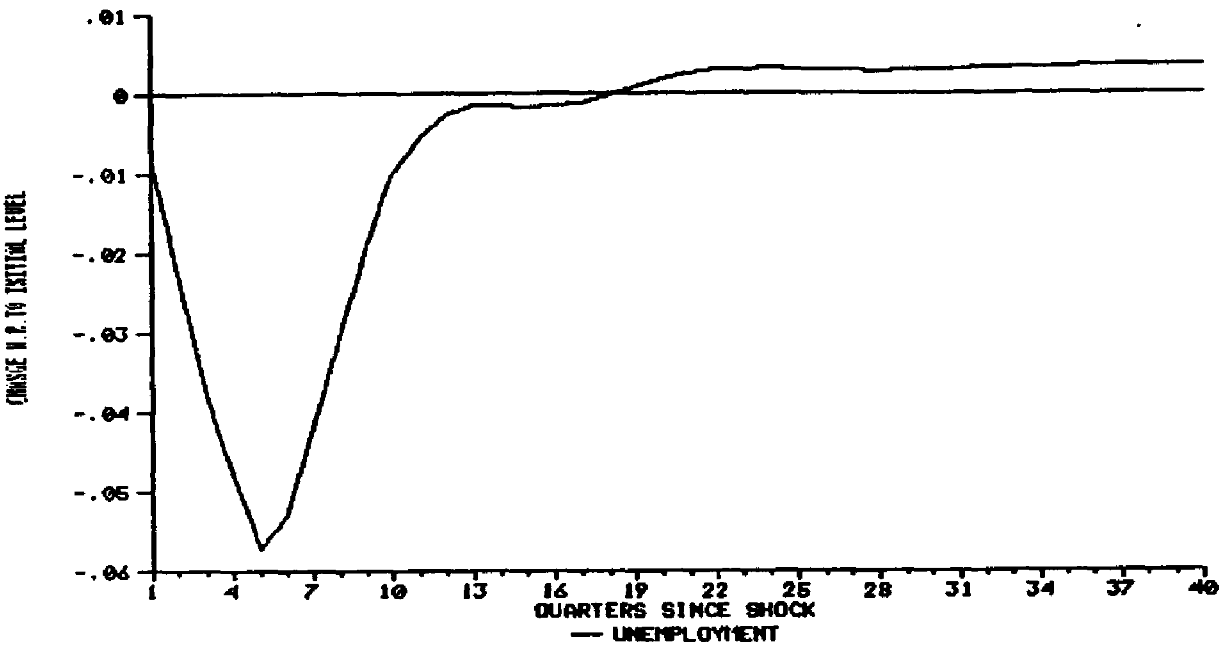
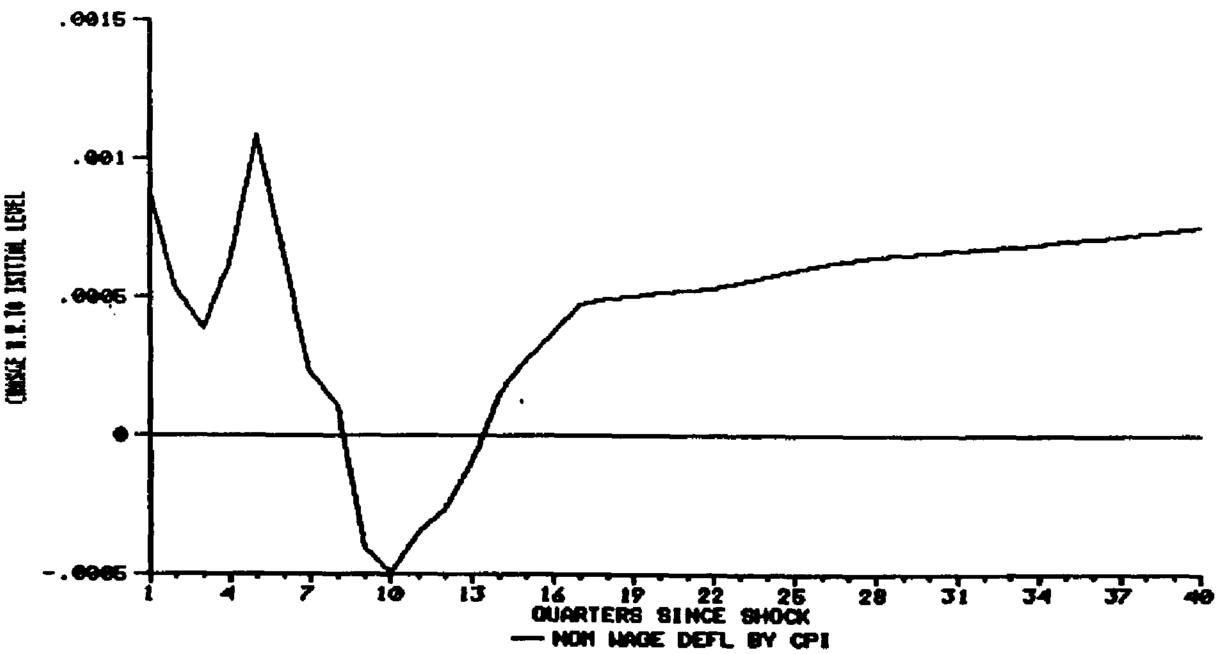


FIGURE 13-F
SHOCK TO MONEY - SYSTEM: PROD,M,Y,U,WC



THE IMPACT OF A BAN ON LEGALIZED ABORTION ON ADOLESCENT CHILDBEARING IN NEW YORK CITY

I. INTRODUCTION

There is great speculation in the popular press that the U.S. Supreme Court will overturn the 1973 decision in Roe versus Wade, the case which legalized abortion across the United States. A possible outcome of such a reversal is that the authority to regulate abortion will be given to individual states. Under this scenario, even states where the sentiment towards abortion is favorable may experience a period in which legalized abortion is unavailable while new legislation is written and debated. A ban on legalized abortion, or even an interruption, could have a profound effect on the incidence of unwanted childbearing.

The objective of this paper is to examine the probable changes in teenage childbearing among New York City residents following a ban on legalized abortion. To this end, we first fit a model that estimates the change in the number of births to adolescents following the 1970 New York State law which liberalized abortion. We argue that the percentage decline in teenage childbearing between 1970 and 1971 is a good approximation, but in reverse, to what would occur today if legalized abortion were no longer available.

A number of studies have noted the decline in births after the New York State Law which liberalized abortion became effective (Pakter et al. 1973, Tietze 1973, Kramer 1975). However, the results in each of the studies were based on annual changes over a very short time span. Such unrefined estimates provide only a crude understanding of what might occur if the legalization of abortion were reversed. Pooled time-series, cross-sectional studies have examined the effect of liberalized abortion laws on annual, age-specific fertility rates within and across states (Sklar and Berkov 1974, Bauman et al. 1977, Quick 1978). The findings suggest that the legalization of abortion had an important impact on fertility. Again, the variation over time was limited to at most seven years.

This study differs substantially in that it is a time-series analysis with monthly data that spans 25 years. The large number of observations allows for a more sophisticated means of fitting the data. For example, we use the Box-Jenkins time-series methodology to predict the monthly number of births to New York City adolescents that would have been observed had abortion not been liberalized in 1970 (Box and Jenkins 1976). Comparing these estimates to the actual number of births between 1970 and 1971 yields a first approximation of the number of births that were averted by the change in the law. The estimates are refined by means of intervention analysis: a technique for determining the magnitude, form, and statistical significance of the change in a time-series following a major external event (Mc Cleary and Hay 1980, Vandaele 1983). Finally, we use the model obtained by the intervention analysis to forecast the time-path of teenage births following a ban on legalized abortion.

We have limited our analysis to adolescents for several reasons. First, the medical, social and economic consequences of teenage pregnancy are staggering (National Research Council 1987a,b). Second, the proportion of teenagers who become pregnant has changed relatively little over the past 15 years. Approximately one out of ten adolescents aged 15-19 becomes pregnant every year and in 1988 about 86 percent of the pregnancies were unintended (Jones et al. 1988). Third, adolescents are disproportionate users of abortion and thus, would be more affected by an overturn of Roe versus Wade. Nationally, 30 percent of all abortions performed in the United States are to teenagers and 40 percent of all adolescent pregnancies are terminated by induced abortion (National Research Council 1987a).

The analysis was restricted to New York City because we needed an area in which the availability of legalized abortion would be as radically altered by a ban as it had been by the liberalization of abortion laws. The 1970 New York State abortion statute had the fewest restrictions of any state law in the country. Before July, 1970 there had been few legal abortions performed in New York City (Erhart et al. 1972).More importantly, prior to the change in 1970, pregnant adolescents had essentially no access to legal abortion. Although Hawaii, Alaska and Washington state had laws similar to New York's by the end of 1970, each of the states had residency requirements. Consequently, the magnitude of the change in adolescent childbearing among New York City residents after the 1970 law became effective was not diminished by migration to other states. This was not true for residents of other states after the passage of the New York State law. Between July 1970, and June 1971, 75.4 percent of the 33,964 abortions performed on adolescents in New York City were to out-of-state residents (Pakter et al. 1973). The upshot is that New York City offers a unique setting from which to understand the changes in adolescent childbearing that followed the liberalization of abortion, and what is likely to occur if legal abortion is banned.

II. METHODS

DATA

Monthly figures on the number of black and white live births to New York City residents less than 20 years of age are from vital statistics maintained by the New York City Department of Health. Each year of individual birth records has been aggregated by month for blacks and whites separately. The number of birth records with unknown age in any one year was less than .03 percent. These records were deleted from the aggregation. The final series consisted of 300 monthly observations for whites and blacks from January, 1963 to December, 1987. Plots of the race-specific series are shown in Figures 14 and 15.

The analysis was limited to whites and blacks because ethnicity was not identified on New York City birth certificates until 1978. However, a substantial proportion of the adolescents who are white are of Hispanic origin or descent. In 1984, the first year in which data on births to Hispanic women were published, 75 percent of the white adolescents who gave birth were of Hispanic origin (NYC Department of Health 1985). How this proportion has varied over time is unknown. Data from the 1970 Census indicates that 13.4 percent of all women 15 to 19 years of age were of Puerto Rican descent (Bureau of Census 1973). The number of Hispanic adolescents that were not Puerto Rican is unknown. By 1980, the percent of Puerto Rican adolescents 15 to 19 years of age had risen to 17.0 percent while the total proportion of Hispanic adolescents of the same age stood at 25.9 percent (Bureau of Census 1984). The potential impact on the results of the apparent shifting ethnic composition of white adolescents who gave birth is discussed below.

The number of births as opposed to birth rates were analyzed because monthly population figures for New York City over the period under study were unavailable. Census data could have been used to estimate monthly population figures between the census years, but such crude estimates would only have introduced measurement error. Furthermore, month-to-month changes in the population are minor compared to the 10 to 20 percent drop in the number of adolescent births that were observed between 1970 and 1971.

STATISTICAL ANALYSIS

The data were analyzed by means of an Autoregressive Integrated Moving Average (ARIMA) model. The methodology is often referred to as the Box-Jenkins approach to time-series modeling (Box and Jenkins 1976). The objective is to use the observed data to describe the underlying time-series process in as concise a manner as possible.

The first step is to decompose the stationary portion of the series into its autoregressive and moving average parts. Once a tentative representation has been established, the parameters of the model are estimated and then diagnosed to insure that the unexplained portion of the model, the residuals, are a random or white noise process. If the checks prove satisfactory, then forecasts can be generated.

A time series may undergo a dramatic change due to an external event that alters the underlying behavioral process generating the series. Such interventions can be incorporated into the Box-Jenkins methodology in order to evaluate the form and the magnitude of the change. In particular, a binary variable is used to capture the impact of the intervention. The variable equals zero prior to the change and one thereafter. The binary variable is appended to the specification and a t-test is used to determine whether the estimated coefficient on the binary variable is statistically significant. The methodology has become known as intervention analysis and has been widely applied in the social sciences (McCleary and Hay 1980, Vandaele 1983).

To successfully apply the intervention analysis, it is necessary to know the starting point of the event as well as the general shape of the response of the series to the event. The hypothesis maintained in this study is that the 1970 New York State Law which liberalized abortion had an important impact on the number of adolescent pregnancies that resulted in live births. The law became effective on July, 1 1970. However, because the law did not apply to pregnancies greater than 24 weeks, the effect of the law on the number of live births would not be observed for at least 16 to 20 weeks later. Thus, November, 1970 became the starting point of the intervention. Furthermore, the full impact of the law on adolescents births would not be possible until April of 1971 when the pregnancies of the first cohort of adolescents who conceived on or after July 1, 1970 reached term. Consequently, the intervention variable was specified in such a manner that the law's impact on the number of adolescent births increased gradually from November, 1970 through April, 1971.

The rate at which the law's impact grew between November and April was based on the distribution abortions to New York City residents the first year the law was in effect. For example, 6.1 percent of all abortions to New York City residents were to women whose pregnancies were beyond the twentieth week (Tietze 1973). Assuming the distribution of abortion by gestational age was the same for every month in the first year, the proportion of the law's full impact that would be felt in November was .061. Thirty percent of all abortions performed in the first year were to women whose pregnancies were between 13 and 20 weeks gestation. We assumed that 15 percent were performed between 13 and 16 weeks gestation, and the other 15 percent were performed between 17 and 20 weeks gestation. Thus, in December, the proportion of the law's full impact would be .216. This accounts for the 6.1 percent of the women who aborted in August, 1970 whose pregnancies were greater than 20 weeks gestation and for the 15 percent of the abortions in July, 1970 to women whose pregnancies were between 17 and 20 weeks gestation. Following this algorithm and noting that 64 percent of all abortion were performed in the first trimester, the figures for the remaining months were as follows: .361 in January, .574 in February, .785 in March, and 1.0 in April and all months thereafter.

To estimate the changes in adolescent childbearing following a ban on legalized abortion effective January 1, 1988, we first forecasted the number of births in 1988 and 1989 under the assumption that abortion remained legal. The complete series, including the intervention component was used to obtain the forecasts. Since the future error terms are unknown, they were set to their expected values of zero (Vandaele 1983, Granger and

Newbold 1986). We then multiplied the forecasted births by the absolute value of the parameters from the intervention model. The assumption was that the absolute value of the percentage change in adolescent births after abortion was legalized provided a good estimate of the expected increase in adolescent childbearing if abortion were to be outlawed.

Finally, although we assumed that the relative change in births resulting from a ban on abortion in 1988 would be the same (except for a sign change) as was observed in 1970 and 1971, the rate of change from the pre- to the post-intervention level of births had to be altered. The modification was necessary because the distribution of abortions by gestational age had changed substantially from what was observed in 1970 and 1971. In particular, national data from 1981 indicates that 91 percent of all abortions were obtained in the first trimester, approximately 8 percent were obtained between 13 and 20 weeks, and fewer than one percent were at more than 20 weeks (Henshaw et al. 1985). Using the same algorithm as above, a ban on abortion that became effective January 1, 1988, would realize one percent of its full potential impact in May of 1988, 5 percent in June, 9 percent in July, 40 percent in August, 70 percent in September, and 100 percent by October.

III. RESULTS

Figures 14 and 15 present the monthly number of births to black and white New York City adolescents from January, 1963 through December, 1987. For blacks, the reversal of a seven-year upward trend which occurs between 1970 and 1971 is dramatic. In the case of whites, a relatively stationary series up to 1970 falls substantially between 1970 and 1971 and then continues to trend downward until approximately 1986. Both figures suggest a major alteration in adolescent childbearing that is coincident with New York State's liberalized abortion law which became effective July, 1 1970.

To understand the impact the legislation had on adolescent childbearing, we estimated the ARIMA structure for the pre-intervention series (January, 1963 - June, 1970). The data are expressed as natural logarithms in order to control for non-stationarity in the variance. The sample autocorrelations of the undifferenced series for blacks(B_t) and Whites(W_t) died out only slowly at high lags, suggesting nonstationary behavior. For the series of first differences $(1-B)B_t$ and $(1-B)W_t$, we found high sample autocorrelations at lags that are multiples of 12 which died out slowly. These considerations suggested examination of the first and twelfth order differenced series for blacks and whites [$(1-B)(1-B^{12})B_t$ and $(1-B)(1-B^{12})W_t$]. Based on the sample autocorrelation and partial autocorrelation functions of the transformed series for black and white births, which are displayed in Tables 7 and 8, and Figures 16A to 17B, the models

$$(1-B)(1-B^{12})B_t=(1-\sigma B)(1-\emptyset B^{12})e_t \quad \text{and}$$
$$(1-B)(1-B^{12})W_t=(1-\sigma B)(1-\emptyset B^{12})e_t$$

are decided to be carried forward for estimation. The coefficients of the models are displayed in Table 9.

If the models adequately depict the ARIMA processes governing the series, then the errors of the models should be white noise. The Q-statistics in Table 9 indicate that the residuals from the estimated models are white noise processes. Another approach for determining if the errors are white noise is to evaluate the autocorrelations of the first-differenced residuals. If the errors are a white noise process, then their first difference should follow an MA(1) process with the moving average parameter equal to 1, and the first autocorrelation equal to -.5. For the estimated models in Table 1, the first differenced residuals indicated an MA(1) process. For blacks the MA(1) coefficient was .91 with a t-ratio of 18.78, and the first autocorrelation was -.53. For whites, the MA(1) coefficient was .95 with a t-ratio of 25.81, and the first autocorrelation was -.50. Thus, the analysis of the errors supports the appropriateness of the specifications.

Based on the ARIMA specifications in Table 9, we forecasted the number of race-specific adolescent births 24 months beyond June, 1970. Comparison of the forecasted births with the actual number of births over this 24 month period yields a first approximation of the number of unintended births to New York City adolescents that were averted by the legalization of abortion. The actual and forecasted births for the two races are presented in Figures 5 and 6. Subtracting the actual births from the forecasted births and summing over the 24 months indicates that 4091 black births and 3128 white births were averted by the availability of abortion.

To more formally test whether the 1970 law liberalizing abortion may have caused the precipitous drop in adolescent births, we used all the data to estimate the ARIMA structure of each series. As outlined in the previous section, an additional variable was added to the specification to control for the impact of the law. The results are shown in Table 10. Except for the intervention component, the ARIMA structure is unchanged for blacks. In the case of whites, a second-order seasonal moving average component improved the model's fit. The coefficient on the intervention variable, α, is statistically significant for blacks and whites. Thus, the data reveal that the decline in the level of births after October, 1970 was a change that could not be explained by the normal variation in the series. The Q-statistics indicate the adequacy of the models. Similarly, for both models the autocorrelation and the partial autocorrelation functions of the first differenced residuals demonstrated that they were governed by a MA(1) process. For blacks, the first autocorrelation of the differenced residuals was -.50, and the MA(1) coefficient was .95 with a t-ratio of 51.28. For whites, the first autocorrelation was -.53 and the MA(1) coefficient was .97 with a t-ratio of 70.45.

The magnitude of the change from the pre- to post intervention level of the series can be obtained by exponentiating the coefficient of the intervention variable α and subtracting it from one. Expressed as a percentage, the level of black adolescent births fell 18.7 percent after the liberalization of abortion in July 1970. White births fell 14.0 percent. By using the estimated percentage changes and taking into account the gradual transition between November 1970 and April 1971, one can calculate the total number of adolescent births that were averted in the 24 month period after July 1970. If we assume the average number of births over twelve months prior to July, 1970 is an estimate of the pre-intervention level of births, then 2588 black adolescent births and 1998 white adolescent births were averted in the 24 months after July, 1970. These figures are smaller than the ones obtained from the difference of the forecasted and the actual births over the two year period after July 1970 since they do not take into account the seasonality and trend movements inherent in the data prior to 1970.

To estimate how a ban on abortion effective January 1, 1988 would affect adolescent childbearing in New York City, we forecasted the monthly number of births to black and white teenagers based on the ARIMA specifications in Table 10. Beginning in May, 1988 we multiplied the predicted number of monthly births by .187 in the case of blacks and .140 in the case of whites. The products for May, 1988 through September 1988, were further adjusted to account for the distribution of abortion by gestational age. Summing the adjusted monthly forecasts from May, 1988 through December, 1989, we estimate that if legal abortion were banned January 1, 1988 there would have been 2143 additional black births and 1067 white births to adolescents in 1988 and 1989 above what would have been expected had the laws regarding abortion remained unchanged.

IV. DISCUSSION

Using monthly data on the number of white and black births to New York City adolescents, we found that the liberalization of the New York State abortion law in 1970 had a substantial and statistically significant impact on adolescent childbearing. In particular, we estimate that at least 2500 unintended births to black teenagers and at least 1900 unintended births to white teenagers were averted in the first two years after the legislation became effective. We then estimated the probable impact on adolescent childbearing among New York City residents if legal abortion were unavailable nation-wide beginning January 1, 1988. Our model predicts that there would be over 2000 black and 1000 white unintended births to New York City teenagers in the first two years following the ban.

The forecasts are based on modeling the change in teenage childbearing that followed the liberalization of abortion laws in New York State in 1970. The plausibility of our forecasts depend on a number of assumptions that must be defended explicitly. First, we assume that if Roe versus Wade were overturned and abortion were banned in New York State, then legalized abortion would not be available to New York State residents in any other state. Although such a restrictive outcome is unlikely, the forecasts remain instructive for several reasons. For one, they provide upper-bound estimates based on the most restrictive scenario possible, a nation-wide ban on legalized abortion. This scenario cannot be dismissed lightly given that the President of the United States, the U. S. Attorney General, and the Head of the Department of Health and Human Services have all publicly stated their opposition to Roe versus Wade. A more likely scenario is that U.S. Supreme Court will return the regulation of abortion to individual states. If so, then New York would probably be one of the first states to respond in a "pro-choice" manner. New York is one of only 13 states plus the District of Columbia that funds abortions to Medicaid eligible women and only 1 of 8 states that does so voluntarily (Gold and Guardado 1988). Nevertheless, New York's 1970 law passed the State Senate by five votes and the State Assembly by only one. The debate in both houses was tumultuous (Lader 1973). Without dissecting the present political situation in the state, if Roe versus Wade is overturned, there could be a substantial delay and serious confusion regarding the availability of legal abortion while new legislation is written and debated. Border states such as Connecticut and New Jersey have been less supportive of abortion than has New York, and may be less likely to respond as quickly and as liberally as New

York. Moreover, many teenagers, especially blacks and Hispanics, lack the resources to travel beyond the tri-state area to obtain an abortion. Thus, legal abortion in states like Massachusetts, Michigan, and California may represent inaccessible alternatives to pregnant adolescents in New York City. At the very least, therefore, the overturn of Roe versus Wade could seriously limit, if only temporarily, the options available to pregnant adolescents in New York City.

A second assumption upon which our estimates are based is that the proportion of teenagers in New York City at risk of an unintended pregnancy today is similar to the proportion at risk in 1970. National data indicates that the pregnancy rate among women 15 to 19 years of age has risen from 94 pregnancies per 1000 adolescents in 1972 to 109 pregnancies per thousand in 1984 (National Research Council 1987a). In addition, the percentage of teenagers residing in metropolitan areas who describe their pregnancies as unwanted rose between 1971 and 1979 (Zelnik and Kantner 1980). Consequently, the proportion of teens at risk of an unintended pregnancy has probably risen since 1970 despite the increased use of contraception among adolescents. In this respect, our projections may underestimate the impact of a ban on legalized abortion.

One explanation for the rising pregnancy rate is that the availability of abortion has engendered less effective contraceptive behavior; the more extreme version is that abortion serves as an alternative method of fertility control. Thus, if abortion is banned, the pregnancy rate may fall. There is no evidence to support either explanation. A recent survey of abortion patients reports that less than one percent of the women who used no contraception admitted doing so because they relied on abortion (Henshaw and Silverman 1988). And although little is known about the relationship between abortion availability and the effective use of contraceptives, a teenager who aborted a pregnancy was less likely to become pregnant again over the next 24 months than was a comparable adolescent who carried her first pregnancy to term (Koenig and Zelnik 1982).

As mentioned above, there has been an apparent rise in the proportion of births to whites of Hispanic origin in New York City. Based on the little data that exists, Hispanic adolescents are more sexually active than their white, non-Hispanic counterparts. At the same time, Hispanic adolescents are more likely to be married and less likely to abort (Hayes 1987, Joyce 1988). The upshot, therefore, is that the suspected increase in the proportion of births to white adolescents of Hispanic descent is unlikely to have altered our predictions in any meaningful manner from what they would have been had the proportion of white births of Hispanic descent remained unchanged since 1970.

With respect to the statistical analysis we imposed a number of restrictions that should also be made explicit. For example, instead of fixing the rate at which the law's impact grew between November 1970 and April 1971, an alternative specification would have been to model the intervention component as a first-order transfer function (McCleary and Hay 1980). This allows the data to determine the rate of change between the pre- and post-intervention series. The distribution of pregnancies by gestational age explains in large part why initially, the impact of the law was gradual. But the availability of abortion services in the early months under the new law may also have delayed the full impact of the law from being realized. There was considerable confusion in New York City with respect to the legality of performing abortions in doctor's offices and free-standing clinics (Lader 1973). For instance, the proportion of all abortions performed in free-standing clinics or physicians' offices rose from 45 percent in the first twelve months of the law to 62 percent in the ensuing year (Tietze 1973). Moreover, abortions performed in hospitals were at least twice as expensive as the ones performed in free-standing clinics or physicians' offices. Teenagers, especially minorities, were probably more affected by the price and non-price barriers since they were less likely to have access to private gynecological care. In sum, the step function we imposed assumes the law became fully effective in April, 1971, yet other factors may have further delayed the new law from realizing its full impact.

To examine this proposition, we re-estimated the effect of the 1970 abortion law by using a first-order transfer function model. The results for blacks were essentially unchanged from the estimates reported above. However, the first-order transfer function for whites indicated that the change in the level of the series between the pre- and post-intervention data declined without end. However, there is little evidence to support the hypothesis that the reform of New York's abortion laws initiated a continuous decline in the number of white adolescent births. The more likely explanation is that the white adolescent population in New York City declined. Census data indicates that the white female population 15 to 19 years of age in New York City (including Hispanics) fell from 219,834 in 1970 to 138,023 in 1980, an annual rate of decline of 4.8 percent. As a point of comparison, the black adolescent population of the same age grew from 77,174 to 95,262, an annual growth rate of 2.1 percent (Bureau of Census 1973, 1984). These figures are clearly insufficient to explain the 14 and 18 percent drop in the monthly level of white and black births respectively following the new abortion law. However, the decreasing size of the white teenage cohort is a reasonable explanation for the smooth decline in the number of white births after 1971 (see Figure 1).

In sum, the step function based on the distribution of abortions by gestational age is an appropriate alternative to the first-order transfer function. An additional advantage of the step function is that the conditions in 1970 with respect to the gestational age distribution of abortions and the availability of abortion services are notably different today. The step function allows us to adjust the rate at which a ban on abortion would impact on childbearing based on present distribution of abortions by gestational age. The parameters from the intervention model based on the first-order transfer function would have underestimated this rate of change.

There is voluminous literature on the social and economic consequences of adolescent childbearing. As the most recent and comprehensive review makes clear (National Research Council 1987a,b), adolescents who become parents will complete less schooling, have lower wages, experience greater marital instability, and be more dependent on welfare programs than their adolescent peers who delay childbearing. Moreover, the children of teenage mothers will experience greater health, cognitive, and socioemotional difficulties.

Nor would a ban on legalized abortion be costless to taxpayers. In 1986, 64.6 percent of all adolescent births in New York City were funded by Medicaid. Assuming Medicaid eligibility is a good proxy for AFDC eligibility, then applying this proportion to the number of unintended births reported above indicates that 689 white and 1384 black teenage mothers and their children will receive AFDC in the two years after the ban. Based on the methodology described in Burt (1986), the present discounted cost in terms of Medicaid, AFDC, and food stamps of supporting a family headed by a teenager over a twenty-year period is 5,560 in 1985 dollars above what it would have cost to support the same teenager and her family had she delayed childbearing until after she was twenty years of age. Thus, the total marginal costs of supporting the 689 white and 1384 black births described above would be 11.5 million dollars.

The relative changes in adolescent childbearing among whites and blacks following a prohibition on legal abortion reported in this paper cannot be generalized to other parts of the country because the forecasts are based on a set of circumstances specific to New York City in 1970. However, the number of unintended pregnancies among U.S. adolescents strongly suggests that areas in which legal abortion is prohibited will experience substantial increases in the number of births to teenagers. The magnitude of the change will vary by area because of differences in the use of abortion prior to a ban, the proximity to areas where abortion remains legal, and the availability of illegal abortions.

For those opposed to legal abortion, the increase in unintended childbearing is irrelevant to the debate. Abstinence from premarital sex, it is argued, is a simple solution to unintended childbearing among adolescents. Even supporters of abortion would agree that preventing an unintended pregnancy is the most preferable strategy for averting an unintended birth. Yet, the political situation in the U.S. is such that a concerted commitment to lowering the rate of teenage pregnancy is not part of the national agenda. In this void, legalized abortion remains one of the safest and most effective means of preventing adolescent childbearing and its attendant consequences.

FIGURE 14

MONTHLY NUMBER OF BIRTHS TO BLACK ADOLESCENTS LIVING IN NEW YORK CITY
JANUARY 1963 - DECEMBER 1987

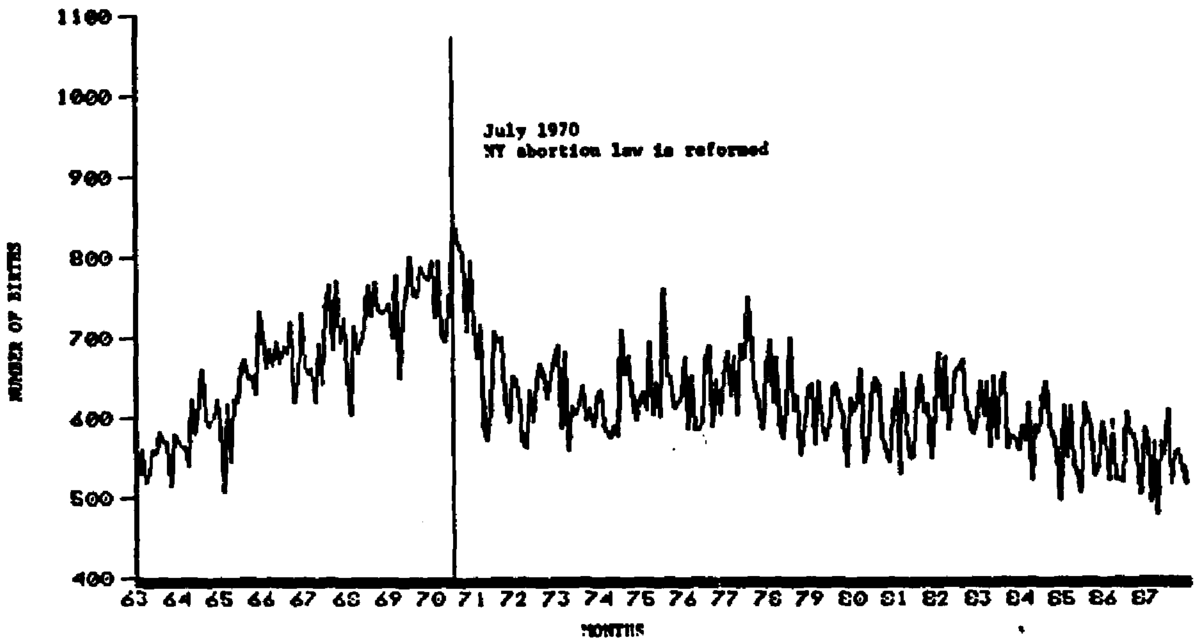


FIGURE 15

MONTHLY NUMBER OF BIRTHS TO WHITE ADOLESCENTS LIVING IN NEW YORK CITY
JANUARY 1963 - DECEMBER 1987

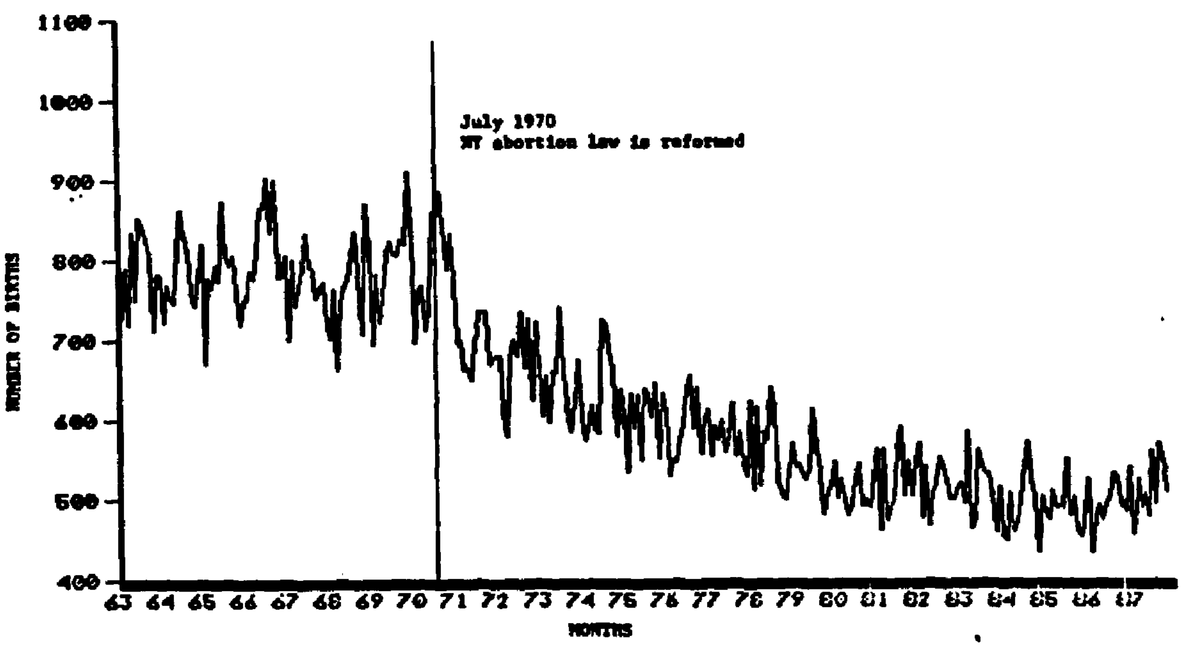


TABLE 7

THE FIRST 40 AUTOCORRELATIONS AND PARTIAL AUTOCORRELATIONS OF $(1-B)(1-B^{12})B_t$
BASED ON SAMPLE 64:2 - 70:6

AUTOCORRELATIONS

1:	-.548317	.069067	.119472	-.175787	.129636	-.034375
7:	-.147902	.225631	-.165273	-.037654	.362716	-.577422
13:	.338930	-.096570	-.007022	.049429	-.063086	.048510
19:	.059708	-.038176	.017450	.073778	-.171865	.132721
25:	-.081016	.040236	-.052282	.084361	-.025343	-.054413
31:	.073819	-.142813	.150025	-.150459	.082874	.047923
37:	-.018445	-.075960	.115288	-.114677		

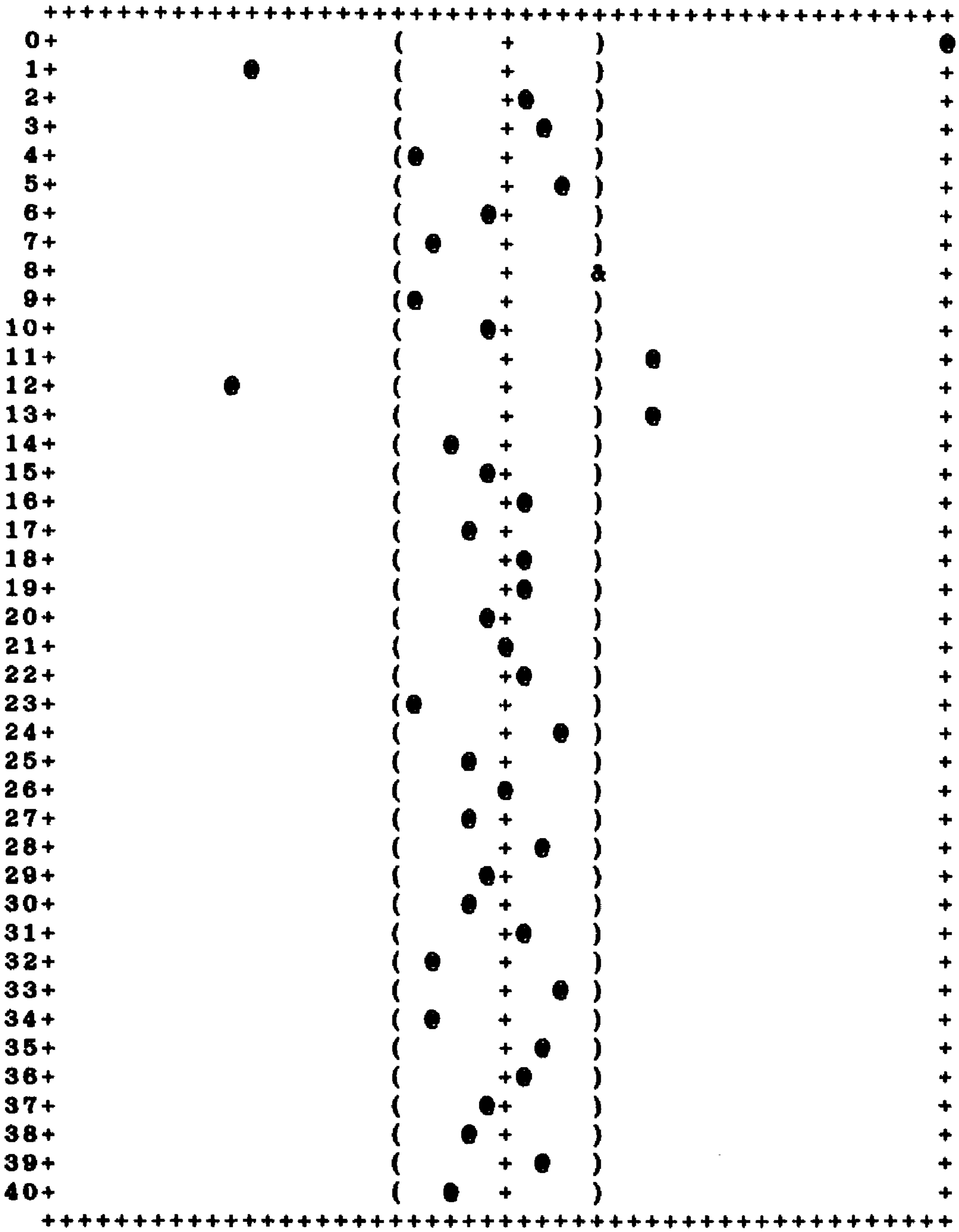
PARTIAL AUTOCORRELATIONS

1:	-.548317	-.331145	-.018773	-.119398	-.019783	.008940
7:	-.198448	.009485	-.038446	-.176976	.324979	-.314800
13:	-.205644	-.204046	-.006123	-.126191	-.055076	.002175
19:	-.123135	.180153	.028930	.078652	.161058	-.198289
25:	-.104023	-.108541	-.044226	.031055	.105742	-.073839
31:	-.004784	-.025742	.026052	-.021069	.010328	-.059834
37:	.033342	-.146720	-.131580	-.041831		

For the plots of the above data along with 95% confidence intervals see Figures 16A and 16B.

FIGURE 16A

THE PLOT OF THE FIRST 40 AUTOCORRELATIONS FOR FIRST AND TWELFTH-ORDER DIFFERENCED BLACK SERIES
1964:2- 1970:6



The parentheses denote 95% confidence intervals.
An & indicates that the value of the autocorrelation function coincides with the upper or lower confidence limit.

FIGURE 16B

THE PLOT OF THE FIRST 40 PARTIAL-AUTOCORRELATIONS
FOR FIRST- AND TWELFTH-ORDER DIFFERENCED BLACK SERIES
1964:2- 1970:6

```
  +++++++++++++++++++++++++++++++++++++++++++++++++++
 0+                  (     +     )                  ●
 1+          ●       (     +     )                  +
 2+               ●  (     +     )                  +
 3+                  (    ●+     )                  +
 4+                  (  ●  +     )                  +
 5+                  (    ●+     )                  +
 6+                  (     ●     )                  +
 7+                  (●    +     )                  +
 8+                  (     ●     )                  +
 9+                  (    ●+     )                  +
10+                  (●    +     )                  +
11+                  (     +     ) ●                +
12+                ● (     +     )                  +
13+                  &     +     )                  +
14+                  (●    +     )                  +
15+                  (    ●+     )                  +
16+                  ( ●   +     )                  +
17+                  (   ● +     )                  +
18+                  (     ●     )                  +
19+                  ( ●   +     )                  +
20+                  (     +    ●)                  +
21+                  (     ●     )                  +
22+                  (     +●    )                  +
23+                  (     +  ●  )                  +
24+                  (●    +     )                  +
25+                  (  ●  +     )                  +
26+                  (  ●  +     )                  +
27+                  (   ● +     )                  +
28+                  (     ●     )                  +
29+                  (     + ●   )                  +
30+                  (   ● +     )                  +
31+                  (    ●+     )                  +
32+                  (    ●+     )                  +
33+                  (     ●     )                  +
34+                  (    ●+     )                  +
35+                  (     ●     )                  +
36+                  (   ● +     )                  +
37+                  (     ●     )                  +
38+                  ( ●   +     )                  +
39+                  ( ●   +     )                  +
40+                  (   ● +     )                  +
  +++++++++++++++++++++++++++++++++++++++++++++++++++
```

The parenthesis denote 95% confidence intervals.
An & indicates that the value of the partial-autocorrelation function coincides with the upper or lower confidence limit.

TABLE 8

THE FIRST 40 AUTOCORRELATIONS AND PARTIAL AUTOCORRELATIONS OF $(1-B)(1-B^{12})W_t$ BASED ON SAMPLE 64:2 – 70:6

AUTOCORRELATIONS

1:	-.521135	-.015880	.026787	.151823	-.143104	.058709
7:	.105457	-.194492	-.010597	.089386	.220852	-.508432
13:	.401188	-.190935	.110727	-.110273	.123707	-.138370
19:	.039491	-.013971	.068167	-.073533	-.065546	.201239
25:	-.205721	.154477	-.186857	.277349	-.247124	.077329
31:	.064993	-.022362	-.138505	.143412	.043412	-.146166
37:	.095067	-.026572	.069843	-.152387		

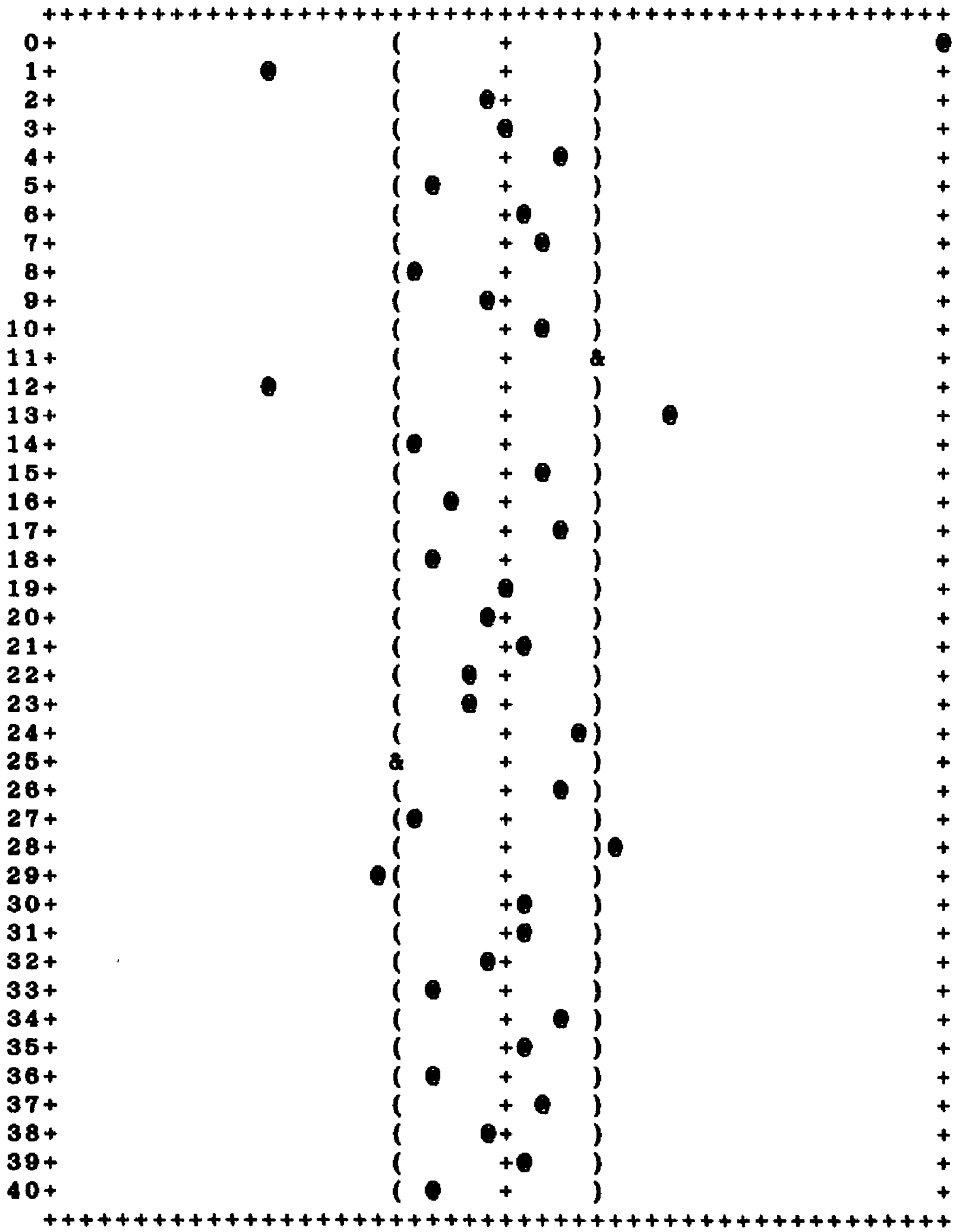
PARTIAL AUTOCORRELATIONS

1:	-.521135	-.394637	-.309626	.006057	.000588	.071086
7:	.264946	.018234	-.207360	-.270107	.228388	-.227423
13:	.119911	-.093379	.057940	-.002848	-.037309	-.138098
19:	.013949	-.224238	-.083719	-.106579	.029075	.035285
25:	.131935	-.010862	-.170941	.047918	-.078680	-.181915
31:	.135543	-.016851	-.052404	-.088026	-.000087	.032844
37:	.064643	-.034638	-.081734	.084131		

For the plots of the above data along with 95% confidence intervals see Figures 17A and 17B.

FIGURE 17A

THE PLOT OF THE FIRST 40 AUTOCORRELATIONS FOR
FIRST AND TWELFTH-ORDER DIFFERENCED WHITE SERIES
1964:2- 1970:6



The parenthesis denote 95% confidence intervals.
An & indicates that the value of the autocorrelation function coincides with the upper or lower confidence limit.

Figure 17B

THE PLOT OF THE FIRST 40 PARTIAL-AUTOCORRELATIONS
FOR FIRST- AND TWELFTH-ORDER DIFFERENCED WHITE SERIES
1964:2- 1970:6

```
  +++++++++++++++++++++++++++++++++++++++++++++++++++
 0+                  (     +    )                   ●
 1+           ●      (     +    )                   +
 2+              ●   (     +    )                   +
 3+                ● (     +    )                   +
 4+                  (     ●    )                   +
 5+                  (     ●    )                   +
 6+                  (     +●   )                   +
 7+                  (     +    )●                  +
 8+                  (     ●    )                   +
 9+                  &     +    )                   +
10+                 ●(     +    )                   +
11+                  (     +    &                   +
12+                  &     +    )                   +
13+                  (     + ●  )                   +
14+                  (  ●  +    )                   +
15+                  (     +●   )                   +
16+                  (    ●+    )                   +
17+                  (    ●+    )                   +
18+                  ( ●   +    )                   +
19+                  (     ●    )                   +
20+                  &     +    )                   +
21+                  (  ●  +    )                   +
22+                  (  ●  +    )                   +
23+                  (     ●    )                   +
24+                  (     ●    )                   +
25+                  (     +  ● )                   +
26+                  (    ●+    )                   +
27+                  (●    +    )                   +
28+                  (     +●   )                   +
29+                  (   ● +    )                   +
30+                  (●    +    )                   +
31+                  (     +  ● )                   +
32+                  (    ●+    )                   +
33+                  (   ● +    )                   +
34+                  (  ●  +    )                   +
35+                  (    ●+    )                   +
36+                  (     ●    )                   +
37+                  (     +●   )                   +
38+                  (    ●+    )                   +
39+                  (  ●  +    )                   +
40+                  (     + ●  )                   +
  +++++++++++++++++++++++++++++++++++++++++++++++++++
```

The parenthesis denote 95% confidence intervals.
An & indicates that the value of the partial-autocorrelation function coincides with the upper or lower confidence limit.

TABLE 9

ESTIMATED ARIMA EQUATIONS FOR NYC BIRTHS TO BLACK AND WHITE ADOLESCENTS 1964:2-1970:6

BLACKS	θ_1	$\emptyset_1$	$Q^*(24)$	R^2
$B_t = -\theta_1 e_{t-1} - \emptyset_1 e_{t-12} + e_t$	.735 (9.16)	.801 (9.24)	15.01 .92	.63
WHITES	θ_1	$\emptyset_1$	$Q(24)$	R^2
$W_t = -\theta_1 e_{t-1} - \emptyset_1 e_{t-12} + e_t$	.772 (9.38)	.644 (5.70)	19.41 .73	.56

B_t and W_t are the natural logarithms of the Black and White births respectively. θ_1 and $\emptyset_1$ are the coefficients and e_{t-1} is the error term. The numbers in the parenthesis are the t-ratios.

* The Ljung-Box Q statistic determines the randomness in autocorrrelations of residual errors, and has a Chi-square distribution (Ljung and Box 1978). The numbers below the Q-statistics are the marginal significance levels; i.e. the probabilities of the null hypothesis that the autocorrelations of the errors are not different from zero.

TABLE 10

ESTIMATED ARIMA EQUATIONS FOR NYC BIRTHS TO BLACK AND WHITE ADOLESCENTS WITH THE INTERVENTION COMPONENT 1964:2-1987:12

BLACKS $B_t = -\theta_1 e_{t-1} - Ø_1 e_{t-12} + e_t + \alpha I_t$

θ_1	$Ø_1$	α	$Q^*(48)$	R^2
.776	.838	-.207	34.20	.64
(20.68)	(24.87)	(-.5.32)	.93	

WHITES $W_t = -\theta_1 e_{t-1} - Ø_1 e_{t-12} - Ø_2 e_{t-24} + e_t + \alpha I_t$

θ_1	$Ø_1$	$Ø_2$	α	$Q(48)$	R^2
.824	.740	.122	-.150	42.16	.61
(23.94)	(12.04)	(1.95)	(-4.10)	.71	

B_t and W_t are the natural logarithms of the Black and White births respectively. θ_1 and $Ø_1$ are the coefficients and e_{t-1} is the error term. I_t is the dummy variable for the intervention; α is its coefficient. The numbers in the parenthesis are the t-ratios.

* The Ljung-Box Q statistic determines the randomness in autocorrrelations of residual errors, and has a Chi-square distribution (Ljung and Box 1978). The numbers below the Q-statistics are the marginal significance levels; i.e. the probabilities of the null hypothesis that the autocorrelations of the errors are not different from zero.

FIGURE 18

MONTHLY NUMBER OF ACTUAL AND FORECASTED BIRTHS TO BLACK ADOLESCENTS LIVING IN NEW YORK CITY
JULY 1970 - JULY 1972

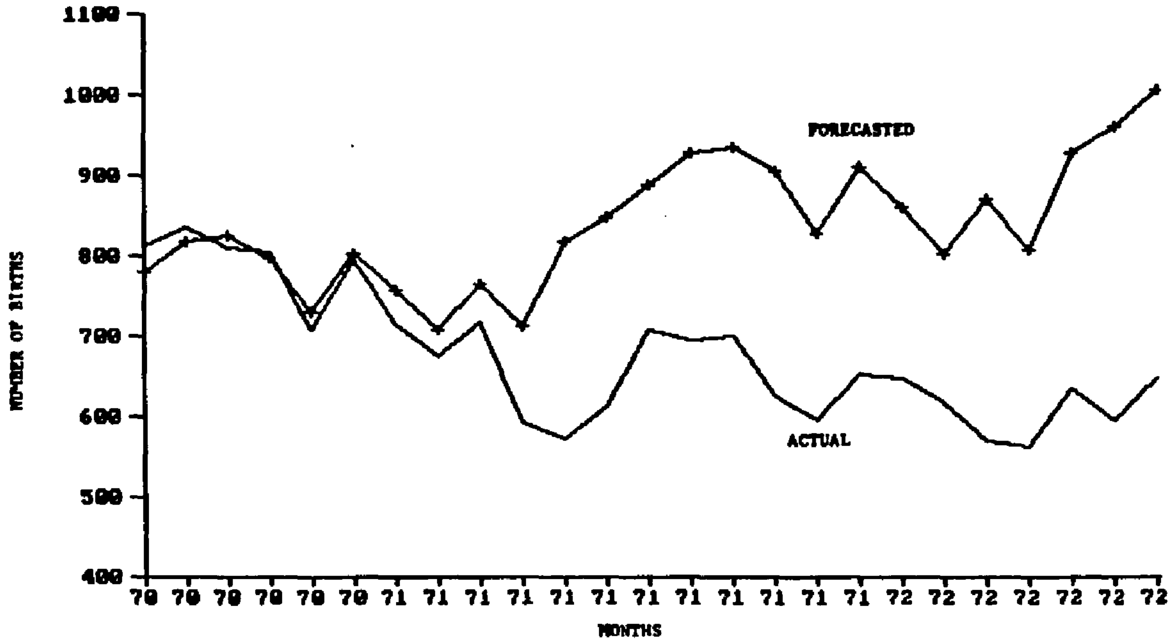
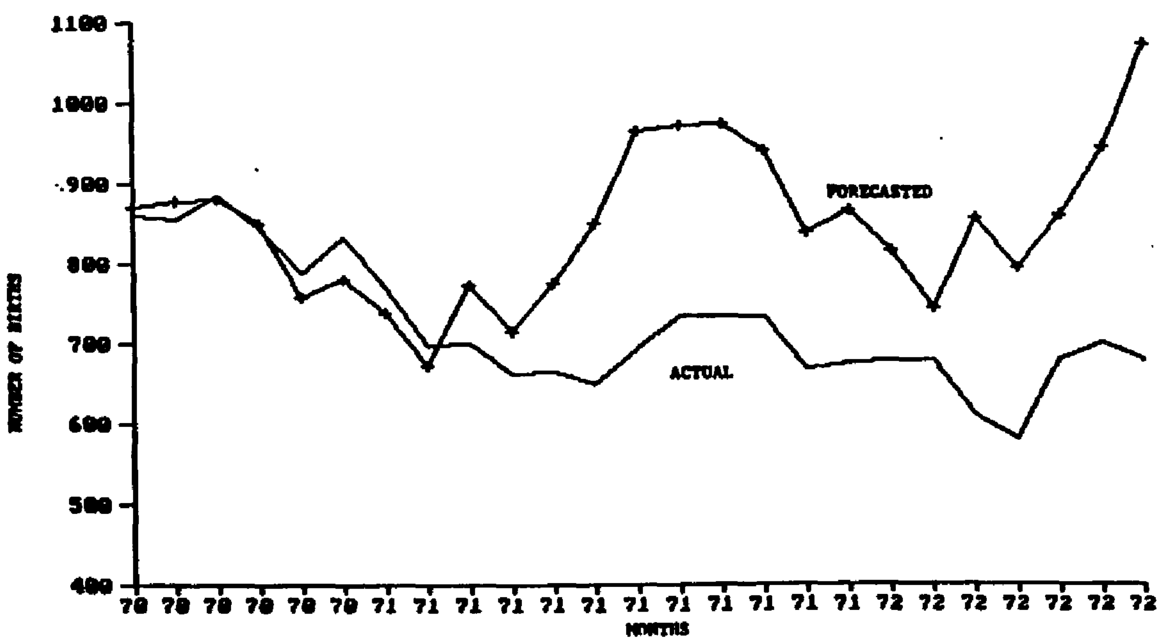


FIGURE 19

MONTHLY NUMBER OF ACTUAL AND FORECASTED BIRTHS TO WHITE ADOLESCENTS LIVING IN NEW YORK CITY
JULY 1970 - JULY 1972



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