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A TIME-SERIES ANALYSIS OF UNEMPLOYMENT AND HEALTH: THE CASE OF BIRTH OUTCOMES IN NEW YORK CITY

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ABSTRACT

The paper presents an aggregate time-series analysis of unemployment and infant health that improves on previous work in several ways. First, the data is monthly as opposed to annual and pertains to New York City from January, 1970 to December, 1986. Second, a structural production function is estimated in which the race-specific percentage of low-birthweight births is the health outcome. Because we are able to control for the race-specific percentage of women who begin care in the first trimester as well as the percentage of births to unmarried mothers, the unemployment rate as a proxy for maternal stress enters the production function as one among a set of well-defined health inputs. Third, because a pregnancy is limited to at most ten months, we can specify a lag length with confidence. Fourth, the data is tested for stationarity and the production function is estimated in levels as well as in deviations from trend. We find no cyclical variation in the percentage of low-birthweight births. The results are insensitive to changes in lag length, the omission of relevant inputs, and the functional form of the coefficients on the distributed lag.

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I. Introduction

Social scientists continue to investigate the aggregate time-series relationship between unemployment and health.[1] One attraction is that time-series analysis is an inexpensive means of investigating relationships in less than real time. Although cross-sectional analysis has similar advantages, the lack of temporal variation prevents a clear determination of whether poor health causes unemployment, or whether unemployment worsens health. Epidemiological research, on the other hand, has focused on work as a risk factor for various diseases and injuries. However, as Catalano and Serxner(1987) point out, policy makers are primarily concerned with the net effects of the business cycle on health. Thus, time-series analysis with aggregate data has the potential to complement the work from micro-level studies.

To date, aggregate time-series has failed to achieve a consensus regarding the effect of economic downturns on morbidity and mortality. The lack of clear conclusions has been attributed to a host of methodological shortcomings.[2] This paper attempts to overcome a number of the most serious weaknesses by estimating a structural model of infant health. In particular, we use monthly data from New York City to examine the impact of economic fluctuations on race-specific variations in the percentage of low birthweight births. An important feature is that we are able to control for the utilization of prenatal care and the incidence of out-of-wedlock childbearing. Thus, if there is an effect of unemployment on health, we can shed light on whether economic downturns

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[1] For the most recent work see Brenner (1987a, 1987b, 1983), Catalano and Serxner (1987), Forbes and McGregor (1984), McAvinchey (1984).

[2] See Gravelle (1984) for an insightful summary.

have a direct effect on birth outcomes by increasing stress, or an indirect effect through decreased consumption of prenatal care and increased marital instability.

Only Brenner (1973, 1979, 1983) has studied the time-series relationship between unemployment and infant health. In his earliest work, annual changes in the United States unemployment rate were inversely related to annual changes in the rate of fetal and infant mortality. In the latter two studies, changes in the infant mortality rate in England, Scotland, and Wales were positively associated with movements in the unemployment rate. Although the more recent work controlled for other factors and used a polynomial distributed lag specification, the robustness of the results has been challenged (Gravelle, Hutchinson and Stern 1981). Specifically, the authors found that Brenner's results were sensitive to the time periods under study, the variables used, and the functional form of the distributed lag.

This study represents a more rigorous test of the time-series relationship between unemployment and health than has previously been possible. First, the data is monthly and pertains to only one city (albeit a large one). Aggregation over time and region can smooth out local variations in unemployment and health. Second, the measures of health and medical care utilization are race-specific. Business cycle variations may impact differently on white and black birth outcomes. Third, we estimate an infant health production function derived from a behavioral model of health. Previous studies have estimated a "hybrid" function in which reduced form determinants are employed as proxies for structural inputs. Such approaches yield biased estimates of the health technology (Rosenzweig and Schultz 1983). Fourth, our indicator of health, the percentage of low birthweight births, is a well-measured predictor of infant mortality and childhood morbidity (McCormick 1985). Yet, unlike infant mortality, the modest decline in low birthweight over the past twenty years cannot be attributed to technological change, an exceedingly difficult variable to control for in time-series analysis. Fifth, we test the data more rigorously for stationarity. Recent work in time-series econometrics has shown that inappropriate detrending can yield highly spurious results (Granger and Newbold 1974; Nelson and Kang 1984).

II. Analytical Framework

Following Rosenzweig and Schultz (1983), Joyce (1987), and Corman, Joyce, and Grossman (1987), we assume the parent's utility function depends on their own consumption, the number of births, and the health of each child at birth. Both the number of births and the infant's health at birth are choice variables. In particular, birth outcomes depend on such factors as the quality and quantity of medical care, nutrition, stress, and the own time of the mother. They also depend on the mother's reproductive efficiency which includes an unobserved biological component, or health endowment, as well as other aspects of her efficiency in household production.

The preceding ideas are formalized in the following model:

$$h = f_1(m,d,u,e) \qquad (1)$$

$$m = f_2(p,w,y,e) \qquad (2)$$

$$d = f_3(p,w,y,e) \qquad (3)$$

$$u = f_4(p,w,y,e) \qquad (4)$$

Equation (1) is the structural production function for it relates the health of the infant at birth (h) to medical care utilization (m), demographic factors (d), stress related inputs (u), and the mother's health endowment (e). Equations (2) through (4) are the input demand equations. They relate the use of an input to measures of price (p), income (y), the value of time (w), and again the mother's health endowment (e).

The health endowment is included in the input demand equations because family history, previous pregnancies, and more recently, diagnostic procedures such as amniocentesis can warn the mother and physician about genetic abnormalities or other conditions that could complicate the pregnancy and birth. Such knowledge, generally unknown to the researcher, can motivate women with less favorable endowments to consume more care, to avoid stressful situations, and to even alter the timing and number of their births. Cross-sectional studies have shown that estimates of the production function can be seriously biased if this form of simultaneity is not controlled (Rosenzweig and Schultz 1983; Corman, Joyce, and Grossman 1987). The implications for aggregate time-series are discussed in Section III.

Following Brenner (1973), unemployment enters the production function as a proxy for maternal stress. Although several studies have found significant associations between stressful events and preterm births, the results are difficult to generalize (Norbeck and Tilden 1983; Newton et al. 1979; Gorsuch and Key 1974;). One reason is the different instruments used to measure stress. Another reason is the lack of control for other factors that can mitigate the impact of stress. For example, there is evidence that pregnant women with strong social networks are less affected by stressful events than women with few such "psychological assets" (Nuckolls, Cassel, and Kaplan 1972). To control for a lack of social support, we will include the percent of births to unmarried women, a well-documented risk factor for low birthweight (Institute of

Medicine 1985).

A difficulty of using unemployment to proxy stress is that pregnant women who work beyond the second trimester, or women whose jobs are physically demanding, may be at risk of a preterm or low-birthweight birth (Naeye and Peters 1982; Mamelle, Laumon and Lazar 1984). In other words, adverse birth outcomes may vary procyclically for a segment of the pregnant population. The magnitude of this offsetting effect is unknown, but it may have a greater impact on black birth outcomes since a greater proportion of black women are more apt to work in low-wage and physically strenuous settings.[3]

Previous time-series analyses of unemployment and health have lacked an adequate measure of medical care utilization. Its omission would bias upwards the effect of unemployment on health if the demand for medical care varied procyclically. In the birth outcome literature, it is widely accepted that appropriate prenatal care has a favorable impact on birth outcomes (Institute of Medicine 1985). If preventive care is a normal good, then the demand for prenatal care will fall during a recession (Elliot, Logerfo, and Daling 1985). Whether it will fall sufficiently to affect aggregate birth outcomes is unclear because women with health insurance that includes maternity benefits may be unaffected by a recession. However, until the Pregnancy Discrimination Act of 1978, only 57 percent of the employees covered by health insurance in 1977 had policies which included maternity benefits (Health Insurance Association of America 1982). Although the percent of

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[3] In 1980, white women in the New York Labor Market Area were more likely than black and other women to work in white collar jobs (76.8 percent versus 59.7 percent) and were less likely to be service workers (11.5 versus 28.0) [U.S. Department of Labor 1981].

employees with maternity benefits rose to 89 percent by 1982, Gold and Kenney (1985) estimated that 25 percent of the women in the U.S. 18 to 24 years of age lacked any public or private health insurance in 1984. Women 18 to 24 account for 40 percent of all births.

Changes in birth rates over the business cycle represent another mechanism by which unemployment may impact on infant health. Specifically, increased fertility may worsen birth outcomes if the women at the margin of the labor force are from groups with above average risks of adverse birth outcomes. Teenagers and unmarried women are two examples.[4] As with prenatal care, failure to control for the changes in the distribution of births among risk groups, may bias the estimated effect of unemployment on infant health.[5]

Whether fertility varies procyclically or countercyclically is still disputed. Butz and Ward (1979) argue that recessions lower the opportunity costs of a woman's time which increases birth rates. With annual data from the United States, Butz and Ward show that the cyclical movement in female wages can explain the countercyclical pattern of U.S. fertility since the mid 1960's. However, Macunovich and Easterlin (1988) report a negative effect of unemployment on fertility. The authors apply a Granger-Sims causality test to monthly U.S. data. They find causality to be unidirectional from unemployment to fertility and that the

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[4] Wilson (1987) argues that the rise in out-of-wedlock childbearing among blacks can be attributed to a decline in the labor force participation of black men. Although Wilson was referring to more secular trends, recessions exacerbate the problem.

[5] It should be noted, however, that lower time costs could have a positive effect on birth outcomes, if the demand by pregnant women for more time-intensive inputs such as exercise and rest were to rise.

relationship has grown stronger since 1973.

In sum, supporters of the stress hypothesis, of which Brenner is the most prominent, argue that unemployment worsens birth outcomes directly by increasing maternal stress, and indirectly by altering the consumption of health inputs. However, if work among pregnant women is a risk factor for prematurity, and if fertility varies over the business cycle, then the net effect of unemployment on infant health is ambiguous a priori.

III. Empirical Test

Data

The data on low birthweight, prenatal care and out-of-wedlock childbearing are from micro-level vital statistics for New York City. Each year of data has been aggregated by month for blacks and whites separately from January, 1970 through December, 1986 (N=204). The data for whites contain a large component which is Hispanic. Ethnicity was not identified on New York City birth certificates until 1978. In 1984, the first year in which data on Hispanics were published, 50 percent of all white women who delivered a live birth were of Hispanic origin; 85 percent of the Hispanic mothers were white (New York City Department of Health 1985).

New York City averaged approximately 100,000 resident live births over the time span under study. The percentage of births to white and blacks has varied between 69 percent white and 29 percent black in 1970 to 58 percent white and 37 percent black in 1985. Because the absolute number of births was higher in the earlier years, the number of births by race has remained relatively stable, approximately 6000 white births and 3000 black births monthly.

The race-specific percentage of low birthweight births is the indicator of infant health. Low birthweight is a major risk factor for infant mortality and childhood morbidity (Institute of Medicine 1985). It is also well-reported on birth certificates. In New York City less than .01 percent of the birth records lacked data on birthweight in any one year.

Prenatal care is measured by the race-specific number of births to women who began care in the first trimester over the total number of births to women who had any care at all. The other health input is the percentage of births to unmarried women. Prior to 1978, women with no care were coded the same as women whose prenatal care was unknown. From 1978 and on, women with no prenatal care and women whose care was unknown were coded separately. To maintain a consistent series we ignored the distinction from 1978 to 1986. Thus, the denominator is the number of women with at least some care.

The percentage of women with no care or whose care was unknown has remained relatively stable for blacks but has increased among whites. In 1970, 13 percent of all black mothers and 6 percent of white mothers who delivered a live birth reported no care or their care was unknown. By 1980, the figures were 16 percent for blacks but 16.5 percent for whites. Of the 16 percent, approximately two-thirds represented women who had no care. The large increase for whites is probably due to the declining proportion of births to white, non-Hispanic women in New York City. Census data combined with unpublished estimates by the New York City Department of Planning indicate that in 1970, 25.7 percent of all whites, males and females, were of Hispanic origin. By 1980, that

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[6] Personal communications with Lori Banks, New York City Department of Planning.

proportion had risen to 38.0 percent.[6] Because of the shifting ethnicity among whites, it is unknown whether the proportion of women who had no care varied over the decade. This could introduce measurement error which should be kept in mind when interpreting the results. Marital status, on the other hand, has been consistently well-reported; less than 1 percent of the annual birth records lack data on marital status.

The New York City unemployment rate is from the Current Population Survey. Because of small sample sizes, the unemployment rate is unavailable by sex, race, or age.[7] The employment-population ratio is also available for New York City. The ratio is the total number of persons employed over the noninstitutional population 16 years and over. Although the unemployment rate allows for greater comparison with other published work, the employment-population ratio may be more sensitive to variations in discouraged workers. This may more accurately reflect the labor market status of teenagers and minorities. The simple correlation between the two labor market measures is -.76.

Specification and Estimation

Visual inspection of the data revealed rather mild trends for low birthweight and unemployment (Figure 1) and more dramatic trends in prenatal care and illegitimacy (not shown). Recent developments in time-series econometrics have underscored the importance of differentiating

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[7] As a proxy for stress, it is questionable whether a sex-specific unemployment rate would be appropriate. First, it might fail to capture the loss of work by a male head of household. Moreover, the female unemployment rate would increase the potential for simultaneous equation bias between fertility and labor market participation. However, a recent time-series analysis of fertility and employment found that the unemployment rate Granger-caused fertility (Macunovich and Easterlin 1988).

between difference-stationary processes (DSP) and trend-stationary processes (TSP) [Nelson and Plosser 1982; Maddala 1988] Failure to distinguish between the two can generate seriously misleading results (Nelson and Kang 1984; Stock and Watson 1988). The Dickey-Fuller (1979, 1981) test was applied to the natural logarithm of each series. The results are presented in Table 1. With respect to the logarithms of the percentage of low birthweight births and the percentage of white women who begin prenatal care in the first trimester, we reject the null hypothesis of a DSP specification; for the natural logarithms of the unemployment rate, the employment-population ratio, the percentage of illegitimate births and the percentage of black women who begin care in the first trimester, we cannot reject the null that the series have a unit root.

Based on the Dickey-Fuller tests we specify two sets of regressions. The first set focuses on cyclical changes in each of the variables. Thus, we use deviations around a linear time trend for the natural logarithms of low birthweight and prenatal care for whites, and first differences of the labor market measures, illegitimacy, and prenatal care among blacks.[8]

The second set of regressions estimates the infant health production functions in the levels of the variables. The coefficients can be interpreted as "long-run" estimates. Although we can not reject the null hypothesis that the labor market measures, illegitimacy, and prenatal care among blacks are integrated processes, the augmented Dickey-Fuller test suggested that for blacks, prenatal care and illegitimacy are co-

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[8] The deviations are the residuals from a regression on a constant and a linear time trend.

integrated (Engle and Yoo 1987).[9] Put differently, the series share a stochastic trend such that a linear combination of the series yields a stationary process. Following the "rules of thumb" outlined by Stock and Watson (1988) when estimating time-series regressions with integrated variables, the coefficients on illegitimacy among whites and the coefficients on the lagged unemployment rate should also be consistent but they will have a nonstandard asymptotic distribution.

The functional form of the production function is assumed to be linear (in logs). The unemployment rate is specified as a distributed lag. The maximum duration of pregnancy provides a strong rationale for truncating the lags after the tenth month. The assumption is that stress induced by business-cycle changes has no impact on the fetus prior to conception. To our knowledge, there is no epidemiological evidence to suggest otherwise. We estimate an unconstrained distributed lag because there is little a priori information regarding its functional form.

As mentioned in Section II, the utilization of health inputs may be conditioned on the anticipated birth outcome. For instance, women who have experienced a preterm delivery in the past may seek out more prenatal care than women with no such history. As a result, ordinary least squares may yield biased estimates of the health technology if the

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[9] The augmented Dickey-Fuller test is based on the residuals from the regression of prenatal care on illegitimacy (in the levels of the logs of the variables). Let R_t denote the residuals and ΔR_t the first difference. The augmented Dickey-Fuller test is the t-statistic on the coefficient of R_{t-1} in the following regression:

$$\Delta R_t = b_0 R_{t-1} + \sum_{i=1}^{p} b_i \Delta R_{t-i}$$

We obtained a t-statistic of -4.86; the critical value at the 5% level with 200 observations is -3.25 (Engle and Yoo 1987, p.158, Table 3). We set p=6 given the residuals seemed to follow an MA(1) process.

inputs are correlated with the unobserved factors embedded in the residuals. Researchers who use cross-sectional data have employed two-stage least squares to control for the simultaneity (Rosenzweig and Schultz 1983, 1988; Joyce 1987). In this study, lagged birthweight will be used as a control for unobserved health and OLS will be used to estimate the production functions. A careful examination of the residuals should reveal the adequacy of the approach.

IV. Results

Black and white production functions estimated with detrended data are presented in Table 2. Not shown in each specification are the coefficients on the eleven seasonal dummies. Overall the fit is rather weak even for detrended data. The adjusted R-square is .29 for blacks and .09 for whites. The residual correlogram has no discernible pattern for blacks, but is more suspect for whites. A Lagrange multiplier (LM) test was applied to the residuals.[10] We tested for first-order and twelfth-order autocorrelation. The latter was a check on the adequacy of the seasonal dummies. In neither case could we reject the null hypothesis that the residuals were white noise. A joint test of the first six autocorrelations indicated that for whites, the null hypothesis could be rejected at the .05 level but not at the .01 level. For blacks, the null hypothesis of white noise residuals could not be rejected.

There is no evidence that changes in the unemployment rate have a

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[10] The Durbin-Watson statistic is not appropriate when the specification includes a lagged dependent variable; nor, strictly speaking, is the Box-Pierce Q-statistic. The LM statistic provides a general test for higher-order autocorrelation and is valid when the set of regressor includes lagged dependent variables (Breusch 1978; Godfrey 1978).

negative impact on infant health. The sum of the coefficients on unemployment are negative and insignificant at conventional levels. If anything, the rate of low birthweight among blacks falls as unemployment declines. The coefficients reveal a relatively smooth quadratic pattern with the greatest impact in the third trimester (lags 0 through 3).

To check whether the lack of statistical significance may have been caused by multicollinearity, we reran the regressions using an Almon distributed lag. The coefficient on the contemporaneous lag was unconstrained and the lag on the eleventh month was assumed to be zero. The maximum length of a pregnancy justified the endpoint restriction. We imposed a second-order polynomial on the coefficients based on the pattern of the unconstrained estimates. The results were unchanged (not shown). The sum of the lags was -.38 for blacks (t-ratio = -1.46) and -.03 for whites (t-ratio = -.10); in neither case were the sum of the coefficients on the unemployment rate statistically different from zero.

The negative coefficients on the unemployment rate in the black regressions suggest that low birthweight may vary procyclically.[11] As discussed in Section II, there is epidemiological evidence that pregnant women who work in stand-up jobs and who continue to work into the third trimester may be at risk of a preterm or low birthweight birth (Naeye and Peters 1982; Mamelle, Laumon, and Lazar 1984). Moreover, it was anticipated that a procyclical effect would be more likely for blacks given their disproportionate representation in such jobs. To further examine this proposition, we reduced the number of lags from 10 to 3. There was little change. The sum of the coefficients was -.17 with a t-

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[11] Even if the results had been statistically significant, they would have been suggestive at best. Employment as a risk factor for preg-

ratio of -1.43.[11]

To test whether the results in Table 2 were sensitive to the measure of employment, we substituted the logarithm of the employment-population ratio (in first differences) for the unemployment rate. Changes in the ratio have no impact on the rate of low birthweigth births. The sum of the coefficients on the employment-population ratio was .636 for blacks (t-ratio = .36) and -2.21 for whites (t-ratio =-1.17).

Lagged birthweight is an important predictor of current birthweight among blacks. It is less important among whites. Prenatal care has the correct sign in the black regression but is statistically insignificant at conventional levels. In the white regression, the prenatal care coefficient has the wrong sign and is insignificant. The results imply that deviations around the trend in prenatal care have little impact on similar deviations in low birthweight. Measurement error may also explain the weak performance of prenatal care. Finally, changes in the percentage of illegitimate births have no impact on cyclical variations in low birthweight.

A major shortcoming in aggregate time-series studies of unemployment and health is the inability to control for medical care utilization. If the use of care varies procyclically, then its omission would bias upwards the effect of unemployment. To test this, we excluded prenatal care and illegitimacy from the specification in Table 2. The results (not shown) change in a trivial manner. Put differently, there is no evidence that unemployment has an indirect effect on low

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nant women would have to be measured with much greater precision before more substantive conclusions could be inferred.

birthweight through prenatal care or out-of-wedlock childbearing.

Estimates of the production function in the levels of each of the variables are presented in Table 3. With few exceptions there are no qualitative differences from the results in Table 2. If unemployment has any impact on low birthweight it is restricted to blacks in the third trimester. The LM tests indicate that the residuals are white noise. The stability of the results between different forms of the data, and the absence of autocorrelation suggest that a lack of stationarity among some of the regressors is not a problem (Plosser and Schwert 1978).

The impact of early prenatal care on the percentage of low-birthweight births among blacks is statistically significant. The coefficient indicates that a ten-percent increase in the percentage of women who begin care early would lower the rate of low-birthweight births among blacks by 1.5 percent. Effects of similar magnitude from aggregate cross-sectional data have been reported by Joyce (1987). There is also evidence that blacks benefit more from early prenatal care than do whites (Murray and Bernfield 1988). However, as with the detrended data, measurement error may have obscured the effect of prenatal care among whites.

Conclusion

The results from this study call into question the findings from previous aggregate time-series analyses that unemployment worsens infant health (Brenner 1973, 1979, 1983, 1987a, 1987b). With monthly data as opposed to annual data, and with the level of aggregation greatly reduced, we were unable to uncover any deleterious effect of changes in the unemployment rate on the race-specific percentage of low birthweight births. The results were the same regardless of whether the data were detrended or in levels. The findings were also insensitive to lag-length manipulations, the omission of relevant regressors, and the imposition of a functional form on the coefficients of the distributed lag.

Our results say little about the impact of unemployment on individual birth outcomes. What they do say is that the net effect is insignificant. However, the results do raise doubts about the findings from other studies that estimated the relationship between unemployment and mortality with annual data and national aggregates. This is not the first time such doubts have been raised (Kasl 1979). But it is the first time that the misgivings with respect to large-scale temporal and spatial aggregation have been tested with more refined data and in a more consistent analytical framework. If aggregate time-series analysis is to be useful, then better modeling and more detailed data are essential.

FIGURE 1

Unemployment and Low Birthweight

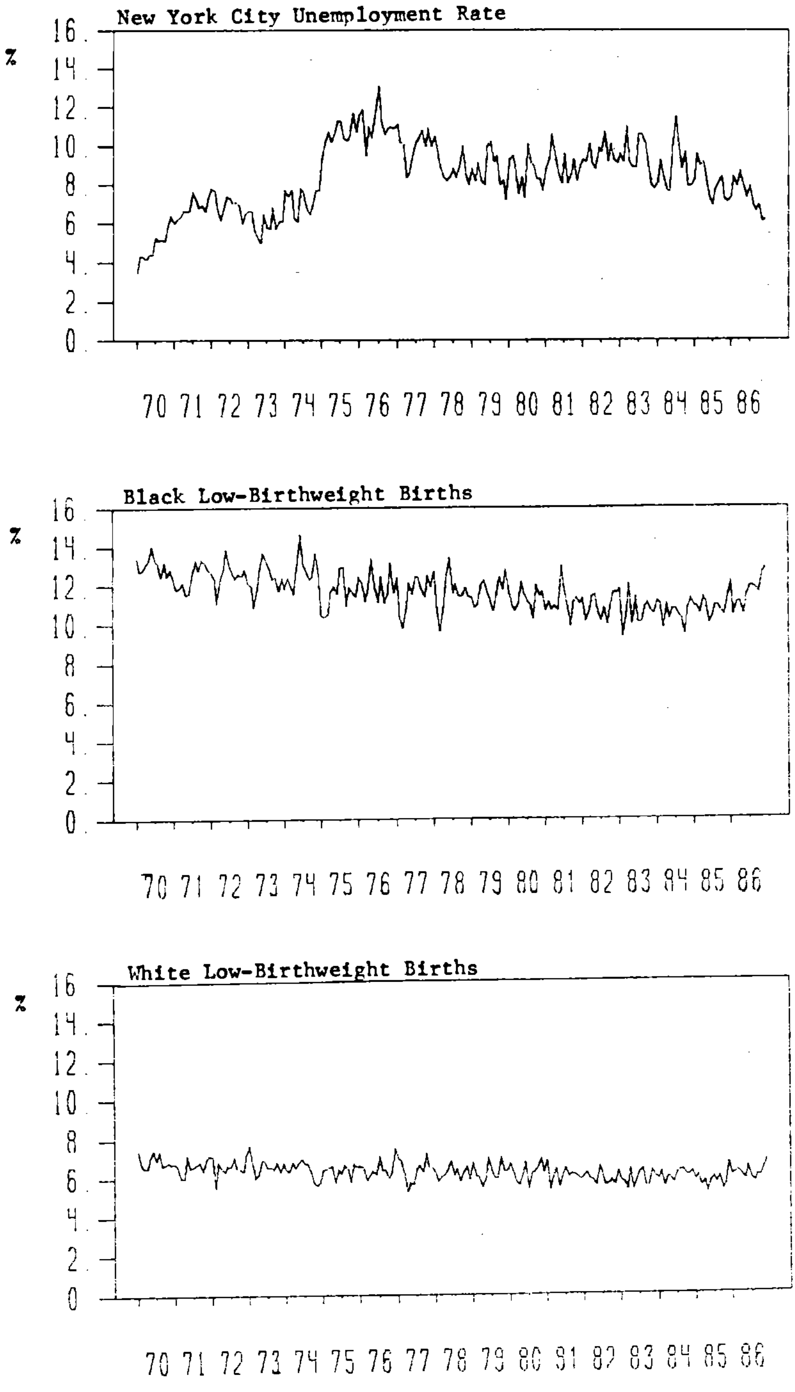

Table 1

Dickey-Fuller Test for Unit Roots[a]

	Whites	Blacks
Ln % low birthweight	15.75	7.53
Ln % early prenatal care	9.96	4.26
Ln % out-of-wedlock	2.46	3.79
Ln unemployment rate	3.41	3.41
Ln employ-population ratio	2.79	2.79

[a]Figures represent F-statistics under the null hypothesis $a_0=0$, $b_1=0$, in the regression

$$(y_t - y_{t-1}) = b_0 + a_0T + b_1y_{t-1} + \sum_{i=1}^{p} b_i(y_{t-i} - y_{t-i-1})$$

where T is a linear trend term. The critical values at the 5% level are 6.49 for a sample of 100 observations and 6.34 for a sample of 250 oberservations (Dickey and Fuller 1981, p. 1063, Table VI). Except for low birthweight and prenatal care among whites, the first difference of each of the series (in logs) was characteristic of an MA(1) process. Thus, we used an autoregressive representation (p=8) to approximate the MA(1) process.

Table 2

Ordinary Least Squares Estimates of Race-Specific Infant Health Production Functions from Detrended Data: January, 1970 -- December, 1986 (N=204).[a]

	Ln % Low Birthweight Births	
	Whites	Blacks
Ln unemployment rate (first diff)		
Lag (months) 0	-.031	.019
	(-.57)	(.37)
1	.014	-.028
	(.25)	(-.54)
2	-.016	-.152
	(-.28)	(-2.97)
3	-.004	-.085
	(-.07)	(-1.57)
4	.052	-.076
	(.89)	(-1.42)
5	-.020	-.013
	(-.35)	(-.19)
6	-.047	-.0001
	(-.82)	(-.001)
7	-.052	-.015
	(-.91)	(-.27)
8	.098	-.032
	(1.71)	(-.61)
9	.015	.058
	(.26)	(1.12)
10	-.061	.006
	(-1.14)	(.12)
Sum of lags	-.053	-.315
	(-.19)	(-1.22)
Ln % low birthweight		
Lag (months) 1	.099	.336
	(1.28)	(4.49)
Ln % early prenatal care	.012	-.124
	(.07)	(-1.28)
Ln % illegitimate (first diff)	-.089	.087
	(-.79)	(.49)
Constant	.055	.007
	(3.20)	(.41)
Adjusted R^2	.091	.291
LM-test, Chi-square[b]		
Ho: $p_1=0$	2.35	.20
Ho: $p_{12}=0$	.02	.17
Ho: $p_1 \ldots p_6=0$	12.89	5.11
Residual Correlogram[c]		
r_1	-.016	.012
r_2	.011	-.064
r_3	.045	.095
r_4	.138	-.072
r_5	.032	.090
r_6	-.167	.017

[a] T-statistics are in parentheses. Ln indicates natural logarithms. Birthweight and early prenatal care among whites represent deviations from a linear trend; black prenatal care, unemployment and illegitimacy are in first differences.

[b] In the LM test for serial correlation, the residuals from the equations shown above are regressed on the complete set of independent variables and a set of lagged residuals. The number (k) and order (p) of the lagged residuals depends on the degree of autocorrelation being tested. An F-statistics is computed for the coefficients on the lagged residuals. Mutiplying F by k yields the Chi-square statistic with k degrees of freedom. The critical value for Chi-square at the .05 level for 1 degree of freedom is 3.841 and for 6 degrees of freedom is 12.592.

[c] The sample autocorrelations are given by r_i.

Table 3

Ordinary Least Squares Estimates of Race-Specific Infant Health Production Functions with Data in Levels: January, 1970 -- December, 1986 (N=204).

	Ln % Low Birthweight Births	
	Whites	Blacks
Ln unemployment rate		
Lag (months) 0	-.023	.019
	(-.42)	(.37)
1	.043	-.047
	(.67)	(-.80)
2	-.041	-.132
	(-.63)	(-2.23)
3	-.021	.069
	(-.33)	(1.44)
4	.053	.003
	(.82)	(.05)
5	-.083	.628
	(-1.27)	(1.05)
6	-.019	.001
	(-.29)	(.02)
7	.003	-.023
	(-.05)	(-.39)
8	.142	-.016
	(2.18)	(-.27)
9	-.065	.083
	(-.99)	(1.40)
10	-.034	-.040
	(-.62)	(-.78)
Sum of lags	-.002	-.020
	(-.06)	(-.65)
Ln % low birthweight		
Lag (months) 1	.106	.347
	(1.36)	(4.74)
Ln % early prenatal care	-.045	-.148
	(-.29)	(-2.17)
Ln % illegitimate	-.120	.013
	(-2.52)	(.08)
Constant	2.242	2.126
	(4.22)	(4.50)
Adjusted R^2	.327	.518
LM-test, Chi-square[a]		
Ho: $p_1=0$	.21	.08
Ho: $p_{12}=0$	.01	1.03
Ho: $p_1 \ldots p_6=0$	9.84	6.58
Residual Correlogram[b]		
r_1	.005	.010
r_2	.025	-.047
r_3	.067	.122
r_4	-.119	-.046
r_5	.049	.122
r_6	-.149	.013

[a] See footnote (b) in Table 2.

[b] See footnote (c) in Table 2.

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