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Illicit Drug Use and Health: Analysis And Projections of New York City Birth Outcomes Using a Kalman Filter Model*

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I. Introduction

As the use of illicit drugs persists as a major social problem facing urban America, the clinical evidence linking prenatal exposure to illicit drugs, in particular cocaine, and adverse birth outcomes has been mounting rapidly [8; 21; 38]. Newborns exposed prenatally to cocaine appear to be more likely to suffer intrauterine growth retardation, low birthweight and preterm delivery than infants unexposed. All three outcomes are strongly linked to infant mortality and excess morbidity [16; 22]. Furthermore, low birthweight infants who survive are more likely to experience serious developmental, health and learning problems later in life. The marginal costs of treating exposed as compared with unexposed infants has been estimated at between \$9,313 and \$13,225 in 1993 dollars for the initial hospitalization only [1, 29].

The extent of the problem and its progression over time, however, are not well known. Estimates on the number of infants exposed to illicit substances in the United States range from 350,000 to 739,000 annually [9]. Determining the prevalence of illicit drug use in a free-living population is difficult and costly. The reporting biases inherent in surveys are substantial. More controlled studies based on urine toxicologies are limited in size as well as generalizability; and they provide little insight as to the changes over time [4].

Indirect evidence suggests that the problem may be more widespread and more dynamic than anticipated. A univariate analysis of monthly time-series data from New York City birth certificates revealed a dramatic increase in the rate of low birthweight births among Blacks be-

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ginning in 1984. The rate of low birthweight rose from 10.6 percent in 1984 to 13.1 percent in 1988: an unprecedented rise which over four years reversed a 22-year decline [19]. The upturn in low birthweight among Blacks in New York City was coincident with anecdotal evidence on the introduction of crack cocaine to the city. Without time-series data on prenatal drug use, however, the nexus between low birthweight and cocaine was speculative.

A more recent study based on New York City birth certificates addressed these shortcomings. Joyce, Racine and Mocan [18] used a pooled times-series, cross-sectional design to examine the link between illicit drug use and low birthweight. The study, based on annual data from 1980 to 1989 across health districts in New York City, reported that prenatal drug use was the most plausible explanation for the upturn in low birthweight among Blacks in the mid-1980s after controlling for prenatal care, smoking, and out-of-wedlock childbearing.

Since early 1989, however, both prenatal drug use and low birthweight among Blacks have turned sharply downwards. The rate of low birth weight has fallen from a peak of 14.1 to approximately 12.2 percent in 1990. The observed rate of prenatal illicit drug use has fallen from 6.0 percent to 5.1 over the same time span. The turnaround in both low birth weight and prenatal drug use is certainly encouraging and it raises several questions. What is the most likely trajectory of prenatal drug use and its impact on the rate of low birthweight in the near term? What are the short-term opportunity costs of not pursuing an aggressive policy that reduces the prevalence of prenatal drug use to its pre-crack epidemic level? What would be the consequences and costs of an unexpected upturn in prenatal illicit drug use?

To address these questions we fit a structural time-series model to race-specific monthly rates of low birthweight in New York City from 1963 through 1990. We first examine whether the upturn in 1984 is consistent with a structural shift in the long-term trend of low birthweight. We then project the rate of low birth weight that would have been observed had the upturn in 1984 not occurred. Finally, we add data on race-specific rates of prenatal illicit drug use, prenatal care and smoking that are available from 1978 to determine the relative contribution of each. With this model we provide answers to consequences and costs of various trajectories in illicit drug use and low birthweight.

II. Analytical Framework

Economic models of infant health emphasize the distinction between the health production function and the input demand functions [5; 17; 32]. The former represents the technical relationship between the birth outcome and the health inputs, whereas the latter focuses on the factors which determine the use of the health inputs. To illustrate, consider the following system of equations.

$$\text{Birth Weight} = f_1(\text{Health Inputs}, \text{Substance Abuse}) \quad (1)$$

$$\text{Prenatal Care} = f_2(\text{Prices and Availability}, \text{Income}) \quad (2)$$

$$\text{Substance Abuse} = f_3(\text{Prices and Availability}, \text{Income}) \quad (3)$$

In this paper, the birth weight is considered as a proxy for infant health, and the main health input is the consumption of prenatal care. Substance abuse stands for the consumption of deleterious substances such as tobacco and cocaine. The aim of the paper is the estimation of equation (1), the infant health production function. In particular, we seek to determine the contribution of illicit

drug use to the rise in the rate of low birth weight. Note that substitution of equations (2) and (3) into equation (1) would yield a reduced form production function, where the rate of low birth weight becomes a function of input prices and income. Although clearly of interest, estimation of the reduced form is problematic, mainly because of unavailability of the price of illicit drugs, especially on a monthly basis.

III. Empirical Implementation

The data include all singleton live births to Blacks and Whites residents of New York City between 1963 and 1990. Data on illicit drug use, however, are only available since 1978. New York City birth certificates are the *only* population based data source in the United States that has routinely reported information on prenatal illicit drug use for over a decade. The size of city, its racial diversity and the magnitude of the illicit drug problem make New York City a unique setting from which to address the time-series relationship between low birth weight and illicit drug use. The birth certificates contain information on prenatal substance abuse, which is based upon a combination of self-reports to physicians and positive toxicology screens applied at delivery. The potential consequences of this measurement are discussed in the results section.

Individual birth certificates have been aggregated by month and race. Our measure of infant health is the race-specific rate of low birth weight births. Low birth weight is superior to infant mortality as a measure of health in time-series context, because there is less potential confounding in low birth weight due to technological change. The rapid decline in infant mortality in the United States over the past 20 years has been attributed to advances in management of newborn care. By contrast, the rate of low birth weight has shown only a slight improvement, which has been attributed to the increased utilization of appropriate prenatal care, better nutrition, and a declining proportion of births to adolescents [20; 27]. It is hypothesized that these favorable trends have been offset by the prenatal consumption of illicit drugs. Our measure of illicit drug use variable is the percentage of women whose pregnancies are complicated by the use of cocaine, heroin, methadone and barbiturates.¹ The production function estimated from 1978–1990 also includes the percentage of women who smoked during pregnancy, and the percentage of women who started consuming prenatal care during the first trimester. In preliminary analyses we experimented with other correlates of low birth weight such as percentage of births out-of-wedlock, the percentage of births to women with a high school education, and the percentage of first births; but these added little explanatory power to our model and were thus excluded from further specifications.

Figure 1 presents the behavior of the Black and White rates of low birth weight. They span the years 1963 to 1990 and include 336 monthly observations. To highlight the underlying secular trends, both series are exponentially smoothed. The Black rate of low birth weight exhibits a downward trend between 1967 and 1984; it goes down from a monthly average of 14.66 in 1967 to 10.59 in 1984. Starting in 1985, the downward trend of 1967–84 is reversed, and the Black rate of low birth weight rises steadily until 1988, where the average reaches 12.95. The trend is reversed again in 1989, and the average value becomes 12.20 in 1990. Even though the White rate of low birth weight exhibits similar trends, the magnitudes of the changes are not as significant as those pertaining to Blacks. The monthly average rate of low birth weight for Whites is 7.23 in 1967, and 5.99 in 1984. It reaches 6.32 in 1988 and falls to 6.15 in 1990.

1. We exclude marijuana because it was added specifically to the birth certificate in 1988.

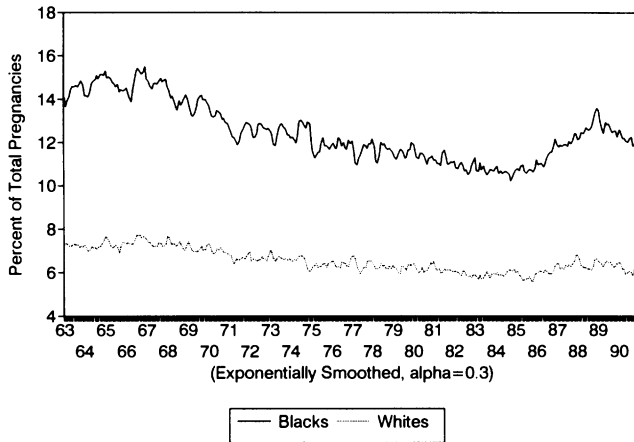


Figure 1. Rate of Low Birth Weight

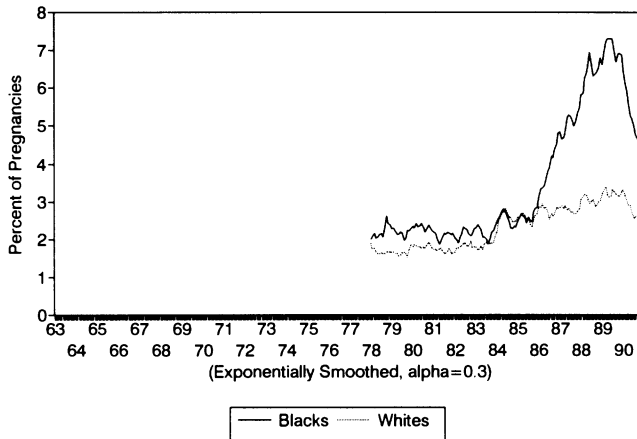


Figure 2. Percentage of Pregnancies Complicated by Drug Use

The variables pertaining to prenatal care consumption, and the use of drugs and tobacco are displayed in Figures 2–4. Each variable spans the years 1978 to 1990, consisting of 156 observations. Figure 2 displays the percentage of pregnancies complicated by drug use for both races between 1978 and 1990. To make the graph comparable to the one presented in Figure 1, the years 1963–77 are left blank. The percentage of Black pregnancies complicated by drugs remains stationary between 1978 and 1983 (the mean is 2.35 in 1983–84), but we note an upswing starting with 1985, and a downturn in 1989 (the monthly average of 1988 is 6.67, and 4.87 in 1990). These dates coincide with the turning points of the Black rate of low birth weight series. For Whites, the corresponding averages are 2.27 in 1983–84, 3.08 in 1988, and 2.69 in 1990.

Figure 3 shows the percentage of Black and White pregnancies complicated by tobacco use. Figure 4 reports percentage of pregnancies where the consumption of prenatal care started in the first trimester. The graphs for Black and White pregnancies complicated by the use of tobacco exhibit similar patterns, the one for Blacks lying almost three percentage points above that of the Whites. In Figure 4 one observes that the average percentage of pregnancies where the consumption of prenatal care started in the first trimester is 29.47 for Blacks during 1978. It rises

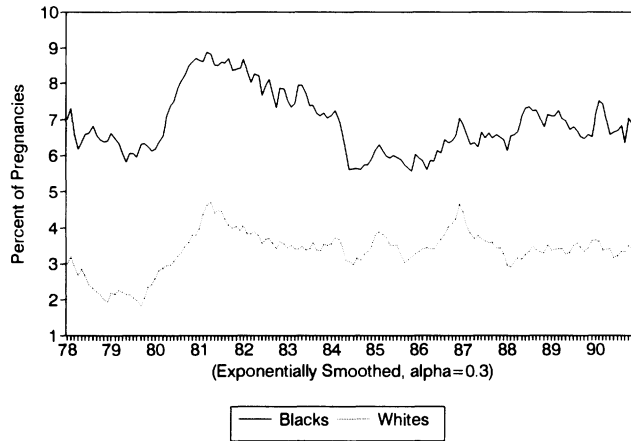


Figure 3. Percentage of Pregnancies Complicated by Tobacco

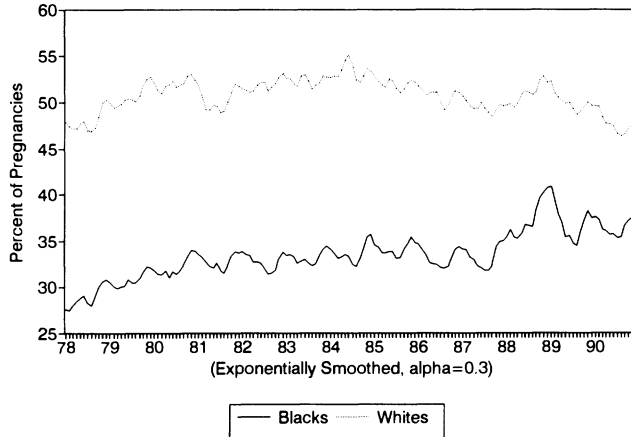


Figure 4. Percentage of Pregnancies with First Trimester Prenatal Care

steadily and levels off with an average of 36.36 in 1989–90. For Whites, it rises smoothly from 48.38 percent in 1978 to 52.57 percent in 1984–85.

As displayed later in the paper, births exhibit a seasonal cycle, with the number of babies born rising during the summer months (see Figure 13). This pattern indicates that at least some portion of prospective mothers plan the timing of their conception and delivery. A similar seasonal pattern is observed for early care pregnancies: the rate of pregnancies where the prenatal care consumption started in the first trimester increases in the fall season (Figure 4). To the extent that the timing of the pregnancy is correlated with the consumption of early care, Figure 4 also suggests the (partially) nonrandom nature of birth timing. This seasonal pattern in the number of births translates into a seasonal pattern in the rate of low birth weight births. There is no significant seasonal pattern in the use of illicit drugs and tobacco.

IV. Methodology and Estimation

Consider the time series H_t which is generated by the following stochastic process

$$H_t = \alpha + \beta t + u_t \tag{4}$$

$$u_t = \gamma u_{t-1} + e_t, \tag{5}$$

where e_t is a covariance stationary process with mean zero, t is a time trend, and α , β , and γ are the parameters. If $\gamma < 1$ the model depicted in (4) and (5) represents an asymptotically stationary AR(1) process with a linear time trend. If $\gamma = 1$, the model is a random walk around a linear trend. Substitution (5) into (4) and rearranging yields the reduced form

$$H_t = \delta_0 + \delta_1 t + \gamma H_{t-1} + e_t, \tag{6}$$

where $\delta_0 = [\alpha(1 - \gamma) + \gamma\beta]$ and $\delta_1 = \beta(1 - \gamma)$.

Equation (6) is said to have a unit root if $\gamma = 1$. The emphasis on unit roots has grown enormously during the past decade after Dickey and Fuller [7] suggested testing the unit root hypothesis using equation (6). Since then, researchers modified and proposed alternative versions of the original Dickey-Fuller test [30; 31; 34]. Testing for unit roots is important because it determines the proper method of de-trending. It has been widely recognized that regressions based on inappropriately de-trended series can result in highly misleading conclusions [3; 28; 36]. Trend elimination should be done by taking the first-differences of the series in which a stochastic trend is evidenced by some unit root test statistic. If the hypothesis of a unit root is rejected, the series should be regressed on some polynomial of time for trend removal.

Recently, however, the power of the unit root tests has been questioned. Evidence has been provided indicating that the unit root tests are not resilient against the trend-stationary alternatives; and the classical unit root asymptotics is asserted to be of little practical value [6; 33; 35]. This paper employs a structural time series model that enables us to avoid the current debate surrounding the unit root testing issue. More precisely, the model described below obviates the need to specify whether the trends are deterministic or stochastic prior to the analysis, and enables us to capture the dynamics of the underlying trend more accurately than other methods. In fact, taking the first differences or regressing on a deterministic time trend are special cases of the flexible trend model we estimate. Thus, it becomes much less likely that spurious estimates will be obtained due to misspecification of the trend. Following Harvey and Jaeger [11], Harvey [14], Harvey and Todd [15], Mocan [23], and Mocan and Topyan [24], we hypothesize that the dynamics of the rate of low birth weight can be formulated as

$$LBW_t = \mu_t + \psi_t + \epsilon_t \tag{7}$$

where LBW_t is the natural logarithm of the rate of low birth weight at time t , and μ_t , ψ_t , and ϵ_t are the trend, seasonal and irregular components, respectively. Within this framework, one can specify a locally linear trend where the level and the slope of the series are governed by random walks as follows:

$$\begin{aligned} \mu_t &= \mu_{t-1} + \beta_{t-1} + \eta_t \\ \beta_t &= \beta_{t-1} + \xi_t \end{aligned} \tag{8}$$

where η_t and ξ_t are white noise disturbance terms that are serially uncorrelated, and uncorrelated with each other with variances σ_η^2 and σ_ξ^2 , respectively. Seasonality is assumed to be generated by the following stochastic trigonometric process, which is allowed to evolve over time [12; 13].

$$\psi_t = \sum_{j=1}^{s/2} \psi_{jt} \tag{9a}$$

$$\begin{aligned} \psi_{jt} &= \psi_{j,t-1} \cos \lambda_j + \psi_{j,t-1}^* \sin \lambda_j + \omega_{jt} \\ \psi_{jt}^* &= -\psi_{j,t-1} \sin \lambda_j + \psi_{j,t-1}^* \cos \lambda_j + \omega_{jt}^* \end{aligned} \tag{9b}$$

where $j = 1, 2, \dots, [s/2]$, ω_{jt} and ω_{jt}^* are zero mean white noise disturbances which are uncorrelated with each other, and ψ_{jt}^* appears by construction [10; 12, 40–49].

The trend in (8) is equivalent to an ARIMA(0,2,1) process. If $\sigma_\xi^2 = 0$, the trend reduces to a random walk with a drift; i.e., LBW_t is stationary in first differences, (integrated of order one). If $\sigma_\eta^2 = 0$, but $\sigma_\xi^2 > 0$, the trend is still integrated of order two as the original case. If $\sigma_\xi^2 = \sigma_\eta^2 = 0$, the model collapses to a standard regression model with a deterministic trend; i.e., $\mu_t = \mu_0 + \beta t$.

The structural model depicted in (7) can be extended by adding exogenous explanatory variables X_t to the right hand side, which gives²

$$LBW_t = \mu_t + \psi_t + \delta X_t + \epsilon_t. \tag{10}$$

V. Results

We estimated the Black rate of low birth weight and the White rate of low birth weight using Kalman Filter for the years 1963 to 1980, and computed normalized recursive residuals from January 1981 through December 1990 as a means of diagnosing a structural change. To illustrate, we estimated the model based on data from January 1963 through December 1980. We then predicted the rate of low birth weight for January 1981. The difference between the actual rate and predicted rate for that month divided by the standard deviation of the residuals is the normalized recursive residual for January 1981. The model was re-estimated with data from January 1963 through January 1981 and the recursive residual for February 1981 is calculated. The process of updating the series, re-estimating the model, and generating new predictions was repeated for each month through December 1990. If the series has not undergone a structural change, then the recursive residuals should be randomly distributed around zero; thus, the cumulative sum of the residuals should fluctuate around zero. If there emerges a secular pattern to the cumulative sum of the residuals, it would imply a structural change, and the point where the change in the cumulative sum of residuals begins would suggest an approximate date. A plot of the cumulative sum of the residuals, therefore, is a means of diagnosing structural shift. It is particularly useful in the present context since the precise date of a specific intervention is not known.

Figure 5 displays the cumulative sum of the residuals for Blacks and Whites between 1981–90. For Blacks it begins to rise in 1985, and it declines after 1989. Thus, 1985 and 1989 seem to be the years in which some change took place which altered the dynamics of Black low birth weight

2. For more on estimation, see the Appendix.

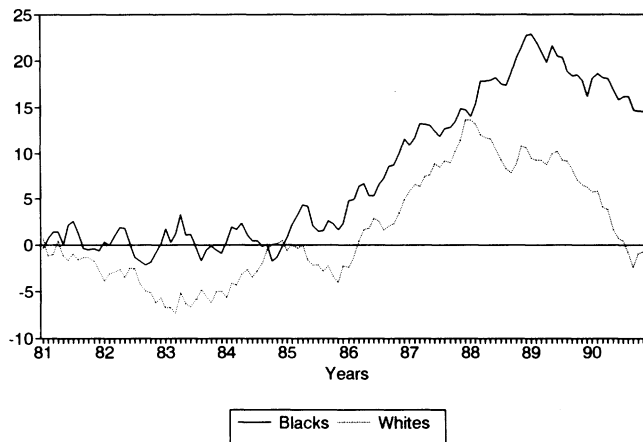


Figure 5. Cumulative Sum of the Residuals

series. The pattern of the cumulative sum of the residuals is consistent with Figure 1, where we observe two turning points in Black low birth weight series in 1985 and 1989. For Whites the cumulative sum of residuals exhibits an increase starting in 1986, and a decline after 1988. The magnitude of the increase in residuals is bigger for Blacks than Whites, which is again consistent with Figure 1 which indicates that shifts in the Black series are more notable than those in Whites.

As a second exercise, we estimated the Black low birth weight and the White low birth weight between 1978 and 1983, and obtained one-step-ahead forecasts. If the forecasts would capture the 1985 upswing and the 1989 downturn, this would support the hypothesis that there was no structural shift in 1985 or 1989. Put differently, it would imply that the own dynamics of the rate of Black and White low birth weight, rather than the influence of other factors was responsible for the observed behavior.

Figure 6A displays Black low birth weight along with the forecasts for 1984–90, which are based on the estimation using the sample 1978–83. As can be seen, neither the upswing in 1985, nor the downturn in 1989 is captured by the forecasts. In order to investigate how the inclusion of the percentage of Black pregnancies complicated by tobacco, the percentage of Black pregnancies where the consumption of prenatal care started in the first trimester, and the percentage of Black pregnancies complicated by drug use improve the forecasts, we added the logged values of each of these variables to the Black low birth weight equation separately, estimated the models between 1978 and 1983, and obtained the forecasts for 1984–90. Adding the percentage of Black pregnancies complicated by tobacco as a regressor did not improve the forecast accuracy. When we included the prenatal care variable, we noticed an improvement in forecasts. In particular, both turning points were captured by the forecasts of that model. Inclusion of the drug use variable as an explanatory variable did improve the forecasts more than that of prenatal care. Figure 6B presents the actual and predicted values of Black low birth weight where the model includes the percentage of Black pregnancies complicated by drug use as a regressor.³ A comparison of Figure 6B with Figure 6A demonstrates that inclusion of the drug use variable improves the forecasts of Black rate of low birth weight, and both turning points of low birth weight are captured

3. The signs of the tobacco, drug use and the prenatal care variables were positive, positive, and negative, respectively, as expected. None of them, however, was significant.

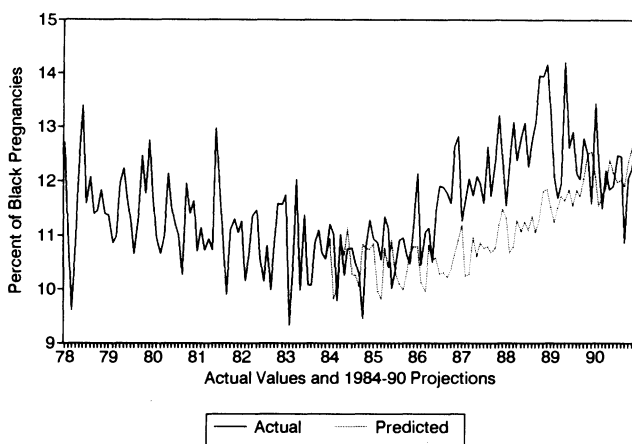


Figure 6A. Black Rate of Low Birth Weight

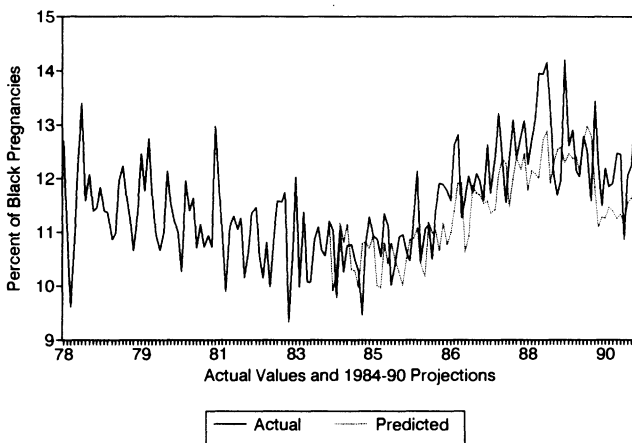


Figure 6B. Black Rate of Low Birth Weight with Drug Use as a Regressor

with precision. To sum, there is evidence that structural changes took place in 1985 and 1989 that altered the dynamics of Black the rate of low birth weight series. Consequently, one-step-ahead forecasts of the rate of low birth weight are not precise and do not capture the turning points. Adding the drug use variable to the model as a regressor increases the forecasts accuracy and enables the forecasts of the rate of Black low birth weight capture the turning points. Figures 7A and 7B display the same information for Whites. Unlike Blacks, however, modeling the White rate of low birth weight between 1978 and 1983 generates accurate forecasts for 1984–90 (see Figure 7A). A comparison of Figures 7A and 7B demonstrates that adding the rate of White pregnancies complicated by drugs as an explanatory variable to the model does not improve the accuracy of the forecasts significantly.

We also obtained the cumulative sum of residuals of the low birth weight equations for both races, where the drug use and the prenatal care variables were included as regressors. Since both explanatory variables are available after 1978, the models are estimated between 1978 and 1984, and the cumulative sum of residuals is calculated between 1985–90. They are displayed in Figure 8. A comparison with Figure 5 reveals that the inclusion of the regressors did improve

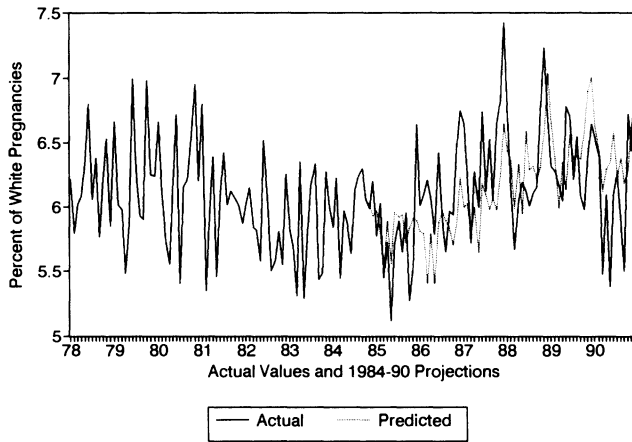


Figure 7A. White Rate of Low Birth Weight

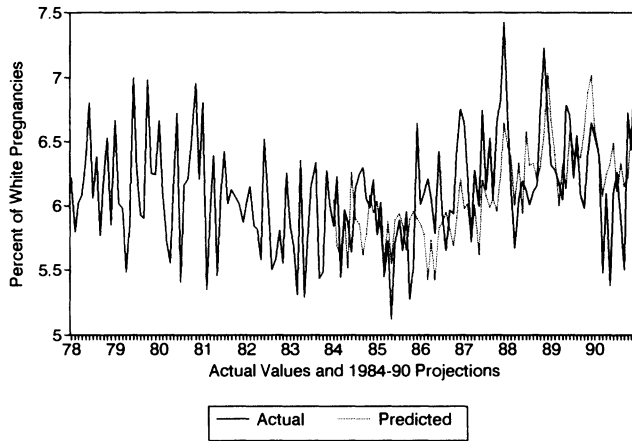


Figure 7B. White Rate of Low Birth Weight with Drug Use as a Regressor

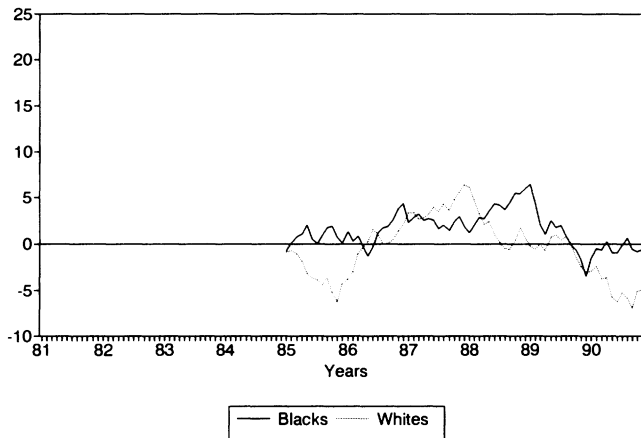


Figure 8. Cumulative Sum of the Residuals: Model with Prenatal Care and Drug Use

the cumulative sum of the residuals for Blacks, but it helped relatively little in case of Whites, which is consistent with the information provided by Figures 6 and 7.

Significance of the Drug Use and Projections for 1991 and 1992

To more formally test the impact of the use of drugs, smoking, and early initiation of prenatal care, we estimated the Black rate of low birth weight and the White rate of low birth weight using the Kalman Filter between 1978 and 1990 including drug, tobacco and prenatal care use as regressors. Table I reports the regression results for Blacks and Whites. To investigate the potential endogeneity of the regressors we performed a Wu-Hausman test as described by Nakamura and Nakamura [26]. We could not reject the null hypothesis that the explanatory variables are exogenous.⁴

Table I demonstrates that none of the explanatory variables has a statistically significant impact on the White rate of low birth weight. This outcome was expected given the ability of White rate of low birth weight to predict the 1984–90 values based on its dynamics only (Figure 7A), and is consistent with the results reported by Joyce, Racine and Mocan [18]. The lack of a relationship between the inputs and the White rate of birth weight is not surprising because unlike Blacks, the White rate of low birth weight does not exhibit significant variation during the period studied (see Figure 1). Furthermore, Figures 2, 3 and 4 indicate that the explanatory variables for the White regressions do not have much variation over time either, which would tend to inflate the standard errors of the estimated coefficients. This result, however, does not imply that the use of drugs or smoking is not a risk factor for White pregnant women at the individual level. It merely highlights the lack of an important association at the aggregate level. One explanation would be differential reporting between races. If White mothers are less likely to be tested for deleterious substances than Blacks, then it will be more difficult to find a relationship between substance abuse and the rate of low birth weight in the White population.

Table I also illustrates that early prenatal care is not a significant determinant of the Black rate of low birth weight. A one percent increase in the Black pregnancies complicated by illicit drugs generates a 0.10 percent increase in the Black rate of low birth weight holding constant the consumption of tobacco and medical care. The rate of Black pregnancies complicated by illicit drugs increased by 107 percent between 1983–84 and 1990 (see Figure 2). This implies that this particular rise in drug use generated a 10.4 percent increase in the rate of Black rate of low birth weight keeping the influences of tobacco and prenatal care constant. As Figure 1 demonstrates, the rate of Black rate of low birth weight increased by 14.7 percent between 1983–84 and 1990. Thus, the increased drug use among Black pregnant women in New York City between 1983–84 and 1990 accounts for 71 percent of the increase in the rate of low birth weight for Blacks.

For both races, we modeled the rate of pregnancies complicated by tobacco and the rate of pregnancies where the consumption of prenatal care started in the first semester using the Kalman Filter as depicted by (7)–(9), and obtained their 1991 and 1992 projections. Figures 9 and 10 present the actual values (1978–90) and the 1991–92 projections of these variables. The same is done for drug use, and the forecasts of 1991 and 1992 are displayed in Figure 11 along with the actual values. For both Whites and Blacks, Figure 11 illustrates three separate paths for 1991–92.

4. The Wu-Hausman test analyzes the correlation of the regressors with the residuals [37]. Instrumental variables were i) the maximum allowable benefits in constant dollars for a family of two under the Aid to Families with Dependent Children (AFDC) program in New York City, ii) the New York City unemployment rate, iii) the real minimum wage, iv) percentage of self financed births v) percentage of births financed by Medicaid.

Table I. Estimated Low Birth Weight Equations 1978–1990^a

| Explanatory Variables ^a | Dependent Variable | |
|------------------------------------|--------------------------------|--------------------------------|
| | Black Rate of Low Birth Weight | White Rate of Low Birth Weight |
| TOBACCO | 0.1074 (0.033) | 0.0012 (0.029) |
| MEDICAL CARE | 0.0450 (0.095) | -0.0435 (0.180) |
| DRUG | 0.0972 (0.026) | 0.0309 (0.039) |
| $\sigma_{\eta}^2 \times 10^3$ | 0.0000 | 0.6840 |
| $\sigma_{\xi}^2 \times 10^3$ | 0.0001 | 0.0000 |
| $\sigma_{\omega}^2 \times 10^3$ | 0.0002 | 0.0017 |
| $\sigma_{\epsilon}^2 \times 10^3$ | 2.5750 | 2.7700 |
| Normality | 3.576 | 3.746 |
| $H(47)$ | 0.858 | 0.845 |
| $Q(13)$ | 9.773 | 9.190 |
| R^2 | 0.588 | 0.268 |
| Wu-Hausman $F_{3,123}$ | 1.129 | 1.839 |
| Number of Observations | 156 | 156 |

a. Variables are in natural logs.
 $R^2 = 1 - [T^*p^2 / \sum(y_i - y_m)^2]$, where $T^* = T - d$, d is the degree of differencing, p^2 is the estimated one step ahead prediction error variance, and y_m is the mean of y_i .
 Normality = $(T^*/6)b_1 + (T^*/24)(b_2 - 3)^2$, where b_1 is the square of the third moment of the standardized residuals about the mean, b_2 is the fourth moment.
 $H(h)$ is a test for heteroscedasticity defined as

$$H(h) = \frac{\sum_{i=T-h+1}^T \nu_i^2}{\sum_{i=d+1}^{d+1+h} \nu_i^2}$$

where $h = T^*/3$, and ν_i is the standardized residuals. $H(h)$ has an $F(h, h)$ distribution.

Q is Box-Ljung Q-statistic for the joint significance of the first fifteen residuals. It is distributed Chi-square with the degrees of freedom in parentheses.

The numbers in parentheses are the standard errors.

The branches labeled *pw* and *pb* are the Kalman Filter projections of the rate of White pregnancies complicated by drug use, and the rate of Black pregnancies complicated by drug use, respectively. For Blacks, one observes another turning point in 1990. Put differently, the projection implies that the observed 1989 downturn in the rate of Black pregnancies complicated by drug use will be countered by another upturn in 1990, and the projected rate of Black pregnancies complicated by illicit drug use will reach 5.60 at the end of 1992. The Kalman Filter projection of the rate of White pregnancies complicated by illicit drug use (*pw*) demonstrates that it remains steady during 1991–92 with an average of 2.76. The branches labeled *h1* and *h3* in Figure 11 portray hypothetical paths of the same variable for Whites and Blacks, respectively. They are constructed so that at the end of 1992 the values of the drug use variables go down to their respective 1978 averages. In other words, they are representations of an imaginary scenario in which the rate of Black and White pregnancies complicated by drug use are brought back to their 1978 levels. The branches labeled *h2* and *h4* demonstrate the scenario where the percentage of pregnancies complicated by drug use for Whites and Blacks gradually increase to their respective 1989 levels.

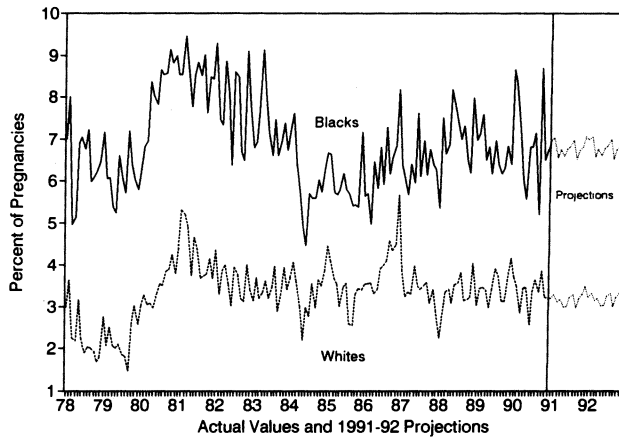


Figure 9. Percentage of Pregnancies Complicated by Tobacco

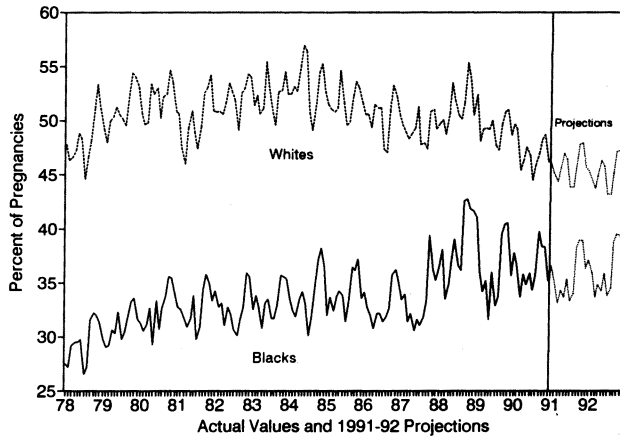


Figure 10. Percentage of Pregnancies with First Trimester Prenatal Care

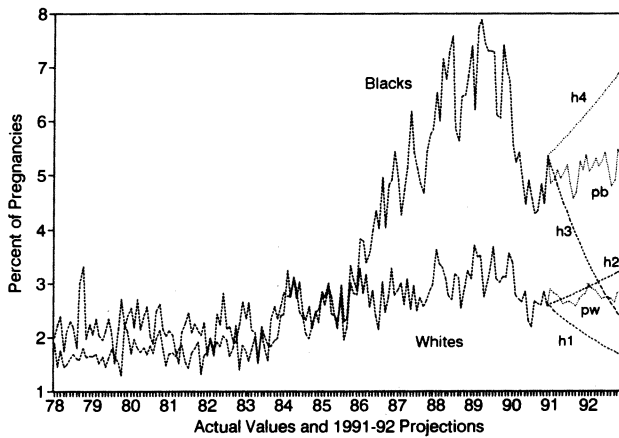


Figure 11. Percentage of Pregnancies Complicated by Drug Use

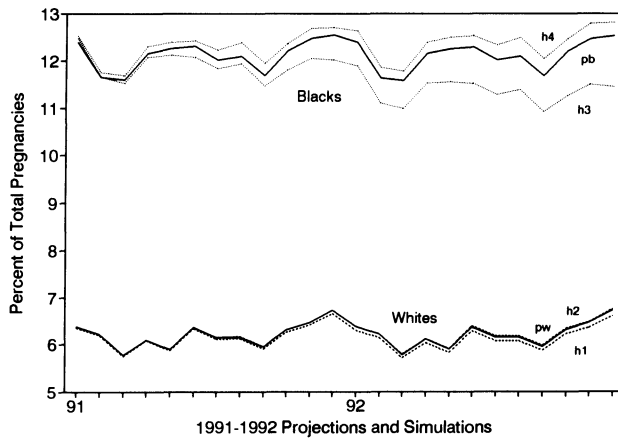


Figure 12. Rate of Low Birth Weight

We estimated the birth weight equations between 1978 and 1990, including the percentage of pregnancies complicated by tobacco, the percentage of pregnancies where the consumption of prenatal care started in the first trimester, and the percentage of pregnancies complicated by illicit drug use as regressors. Using the estimated coefficients and the 1991-92 projections of tobacco, drug and prenatal care variables, we obtained the forecasts of the rate of low birth weight for both races for 1991 and 1992. Figure 12 reports the three projections for each race. The projections *pw* and *pb* are obtained by using the 1991-92 Kalman Filter forecasts of the explanatory variables displayed in Figures 9-11. *h1-h4* of Figure 12 are the simulated paths for the White and Black rate of low birth weight, which are obtained by employing the same projected values of the tobacco and prenatal care, but using the hypothetical paths of drug use (*h1 through h4* in Figure 11). Therefore, the difference between projections *pb* and *h3* of Figure 12 represents the decline in the rate of Black rate of low birth weight that would have occurred, had the rate of Black pregnancies complicated by drug use gone down to its 1978 level instead of following its normal path, keeping the paths of the use of tobacco and prenatal care intact. Similarly, the difference between projections *pb* and *h4* illustrates the increase in the rate of Black rate of low birth weight that would have taken place, had the rate of Black pregnancies complicated by drug use gone up to its 1989 level by the end of 1992. For Whites, we notice that the branches *pw*, *h1* and *h2* are virtually the same in Figure 12. This means that the decrease of the rate of White pregnancies complicated by drug use to its 1978 level or its increase to the level prevailed in 1989 would not have an impact on the behavior of the White rate of low birth weight during 1991-92.⁵

According to Figure 12, employing the Kalman Filter projections of the rate of pregnancies complicated by tobacco, the rate of pregnancies where the consumption of prenatal care started in the first trimester, and the rate of pregnancies complicated by drugs for 1991-92, the predicted 1992 average of Black rate of low birth weight becomes 12.10 (branch *pb*). Using the same model and the same projections of tobacco and prenatal care, but forcing the drug use variable to go down to its 1978 level generates a reduction in Black rate of low birth weight, where the 1992 average becomes 11.36 (branch *h3*). This implies that if the drug use among Black pregnant women in 1992 went down to the level prevailed in 1978, the Black rate of low birth weight would be 0.74

5. Once again, this result is not surprising given the lack of a relationship between the rate of White low Birth weight and the health inputs. Nevertheless, we kept Whites for completeness.

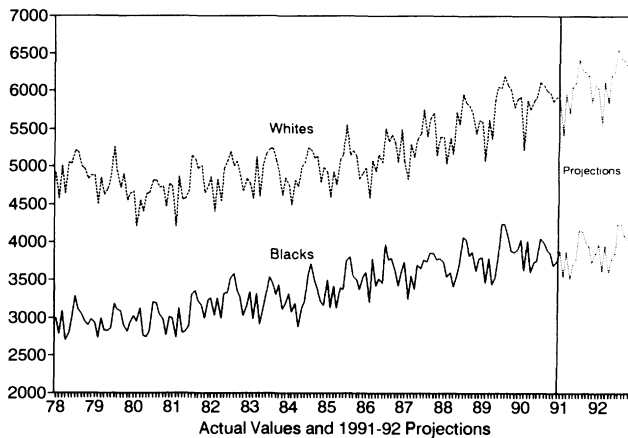


Figure 13. Number of Births

percentage points less than the one to be observed in 1992. On the other hand, all else being constant, if the rate of Black pregnancies complicated by drugs went up to its 1989 level (branch *h4* in Figure 11), this would generate the simulated path *h4* of Figure 12 for the Black rate of low birth weight, where the 1992 average becomes 12.39.

We modeled the monthly number of Black and White births using the procedure discussed earlier and obtained their 1991–1992 projections which are displayed in Figure 13. Multiplying the 1991–92 projections of the number of births by the projected rate of low birth weight rates yields the projected number of low birth weight births for the period 1991–92. Using this algorithm we find that the projected total number of low birth weight births between January 1991 and December 1992 is 11,289 for Blacks and 9,068 for Whites. If by the end of 1992 the rate of Black pregnancies complicated by drug use could be pulled down to its level prevailed in 1978 (path *h3* of Figure 11), the resultant outcome would be a total of 10,845 low birth weight Black births for the period 1991–92. Thus, 444 Black low birth weight births would be averted during the period of 1991–92 by reducing the maternal drug use to its 1978 level.

It should be noted that this amount is the result of a gradual decline in the drug use over a period of two years. As Figure 12 demonstrates, the decline in the Black rate of low birth weight increases as the drug use variable approaches to its pre-crack epidemic level. The average monthly number of low birth weight births during the period of October 1992–December 1992 is 447 given that the drug use reaches its pre-epidemic level in these months. The projected monthly average Black low birth weight births for the same period is 487 given that the drug use follows its normal (status quo) path. Thus, an average of 40 Black low birth weight births would be averted per month with the drug use being reduced and kept at its 1978 level. This implies that the monthly number of averted Black low birth weight births would approximately be 8% of total Black low birth weight births (40/487).

The marginal cost of treating infants who are exposed to drugs is between \$9,313 and \$13,225 in 1993 dollars for the initial hospitalization [1, 29]. This implies that in the first two years \$4.1 to \$5.9 million in initial hospitalization cost could be saved in New York City by lowering the drug use among Black pregnant women to its 1978 level. After the first two years, where the intervention has pulled the Black pregnancies complicated by drug use to the level prevailed in

1978, annual savings in initial hospitalization costs would be between \$4.5 and \$6.3 million in 1993 dollars.

The total social cost of low birth weight births is higher than portrayed by the ones associated with initial hospitalization. For example, Chaikind and Corman [2] show that low birth weight babies have a higher probability of needing special education because of learning disabilities and emotional problems. More precisely, they show that low birth weight babies are almost fifty percent more likely to be enrolled in any type of special education than babies who are of normal weight at birth, controlling for the individual, family and regional characteristics. This implies that 222 Black children who were born in 1991–1992 and an additional 240 every year after that could avoid special education if the maternal drug use went down to its pre-crack epidemic level. Chaikind and Corman cite a study by Moore et al. [25] where excess cost of special education (the total cost required to educate a special education student minus the costs to educate a regular student) are estimated as \$4,350 per pupil in 1989–90 dollars. Thus, an additional \$1.1 million (in 1993 dollars) in terms of avoided special education costs could have been saved by decreasing the maternal drug use to its 1978 level. Therefore, \$5.2 to \$7.0 million in 1993 dollars could have been saved in terms of avoided initial hospitalization and special education costs pertaining to Black rate of low birth weight babies born in New York City between 1991–92 if an intervention program could bring the maternal drug use among Black pregnant down to its level prevailed in 1978. After the prevalence of prenatal illicit substance use reached and remained at its 1978 level, the initial hospitalization and special education savings would amount to \$5.7 to \$7.5 million annually (in 1993 dollars) with respect to status quo.

VI. Conclusion

Using monthly data from New York City that span the years 1978–90, we investigate the relationship between the incidence of drug use during pregnancy and the rate of low birth weight. Using Kalman Filter we model the dynamics of the White and Black rate of low birth weight between 1963 and 1980, and obtain the cumulative sum of the residuals which signal structural shifts in 1984 and 1989. When we estimate the model between 1978 and 1983, and obtain forecasts for 1984–90, we get accurate forecasts for Whites, but inaccurate ones for Blacks. Including the percentage of pregnancies complicated by the use of illicit drugs as a regressor to the model increases the forecast accuracy significantly. More precisely, the upturn in the Black rate of low birth weight that took place in 1984 and the downturn of 1989 are captured. Estimating the models between 1978 and 1990 reveals that smoking during pregnancy, early initiation of prenatal care or the use of drugs during pregnancy have no impact on the rate of White low birth weight. This result does not imply that the weight of White babies is not influenced by the use of deleterious substances during pregnancy. It indicates that the variations in maternal drug use, smoking and the use of prenatal care are not serious enough to generate a change in the White rate of low birth weight. The lack of variation in these variables, however, may be due to differential reporting between Whites and Blacks, where Black mothers may be subjected to drug tests at a higher rate than their White counterparts [4]. Smoking and drug use are significant determinants of low birth weight for Blacks at the aggregate level. We find that the rise in pregnancies complicated by drugs accounts for 71 percent of the increase in the rate of Black low birth weight that took place between 1983–84 and 1990. Furthermore, our simulations reveal that given the projected paths

of smoking and prenatal care, if drug use among Black pregnant women had been reduced back to its 1978 level between January 1991 and December 1992, we would have had 444 less Black low birth weight babies during the same period. This implies that between \$5.2 and \$7.0 million in 1993 dollars would have been saved in initial hospitalization and special education costs for that period. Furthermore, after the first two years if the level of maternal drug use is kept at its pre-crack epidemic level, this reduces the number of Black low birth weight babies by 8%, or 40 births per month with respect to the level that would have observed in the absence of any intervention. Thus, \$5.7 to \$7.5 million (in 1993 dollars) per year would be saved in terms of initial hospitalization and special education costs as a result of reducing the prevalence of prenatal illicit substance use to its pre-crack epidemic level.

Appendix

The model described by equations (7)–(10) can be put into state space form which consists of the following equations.

$$LBW_t = \mathbf{z}'_t \boldsymbol{\alpha}_t + \epsilon_t \quad t = 1, \dots, T \quad (11a)$$

$$\boldsymbol{\alpha}_t = \mathbf{A}_t \boldsymbol{\alpha}_{t-1} + \boldsymbol{\nu}_t \quad t = 1, \dots, T \quad (11b)$$

Equations (11a) and (11b) are the observation and transition equations, respectively. $\boldsymbol{\alpha}_t$ is an $(m \times 1)$ unobservable state vector, \mathbf{Z}_t is an $(m \times 1)$ fixed vector, \mathbf{A}_t is a non-stochastic $(m \times m)$ matrix, ϵ_t is a serially independent, normally distributed irregular component with mean zero and variance σ_ϵ^2 . Equation (11b) demonstrates that the state vector is updated each period, but it is also subject to some serially uncorrelated random distortions with zero mean and covariance matrix $\boldsymbol{\Omega}_t$, represented by the $m \times 1$ vector $\boldsymbol{\nu}_t$.

After expressing the model in terms of state space representation [Equations (11a) and (11b)], maximum likelihood estimates of the parameters of the structural model are obtained in the time domain and the Kalman filter is used for updating the unobserved components. If \mathbf{a}_{t-1} is an estimate of $\boldsymbol{\alpha}_{t-1}$, and \mathbf{P}_{t-1} is its covariance matrix, then the optimal (minimum mean square error) linear projections of \mathbf{a}_t and \mathbf{P}_t at time $t - 1$ are

$$\mathbf{a}_{t|t-1} = \mathbf{A} \mathbf{a}_{t-1} \quad (12)$$

$$\mathbf{P}_{t|t-1} = \mathbf{A} \mathbf{P}_{t-1} \mathbf{A} + \boldsymbol{\Omega}_t. \quad (13)$$

The Kalman filter updates the already available optimal predictor $\mathbf{a}_{t|t-1}$ with the new information contained in LBW_t , as follows [12].

$$\mathbf{a}_t = \mathbf{a}_{t|t-1} + \mathbf{P}_{t|t-1} \mathbf{Z}_t (LBW_t - \mathbf{Z}'_t \mathbf{a}_{t|t-1}) / f_t \quad (14)$$

$$\mathbf{P}_t = \mathbf{P}_{t|t-1} - \mathbf{P}_{t|t-1} \mathbf{Z}_t \mathbf{Z}'_t \mathbf{P}_{t|t-1} / f_t \quad (15)$$

where

$$f_t = \mathbf{Z}'_t \mathbf{P}_{t|t-1} \mathbf{Z}_t + \sigma_\epsilon^2.$$

Note that the term in the parentheses in equation (14) is the prediction error. Thus, equation (14) demonstrates that the predictor $\mathbf{a}_{t|t-1}$ is updated by incorporating the prediction error, weighted by $\mathbf{P}_{t|t-1} \mathbf{Z}_t / f_t$, which is the Kalman gain. Similarly, the new covariance matrix \mathbf{P}_t in equation (15) is equal to the prior covariance matrix minus $\mathbf{Z}'_t \mathbf{P}_{t|t-1}$ weighted by the Kalman gain.

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