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Marital Status and Infant Mortality

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Abstract

This paper examines the relationship between marital status and infant mortality in Jamaica. Discrete-time hazard models that account for unobserved heterogeneity are estimated using the demographic histories of the 1975/76 Jamaican Fertility Survey. The analysis indicates that marital status is an important factor in explaining differences in infant mortality. More specifically (and contrary to what is found unconditionally), infant mortality is higher in common-law (and visiting) unions compared to marriage, after other factors thought to affect infant mortality are held constant.

Résumé

Cet article examine le rapport entre l'etat matrimonial et le taux de mortalitw infantile. En utilisant l'analyse demographique des biographies de l'enquête de fertilité jamaquaine, on a estimé des modéles de temps-discretedes quotients instantanes qui tiennent en compte l' hétérogénéité non observé. L'analyse a revelé que l'etat matrimonial est un élément important dans l'explication des differences des taux de mortalité infantile. Plus précisément, quand on tiene en constant les autres factures qu'on pense peuvent avoir un effet sur le taux de mortalitw infantile, on voit que le taux de mortalité infantile est plus haut dans le concubinage et l'union visitantes que dans le marriage. Ce resultat est significant, parce que c'estau contraire de ce qu'on a observé inconditionnellement.

***Key Words**: infant mortality, martial status, hazard models*

Introduction

Studies of family patterns in the British Caribbean usually distinguish three types of fertile sexual unions (Harewood, 1964). The first type is marriage – a union in which a man and a woman are legally married and are living together in the same household. The second type is a common-law union – a union in which a man and a woman are living together but are not legally married. And the third type is a visiting union – a union in which a couple do not live together but do have regular sexual relations. There is a substantial amount of movement between these different union types (see for example, Ebanks, George and Nobbe, 1974; Wright, 1989). For example, a sexual relationship may begin as a visiting union, proceed to a common-law relationship, and then on to marriage. Many other combinations are possible, and of course, any of these relationships may end with the partnership terminating. On average, visiting unions are the shortest in duration, followed by common-law unions and then marriages (Burch, 1983). Therefore, in terms of temporal stability, visiting unions are the most unstable form; marriages are the most stable; and common-law unions fall in an intermediate position.

Little attention has been directed towards evaluating the impact of marital status on infant mortality, despite the fact that the determinants of infant mortality have been intensively researched both theoretically and empirically (see for example the comparative empirical studies of Hobcraft, McDonald and Rutstein, 1983, 1984; and the recent theoretical review by Wolpin, 1997). This neglect is surprising given that there are a variety of mechanisms by which family patterns of the type described above could have an impact on infant mortality. For example, it is well documented that there is a relationship between marital status (or more generally sexual union status) and fertility in Jamaica, with fertility varying considerably in the three main types of sexual unions (see Ebanks, 1985; Wright and Madan, 1988). Likewise, there appears to be a relationship between infant mortality and fertility, with higher infant mortality supporting higher fertility (see Ebanks, 1985). Given there is relationship between fertility and infant mortality in Jamaica, and given there is relationship between fertility and sexual union status in this country, it seems reasonable to hypothesize that there should also be a relationship between sexual union status and infant mortality.

A second possible reason for expecting a relationship between sexual union status and infant mortality concerns the extent to which the amount of resources flowing into the household might differ by sexual union type. As mentioned above, in visiting unions the couple do not live and these unions are (on average) the shortest in terms of duration. This suggests that the couple's level of 'commitment' in visiting unions (for whatever reasons) is lower compared to common-law unions and marriage (where the couple are living together). This lower level of commitment may result in the male partner devoting fewer resources (measured in terms of money and time) towards the upkeep of the household and the rearing of children. If this is the case, the woman will likely be required to work more in order to insure that sufficient material resources are available (e.g. work longer hours, hold more than one job, etc.). This hypothesis is consistent with the observation that labour supply varies considerably by sexual union type, with participation rates being highest for women in visiting unions compared to marriage and common-law unions (see Wright, 1988). However, if the woman is working more, *ceteris paribus*, she has less time to devote to the rearing of her children. Since parental time and money are thought to be the main investments in 'child quality' (which is thought to be strongly correlated with infant mortality), if such investments differ by sexual union type, infant mortality should also vary by sexual union type.

A third possible explanation for a relationship between sexual union status and infant mortality is based on assumption that entry into different types of sexual unions is a selective process (1). That is, social and economic characteristics that make women and men better suited to marriage (for example), may be the same characteristics that result in lower infant mortality amongst their children. Put slightly differently, individuals who are selected into marriage may also be selected into reproductive outcomes that are more favourable, one of which is lower infant mortality (see Goldman, 1993). In this sense there is nothing about marriage 'per se' that affects infant mortality, since both 'outcomes' are caused by a common set of factors. Although the analysis carried out in this paper cannot distinguish between the three possible explanations outlined here (and therefore cannot comment on their relative importance), taken together they suggest that there should be a relationship between sexual union status and infant mortality.

Ebanks (1985) is the only study that I could find that has examined empirically the relationship between sexual union status and infant mortality. He calculated infant mortality rates, using data from the World Fertility Survey, for Jamaica, Guyana and Trinidad and Tobago. For Jamaica he finds an infant mortality rate for marriages of 50.3 per 1,000 births; for common-law unions a rate of 57.3; and for visiting unions a rate 46.5, suggesting that infant mortality is highest in common-law unions and lowest in visiting unions (Ebanks, 1985: 58). The rates were calculated by comparing current union status (or the last union status for women not currently in a union) with past fertility/mortality events. The problem with this approach is that it is prone to measurement error because there is a considerable amount of movement between these different types of unions. For example, a woman may currently be in a marriage (e.g. at the time of the survey), but her first birth may have occurred in a visiting union. Because of this potential mismatching, these estimates should be viewed with some caution. To avoid this problem, the approach that this paper adopts (as described below) is to include the sexual union status variables as time-varying covariates in a hazard regression model of infant mortality.

The purpose of this paper is to examine empirically relationship between sexual union status and infant mortality using data collected in the 1975-76 Jamaican Fertility Survey. The remainder of this paper is organized as follows. In the next section, the data used in the analysis is described. In the second section, the statistical models are presented. The third section describes the variables included in the models. In the fourth section, the results are presented and discussed. A brief conclusion follows in the last section.

Methods and Materials

Data

The data set used in this study is the 1975-76 Jamaican Fertility Survey (JFS), which was carried out as part of the World Fertility Survey Programme (see Singh, 1982). A probability sample of women between the ages of 15 and 49 was interviewed (N = 3,096) and detailed demographic and socio-economic information was collected. Of particular importance, from the point of view of this paper, was the inclusion of questions concerning past union statuses and fertility behaviour. The union status questions included the order of the union (i.e. first union, second, union, etc); the type of union (visiting, common-law or marriage); the starting date of the union (measured in calendar month time); and the finishing date of the union (if applicable). The fertility questions included the order of the birth; the date of the birth (measured in calendar month time); and the date of death of the child (if applicable). Given this event history information, it is relatively easy to determine whether an infant death occurred in a particular month and then relate that event to the type of union the women was member of in that month.

There are some problems concerning the quality of event history data collected by retrospective questioning, since individuals are asked to supply information about events that occurred in the past. Either intentionally or unintentionally, date misreporting occurs and events are forgotten and go unrecorded. The seriousness of these errors usually increase the farther back in time the event took place. In an empirical assessment of the JFS, Singh (1982) concluded that the data are 'quite good' and the union status and fertility histories are 'more complete' than similar data collected in censuses or through vital registration. Furthermore, a great deal of time and effort was directed towards checking the dates in the JFS, obvious discrepancies were corrected, and in some cases missing dates were imputed (see Trussell, 1987). Nevertheless, despite these problems and age of the data, the JFS is still the best available nationally-representative data set in which to examine empirically the relationship between sexual union status and infant mortality.

One of the clear advantages of the JFS is the relatively large sample size. Of the 3,092 women interviewed in the JFS, 2,457 had at least one birth (see Table 1). In total, there were 9,887 births recorded in the demographic histories (probably a slight under-estimate) and 523 infant deaths (also probably a slight under-estimate).

Statistical Model

In order to analyze the effect that sexual union status has on infant mortality, the risk of a child dying in each of the first twelve months of life (i.e. $t = 1, 2, ..., 12$) is modeled as a discrete-time hazard rate:

$$(1) \qquad P_{it} = \text{Prob}\ (T_i = t \mid T_i \geq t)\ ,$$

where T is a random variable giving the uncensored time of the child's death. The hazard rate is simply the conditional probability of the child dying at duration t (i.e. in a particular month), given that the child has not already died (see, Allison, 1982; Wright, 1998).

One popular way of relating the hazard rate to a vector of explanatory covariates, X_{it}, including a duration dependence specification, is through the so-called proportional hazard model. Such models are now routinely used to analyze the determinants of infant (and child) morality (see for example, Trussell and Hammerslough, 1983; Wolpin, 1997). In discrete-time, a close approximation to this model is:

$$(2) \qquad P_{it} = \frac{\exp\ (\alpha X_{it})}{1 + \exp\ (\alpha X_{it})}\ ,$$

where α is a vector of unknown parameters to be estimated (see, Allison, 1982). This is simply a logistic regression model. Therefore the associated likelihood function is straightforward and can be maximized using most standard statistical packages.

One problem with this simple model is that it is based on the assumption of 'perfect specification,' in the sense that it is assumed that all the relevant differences across individuals are captured by the covariates included in the model. If this is not the case, then parameter estimates may be biased and incorrect patterns of duration dependence may be observed. For example, it is well known that this problem of 'unobserved heterogeneity' tends to bias estimates of duration dependence downwards, which is especially problematic if one is concerned with testing hypotheses about the nature of duration dependence.

Despite the obvious importance of unobserved heterogeneity, there is considerable disagreement concerning what is the preferred way in which to 'control' for it in hazard models of mortality; there is even less agreement on what statistical methods should be employed (see Heckman, J.J. and B. Singer, 1984; Trussell and Richards, 1985; Trussell and Rodriguez, 1990;

Vaupel, Manton and Stallard, 1979). One way to model the effects of unobserved heterogeneity is to include an individual-specific nuisance parameter, ε_i, as a linear predictor in Equation (2):

$$\text{(3)} \qquad \text{Prob}\,(S_i \mid X_{it}) = \prod_{t=1}^{T_i} \frac{[\exp(aX_{it} + \varepsilon_i)]^{Y_{it}}}{1 + \exp(aX_{it} + \varepsilon_i)}$$

Assuming ε_i are drawn from a single mixing distribution with density $g(\varepsilon)$ and are independent of the covariates X_{it}, integrating out e_i gives the likelihood function:

$$\text{(4)} \qquad L_i(\alpha) = \int \prod_{t=1}^{T_i} \frac{[\exp(aX_{it} + \varepsilon_i)]^{Y_{it}}}{1 + \exp(aX_{it} + \varepsilon_i)}\, g(\varepsilon_i)\, d\varepsilon_i$$

Assuming a Normal parametric form for $g(\varepsilon)$, the Gaussian quadrature method for the numerical evaluation of the integral in Equation (4) may be used. The associated likelihood is identical to the logistic variance component model of Anderson and Aitkin (1985):

$$\text{(5)} \qquad L_i = \sum_{j=1}^{Q} \left[\prod_{t=1}^{T_i} \frac{[\exp(\alpha X_{it} + \omega z_i)]^{Y_{it}}}{1 + \exp(\alpha X_{it} + \omega z_i)} \right] \beta_j$$

where z_j are the fixed quadrature location points; β_j are the corresponding probabilities; and ω is the (unknown) standard deviation of the mixing distribution. Clearly if ω is small (i.e. not statistically significant to zero), then given the assumptions of the model, degrading unobserved heterogeneity is not present and the included covariates appear to capture the important heterogeneity across individuals.

There are various ways of maximizing this likelihood function. All are computational expensive. In this paper, the method used combines the Newton-Raphson algorithm with the Berndt, Hall, Hall and Hausman (1974) method of estimating the second derivatives from the variance-covariance matrix of first derivatives. The software package SABRE was used to perform the estimation (see Barry, Francis and Davies, 1990; Wright, 1991).

Covariates

The variables included in the models are summarized in Table 1. Sexual union status is measured by four time-varying binary covariates. Three of these variables correspond to each of three main types of sexual unions already discussed: visiting, common-law and marriage. However, the fourth category – 'not in union' – is a residual category. This variable is coded '1' if the woman is not married, not in a common-law union and not in a visiting union and coded '0' otherwise. It is important to point out that the 'not in union' category does not necessarily mean that the woman is not having sexual relations (although this possibility is included in this category) and therefore she may very well be exposed to the risk of pregnancy. It is simply the category that she is placed in if she is not a member of one of the three main union types.

As discussed above, there is considerable movement between these different types of sexual unions. Therefore, the four variables used to measure sexual union status are included as time-varying covariates in the sense that they are allowed to change in value on a month-to-month basis. The importance of including sexual union status in such a way is best illustrated using an example. Assume a child is born when the woman is a member of a visiting union. Assume further that the child dies in his/her twelfth month of life (i.e. an infant death is recorded). In the child's fourth month of life, the man moves into the household of the woman and the relationship becomes a common-law union. In the child's eighth month of life, the relationship between the father and mother ends and the man leaves the household. Until the child's death in the twelfth month, the woman is not having sexual relations and therefore she is 'not in union.' In this example, the child spent 4 months exposed to the risk of dying in a visiting union, 4 months exposed to the risk of dying in a common-law union, and 4 months exposed to the risk of dying when the mother was not in a union. The use of time-varying covariates therefore allows one to accurately allocate the exposures-to-risk associated with each of the different sexual unions types.

Sexual union status is not the only determinant of infant mortality and other variables are included in the models. These variables are included primarily as control variables and are not of direct substantive interest given the focus of this paper. Their inclusion simply allows one to examine the relationship between sexual union status and infant mortality holding constant other known determinants of infant mortality. These variables include: (1) the age of the women at time of birth along with its square; (2) the woman's education measured as the number of years of schooling completed; (3) the calendar year of the birth; and (4) the order of the birth. All these variables are by nature or by choice fixed covariates.

The models are estimated based on the assumption that the underlying pattern of duration dependence follows a Waybill distribution. With this assumption, the hazard rate is allowed to increase or decrease monotonically

Table 1
Variables Included in Models

Sexual Union Status Variables:

Not in union[1]	Binary variable ($\bar{X}$= 11.2%).
Visiting union[1]	Binary variable ($\bar{X}$ = 28.9%).
Common-law union[1]	Binary variable ($\bar{X}$ = 11.2%).
Marriage[1]	Binary variable reference category ($\bar{X}$ = 24.9%).

Control Variables:

Age at birth[2]	Age of the women at time of the birth measured in years ($\bar{X}$ = 24.9, σ = 6.3).
Age at birth squared[2]	($\bar{X}$ = 661.1, σ = 342.2).
Years of schooling[2]	Number of years of schooling completed by women ($\bar{X}$ = 7.0, σ = 2.1).
Year of birth[2]	The calendar year of the birth ($\bar{X}$ = 64.8, σ = 7.2, Range = [19]39-[19]76).
Birth order[2]	The order of the birth ($\bar{X}$ = 3.5, σ = 2.5).

Number of women	2,457
Number of births	9,887
Number of infant deaths	523
Months of exposure	111,677

Notes. 1. Time varying covariate.
2. Fixed covariate.

Source. 1975-76 Jamaican Fertility Survey through time. In discrete-time this form of duration dependence is modeled by simply including the natural logarithm of duration (time) as an additional variable in the vector of explanatory covariates i.e. $\gamma\, ln(t)$. If $\gamma < 0$ there is negative duration dependence and the underlying risk of dying decreases with time. If $\gamma > 0$ there is positive duration dependence and the risk of dying increases with time. If $\gamma = 0$ then the Waybill distribution reduces to the exponential distribution and there is no duration dependence. It is worth noting that other forms of duration dependence where considered (e.g. piece-wise constant and quadratic). However, the estimates were not sensitive to these alternative specifications. Therefore, for brevity and simplicity, only the estimates based on the Waybill specification are presented.

Results

Tables 2 and 3 report the estimates of the models. Table 2 presents the parameter estimates of the model, along with standard errors and other summary statistics. Table 3 expresses the parameters of the union status variables as ratios of relative risk, which are simply the exponential of the parameters given in Table 2. Since the excluded union status variable is marriage, the relative risk associated with this type of union is 1 [i.e. *exp*(0) = 1], to which the risks associated with the other types of unions can be easily compared. The first two columns of Table 2 are the estimates that do not control for unobserved heterogeneity. The last two columns are the estimates that attempt to do so. Columns 1 and 3 are the estimates of models where only the union status variables are included. Columns 2 and 4 are the estimates where the control variables are added.

Turning first to the estimates that do not control for unobserved heterogeneity, when no control variables are included in the model, a clear gradient is observed with respect to the impact of union status on infant mortality (Column 1). The risk is highest for women 'not in union' (1.46) and lowest for women in marriage (1.0). The risks associated with women in visiting and common-law unions are very similar and fall between the risks associated with women who are married and not in union (i.e.1.30 and 1.32, respectively). These estimates suggest that risk of experiencing an infant death is lower in marriage and higher in the other three union statuses.

The picture changes somewhat when the control variables are included (Column 2). The risks associated with women in visiting and common-law unions and not in union are not significantly different from one another (1.26, 1.24 and 1.24 – confirmed by a likelihood ratio test) and are higher than for women in marriage (risk = 1.0). These estimates suggest that when other variables that are known to affect infant mortality are taken into consideration, the risk of experiencing an infant death is still lower in marriage, compared to other types of unions.

Table 2
Parameter Estimates
of Infant Mortality Hazard Models, Jamaica
[Standard Errors in Parentheses]

Variable	Model (1)	(2)	(3)	(4)
Not in union	0.380	0.213	0.446	0.250
	[0.152]	[0.167]	[0.161]	[0.176]
Visiting union	0.265	0.234	0.332	0.266
	[0.129]	[0.141]	[0.135]	[0.151]
Common-law union	0.278	0.215	0.319	0.242
	[0.124]	[0.129]	[0.129]	[0.130]
Age at birth	---	- 0.094	---	- 0.093
		[0.048]		[0.053]
Age at birth /100	---	0.140	---	0.150
		[0.089]		[0.100]
Years of schooling	---	- 0.052	---	- 0.062
		[0.021]		[0.022]
Year of birth	---	-0.043	---	- 0.045
		[0.006]		[0.007]
Birth order	---	0.088	---	0.074
		[0.027]		[0.026]
α	- 3.923	0.311	- 4.326	0.135
	[0.108]	[0.737]	[0.140]	[0.797]
γ	- 1.385	- 1.390	- 1.373	- 1.380
	[0.056]	[0.056]	[0.046]	[0.046]
ω	---	---	0.850	0.795
			[0.089]	[0.096]
-2. lnL	5959.3	5884.2	5920.4	5855.8

Notes: Number of women = 2,457; number of deaths = 523; number of Births = 9,887; and months of exposure = 111,677.

Source: 1975-76 Jamaican Fertility Survey

Table 3.
Relative Mortality Risks Associated with Sexual Union Status, Jamaica

	Model			
Union type	(1)	(2)	(3)	(4)
Not in union	1.46	1.24	1.56	1.28
Visiting union	1.30	1.26	1.39	1.30
Common-law union	1.32	1.24	1.38	1.27
Marriage	1	1	1	1

Notes: Relative risks are exponential of the sexual union status parameters given in Table 2. Since marriage is the excluded category, the risk for this category is one i.e. exp(0) = 1.

The estimates that attempt to control for unobserved heterogeneity reveal a similar pattern with respect to union status (Columns 3 and 4). The model that includes only the union status variables (Column 3) confirms that risk is highest for women 'not in union' (1.56) and lowest for women in marriage (1.0). Likewise, the risks associated with women in visiting and common-law unions are very similar and fall between the risks associated with women who are married and not in union (i.e. 1.39 and 1.38, respectively). However, there appears to be significant unobserved heterogeneity present given the large value of magnitude of ω and the small value of its standard error (0.850 and 0.089 respectively). However, it is very encouraging that despite the presence of unobserved heterogeneity, the parameter estimates are very similar.

When the control variables are added to the model that controls for unobserved heterogeneity (Column 4), the estimates are again very similar to what was found when no control was attempted (Column 2). Again the risks associated with women in visiting and common-law unions and not in union are not significantly different to one another (1.30, 1.27 and 1.28), but are still much higher than the risk associated with women in marriage (1.0). In addition, despite the inclusion of the control variables, there still seems to be considerable unobserved heterogeneity given the large value of the ω parameter and its small standard error (0.795 and 0.096, respectively).

Turning to the impact of the control variables, the estimates are in agreement with what has been found in most empirical studies of infant mortality. There appears to be a non-linear relationship between the age of the women at the time of the birth and infant mortality, with the risk of experiencing an infant death being higher for 'younger' and 'older' women and reaching a minimum around age 33. Women's education is also important. Children of women who have higher levels of schooling have lower mortality risks. The year of birth variable is negative, suggesting that infant mortality decreased considerably across the period covered by the JFS data (i.e. 1939 to 1976). It appears that birth order has a positive impact on infant mortality, with the risk of dying being higher for higher order births. Unlike some other studies, we find no evidence for a J-shaped relationship between birth order and infant morality Finally, It is important to note that negative duration dependence is observed. In all the models, $\gamma < 0$ and is statistically significant. This suggests that the risk of a child dying declines as the child ages. This finding of negative duration dependence is consistent with most other hazard rate models of infant mortality (see Wolpin, 1997).

Conclusion

This paper has examined the relationship between sexual union status and infant mortality in Jamaica using data collected in the Jamaican Fertility Survey. Discrete-time hazard models that account for unobserved heterogeneity are estimated using the demographic histories of the 1975/76 Jamaican Fertility Survey. The analysis indicates that marital status is an important factor in explaining differences in infant mortality. The main finding is that infant mortality is higher in common-law and visiting unions compared to marriage, after other factors thought to affect infant mortality are held constant. The risk of a child dying in the first twelve months of life is clearly lower when that child is born in marriage.

There are numerous ways in which the analysis carried out in this paper can be extended in order to more rigorously evaluate the relationship between sexual union status and in mortality. The analysis could be replicated for other countries that have family patterns similar to those observed in Jamaica. Such patterns occur in Guyana and Trinidad and Tobago (amongst individuals of African descent) and both these countries participated in the World Fertility Programme. The data used in this paper are now quite old, collected in 1975/76. Although the family patterns described in this paper persist in Jamaica, infant mortality has continued to decline over the past two and half decades. It would be useful to examine this relationship using more recent data in order to see if the differential documented in this paper is still observed. However, I am unaware of any recently collected data-set capable of supporting such an analysis.

This paper has focused on infant mortality. The analysis could also be easily extended to child mortality (e.g. mortality up until the age of five). Likewise, an analysis could be carried out focusing separately on neonatal mortality (i.e. deaths occurring in the first month of life) and post-neonatal mortality (i.e. deaths occurring in months two to twelve). Such an analysis would be useful since it is often argued that the social and economic factors are more important determinants of post-neonatal mortality compared to neonatal mortality where it is thought that biological factors are more important. Such an analysis, although straightforward to carry out, would be hampered by issues relating to sample selection bias resulting from splitting the sample into separate groups based on the variable you are interested in modeling (i.e. mortality) and the well-known problems associated with using data collected in demographic histories to examine neonatal mortality (see Goldman, Pebbly and Lord, 1984). Finally the models could be estimated separately for each parity since there appears to be a relationship between birth order and infant mortality. However, controlling for unobserved heterogeneity is more difficult in the single-spell models that would be used in a parity-specific analysis.

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Endnotes:

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