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ABSTRACT

Primary health care programs, including childhood immunizations, are intensively utilized by the majority of Latin American households. During 1985-1991, governments used national PHC infrastructures to implement the Regional Polio Eradication Initiative. Unprecedented vaccine coverage levels were attained through a combination of mass campaigns, house-to-house vaccinations and improvements in routine immunization services. Little is known, however, about the effects of these interventions on immunization demand; whether they reached so-called high-risk households and, if so, whether program effects were sufficient to offset the negative risk factors. Using the household production of health model, this paper examines the probability and timing of full immunization over this period in one case country, Colombia. Information on the immunization status of 3609 vaccine-eligible children born 1985-90 was extracted from Colombia's 1990 Demographic and Health Survey. Annual immunization coverage estimates from the Colombian Ministry of Health for 1985-90 for 148 sample *municipios* were appended to each child record, along with household-level data. Initial non-parametric regressions showed that five of six observed risk factors (maternal age <20, no maternal education, short previous birth interval, lack of a safe toilet, female headship and loss of a previous child) negatively influenced full immunization probability. Multivariate logit models showed that high-risk households were significantly less likely to obtain immunization cards (a proxy for exposure to the routine immunization program), despite rising cardholdership rates over the period. Based on recall or card data, high-risk children without cards were significantly less likely to be fully immunized than were high-risk children with cards. A treatment effects model of full immunization probability showed that cardholdership effectively offset all but one risk factor (previous birth interval <24m). Among 1376 cardholders, waiting times to full immunization fell monotonically over the period. Local program coverage of 80% or higher and prior prenatal use both increased the hazard of full immunization. However, three of five maternal occupational categories decreased the hazard, as did no maternal education, consensual versus legal union and lack of a safe toilet. The results show that demand for routine immunizations rose over the period; that high-risk households were less likely to use the routine program but that exposure to the routine program effectively increased their demand for immunizations. To maximize health benefits, future interventions should aim to selectively recruit high-risk households into the ranks of routine immunization users.

BACKGROUND AND PROBLEM STATEMENT

Mortality and fertility rates have fallen faster in Latin America in recent decades than anywhere else in the developing world (Hill and Pebley, 1989; McNicoll, 1992; Sullivan, 1991). This rapid transition, many argue, has been propelled more by public health programs than by improvements in living standards (Preston, 1985; Bahr and Wehrhahn, 1993). Well before the 1978 Alma-Ata Conference, comprehensive primary health care (PHC) services, including immunizations, growth monitoring, prenatal care, health education, inexpensive treatment for common illnesses and environmental monitoring had been available to the majority of the region's households (PAHO, 1990). In certain areas, however, utilization of those services has lagged, despite various outreach and promotional efforts by health workers (Lechtig et al. 1983; Askew, 1988; WHO, 1991; Mora and Yunes, 1993).

To maximize health impacts, Latin American health policymakers began emphasizing a smaller set of specific PHC programs in the 1980s, including the Expanded Program on Immunization (EPI), programs for the control of diarrheal and acute respiratory diseases, nutritional monitoring and, more recently, maternal health (Walsh and Warren, 1979; Grant, 1991; Chelala, 1992). In the past decade, governments and donor agencies alike have spent unprecedented amounts on these specific PHC programs. An example was the Regional Polio Eradication Initiative, launched in 1985 through the Pan American Health Organization Regional EPI Program. At the onset of the polio effort, routine immunization services were accessible to at least 70% of the Region's families but coverage levels hovered around 50% (de Quadros et al. 1991). As the eradication effort progressed, ministries of health began continuously monitoring the percentage of children fully immunized down to the *municipio* level, enabling them to target inputs to high-risk areas. Over half of the region's approximately 10,000 *municipios* had passed the 75% coverage mark by 1988 but a small proportion—around 500 mostly low-income urban communities—had not yet reached 50% coverage. By 1991 special "mop-up" teams of house-to-house vaccinators had pushed up coverage levels to over 75% in these laggard *municipios* as well (MMWR 1994). Regional polio eradication was formally certified in October 1994.

A recent six-country evaluation found that the Regional Polio Effort had strengthened delivery capacities for routine immunization services and other PHC programs (PAHO, 1995). The evaluation, however, did not address demand for routine PHC services. It seems likely that such a large-scale, sustained intervention must have affected at least immunization demand (Wright, 1995). The key policy question is whether any increase in demand will be sustained after mass campaigns are discontinued. A sustained increase would lead to higher routine immunization service utilization and would reflect long-term behavioral change at the household level. Immunization demand might rise due to a range of factors: parents' heightened awareness of the need; increased availability and quality of PHC services; increased income and changes in other household constraints; sanctions and other social effects. A positive program treatment effect would cause them to complete the immunization schedule because they understand the importance of disease prevention. The polio initiative has been criticized for emphasizing expediency over education (Nichter, 1995; Wright, 1995). A second policy question is whether or not immunization demand increased uniformly across various socioeconomic strata and risk groups (Nichter, 1995).

In this paper we examine immunization demand over the period 1985-1990 in Colombia, one of the six countries evaluated and one of the last to eliminate polio. Data provided by the Colombian Ministry of Health show that the proportions of children under age one who had received at least three doses of polio vaccine rose monotonically over the period 1983-1990

(Table 1). Complementing the ministry's data are the household-level immunization data reported in Colombia's 1986 and 1990 Demographic and Health Surveys (DHS), summarized in Table 2. These data suggest that coverage rose for children ages 12-23 months, that the mean age of immunization declined over the interval but that the proportion of children with immunization cards did not increase. An exact comparison of full immunization likelihood cannot be made, however, because different methods were used to estimate the age-specific coverage levels. The 1981-85 period rates are derived from immunization cards with no maternal recall while the 1985-89 rates in the DHS II combine card and maternal recall data (Macro International 1990). As Goldman and Pebley (1994a) showed using DHS data from Guatemala, disregarding the reporting method may lead to biased estimates. Secondly, direct comparison may be misleading because the surveys represent independent probabilistic samples.

Conceptualizing our models is difficult for both practical and methodological reasons. Under the changing immunization regime, a child could have been immunized through routine services, mass campaigns, house-to-house teams or a combination of the three. We assume that parents' health-seeking behaviors would most likely be favorably affected by repeated contact with routine PHC providers, and less so by participation in the campaigns. Those parents who did obtain immunization cards presumably chose to utilize the routine health services; exposure to those services may have produced behavioral changes favoring full and timely immunization. In contrast, parents whose children were reached via the aggressive campaigns and house-to-house mop-up strategies may not have been in contact with the routine program, and so would not have received immunization cards. Rather than an outcome of choice and behavioral change, their immunization decisions would have been exogenously determined. Simply comparing the conditional probabilities of full immunization of cardholders to non-cardholders could therefore be misleading.

A second problem is that obtaining a card and having one's child fully immunized may not be independent processes. Parents with a high health-seeking propensity (an unobservable latent variable) might have self-selected into the ranks of routine program users. The decision to obtain a card would be confounded with any observed routine program effect on immunization behavior. To infer a program effect on immunization outcomes would be spurious in such a case (Maddala, 1983).

There is empirical evidence that parents indeed have varying propensities to seek health care. Recent studies have shown that as few as six percent of households produce as much as two-thirds of infant morbidity and mortality in developing areas (Das Gupta 1990; Desai 1992; Curtis et al. 1993; Guo 1993). Given that immunization is an important proximate determinant of infant and childhood mortality, one would expect so-called "high-risk" households to be among the least likely to have cards or fully immunized children (Rutstein et al. 1990; Mosley and Chen 1984). Among observable mortality risk determinants are: poor sanitary conditions, low maternal age and education, short birth intervals, high fertility, female household headship and a history of previous childhood deaths in the household. Das Gupta (1990) has speculated that unobserved causes of low "maternal competence" are more important mortality predictors than these observed risk covariates. It remains to be seen whether exposure to the routine program altered the effects of both these observed and unobservable risk factors.

With the above in mind, we pose three questions in this paper: (1) Did Colombia's changing immunization regime recruit high-risk households into the ranks of regular immunization service users or did it merely reach them through the periodic campaigns? (2) Once recruited, were high-risk parents as likely as "low-risk" parents to complete their children's immunization? (3) Given routine program exposure, did the pace or timing of immunization vary among high and low-risk children?

MODEL SPECIFICATION

Dependent variables

We examine household- and community-level determinants of three individual immunization outcomes: immunization cardholdership, immunization status and the waiting times to full immunization. We interpret cardholdership as an indicator of exposure to the routine EPI program. Full immunization status is our dichotomous measure of whether or not a child has received all four WHO-recommended antigens (BCG, DPT, OPV and measles). The waiting times, in months, indicate how long it took parents to fully immunize an index child with exposure starting at 9 months of age, the minimal age at which WHO recommends the last antigen, measles, should be administered.

Independent variables

We incorporate six observed health risk covariates as predictors in our models: maternal age less than 20 years, no formal maternal education, female household headship, no toilet in the household, preceding birth interval less than 25 months, and a history of a prior child death. We also include controls for maternal occupation, number of living children, metropolitan vs. nonmetropolitan area residence, mother's union status, prior use of prenatal services and local vaccine (OPV3) coverage estimates for *municipios* of residence. The latter, a measure of local immunization effort, is described in more detail below.

Causal framework

We adapt the household production of health model, considering immunization as a specific input to a more general household health production function (Schultz, 1984; DaVanzo and Gertler, 1990; Akin et al. 1992). Subject to income and other constraints, parents choose immunizations over other behavioral alternatives in order to maximize the utility healthy children affords them. Under an optimally efficient household health production function, a child is fully immunized at the lowest possible age—9 months—thereby minimizing her exposure to the target vaccine-preventable diseases. The efficiency of a production function is affected by both cost and socioeconomic factors. A positive treatment effect would increase the efficiency of a household health production function, resulting in higher probabilities of complete immunization and shorter waiting times among routine program users. The mechanism for this effect might be endogenous (by causing parents to reorder preferences) or exogenous (by reducing the time and other costs or adding incentives to obtain immunizations). For high-risk households a strong treatment effect might offset the negative effects of the observed high-risk predictors.

METHODS

Our first step is to examine the data nonparametrically for evidence of a program treatment effect. For this, we use the BOUNDS program (Manski and Shen, 1993) to regress immunization status on each of the six dichotomous risk status covariates, stratifying by cardholdership. By the law of total probability, the treatment effect can be bounded non-parametrically:

$$P[Y(T)=1|X,Z=1]*P(Z=1|X) - P[Y(N)=1|X,Z=0]*P(Z=0|X) - P(Z=1|X) \leq P[Y(T)|X] - P[Y(N)|X] \leq P[Y(T)=1|X,Z=1]*P(Z=1|X) + P(Z=0|X) - P[Y(N)=1|X,Z=0]*P(Z=0)$$

where:

$Y(T)$ = probability of full immunization for cardholders
$Y(N)$ = probability of full immunization for non-cardholders
X = observed risk covariates
Z = a dummy variable indicating cardholdership

Numerical limitations require us to first select a 50% random sample of the dataset. We then stratify the sample into the twelve sub-samples. Using the kernel method, we then estimate bootstrap probability bounds for full immunization for pseudosamples from each stratum, assuming exogenous (random) selection (Manski and Shen, 1993). Should the bounds fail to overlap for any given risk status covariate, controlling for cardholdership, or for cardholders versus non-cardholders with the same risk status, the empirical distributions can be said to differ significantly. If the probability bounds for high-risk strata are lower than those for low-risk strata among non-cardholders, but do not differ significantly among cardholders, we have necessary but not sufficient evidence of a treatment effect.

Binary models

We model immunization utilization as a two-step endogenous process: parents first obtain cards, then either do or do not go on to complete their child's immunization. Their propensity to do either is a function of an unobservable health-seeking variable. A positive shift in this variable could have been caused by improvements in local PHC program effort, better information, cognitive change and/or endogenous social effects (Manski 1994). A specific immunization program effect, however, would be limited to behavioral changes among cardholders. In bivariate models with cross-sectional data, a program treatment effect on high-risk households would be indicated by positive interactions between cardholdership and the observed risk covariates.

The local supply of immunizations may also affect individual outcomes. To control for this possibility we include in our models a contextual measure for local immunization program performance: the Colombian Ministry of Health's estimate of the percentage of children below age 2 who had received their third dose of polio vaccine in each sample *municipio* in 1990.

We first model the probability of obtaining an immunization card, conditional on the simultaneous effects of observed risk status and the other control variables. Since our aim is to estimate the controlled effects of these risk covariates on the likelihoods of cardholdership we estimate only a direct effects logit model of the form:

$$\log (\Pr(Y=1)/[1-\Pr(Y=1)] = \propto + ß_1 x_1 + ß_2 x_2$$

where

Y = 1 if child has a card
0 if otherwise
$\propto$ = a constant
x_1 = a vector of household-level covariates
$ß_1$ = a vector of parameters for those covariates
x_2 = a vector of 1990 vaccine coverage estimates
$ß_2$ = a vector of parameters for the coverage effect

If the effects of the risk covariates in $ß_1$ are negative and significant, this is evidence that households less-at-risk were differentially recruited into the routine program. If $ß_2$ is significant, then local immunization supply affects individual immunization outcomes over and above household-level factors.

Regarding the likelihood of full immunization, if routine program participation conferred a positive treatment effect then the effects of the observed risk covariates on the conditional probability of full immunization ought to be less negative among cardholders than non-cardholders. We test these hypotheses in heuristic fashion by separating cardholders from non-cardholders and estimating logit models for full immunization, of the same form described above, for each group. Further, if exposure to the routine program caused behavioral change among high-risk parents, then a dummy variable for cardholdership, our proxy for routine program exposure, ought to interact positively with the observed risk factors. We estimate such an interactive model, and, for comparison, a direct effects model, on the full dataset.

The logit models still do not allow us to rule out a possible endogeneity bias. The coefficient estimates in our immunization logit models will be inconsistent to the extent this endogeneity problem exists. To purge any endogeneity bias between cardholdership and immunization behavior we estimate a treatment effects model of the form:

(1) $y_i = ß'x_i + \delta'z_i + \epsilon_i$

(2) $z_i = \gamma'w_i + u_i$

where, for child i:

y_i = 1 if fully immunized, 0 if otherwise
x_i = a set of observed predictors of full immunization status
ß = the effects of predictors x_i on full immunization probability
ϵ_i = unmeasured determinants of immunization
δ = the treatment effect
z_i = 1 if child has a card, 0 if otherwise
w_i = a set of instrumental variables predicting cardholdership
γ = the effects of instruments w_i on cardholdership probability
u_i = a stochastic disturbance associated with child i but independent of w_i

Equation (1) corresponds to a specific household health production function for immunization while Equation (2) represents the derived demand for one input to that function, enrollment into the routine immunization program (Schultz, 1984). Ordinary least squares would not be appropriate for solving these simultaneous equations because $E[cov(x_i,\epsilon_i)] \neq 0$ and $E[cov(u_i,\epsilon_i)] \neq 0$. Instead, we use two-stage least squares. We first impose a normal distribution on Equation (2) and regress cardholdership z_i on a set of exogenous covariates w_i. We then substitute the predicted probit values, $\hat{z}_i$, into Equation (1) and solve it using ordinary least squares (Greene, 1993). This renders $E[cov(z_i,\epsilon_i)]=0$ and shows the unbiased effect of cardholdership on full immunization probability, controlling for the effects of covariates x_i.

We recognize that recall and reporting bias may well affect our results. To the extent mothers without cards tended to overstate their children's immunization statuses, the effects of cardholdership on full immunization probability will be underestimated. A second source of noise is that a given child with a card could have been at least partially immunized outside of the routine program. We make no effort to control for these biases in our bivariate models. Instead, we indirectly test the validity of our bivariate results by comparing the effects of the same covariates on the hazard of full immunization.

Hazard models

A monotonic decrease in waiting times to full immunization across successive cohorts would be compatible with a positive program effect. A convergence in waiting times for children from high-risk and low-risk households would offer further evidence that exposure to the routine services changed behaviors of both risk groups. In the logit models we can test the program effect directly by examining the interactions of cardholdership with the observed risk covariates. This strategy is not possible in our hazard models because only cardholders are in the risk set. However, if the observed risk covariates have no significant effect on the hazard of full immunization, a necessary but not sufficient condition for inferring a positive treatment effect will be met.

To examine timing to full immunization among cardholders we estimate a two-state, continuous-time piecewise hazard model using standard maximum likelihood procedures[1]. The model takes the form:

$$u(t) = u_0(t)*\exp(X_1B_1 + X_2(t)B_2)$$

where

$u(t)$ = the hazard function at time t
$u_0(t)$ = the baseline hazard at time t
B_1 = a vector of fixed covariate parameters

[1] We used CTM (Yi, Honore and Walker 1987).

B_2 = a vector of time-varying parameters
X_1 = a vector of fixed household and individual covariates
X_2 = a vector of time-varying local coverage estimates.

We measure exposure to the hazard of full immunization in months. Each "piece" in the model corresponds to some segment of the waiting time distribution. The segment 1-3 months, for instance, captures the baseline hazard for children immunized between ages 9-11 months. The hazard of full immunization within each piece is assumed to be constant. Based on preliminary results (not shown), we collapsed the duration structure from 68 to only 10 segments.

In the bivariate models our community-level proxy for vaccine supply consists of a point estimate for 1990, the year the data were collected. In the hazard models we reparameterize this measure as a time-varying covariate so that a child aged 5 in 1990 might have five different local coverage values, one for each year of exposure.

Further exploration revealed that one fixed covariate, use of prenatal services, interacted significantly with duration of exposure. The model we report includes these interactions.

It is quite possible that frailty, differential endowments or other covariates not specified in our model systematically affected the hazard of full immunization. Not accounting for the effects of such unobserved heterogeneity could seriously bias coefficient estimates and lead to misinterpretation of the duration dependence structure (Heckman and Singer 1982; Vaupel and Yashin 1985). Hazard models are also sensitive to the functional forms imposed on the hazard (Trussell and Richards, 1985). To test for this distributional sensitivity we first compared our direct-effects piecewise exponential model using Gompertz and Cox nonparametric proportional hazard specifications. We then re-run the Gompertz model and test for the effects of unobserved heterogeneity using the Heckman-Singer (1982) random effects approach.

DATA

Colombia's DHSII survey obtained detailed data for 8644 reproductive-age women and their 4215 surviving children born between January 1985 and September 1990. Included are several household health risk variables as well as exact immunization data for the 42% of children whose immunization cards were seen by interviewers (Table 3). To each of the individual child health records we attached the corresponding data from the DHS household module. To create the time-varying covariate for local vaccine supply, we merged these records with a file containing the annual *municipio*-level ministry of health immunization coverage (OPV3) estimates for 1985-1990. To each child record we attached the vector of local coverage estimates beginning with the value obtained in the year that child reached 9 months of age. We then converted this continuous measure into a categorical variable.

Matches were made with 122 of the 148 *municipios* represented in the 754 DHSII sample clusters. For the remaining unmatched DHS *municipios* we assigned the coverage data from the ministry's next largest reporting units (*departmentos*). On average, there were 25 children born on or after January 1985 in each *municipio* and about six per sample cluster. Six *municipios*, representing 17% of the total children in the sample, had system-missing or zero reported coverage values for at least one of the five years. Correspondence with the Colombian Ministry of Health established that the system-missing values referred to *municipios* where regular immunization services were not operating that year while the zeroes referred to years where *municipios* which should have reported their coverage failed to do so. Rather than impute values we created and later estimated a separate category for these non-reporters.

For 1946 children with cards we computed durations, in months, beginning at age nine months to the date of the survey (exposure) and to the date on which each of four immunizations (BCG, DPT3, OPV3 and measles) was administered. We used the longest of these durations to represent the waiting time to full immunization for 1082 children who had received all four antigens. Those missing one or more immunizations we censored at the time of the survey. In cleaning the data we rejected 44 cases for whom we could not impute dates and 17 cases for whom card status could not be determined. Additionally, there were missing value codes for 167 of 783 incompletely immunized children (21%) wherein a date was not entered or the mother reported an immunization not noted on the card. These we censored at the longest computable duration.

We created dichotomous variables for cardholdership and for full immunization using both card and recall data. Based on frequency distributions we categorized other household variables of interest, computed preceding and following birth intervals and created an annual birth cohort variable. We next eliminated children below nine months of age who, according to WHO and ministry norms, were too young to be fully immunized. We also dropped 24 cases whose previous birth intervals were coded zero. These, presumably, were twins. The final dataset thus contained 3609 child records with information on immunization status, cardholdership, fixed and time-dependent covariates and local immunization program performance. By either card or recall, 2687 (74.5%) of the children were fully immunized. Additionally, exposure times, state and censoring codes were computed for 1376 children with immunization cards, of whom 1038 (75.3%) were fully immunized.

RESULTS

Non-parametric bounds

In Table 4, we report 5, 50 and 95% bootstrap quantiles for full immunization probabilities among cardholders and non-cardholders for each of six risk status covariates. For each stratum we set bandwidths for the kernels at 1/10 of the standard errors of each of the six covariates

(Manski and Shen, 1993). The bandwidths are represented by the variable lambda in Table 4. Also, the numbers of observations in each stratum are indicated in parentheses. We found that 100 iterations were sufficient to generate stable estimates for each stratum. For five of the six risk covariates, non-cardholding children were significantly less likely to be fully immunized than those whose parents presented cards to the interviewers. Among non-cardholders, the bounds for all risk variables except female headship did not overlap with their respective reference categories. This indicates that immunization probabilities for children from non-cardholding households with these risk factors were significantly lower. Among cardholders, in contrast, the bounds do overlap for each risk variable, indicating no significant differences. Either a true positive program treatment effect or self-selection of the lesser-at-risk into the routine program could explain this pattern.

Logit models

Our multivariate models revealed more about the processes underlying immunization behavior. Table 5 shows that low maternal age, high parity, lack of any union and a prior child death all significantly reduce the log-odds of cardholdership. Families with these risk factors remained less likely to participate in the routine program, despite the intensifying immunization regime. A mother employed in a relatively high-status occupation had more difficulty obtaining a card than did a mother not working in the paid labor force, evidence that such mothers may have faced higher opportunity costs of participation. Interestingly, mothers in manufacturing jobs were 31% more likely than non-employed mothers to have cards. Strong positive cohort effects all but overshadowed the fixed covariate effects. The likelihood of a child born in 1989 having a card was nearly three times that of a child born before 1987, despite her much shorter period of eligibility. Local vaccine coverage levels for 1990 had no significant effect on the likelihood of cardholdership.

Table 6 reports separate logit regressions for full immunization for cardholders and non-cardholders. Among the latter, all risk factors except short previous birth interval exert significant negative effects on the likelihood of full immunization. Among cardholders, in contrast, female household headship and short previous birth interval reduced the odds of full immunization by 28% and 39%, respectively. Additionally, cardholding mothers who worked in professional jobs or who had above high school educations were less likely to have fully immunized children as compared to those who worked only at home or had less formal educations. These effects suggest either that parents from higher socioeconomic levels may have self-selected into the routine program or that the effects of the risk factors changed as a result of cardholdership.

To test whether the routine program affected high-risk households differentially we fitted direct effects and multiplicative models for full immunization for the entire dataset (Table 7). In the direct effects model, having an immunization card significantly increased the log-odds of full immunization while each of the six observed risk covariates had significant negative effects. In the interactive model, three of the risk variables—maternal age below 20 years, no formal maternal education and having had a prior child death in the household—interacted significantly with cardholdership.

In Figure 1 we display the cumulative effects of the three significant interactions on the conditional probabilities of full immunization. All other covariates are set to their omitted categories. The graph shows that having a card effectively offset the negative effects of the three risk factors. A child without a card whose mother was less than 20 years of age, had no formal education and had lost a previous child had only a 56% probability of being fully immunized. The probability of full immunization for a cardholding child with the same combination of risk factors was 86%, virtually the same as that of a cardholding child with no risk factors. As with cardholdership, local vaccine coverage levels did not affect full immunization likelihood.

Treatment effects model

As described earlier, there is a chance these logit model results are biased. The rightmost column of Table 8 shows the treatment effects model for full immunization estimated by two-stage least squares. Next to it is the equivalent probit model for comparison. For parsimony we dropped the local immunization services coverage figures, which were insignificant in all our binomial models. We used as our instruments three covariates significantly correlated with the indicator for treatment exposure, (cardholdership), but not with the probability of full immunization. These instruments were maternal occupations in clerical (SJOB2) and manufacturing (SJOB5) sectors and non-metropolitan residence (METRO2). Each we judged to be exogenous to the parents' immunization behaviors. As shown, the adjusted R-squared for an OLS model regressing cardholdership on this set of instruments was -0.48. Control and risk variables agreed in direction though not in significance in the probit and instrumental variables models. In both, cardholdership increased full immunization probability. In the probit model, only prenatal use exerted a stronger positive effect. When the endogeneity bias is purged, the prenatal effect disappears and that of cardholdership increases dramatically; its effect exceeds that of any other covariate. As reported in Tables 5 and 6, prenatal service utilization increased the odds of cardholdership and of full immunization among non-cardholders by 51% and 74%, respectively. The prenatal effect disappears, however, in the presence of the immunization program treatment effect—evidence, perhaps, that prenatal and immunization use are governed by the same endogenous process. Notably, the effects of only one risk covariate, short previous birth interval, remained significant in the presence of the treatment effect. In behavioral terms one might speculate that short birth intervals exert the least mutable of the observed risk covariate effects.

Survival analysis

To check the correspondence between ministry of health and DHSII data we failed all 1376 cardholding children ages 9-62 months using cohort as strata (Figure 2). Kaplan-Maier estimates (not shown) showed that waiting times to full immunization were not homogenous by either log-rank or Wilcoxon tests. We concluded that waiting times indeed fell monotonically over the period. A plot of the conditional hazard function for all cardholding children shows that the risk of full immunization rose sharply until age 15 months, then declined monotonically (Figure 3).

In our initial estimates low maternal age, female household headship, previous birth interval and a previous child death were insignificant so we dropped these risk factors from our dynamic models. Of the remaining covariates, the results in Table 9 show, no maternal education, consensual union status and lack of toilet facilities all significantly reduced the hazard of full immunization. Employment in blue-collar manufacturing and high-status occupations exerted comparable negative effects.

When modeled as a time-varying covariate, the ministry of health vaccine coverage estimates had significant effects on the hazard of full immunization. All else being equal, the risk of full immunization for a child living in a *municipio* with reported OPV3 coverage of 80% or more was up to 49% higher than that of a child living in a lower-coverage *municipio*.

The hazard of full immunization was also significantly greater for children whose mothers had used prenatal services. The interactions reported in Table 9 show that the salutary effect of prenatal use, moreover, was duration-dependent. Among prenatal users the hazard of full immunization for children age 9 months was no different than that of non-users. At ages 10, 12 and 22-24 months, however, it was higher.

Additionally, the risk of full immunization was higher for more recent birth cohorts, despite falling ages. This positive cohort effect, not distinguishable in the bivariate models, indicates that the routine program's positive treatment effect in fact increased over the period.

As shown in the second and third columns of Table 10, we found little variation in our piecewise exponential and Gompertz coefficients. In the Gompertz model the negative sign of the gamma coefficient indicates that the hazard is falling with time, a result generally compatible with Figure 3.

We thus used the Gompertz specification to test for the effects of unobserved heterogeneity on the hazard of full immunization. The fourth column of Table 10 shows the Gompertz model with the Heckman-Singer non-parametric correction for unobserved heterogeneity. The model converged using two points of support. The factor loading term was positive and significant, indicating that the hazard for the observed group, representing about 8% of the study population, was significantly higher than that of the remaining 92% (omitted category). When the mixed sub-populations are identified the gamma term reverses direction.

The 8% of children on the observed support could be those of parents who made it a point to have their children immunized as early as possible. Although the unobserved heterogeneity effect and the latent variable construct in the treatment effects model are analytically distinct, one would expect that this small subset of parents would have scored highest on the latent health-seeking propensity variable. For the remainder of parents, the hazard of full immunization in fact increased with time, a finding that accords with a general tendency to procrastinate.

When the unobserved heterogeneity is controlled, covariates with negative effects become more negative. No fixed effect covariates lose statistical significance and none reverse direction. We conclude that our piecewise hazard covariate estimates are fairly robust.

ISCUSSION AND CONCLUSIONS

The effects of Colombia's polio eradication effort on both the probability and timing of immunization were overwhelmingly positive. Cardholdership and full immunization probabilities increased while waiting times to full immunization decreased over the period.

Despite the rising force of the immunization effort, high-risk households remained less likely to become routine immunization service users. Mothers below age 20 and those who had previously lost a child stand out among the sub-groups most resistant to entering the routine program. This finding is compatible with that of Curtis et al. (1993), who showed that infant mortality risk was clustered among mothers who had previously lost children in Northeast Brazil. For such mothers the psychic costs of dealing with health care providers or of "publicizing" these earlier traumas may well exceed any offsetting wealth or cognitive effects (Coreil et al. 1994; Berman et al. 1994).

Once enrolled in the routine program, strong positive program effects effectively offset all but one of the observed risk factor effects on the likelihood of full immunization. The survival analysis results suggest, however, that the behavioral changes high-risk families experienced were not absolute. High-risk status generally prolonged the waiting time to full immunization despite the countervailing program, cohort and community supply-side effects. With unobserved heterogeneity controlled, the negative effects of low socioeconomic status, high parity, consensual union and maternal employment became even more negative.

Our results show that the effects of maternal employment on immunization uptake vary considerably by occupation. The logit results show that, relative to mothers not in the paid work force, mothers in higher-status occupations found it costlier to obtain cards. Once they obtain cards, mothers working in the professional, services and manufacturing sectors took longer to complete the immunization schedule than did mothers working only at home. This pattern accords with standard microeconomic models of the family wherein time and other resources are allocated according to the value of each members' time (Becker, 1965; Grossman, 1972a, 1972b). The negative employment effects, however, do not increase monotonically with occupational prestige, suggesting that job autonomy rather than opportunity costs may be the more salient constraint among the lower-status occupations. Rosenberg (1982) documented the lax enforcement of childcare, breastfeeding and other regulations protecting mothers working in manufacturing firms in Bogota. As with the household risk factors, the rising force of the immunization program could not entirely overcome these occupational differentials.

The inclusion of contextual data yielded additional insights. The fact that the local immunization program effort had an independent, positive effect on the hazard of full immunization points up the efficacy of targeted PHC interventions. Our results validate those reported by Colombia's Ministry of Health. The absence of urban-rural differentials implies that geographical accessibility was not a decisive determinant of immunization behaviors. High local coverage rates reflect intensive utilization. Beyond a certain threshold, one might speculate, social "contagion" or normative effects may begin to affect parents' immunization behaviors within a given collectivity. Alternatively, local vaccine coverage levels may be capturing more subtle supply-side effects, such as the quality of services provided. Such effects would only have been salient among cardholders and may have been obscured by the stronger supply-side effects of the heterogenous delivery strategies in our binomial models. In a recent study, Mensch et al. (1994) matched health facility evaluation data to individual records from Peru's 1992 DHS survey. Using multilevel logit models, they showed that contraceptive behaviors were significantly affected by the quality of local family planning services. The latter effect, however, was modest compared to those of maternal education and other household-level indicators.

Use of prenatal services added to both cardholdership and full immunization likelihoods and increased the hazard of full immunization among cardholders. These two health-seeking behaviors are clearly positively correlated, a finding which has been observed elsewhere (Eng, 1988; Chowdhury, 1990; WHO, 1991; Goldman and Pebley, 1993). The positive program effects among cardholders lends support to the notion that long-term behavioral changes are mainly the result of routine PHC service utilization. Campaigns and social mobilization efforts should therefore be designed with these facts in mind.

In future studies we hope to model the more elusive endogenous social effects—the contingent, collective nature of health-seeking behaviors within a collectivity. Recently, Liao outlined such a conceptual model for family planning programs (Liao, 1994). Collective action may help explain why health behaviors tend to cluster at the community level (Entwisle et al. 1989; Sastry et al. 1993) and why, in the course of the Regional polio eradication initiative, some collectivities could not attain high coverage without resorting to house-to-house campaigns. Exogenously induced collective action could cause normative change within collectivities (Hardin, 1982; McQuestion, 1994), a mechanism which would explain in part how individual health-seeking behaviors change and are subsequently reproduced. Targeted campaigns and sustainability of high PHC utilization may therefore not be mutually exclusive as various observers have suggested (Loevinsohn and Loevinsohn, 1987; Turshen, 1989).

A key analytical challenge in fitting such a model will be the separation of endogenous social effects from the artifactual clustering bias inherent in the DHS and other sample surveys (Elo, 1992; Guo, 1993; Goldman and Pebley, 1994b; Rodriguez, 1994; Sastry, 1995; Steele and Diamond, 1995), a bias which, admittedly, may have affected the results we report in the present paper. Uncontrolled clustering bias would cause standard errors to be underestimated. Further, to specify an endogenous collective action effect on individual health-seeking behaviors would require relational data which are not currently available for the DHS survey clusters.

BIBLIOGRAPHY

Akin, J., D. Guilkey, and B. Popkin. 1992. A child health production function estimated from longitudinal data. *Journal of Development Economics* 38: 323-351.

Askew, I. 1988. A comparative analysis of community participation projects in South Asia with policy and programme recommendations for family planning associations. International Planned Parenthood Federation, Occasional Series on Community Participation, 47 pages, London, England.

Bahr, J., and R. Wehrhahn. 1993. Life expectancy and infant mortality in Latin America. *Social Science and Medicine* 36(10): 1373-1382.

Becker, G. S. 1965. A theory of the allocation of time. *Economic Journal* 75(299): 493-517.

Berman P., C. Kendall, and K. Bhattacharyya. 1994. The household production of health: Integrating social science perspectives on micro-level health determinants. *Social Science and Medicine* 38(2): 205-15.

Chelala, C.A. 1992. Prevención de la mortalidad materna en las Americas: Perspectivas para los años noventa. Washington, DC. PAHO, *Comunicaciones para la Salud*, No. 2, 1-30.

Chowdhury, A.M. 1990. Empowerment through health education: The approach of an NGO in Bangladesh. In *Implementing Primary Health Care: Experiences since Alma-Ata*, eds. P. Streetland and J. Chabut. Pp. 113-120. Amsterdam: Royal Tropical Institute.

Coreil, J., A. Augustin, N. Halsey, and E. Holt. 1994. Social and psychological costs of preventive child health services in Haiti. *Social Science and Medicine* 38(2): 231-238.

Curtis, S. L., I. Diamond, and J.W. McDonald. 1993. Birth interval and family effects on postnatal mortality in Brazil. *Demography* 30(1): 33-44.

Das Gupta, M. 1990. Death clustering, mother's education and the determinants of child mortality in rural Punjab, India. *Population Studies* 34: 489-505.

DaVanzo, J., and P. Gertler. 1990. Household production of health: A microeconomic perspective on health transitions. Rand Corporation, Rand Note No. n-3014-RC. Santa Monica, CA.

Desai, S. 1992. Children at risk: The role of family structure in Latin America and West Africa. *Population and Development Review* 18(4): 689-718.

Elo, I. 1992. Utilization of maternal health care services in Peru: The role of women's education. *Health Transition Review* 2(1): 49-69.

Eng, E. 1988. The role of community participation in child health: Water supply, ORT and EPI. In *International Health in the 1990s: Directions in Research and Development*, ed. M. Hefferman. Pp. 15-19. National Council for International Health, Southern Regional Conference, Washington, DC.

Entwisle, B., J.B. Casterline, and H. A-A Sayed, 1989. Villages as contexts for contraceptive behavior in rural Egypt. *American Sociological Review* 54: 1019-1034.

Goldman, N., and A. Pebley. 1993. Family, community, ethnic identity and the use of formal health care services in Guatemala. Paper presented at the Annual Meeting of the Population Association of America, 1-4 April, Cincinnati, OH.

Goldman, N., and A. Pebley. 1994a. Health cards, maternal reports and the measurement of immunization coverage: The example of Guatemala. *Social Science and Medicine* 38: 1075-1089.

Goldman, N., and A. Pebley. 1994b. Childhood immunization and pregnancy-related services in Guatemala. *Health Transition Review* 41: 29-44.

Grant, J.P. 1991. *The State of the World's Children.* New York: Oxford University Press.

Greene, W. 1993. *Econometric Analysis.* Second Edition. New York: Macmillan, Ch. 22.

Grossman, M. 1972a. On the concept of health capital and the demand for health. *Journal of Political Economy* 80: 223-255.

Grossman, M. 1972b. The Demand for Health: A Theoretical and Empirical Investigation. NBER Occasional Papers, No. 119. New York: Columbia University Press.

Guo Guang (1993), Use of sibling data to estimate family mortality effects in Guatemala, *Demography* 30(1): 15-32.

Hardin, R. 1982. *Collective Action.* Baltimore. Johns Hopkins University Press.

Heckman, J.J., and B. Singer. 1982. The identification problem in econometric models for duration data. In *Advances in Econometrics,* eds. J.J. Heckman and B. Singer. Pp. 39-77. Cambridge: Cambridge University Press.

Hill, K., and A.R. Pebley. 1989. Child mortality in the developing world. *Population and Development Review* 15: 657-687.

Lechtig, A., J.W. Townsend, F. Pineda, J.J. Arroyo, R.E. Klein, and R. De Leon. 1983. SINAPS: The Guatemalan Program. In *Advances in International Maternal and Child Health, Vol. 3,* eds. D.B. Jelliffe and E.F. Jelliffe. Pp. 146-173. Oxford: Oxford University Press.

Liao, T.F. 1994. A theoretical framework of collective action for the evaluation of family planning programs. *Population Research and Policy Review* 13: 49-67.

Loevinsohn, B.P., and Loevinsohn, M. E. 1987. Well child clinics and mass vaccination campaigns: An evaluation of strategies for improving the coverage of primary health care in a developing country. *American Journal of Public Health* 77(11): 1407-11.

Macro International. 1990. Colombia: Encuesta de prevalencia, demografía y salud 1990. Columbia MD: Macro International.

Maddala, G.S. 1983. *Limited Dependent and Qualitative Variables in Econometrics*. Cambridge: Cambridge University Press.

Manski, C., and Shen, T. 1993. Non-parametric estimates and bounds on treatment effects. Unpublished document, Dept. of Economics, University of Wisconsin-Madison.

Manski, C. *Identification Problems in the Social Sciences.* 1994. Cambridge: Harvard University Press.

McNicoll, G. 1992. Changing Fertility Patterns and Policies in the Third World. *Annual Review of Sociology* 18: 85-108.

McQuestion, M. 1994. How mass campaigns affected immunization likelihoods in Guatemala and the Philippines. CDE Working Paper 94-09. Madison, Wisconsin: University of Wisconsin Center for Demography and Ecology.

McQuestion, M., and R.L. Jones. 1995. An analysis of recent immunization trends in Colombia. Paper presented at the 1995 Meeting of the Population Association of America, 5-10 April, San Francisco, CA.

Mensch B., M. Arends-Kuenning, and A. Jain. 1994. Assessing the impact of the quality of family planning services on contraceptive use in Peru: A case study linking situation analysis data to the DHS. Research Division Working Paper No. 67. New York: The Population Council.

MMWR. 1994. Progress toward global eradication of poliomyelitis, 1988-1993. *Morbidity and Mortality Weekly Report* 43: 499-503.

Mora, G., and J. Yunes. 1993. Maternal mortality: An overlooked tragedy. *Gender, Women and Health in the Americas*, ed. E. Gomez Gomez. Pp. 62-79. Scientific Publication No. 541. Washington, DC: Pan American Health Organization.

Mosley, W.H., and L.C. Chen. 1984. An analytic framework for the study of child survival in developing countries. *Population and Development Review* 10(Supplement): 25-45.

Nichter, M. 1995. Vaccinations in the Third World: A consideration of community demand. *Social Science and Medicine* 41(5): 617-632.

PAHO. 1990. *Health Conditions in the Americas, Vol I.* Scientific Publication No. 524. Washington, DC: PAHO Press.

PAHO. 1995. The impact of the Expanded Program on Immunization and the Polio Eradication Initiative on health systems in the Americas, Final Report of the Taylor Commission. Washington, DC, Pan American Health Organization, Non-serial Publication No. 1995-000003. 80pp.

Preston, S.H. 1985. Mortality and development revisited. *Population Bulletin of the United Nations* 18: 34-40.

de Quadros, C.A., J.K. Andrus, J.M. Olive, C.M. Da Silveira, R.M. Eikof, P. Carrasco, J.W. Fitzsimmons, and F.P. Pinheiro. 1991. Eradication of poliomyelitis: Progress in the Americas. *Journal of Pediatric Infectious Diseases* 103(March): 222-229.

Rodriguez, G. 1994. Statistical issues in the analysis of reproductive histories using hazard models. *Annals of the New York Academy of Sciences* 709: 266-279.

Rosenberg, M.J. 1982. Female industrial employment and protective labor legislation in Bogota, Colombia. *Journal of International Studies and World Affairs* 24(1): 59-80.

Rutstein, S.O., A.E. Sommerfelt, and J. Schoemaker. 1990. Who uses maternal and child health services? Evidence from the Demographic and Health Services. In *Child Survival Programs: Issues for the 1990s*, ed. K. Hill. Pp. 37-84. Baltimore, Maryland: The Johns Hopkins University School of Hygiene and Public Health, Institute for International Programs.

Sastry, N., N. Goldman, and L. Moreno. 1993. The relationship between place of residence and child survival in Brazil. In: International Population Conference, Montreal, Vol 3, pp. 293-322. Liege, Belgium: International Union for the Scientific Study of Population.

Sastry, N. 1995. *A multilevel hazards model for hierarchically clustered data: Model estimation and an application to the study of child survival in Northeast Brazil.* Santa Monica, CA: Rand Corporation.

Schultz, T.P. 1984. Studying the impact of household economic and community variables on child mortality. *Population and Development Review* 10(Supp): 215-237.

Steele, F., and I. Diamond. 1995. Immunization uptake in rural Bangladesh: A multilevel analysis. Paper presented at the annual meeting of the Population Association of America, 6-8 April, San Francisco, CA.

Sullivan, J.M. 1991. The pace of decline in under-five mortality: Evidence from the DHS surveys. In: DHS. 1991. Demographic and Health Surveys World Conference. Proceedings, Volume I, pp. 25-38. Columbia, Maryland: IRD/Macro International, Inc.

Trussell J.T., and T. Richards. 1985. Correcting for unmeasured heterogeneity in hazard models using the Heckman-Singer procedure. In *Sociological Methodology 1985*, ed. N.B. Tuma. Pp. 242-276. San Francisco: Jossey-Bass.

Turshen, M. 1989. *The Politics of Public Health.* New Brunswick, NJ: Rutgers University Press.

Vaupel J., and A. Yashin. 1985. Heterogeneity's ruses: Some surprising effects of selection on population dynamics. *American Statistician* 39: 176-85.

Walsh, J.A., and K.S. Warren. 1979. Selective primary health care: An interim strategy for disease control in developing countries. *New England Journal of Medicine* 301: 967-974.

WHO. 1991. Community involvement in health development: Challenging health services. WHO Technical Report Series, No. 809. Geneva: World Health Organization.

Wright, P.F. 1995. Global immunization—a medical perspective. *Social Science and Medicine* 41(5): 609-616.

Yi, K-M., B. Honore, and J. Walker. 1987. *CTM: A Program for the Estimation and Testing of Continuous Time Multi-state Multi-spell Models, User's Manual.* Version 50. Chicago: ERC/NORC.

Table 1. Percentage of children having received three doses of polio vaccine (OPV) before age one, 1983-1990, Colombia (source: Ministry of Health)

Year	No. of reporting *municipios*	Mean Percent	Std. Dev.
1983	1017	33.77	25.68
1984	1017	52.25	33.36
1985	1017	53.67	30.76
1986	1017	59.69	28.93
1987	1017	73.78	25.76
1988	1006	78.86	22.15
1989	1022	81.24	20.79
1990	1031	85.63	19.3

Table 2. Proportions of children with cards, fully immunized, ages 1-59 months, DHS I and DHS II, Colombia

	DHS I (1986)		DHS II (1990)		
Ages (months)	Card shown	Fully immunized card only	Card shown	Fully[1] immunized	Fully immunized card only
1-5	.50	0	-	-	-
6-11	.59	.12	-	.43	.47
12-23	.55	.40	.58	.69	.73
24-35	.50	.44	.42	.39	-
36-47	-	-	.46	.27	-
48-59	.41	.39	.37	.32	-

[1] Card + maternal recall

Table 3. Description of independent variables, Colombia 1985-1990 (source: DHS II)

Variable	Description	Frequency	Percent
Maternal occupation			
no work past 5 yr	omitted category	1635	45.3
professional/ high-wage (SJOB1)	mgrs, accountants, teachers	185	5.1
secretarial, clerical (SJOB2)	secondary ed or more/ public or priv sectors	669	18.5
services (SJOB3)	primary ed or less/ cooks, domestics	695	19.3
farm, prim sector (SJOB4)	family worker, owner-operator/mining	107	3.0
secondary sector (SJOB5)	manufacturing/skilled trades	318	8.9
Maternal education			
none (MED1)		177	4.9
primary	omitted category	1605	44.5
secondary	completed	1572	43.6
superior (MED4)	some university/trade	255	7.1
Maternal age			
under 20 (MAGE1)		143	4.0
20-29	omitted category	2052	56.9
30-39		1227	34.0
40+ (MAGE4)		187	5.2
Children now alive			
one (KIDS1)		827	22.9
two	omitted category	1045	29.0
three + (KIDS3)		1737	48.1
Previous child death			
no	omitted category	3218	89.2
yes (HMORT2)	one or more	391	10.8

Table 3, continued, Description of independent variables, Colombia 1985-1990 (source: DHS II)

Variable	Description	Frequency	Percent
Household headship			
male	omitted category	3041	84.3
female (HHEAD2)		568	15.7
Union status			
legal union	omitted category	1682	46.6
consensual (UNION2)	includes no current union	1473	40.8
div/sep/widow (UNION3)		454	12.6
Toilet			
yes	omitted category	3164	87.7
no (TOILET3)		445	12.3
Residence			
metropolitan	omitted category	2958	82.0
non-metropolitan (METRO2)		651	18.0
Previous birth interval			
none	only child	1276	35.4
≤ 24m (PBI2)		760	21.1
more than 24m	omitted category	1573	43.6
Prenatal use			
none, missing	omitted category	726	21.3
some (PREN)	mean 5 visits	2678	78.7
Immunization card			
no	omitted category	1955	54.2
yes (VCARD)		1654	45.8

Table 4. Bootstrap distribution quantiles and non-parametric bounds, full immunization probability regressed on selected health risk indicators, cardholders and non-cardholders, Colombia 1985-1990 (Source: DHSII)

Indicator	**Non-cardholders**			**Cardholders**		
	quantiles			quantiles		
	0.05	0.50	0.95	0.05	0.50	0.95
Maternal age						
below 20y	0.200	0.311	0.422	0.625	0.750	0.875
above 20y	0.724	0.749	0.779	0.739	0.765	0.791
lambda (n)		.03 (961)			.03 (818)	
Maternal education						
none	0.281	0.386	0.474	0.543	0.657	0.771
some	0.711	0.742	0.771	0.748	0.771	0.797
lambda (n)		.02 (996)			.02 (818)	
Female headship						
yes	0.619	0.699	0.744	0.679	0.750	0.804
no	0.708	0.738	0.772	0.741	0.766	0.792
lambda (n)		.03 (961)			.03 (818)	
Toilet						
no	0.455	0.537	0.618	0.660	0.732	0.804
yes	0.717	0.749	0.778	0.766	0.793	0.820
lambda (n)		.02 (996)			.02 (862)	
Previous birth interval						
$\leq$ 24m	0.567	0.622	0.664	0.691	0.768	0.812
none or 25m+	0.727	0.752	0.779	0.768	0.793	0.816
lambda (n)		.05 (996)			.05 (862)	
Prior child death						
yes	0.433	0.520	0.583	0.610	0.695	0.814
no	0.739	0.761	0.789	0.744	0.769	0.796
lambda (n)		.03 (961)			.03 (818)	

Table 5. Logistic regression results: Selected predictors of cardholdership, Colombia 1985-1990, n=3509 (source: DHS II)

Variable	Coefficient Estimate[1]	Std. Error	p-value	Odds Ratio
Intercept	-.60	.15	0.00	0.55
Controls				
SJOB1 prof	-.47**	.20	0.02	0.63
SJOB2 clerk	-.33**	.10	0.00	0.72
SJOB3 service	-.08	.10	0.44	0.93
SJOB4 primary	-.10	.21	0.61	0.90
SJOB5 manuf	.27**	.13	0.03	1.31
KIDS1 one	.19	.13	0.13	1.21
KIDS3 3+	-.17*	.09	0.07	0.84
UNION2 consen	-.10	.08	0.18	0.90
UNION3 d/s/w	-.37**	.13	0.00	0.69
NON-METRO	.18*	.11	0.09	1.20
COHO2 24-35m	.15	.10	0.13	1.16
COHO3 12-23m	.67**	.10	0.00	1.96
COHO4 9-12m	1.06**	.14	0.00	2.90
COVER 80%+	.10	.08	0.26	1.10
COVER missing	-.04	.13	0.78	0.97
PREN some	.41**	.10	0.00	1.51
Risk factors				
TOILET3 none	-.13	.13	0.33	0.83
MAGE1 <20	-.49**	.19	0.01	0.61
MAGE4 40+	.45**	.16	0.01	1.57
MED1 no ed	.10	.17	0.56	1.10
MED4 h.s.+	-.24	.17	0.15	0.79
HHEAD2 female	-.16	.11	0.15	0.86

Table 5, cont. Logistic regression results: Selected predictors of cardholdership, Colombia 1985-1990, n=3509 (source: DHS II)

PBI1 missing	-.01	.13	0.94	0.99
PBI2 $\leq$24m	.16*	.09	0.10	1.17
HMORT 1+ dead	-.42**	.12	0.00	0.66
Model fit[2]	4978.0-4735.5=242.5, 25 df(p<0.0001)			

** significant at p < 0.05 level; * significant at p < 0.10 level

[1] contrast with omitted categories: no work past 5y; mat age 21-39; 2 children; primary, some secondary mat ed; metro; male head; formal union; toilet; ages 36m+; no prenatal; prev birth int >24m; no child death; OPV3 <80%

[2] -2 log likelihood: diff between null, full models distributed chi square

Table 6. Logistic regression results: Selected predictors of full immunization status, Colombia 1985-1990, (source: DHSII)

Variable	**Cardholders**			**Non cardholders**		
	Parameter Estimate[1]	Std. Error	Odds Ratio	Parameter Estimate	Std. Error	Odds Ratio
Intercept	1.96	.31	7.10	1.66	.24	5.28
Controls						
SJOB1 prof	-.87**	.37	0.42	.45	.36	1.57
SJOB2 clerk	.11	.20	1.11	-.20	.16	0.82
SJOB3 service	-.20	.18	0.82	.05	.16	1.05
SJOB4 primary	.13	.43	1.13	.12	.32	1.13
SJOB5 manuf	-.02	.23	0.98	.02	.23	1.02
KIDS1 one	.33	.27	1.39	.10	.22	1.10
KIDS3 3+	-.03	.18	0.98	-.12	.16	0.89
UNION2 consen	-.10	.15	0.91	-.25*	.13	0.78
UNION3 d/s/w	.22	.26	1.24	-.21	.20	0.81
NON-METRO	-.00	.19	0.99	.00	.18	1.00
COHO2 24-35m	-.36*	.20	0.70	-.54**	.16	0.59
COHO3 12-23m	-.68**	.19	0.51	-1.28**	.17	0.28
COHO4 9-12m	-2.98**	.23	0.05	-2.87**	.27	0.06
COVER 80%	-.03	.17	0.97	.16	.14	1.17
COVER missing	-.08	.28	0.92	-.27	.20	0.77
PREN some	.23	.19	1.26	.56**	.15	1.74
Risk factors						
TOILET3 none	-.33	.22	0.71	-.47**	.21	0.62
MAGE1 <20	-.25	.32	0.78	-.94**	.29	0.39
MAGE4 40+	.04	.31	1.04	1.09**	.35	2.96
MED1 no ed	.03	.31	1.03	-.73**	.24	0.48
MED4 h.s.+	-.22*	.33	0.80	.36	.30	1.44

Table 6, cont. Logistic regression results: Selected predictors of full immunization status, Colombia 1985-1990, (source: DHSII)

Variable	**Cardholders**			**Non cardholders**		
	Parameter Estimate	Std. Error	Odds Ratio	Parameter Estimate	Std. Error	Odds Ratio
HHEAD2 female	-.33*	.20	0.72	-.39**	.17	0.68
PBI1 missing	-.02	.27	0.98	-.05	.21	0.96
PBI2 <=24m	-.49**	.17	0.61	-.16	.15	0.85
HMORT 1+ dead	.06	.24	1.06	-1.15**	.16	0.32
N	1654			1955		
Model fit[2]	1825.4-1518.3=307.1,25df			2272.9-1905.8=367.0,25df		

** significant at the p < 0.05 level; * significant at p < 0.10 level

[1] contrast with omitted categories: no work past 5y; mat age 21-39; 2 children; primary, some secondary mat ed; metro; male head; formal union; toilet, ages 36m+; no prenatal; prev birth int >24m; no child death; OPV3<80%

Table 7. Logistic regression results: Selected predictors of full immunization status, Colombia 1985-1990, n=3509 (source: DHS II)

	Direct effects		Multiplicative	
Variable	Coefficient[1]	Std. Error	Coefficient	Std. Error
Intercept	1.56	.19	1.71	.19
Controls				
SJOB1 prof	-.14	.24	-.14	.25
SJOB2 clerk	-.08	.12	-.08	.13
SJOB3 service	-.05	.12	-.05	.12
SJOB4 primary	.08	.25	.15	.25
SJOB5 manuf	-.04	.16	-.02	.16
KIDS1 one	.17	.17	.19	.17
KIDS3 3+	-.05	.12	-.07	.12
UNION2 consen	-.20	.10	-.20**	.10
UNION3 d/s/w	-.06	.15	-.04	.15
NON-METRO	.01	.13	.00	.13
COHO2 24-35m	-.46**	.12	-.46**	.12
COHO3 12-23m	-.99**	.12	-1.00**	.13
COHO4 9-12m	-3.00**	.17	-3.00**	.17
COVER 80%+	.10	.10	.09	.11
COVER missing	-.20	.16	-.22	.16
PREN some	.46**	.11	.43**	.11
CARD has card	.40**	.09	.16	.12
Risk factors				
TOILET3 none	-.41**	.15	-.54**	.18
MAGE1 <20	-.17**	.09	-1.08**	.28
MAGE4 40+	.57**	.22	.56**	.23
MED1 no ed	-.41**	.18	-.76**	.24
MED4 h.s.+	.09	.21	.06	.22
HHEAD2 female	-.33**	.13	-.45**	.16
PBI1 missing	-.02	.16	-.02	.16
PBI2 ≤24m	-.30**	.11	-.16	.14
HMORT 1+ dead	-.75**	.13	-1.13**	.16

Table 7, cont. Logistic regression results: Selected predictors of full immunization status, Colombia 1985-1990, n=3509 (source: DHS II)

	Direct effects		Multiplicative	
Variable	Coefficient[1]	Std. Error	Coefficient	Std. Error
Interactions				
CARD*TOILET3			.32	.25
CARD*CMAGE1			.97**	.40
CARD*CMED1			.78**	.39
CARD*HHEAD2			.18	.23
CARD*PBI3			-.31	.21
CARD*HMORT			1.18**	.28
Model fit[2]	603 (26df)		638 (32df)	

** significant at $p < 0.05$ level; * significant at $p < 0.10$ level

[1] contrast with omitted categories: no work past 5y; mat age 21-39; 2 children; primary, some secondary mat ed; metro; male head; formal union; toilet; ages 36m+; no prenatal; prev birth int >24m; no child death; OPV3 <80%

[2] -2(log likelihood: difference between null, full models) = 35 (6df), distributed chi square, p<0.001.

Table 8. Treatment effects model of full immunization, 1985-1990, (source: DHSII Colombia)

Variable	**Probit**		**Treatment[1]**	
	ß	S.E.	ß	S.E.
Controls				
SJOB1 prof	.02	.13	.14	.09
SJOB2 clerk	-	-	-	-
SJOB3 service	.07	.06	.02	.05
SJOB4 primary	.15	.15	.05	.11
SJOB5 manufacturing	-	-	-	-
KIDS1 one child	.29**	.07	-.04	.05
KIDS3 3+ children	.24**	.05	.09**	.04
UNION2 consensual	.07	.05	.04	.04
UNION3 d/s/w	.13	.09	.19**	.06
METRO2 non-metro	-	-	-	-
COHO2 24-35m	-.06	.06	-.14**	.05
COHO3 12-23m	-.40**	.06	-.48**	.07
COHO4 9-12m	-1.61**	.09	-1.10**	.08
PREN some	.71**	.05	-.08	.06
VCARD has card	.32**	.05	2.06**	.16
Risk factors				
TOILET3 none	-.13*	.08	-.06	.06
MAGE1 <20	-.32**	.12	.11	.10
MAGE4 40+	.41**	.12	-.14	.08
MED1 no formal ed	-.16	.11	-.13	.09
MED4 h.s.+	.14	.12	.20**	.08
HHEAD2 female	-.15**	.07	.03	.05
PBI2 $\leq$24m	-.07	.06	-.11**	.05
HMORT2 1+ dead	-.42**	.05	.05	.06
-2 logL	3612		10540	
χ^2	489.5(19df)		-	
F (2,3606)	-		0.00	
adj R^2 (1st stage)	-		-0.48	
N	3609		3609	

** significant at the p < 0.05 level; * significant at p < 0.10 level

[1] two stage least squares; instruments=SJOB2,SJOB5,METRO2.

[2]Contrast with omitted categories: no work past 5y; mat age 21-39; 2 children; primary, some secondary mat ed; mtero; male head; formal union; toilet, ages 36m+; no prenatal; prev birth int >24m, none; no previous child death; no card.

Table 9. Time-dependent effects of prenatal use on the hazard of full immunization, Colombia 1985-1990, n=1376 (Source: DHSII, MOH)

Variable	Coefficient (S.E.)	Relative Risk
Intercept	1.75 (.22)	-
Fixed covariates		
Controls		
SJOB1 prof	-.46** (.25)	0.63
SJOB2 clerk	-.12 (.10)	0.89
SJOB3 service	-.22** (.10)	0.80
SJOB4 primary	-.27 (.24)	0.76
SJOB5 manuf	-.37** (.12)	0.69
KIDS1 one child	.22** (.09)	1.25
KIDS4 3+	-.13 (.12)	1.20
COHO2 24-35m	.18* (.10)	1.52
COHO3 12-23m	.42** (.10)	0.80
COHO4 9-11m	.36** (.18)	1.43
PREN some	-.26 (.17)	0.77
Risk factors		
TOILET3 none	-.58** (.13)	0.56
MED1 no ed	-.37** (.18)	0.69
MED4 h.s.+	-.23 (.20)	0.88
UNION2 consensual	-.25** (.08)	0.78
UNION3 d/s/w	-.08 (.13)	0.92
Time-varying covariates		
COVER 60-79%	.19 (.16)	1.21
COVER 80-89&	.36** (.16)	1.43
COVER 90%+	.40** (.15)	1.49
COVER missing	.22 (.19)	1.25

Table 9, continued. Time-dependent effects of prenatal use on the hazard of full immunization, Colombia 1985-1990 (source: DHSII, MOH)

Variable	Coefficient (S.E.)	Relative Risk
Durations		
D2=2m	.78** (.29)	2.18
D3=3m	.46 (.35)	1.58
D4=4m	-.20 (.49)	0.82
D5=5m	.48 (.39)	1.62
D6=6m	.61 (.39)	1.84
D7=7-9m	-.17 (.36)	0.84
D8=10-11m	-.88 (.53)	0.42
D9=12-15m	-.33 (.45)	0.72
D10=16-18m	.37* (.21)	1.45
D11=19m+	.02 (.15)	1.02
Interactions		
D2*PREN	.71** (.30)	2.03
D3*PREN	.54 (.37)	1.72
D4*PREN	1.11** (.51)	3.03
D5*PREN	.63 (.41)	1.88
D6*PREN	-.06 (.43)	0.94
D7*PREN	.59 (.39)	1.80
D8*PREN	1.41** (.55)	4.10
D9*PREN	.57 (.50)	1.77
- Log-likelihood	-1548.87	

[1] D2=2 months; D3=3 months; D4=4 months; D5=5 months; D6=6 months;
D7= 7-9 months; D8=10-11 months; D9=12-15 months; D10=16-18 D11=19+months

Table 10. Comparison of direct effects hazard models for full immunization, Colombia 1985-1990, n=1376
(Source: DHSII, Ministry of Health).

	Piecewise[1]	Gompertz	
		no heterogeneity	heterogeneity
Variable	Coefficient (S.E.)	Coefficient (S.E.)	Coefficient (S.E.)
Intercept	1.33 (.19)	.92 (.45)	1.63 (.86)
Gamma		-1.06** (.45)	2.19** (.77)
Fixed covariates			
Controls			
SJOB1 prof	-.45* (.25)	-.44* (.24)	-.59** (.26)
SJOB2 clerk	-.12 (.10)	-.13 (.10)	-.19 (.12)
SJOB3 service	-.21** (.10)	-.23** (.09)	-.23** (.11)
SJOB4 primary	-.26 (.24)	-.27 (.23)	-.43* (.24)
SJOB5 manuf	-.37** (.12)	-.38** (.12)	-.52** (.13)
KIDS1 one	.22** (.09)	.23** (.08)	.23** (.10)
KIDS4 3+	-.13 (.12)	-.13 (.11)	-.29** (.11)
COHO2 24-35m	.18* (.10)	.25** (.10)	.18* (.11)
COHO3 12-23m	.42** (.10)	.53** (.10)	.48** (.11)
COHO4 9-11m	.36** (.18)	.35* (.19)	.32* (.19)
PREN some	.20** (.11)	.22** (.10)	.29** (.11)
Risk factors			
TOILET3 none	-.57** (.13)	-.61** (.13)	-.78** (.13)
MED1 no ed	-.37** (.18)	-.39** (.16)	-.43** (.19)
MED4 h.s.+	-.22 (.20)	-.23 (.18)	-.24 (.21)
UNION2 consen	-.25** (.08)	-.27** (.07)	-.30** (.08)
UNION3 d/s/w	-.08 (.13)	-.08 (.13)	-.14 (.13)
Time-varying covariates			
COVER 60-79%	.20 (.16)	.19 (.16)	.17 (.16)
COVER 80-89%	.36** (.16)	.32** (.16)	.37** (.17)
COVER 90%+	.40** (.15)	.37** (.15)	.38** (.16)
COVER missing	.23 (.19)	.23 (.18)	.24 (.19)
Factor loading			2.59** (.49)
Probability mass			.08** (.02)
Log-likelihood	-1540.03	-1399.38	-1405.23

[1] Coefficients for D2=1m; D3=2m; D4=3m; D5=4m; D6=5m; D7=6-9m; D8=10-12m; D9=13-15m; D10=16-18m; D11=19+m not shown

Fig. 1. Cumulative effects of selected risk factors on the probability of full immunization, controlling for cardholdership, DHSII/Colombia (n=3609)

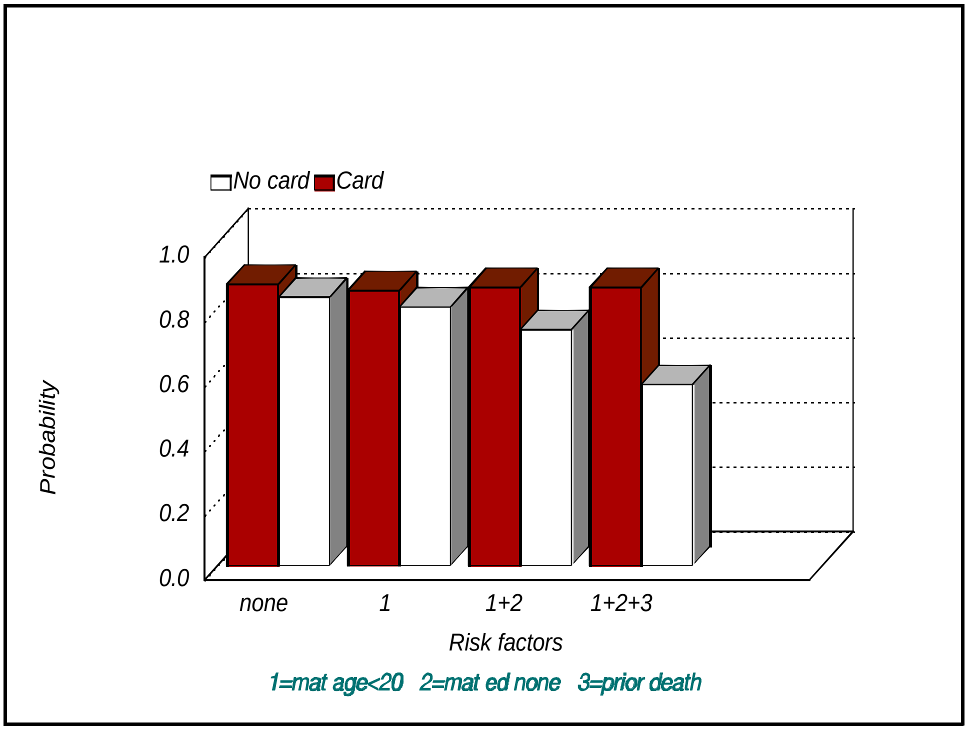
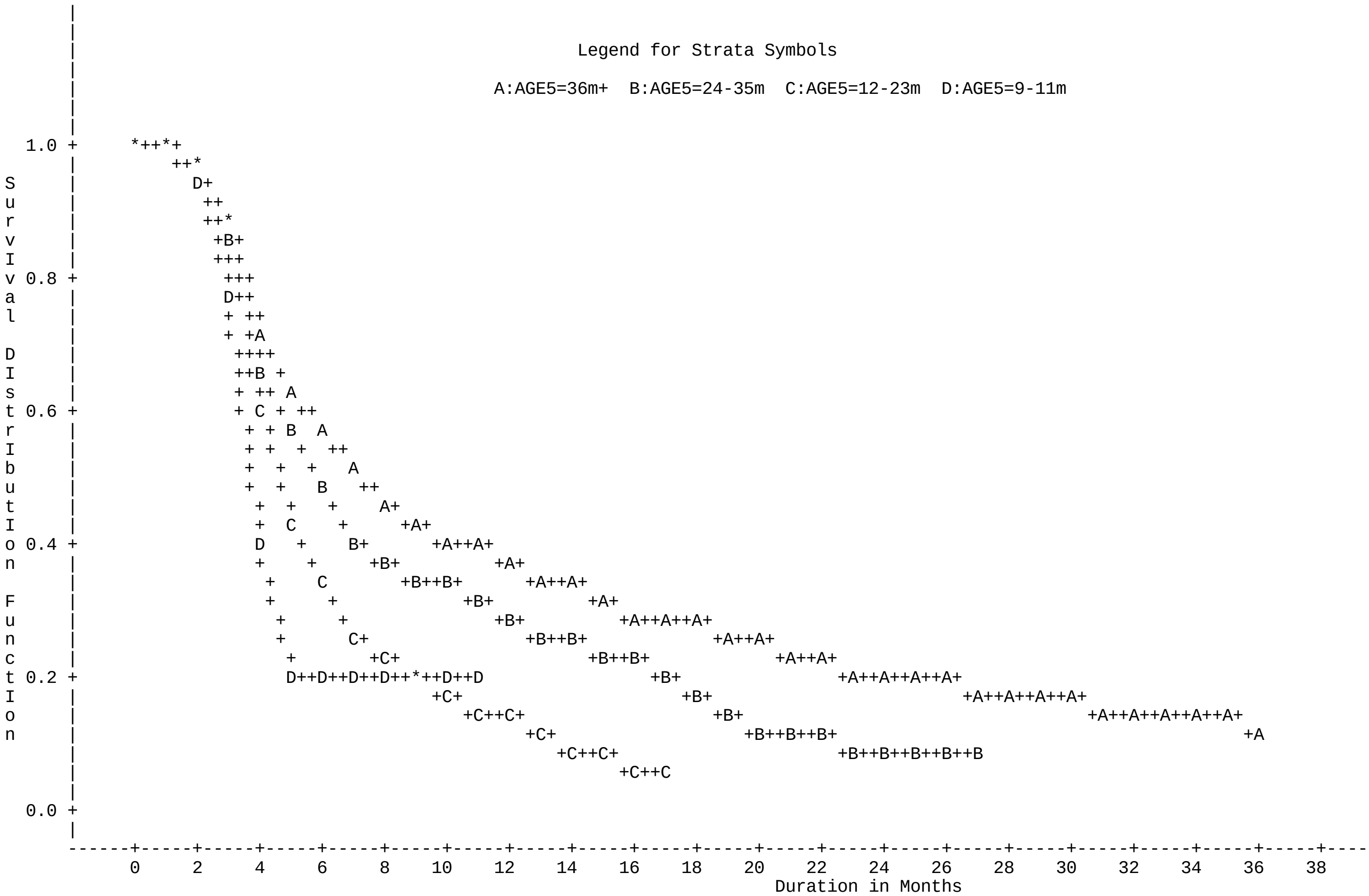
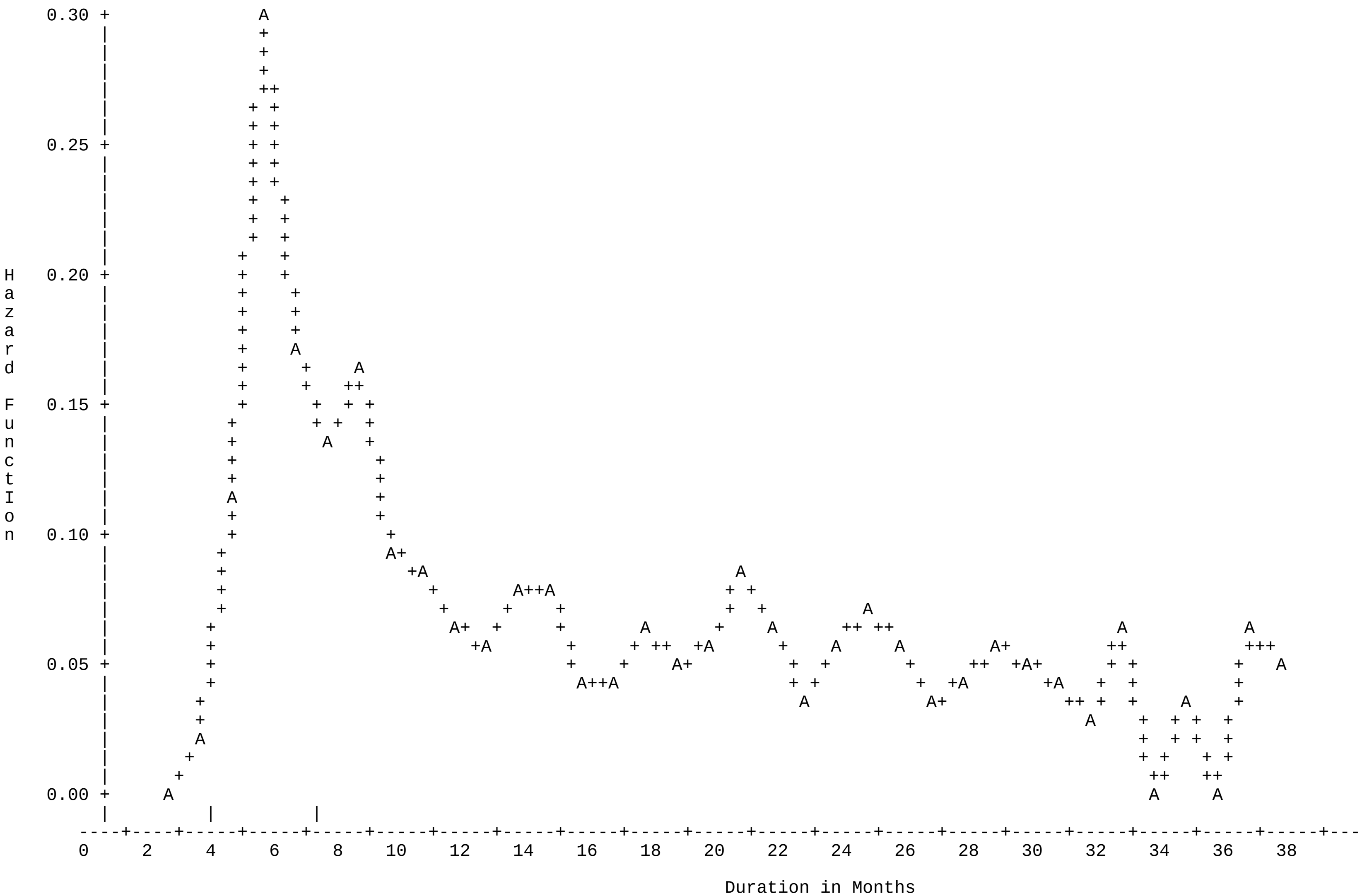


Figure 2. Survival function estimates by birth cohort, full immunization among cardholders (n=1376), DHSII Survey, Colombia, 1985-1990[1]



[1] duration of exposure begins at age 9 months

Fig. 3. Conditional hazard function for full immunization, cardholders ages 9-62 months, DHSII/Colombia (n=1376) [1]



[1] duration of exposure begins at age 9 months

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