# Socioeconomic and Kinship Factors in Infant and Child Mortality in Historical Slavonia

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## PAA Extended Abstract

Slavonia (Croatia) was the borderland between the Habsburgs and the Ottomans from 1698 until 1878. From 1745 to 1871 it was divided into a northern civil zone with an enserfed peasantry under Hungarian control, and a southern military zone under Austrian control with peasant serfs obliged to render perpetual military service. From its liberation by Austro-Hungarian forces in 1683 to the mid-nineteenth century, the region was subject to periodic heavy military mobilization. From the mid-eighteenth century on, it witnessed the collapse of feudalism, increases in wage labor, immiseration of the peasants, overpopulation, and the beginnings of heavy emigration. Households in this region were generally organized along patrilineal extended and fraternal joint lines, where inmarrying women typically lived with their husband and his adult brothers.

Previous research has shown that maternal mortality in this region was strongly affected by social and political conditions that diminished the amount of male labor available to family farming and by the size and composition of a woman's kin networks (Hammel and Gullickson 2004, 2005). This paper extends this model to the survival of offspring in these families.

The particular focus of this research is on the availability of female labor for the care of infants and children. We postulate that infants and children have lower survival chances when labor for their care is diminished. We propose that when male labor is diverted because of military and labor market conditions and women are forced to assume male tasks in subsistence farming, children suffer. However, we also propose that both the size and the structure of the household female labor pool are important. Where more women are available per child, survival should be higher. However, if women other than the mother of a child have competing interests, survival of that mother's child will not benefit as much. These competing interests will include whether the other women are themselves married, how many children they have, and of what age and gender, and whether the other women are themselves nursing. Having other children will diminish the availability of labor as a public good in the household or its kin network. Having older children, so that the familial unit is close to the point of fission from the larger familial aggregate, will diminish the availability of labor. Having a nursing child may actually increase the survival of other children, because it permits surrogate nursing.

The data are family reconstitutions from seven contiguous Catholic parishes in south central Slavonia (Croatia) c. 1750-1900, with scattered evidence from libri status animarum (LSA). Six parishes were in the Military Border of Croatia where male serfs were obliged to render military service and to work in fortress and road construction; one was in Civil Croatia where serfs were obliged to work corvee labor on latifundia. Of approximately 25,000 marriages, 100,000 baptisms, and 90,000 burials, a subset has sufficient linkage information to permit determination of parity and/or reconstruction of the agnatic networks that underlie the patrilineal extended and fraternal joint households common in that time and place. We can determine parity and consanguineal kin for 8,737 mothers in 37,757 births because we have their baptismal and marriage records. We can find the consanguineal kin of fathers in 6,261 of the 8,737 for 27,846 of the 37,757 births because we have the fathers' baptismal records. Where we have occasional LSA, we have found good matching between those and the reconstitutions on the children of conjugal pairs and on sibling sets of coresident brothers.

The model we use is an extension of the piecewise constant hazard model typically used to study infant and child mortality (see Pebley and Stupp (1987) for a discussion of the model and physiological variables). The child's first five years are broken into exposure periods of 0-1 month, 1-6 months, 6-12 months, 12-24 months, and 24-60 months. Within each exposure period, the covariates have a multiplicative effect on a baseline constant hazard rate. The physiological baseline model controls for age, parity, twinning, whether the mother has died, the length of the previous birth interval, the length of the subsequent birth interval, and the gender of the child. Currently, we also include a random frailty term for all children in the mother's affinal network in order to control for any unobserved differences in health shared within these networks. We plan to also include a nested frailty term for the mother herself in the future.

Our interests lie in three types of additional variables not included within the physiological baseline model. First, we are interested in how episodic and endemic male labor withdrawal affected children's health. In order to measure episodic male labor withdrawal during period of war in the Austro-Hungarian Empire, we include a dummy variable indicating whether the period under observation was during a crisis period, in which the men in our study area may have been mobilized. We also expect that the increased labor demands of the military parishes will reduce the availability of male labor, even in times of peace. Therefore, we include parish of residence as a set of categorical variables, with the civil parish as the reference.

Second, we are interested in how the composition and size of the affinal and consanguineal kin network of the mother affected children's health. Our variables here mirror the variables used in our study of maternal mortality. We count the number of lateral kin members in both of these networks, in order to assess the economies of scale available to women. Second, we include a separate count of father's brothers' wives, because we expect these relationships to be more fragile, due to the political divisions surrounding eventual family fission within these types of households. Finally, we include a measure of the relative age rank of the father, because we expect a hierarchical division of labor and resources along these lines.

Third, we are interested in how the composition of existing children in the family network will affect children's health. We expect that large numbers of children may reduce the availability of resources, but also produce economies of scale if these children are of similar ages. As children age, they should become more beneficial as producers, who can help divert more labor to the health care of younger children. However, this relationship will be complicated for fraternal cousins within the household because the threat of family fission and the distribution of inheritance shares may reduce the desire to share care and resources across nuclear sub-units. In order to measure these potentially different effects, as a first step, we include counts of siblings and fraternal cousins in three age groups: 0-5 years, 5-10 years, and 10 years until marriage or censoring at age 25.

Our preliminary results, shown in Table 1, meet some expectations, but also produce some surprising results. In terms of male labor withdrawal, we find that infant and child mortality increased during crisis periods, and was generally higher in the military parishes relative to the civil parish. Both of these results suggest that the withdrawal of male labor place greater labor demands on the other members of the household that ultimately result in less care for young children.

Surprisingly, the kinship structure variables that were important predictors of maternal mortality are not in any case important predictors of infant and child mortality. Neither the size nor composition of affinal and consanguineal groups has any effect on child health. Similarly, the father's relative rank within the household also has no effect.

The composition of children in the household, however, is very important for infant and child health. Among siblings, we find that the strongest effects are for children 0-5 years of age, who reduce the mortality risk of the ego child considerably, suggesting that economies of scale in childrearing outweigh the resource drain of non-producing children. The number of children 5-10 years of age has little effect on the mortality risk of the ego child, but as the number of children greater than 10 years of age increases, the mortality risk decreases substantially, suggesting that as children become more valuable as producers in the household, they have a positive effect on their younger siblings' health.

Among fraternal cousins, the results are very different. In no case does increasing the number of fraternal cousins reduce the mortality risk of the ego child. Indeed as fraternal cousins age from the 0-5 to 5-10 age group, the mortality risk of the ego child increases, suggesting that the aging of children within these joint households may lead to more resource hoarding among the nuclear sub-units. Interestingly, the number of fraternal cousins older than 10 has no effect on mortality risk, whatsoever. It may be that the issue of resource hoarding is outweighed by these children's general productivity. On the other hand, it may indicate that households with many older children have, in effect, already completed the process of family fission and are operating as independent nuclear units.

These results are preliminary and may change as we experiment with more complex coding of family structure. Currently, we are exploring the effects of closeness in age to ego child and the gender composition of existing siblings and fraternal cousins.

The intent of the work is not only substantive – to explain demographic conditions in the past – but in addition both practical and theoretical. In a practical sense, the insights gained into demographic outcomes by the application of historical and ethnographic knowledge in historical Slavonia are, in the presence of similar information, extensible to modern peasant societies, e.g. in Africa, South Asia, and Latin America, especially where extended household structures also exist. In a theoretical sense this work seeks to enrich demographic analysis by showing how ethnographic knowledge of microstructural factors is essential to a close understanding of the demographic behavior of social actors.

**References**

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**Table 1 - Coefficients (log risk ratios) from a constant piecewise hazard model of dying within the first five years of life in seven Slavonian parishes, 1750-1898**

| Variable | Coefficient | SE |
|---|---|---|
| *Baseline Risk* | | |
| 0-1 month | | |
| 1-5 months | -1.481 | 0.035*** |
| 6-11 months | -1.874 | 0.036*** |
| 12-23 months | -2.347 | 0.034*** |
| 24-60 months | -3.315 | 0.034*** |
| | | |
| *Physiological Factors* | | |
| Age | -0.015 | 0.004*** |
| Age squared | 0.001 | 0.0001*** |
| Parity | 0.032 | 0.019 |
| Parity squared | -0.001 | 0.002 |
| Previous birth interval | | |
| First birth | -0.137 | 0.045** |
| Short interval | -0.003 | 0.039 |
| Medium interval (ref.) | - | |
| Long interval | -0.070 | 0.025** |
| Previous child died | -0.305 | 0.032*** |
| Twinning | 0.900 | 0.042*** |
| Mother died | 0.334 | 0.075*** |
| Male child | 0.084 | 0.019*** |
| | | |
| *Socioeconomic Factors* | | |
| Crisis period | 0.190 | 0.021*** |
| Parish | | |
| C (ref) | | |
| B | 0.513 | 0.053*** |
| G | 0.373 | 0.034*** |
| O | 0.200 | 0.044*** |
| P | 0.500 | 0.033*** |
| S | 0.564 | 0.042*** |
| V | 0.393 | 0.036*** |
| year (origin 1715) | 0.011 | 0.001*** |
| | | |
| *Kin Network Factors* | | |
| Number of mother's affinal kin | -0.005 | 0.011 |
| Number of mother's consanguineal kin | 0.003 | 0.006 |
| Number of father's brothers' wives | -0.020 | 0.040 |

| | | |
|---|---|---|
| Father's relative age rank (senior=1, junior=0) | 0.004 | 0.037 |
| no father's brother's wives | -0.035 | 0.050 |
| | | |
| *Child Composition Factors* | | |
| Number of siblings | | |
| 0-5 years old | -0.164 | 0.021*** |
| 5-10 years old | -0.013 | 0.020 |
| 10+ years old | -0.063 | 0.020** |
| Number of fraternal cousins | | |
| 0-5 years old | 0.028 | 0.017 |
| 5-10 years old | 0.076 | 0.018* |
| 10+ years old | -0.017 | 0.011 |
| | | |
| Variance of affinal network frailty | 0.082 | |

* p<.05; ** p<.01; *** p<.001