Recent Trends in Infant Mortality In Kazakhstan: The Result of Discrimination?

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Abstract: In the past decade in Kazakhstan, infant mortality rates have significantly increased over time. In order to understand the determinants of change, I use a modification of the Oaxaca-Blinder decomposition technique for logit models to decompose the shifts in determinants of infant mortality in Kazakhstan into effects due to changes in the relative riskiness of different determinants of infant mortality and to changes in population composition. I examine covariates relating to ethnicity, geographic location, maternal education, household economic status, and characteristics of pregnancies. I find that the most significant changes are driven by the increased likelihood of infant death among non-ethnically Kazakh infants, the decreasing contribution to the birth pool from wealthier households, and the increasing risk to females and infants from multiple births.

Introduction:

Since the early 1990s, as a result of political independence, economic and social conditions in several republics of the former Soviet Union, including Kazakhstan, have fluctuated drastically. Following independence on December 16th, 1991, Kazakhstan's economy contracted by more than 50%, in part due to the loss of approximately 8% of pre-independence GDP that came from transfers from the central Soviet government,[1] as well as a loss of trading partners from the former Soviet Union and the effects of transitioning from a centrally-planned to a rudimentary market economy. This collapse of economic output has had negative effects on population health indicators throughout Central Asia, such as lowered life expectancy and rising adult and infant death rates. Although the macroeconomic situation has stabilized, with inflation falling and the economy growing once again, the short- and long-term consequences for individual health and welfare are not well documented.

One consequence of the political and economic turmoil of the 1990s in Kazakhstan is an increase in infant mortality rates. According to the World Bank,[2] following a decline in infant mortality rates from 50 per 1,000 live births in 1980 to 42 per 1,000 live births in 1990, infant mortality rates then rose to 56 per 1,000 live births in 1995 and to 81 per 1,000 live births in 2001. In general, infant mortality rates throughout the Central Asian countries, including Kazakhstan, have risen from Soviet-era levels, while rates in other formerly Soviet Union countries, including Russia, have fallen.[3] In Kazakhstan, infant mortality levels, both pre- and post-independence, have been relatively high, with Central Asia generally featuring the highest infant mortality rates during the Soviet-era period.[4] Furthermore, while the formerly Soviet Union nations in Eastern Europe and the Caucasus have improved their infant mortality statistics and the

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[1] Falkingham J., Klugman J., Marnie S., and Micklewright J. (1997) "Household Welfare in Central Asia: An Introduction to the Issues." Household Welfare in Central Asia, ed. Falkingham J., Klugman J., Marnie S., and Micklewright J. New York: St. Martin's, 1997.

[2] http://devdata.worldbank.org/hnpstats/HnpAtaGlance.asp?sCtry=KAZ,Kazakhstan

[3] http://devdata.worldbank.org/hnpstats/files/Tab2_19.xls

[4] Jones E. and Grupp F. W. (1983) "Infant Mortality Trends in the Soviet Union." *Population and Development Review*, 9:2, 213-246.

other Central Asian nations have kept levels of infant mortality roughly the same, Kazakhstan's infant mortality rates have increased rapidly.[5]

Table 1 summarizes recent economic and mortality conditions for the fifteen countries of the former Soviet Union, including Kazakhstan. Notice that, although life expectancy at birth has declined for many formerly Soviet countries since 1980, this decline is greater in Kazakhstan than in any of the other fourteen countries. At least some of this decline is likely due to increased infant mortality, which has increased more in Kazakhstan than in any other former Soviet nation (apart from Tajikistan, where infant mortality levels are now higher but are unknown for 1980). Although general underinvestment in Central Asia during the Soviet period can account for poor health conditions and inferior health statistics, including infant mortality rates, in this region in 1980, relative to the other areas of the Soviet Union,[6] Soviet-era underinvestment does not explain why other countries in Central Asia, such as Kyrgyzstan, have managed either to improve infant mortality rates or to keep them at steady levels while Kazakhstan has not, despite inheriting similar health infrastructure.

While the infant mortality levels from both periods are alone worrisome, the increase in infant deaths over the past decade gives particular cause for concern, both because this increase represents a setback to child health efforts of the past twenty years and because it presages serious long-term health consequences in the general population that will result from the arresting of economic development and lowered living standards that followed independence. The Kazakhstani population has endured high rates of infant mortality before, corresponding to time periods such as Stalin's artificial famines during agricultural collectivization in the 1930s and World War II; even after socioeconomic conditions and infant mortality rates improved, the cohort effects on infants and children surviving these harsh periods resulted in lowered life expectancies in adult survivors, decades after the crisis has passed.[7] Therefore, in order to improve infant survival in the present, as well as plan for the long-term consequences of economic and social upheaval,

[5] http://devdata.worldbank.org/hnpstats/files/Tab2_19.xls

[6] Anderson B. A. and Silver B. D. "Patterns of Cohort Mortality in the Soviet Population." *Population and Development Review*, 15:3, 471-501.

[7] Ibid.

it is important to document i.) which segments of the population have suffered the most, in terms of infant mortality, in the past decade, ii.) how the effects of factors that predict infant death are changing, and iii) how these changes affect infant death rates.

Research Questions:

In Kazakhstan since independence, in addition to dramatic macroeconomic changes, there has been significant out-migration of the ethnically Russian population, particularly amongst the more educated and economically well-off; household welfare has declined as individuals have ceased receiving subsidies from the Soviet state; the health system has lost capacity due to the loss of funds from the Soviet government and an out-migration of trained personnel; and the composition of births has changed, with an increasing number of births to rural, ethnically Russian and Kazakh women from higher-risk areas of the country.

These changes might contribute to the fluctuation of infant mortality rates in two main ways. Firstly, changes in factors such as household wealth or the availability and quality of medical care might affect infant mortality rates by changing the risk of infant death to infants born to specific groups, so that an infant born to a Russian mother in 1995 or 1998, for instance, would be born into conditions less favorable for survival than in 1992. These changes that result in a harsher environment could affect all sectors of Kazakhstan's society equally, or, more likely, could disproportionately affect certain ethnic, economic, and regional groups.

Secondly, changes in the economic, political, and social conditions of the country, including the out-migration of non-Kazakh ethnic groups, could affect the composition of the birth pool. The risk of death for an infant with given characteristics may remain static, but higher-risk infants may represent an increasing share of the total number of infants over time as mothers who give birth to higher-risk infants might be more likely either to still be in the country or to select into childbirth as time progresses.

The observed changes in infant mortality rates could also arise from a combination of these two factors: the risk to different individuals could change over time, as well as individuals at relatively higher risk contributing a larger proportion of births to the overall pool. The next section examines specific hypotheses regarding five main factors that affect infant mortality in Kazakhstan:, economic status, maternal education, ethnicity, regional effects, and high-risk pregnancies.

Household Economic Status

Economic factors, such as household wealth, income, and parents' education, are hypothesized to affect infant survival indirectly by determining the quality and quantity of resources available to the infant, in terms of food, housing, medical care, and time and ability to supervise the child.[8] The extent to which income and household wealth matters for infant mortality is the subject of debate: Casterline et. al.,[9] for instance, find that household income does not play a role in infant survival (although it does in early child survival), once other socio-economic characteristics, such as maternal and paternal education and region of residence, are controlled for. Other authors, such as Cramer,[10] find a strong, consistent link between increased infant survival chances and household income. In general, a minimal level of household income seems important to infant survival, although this positive effect may level off after reaching a certain threshold, perhaps because basic changes that can be made at lower income levels, such as improvements to household toilet and water facilities, increase the probability of survival more than comparative luxuries that can be purchased at higher income levels.[11]

Compared to other Central Asian countries, Kazakhstan is not as absolutely poor, mainly due to large reservoirs of gas and oil, especially in the Caspian Sea region. This national wealth, however, has not necessarily translated into higher levels of household wealth for the average citizen, and disparities in household wealth have certainly risen

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[8] Mosley W. H. and Chen L. C. (1984) "An Analytical Framework for the Study of Child Survival in Developing Countries." Population and Development Review, 10: Supplement(Child Survival), 25-45.

[9] Casterline J., Cooksey E. and Ismail A. F. E. (1989) "Household Income and Child Survival in Egypt." *Demography*, 26:1, 15-35.

[10] Cramer J. C. (1995) "Racial and Ethnic Differences in Birthweight: The Role of Income and Financial Assistance." *Demography*, 32:2, 31-47.

[11] DaVanzo J. (1988) "Infant Mortality and Socioeconomic Development: Evidence from Malaysian Household Data." *Demography*, 25:4, 581-595.

since Soviet days[12] (although, in Soviet times, citizens could be 'privileged' for reasons other than income, and thus have better food, housing, and access to medical care than the average citizen).

Also pertinent to the discussion of any relationship between economic status and health in the post-Soviet context is the role of informal payments, or bribes. While state-sponsored health care is nominally free at the point of service in Kazakhstan, there is ample evidence to suggest that bribes to providers are often given out of tradition or gratitude, or demanded prior to service, thus indicating another reason why income and economic status may be important determinants of infant mortality.[13] In addition to Lewis's findings that, in general, the cost of informal charges for medical care has deterred individuals from seeking care, Falkingham[14] finds specifically that forty percent of currently pregnant women in neighboring Tajikistan were not receiving prenatal care, despite the majority having received care for previous pregnancies, citing 'inaffordability' as the reason. She also found that parents of low economic status were finding it increasingly difficult to afford care for sick infants, which affected overall levels of infant and child mortality.

Changes in household wealth and general economic status over time might contribute to changes in the risk of an infant death via a similar mechanism in Kazakhstan, especially as households adjust to the longer-term economic changes such as an end to subsidies from the Soviet government, privatization of state firms, and allowing prices and incomes to be determined in a market economy. These changes could make a lower economic status relatively more risky over time in terms of infant survival as the public safety net shreds or, conversely, higher economic status could become more advantageous over time as individuals are able to purchase more things that aid in infant survival than were available in the past. Alternately, infants from lower-status households

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[12] Falkingham J., Klugman J., and Marnie S. (1997) "Household Welfare in Central Asia: An Introduction to the Issues." In Household Welfare in Central Asia. Ed. Falkingham, J. NY: St. Martin's Press.

[13] Lewis M. (2000) "Who is Paying for Health Care in Central Asia?" International Bank for Reconstruction and Development.

[14] Falkingham J. (2004) "Poverty, Out-of-Pocket Payments and Access to Health Care: Evidence from Tajikistan." *Social Science and Medicine*, 58, 247-58.

may always have been at higher risk of death, and could represent a greater share of births over time. I will look at how average household wealth changes over time, as well as how the effect of household wealth on the probability of infant survival changes.

Maternal Education

Past research has identified maternal educational status as a particularly important predictor of infant mortality. Sandiford et. al.,[15] for instance, estimated that the effect of educating mothers on infant mortality exceeded those from improving sanitation, better water supplies, or increasing income. Sastry[16] finds that the additional knowledge of mothers can even serve as a substitute for sanitation and health facilities in infrastructure-poor areas. Education is hypothesized to have a direct effect on infant mortality by allowing mothers to better incorporate public health knowledge into parenting practices, including information about proper nutrition, sanitation, supervision, and health risks. Maternal education also has an indirect effect, as it enables mothers to serve as better advocates for their children within a given health system, and changes the balance of power within the household towards resource distribution that favors the mother and the children.[17] A mother who completes secondary school increases her child's chances of surviving, and completing higher education yields an even greater protective effect.

In Kazakhstan, a high degree of high school completion for both men and women was one of the primary positive legacies of the Communist government – secondary school completion rates have not fallen much since independence either. Still, there is considerable variation in whether a mother has obtained higher education, and the percentage of women who do is falling slowly over time. Furthermore, recent changes in the educational system, such as a move towards a more Islamic/religious-based education, as well as societal changes following the breakup of the Soviet Union, such as widespread migration and unemployment among women, could affect the ability of

[15] Sandiford P., Cassel J., Montenegro M., and Sanchez G. (1995) "The Impact of Women's Literacy on Child Health and its Interaction with Access to Health Services." *Population Studies*, 49:1, 5-17.

[16] Sastry N. (1996) "Community Characteristics, Individual and Household Attributes, and Child Survival in Brazil." *Demography*, 33:2, 211-229.

[17] Caldwell J. (1979) "Education as a Factor in Mortality Decline: An Examination of Nigerian Data." *Population Studies*, 29, 259-272.

mothers to obtain and exploit the human capital provided by education in ways that are useful for infant survival. I expect that the educational demographics of the birth population will not change much over time, but that the protective effects of more education might erode, as the nature of this education changes and as resource constraints impede individuals' ability to exploit their human capital.

Population Dynamics and Ethnic Subpopulations

Kazakhstan's population in the past eighty years has undergone a series of migrations that have led to a state with several different sub-populations delineated by ethnic group. During the early Soviet period, in order to impose Soviet bureaucracy and ideology on the nation, large numbers of Russians moved into the region, mostly into the professional and managerial positions in cities. These Russians were better educated and wealthier than the native Kazakh population and, although both education levels and the degree of urbanization of the largely nomadic, rural Kazakh peasantry had increased since the 1920s, Russians, prior to independence, still represented an elite socio-economic group. Russians have historically had better health outcomes than the native Kazakh population, including lower infant mortality rates and better child health indicators, an advantage which has persisted throughout the 1990s.[18] Since independence, there has been substantial migration of ethnic Russians back to Russia, with at least 10% of the pre-independence Russian population leaving,[19] and those remaining representing the poorest and less-well educated among the Russian ethnic group, although the remaining Russians still have lower rates of infant mortality than Kazakhs.

Furthermore, there is a substantial non-Kazakh, non-Russian proportion to the population, mainly due to historical forced and non-forced migration of non-Russian peoples from elsewhere within the former Soviet empire. Although representing over 30 different ethnic groups, including Poles, Germans, Uzbeks, Ukrainians, Tatars, and Belorussians, and almost 13% of the pre-independence population, for simplicity's sake

[18] Buckley C. (2003) "Children at Risk: Infant and Child Health in Central Asia." Working Paper #523, William Davidson Institute.

[19] Brubaker R. (1998) "Migrations of Ethnic Unmixing in the 'New Europe.'" *International Migration Review*, 32:4, 1047-1065.

these 'others' are treated as a single group, with an experience that mirrors neither that of Kazakhs nor that of Russians. In general, these other groups can be considered, economically at least, as midway between the two dominant population groups, and have also experienced substantial out-migration since independence, up to 70% for some ethnic groups.[20]

Rather than a biological difference, this ethnic divide in experience of infant mortality is largely a reflection of the generally higher socio-economic status of Russians, their concentration in urban areas, and higher utilization of health services, although the privileged status of Russian citizens, particularly in terms of gaining government jobs and dominance of the Russian language in higher education, has been eroded since independence.[21] Two factors regarding ethnicity likely contribute to the changes in infant mortality: firstly, the effect of being Russian (or Kazakh or other ethnicity) might be changing over time, as higher-educated and economically well-off Russians leave, with those left behind at higher risk of giving birth to an infant that dies. The relative risk to non-Kazakh ethnic groups for those who stay behind will probably increase, as ethnic Kazakhs gain a more privileged status in government and business, relative to the Soviet period. The demographic composition of births, in terms of ethnicity, has also fluctuated since independence, which could also explain fluctuating infant mortality rates.

Regional Variation

Even in Soviet times, there was a high degree of regional variation in infant and adult mortality rates, as a result of differential economic opportunities and access to services. This geographic variation is particularly reflected in an urban-rural divide that results from concentration of educational and economic opportunities in cities, with poorer health facilities, lack of trained medical personnel, and a general pattern of lower investment in public infrastructure (including hospitals, clinics, and roads) characterizing rural areas.[22,23] There exists also a degree of regional concentration of infant deaths, with

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[20] Ibid.

[21] Ibid.

[22] (1988) "USSR: High Infant Mortality Rates Show Slight Decline in 1980s." *Family Planning Perspectives*, 20:3, 145.

a higher proportion of infant deaths from the southern region of the country. Infrastructure deterioration and erosion of the public health system could occur at different rates in different areas of the country, and is more likely to occur more quickly outside of urban areas and Almaty. Falkingham,[24] for instance, finds that, in neighboring Tajikistan, the health infrastructure fell apart much more quickly in poor, rural areas of the country, with rural women turning increasingly to untrained medical personnel for prenatal care and delivering babies in unheated homes without running water as staffing levels fell and medical facilities closed. Infants born in these conditions subsequently suffered higher rates of infant mortality. I will examine whether the relative advantage or disadvantage of being born in a particular region changes over time, as well as how the regional composition of births changes over time.

High-Risk Pregnancies

There are a number of factors pertaining to the mother and the nature of the pregnancy itself that contribute to an infant's risk of mortality, including characteristics such as parity, the preceding birth interval, and overall maternal health status. Generally, a short (two years or less) preceding birth interval, the length of time since last birth, can lead to suboptimal pregnancy outcomes, such as increased infant mortality and low birth weight,[25] particularly in developing countries.[26] These negative outcomes can be explained by the 'maternal depletion' hypothesis, which maintains that giving birth too soon after a previous pregnancy does not allow the mother to recover from the fatigue and stress (mental and physical) of giving birth. Because maternal nutritional resources are not properly restored, the infant can be born too early, or fail to grow properly in

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[23] Buckley C. (1998) "Rural/Urban Differentials in Demographic Processes: The Central Asian States." *Population Research and Policy Review*, 17:4, 71-89.

[24] Falkingham J. (2002) "Poverty and Access to Maternity Care in Tajikistan." Working Paper A03/09, University of Southampton Statistical Sciences Research Institute.

[25] Hobcraft J. N., McDonald J. W., and Rutstein S. O. (1983) "Child-Spacing Effects on Infant and Early Child Mortality." *Population Index*, 49:4, 585-618.

[26] Winikoff B. (1983) "The Effects of Birth Spacing on Child and Maternal Health." *Studies in Family Planning*, 14, 231-45.

utero. There has been some debate about whether this two-year figure is accurate: when Zhu et. al.[27] examined infant mortality in Utah, they found that the 'ideal' birth interval was between 18 and 23 months. Even in the United States, however, others, including Klerman et. al,[28] have found that birth intervals of two years or less are associated with increased risk of poor pregnancy outcomes, particularly among mothers who are either young or poor (or both).

Parity can also affect infant mortality through maternal depletion, with higher-order children experiencing markedly lower survival rates than lower-parity children, although exactly at which parity infants begin to experience lower survival is the subject of debate. Trussell and Pebley,[29] for instance, indicate that infant survival begins to diminish with the fourth child, while Cooksey and Millman indicate the adverse effects of 'too many' children begin with child seven.[30] First-born children also exhibit a higher risk of infant mortality, perhaps because mothers' reproductive systems are still adjusting to pregnancy and delivery, or because these mothers have less experience in caring for infants.[31,32] There is some evidence that, once maternal age is controlled for, the negative effects of parity are alleviated.[33] Additionally, on the household level, parity can affect infant survival through competition for resources and care with other young children in the household, as well as household crowding, which can lead to the rapid spread of

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[27] Zhu B. P., Rolfs R. T., Nangle B. E., and Horan J. M. (1999) "Effect of the Interval Between Pregnancies on Perinatal Outcomes." *The New England Journal of Medicine*, 340:8, 589-594.

[28] Klerman L. V., Cliver S. P., and Goldenberg R. L. (1998) "The Impact of Short Interpregnancy Intervals on Pregnancy Outcomes in a Low-Income Population." *American Journal of Public Health*, 88,1182-5.

[29] Trussell J. and Pebley A. (1984) "The Potential Impact of Changes in Fertility on Infant, Child, and Maternal Mortality." *Studies in Family Planning*, 15:6, 267-280.

[30] Millman S. and Cooksey E. (1987) "Birthweight and the Effects of Birth Spacing and Breast Feeding on Infant Mortality." *Studies in Family Planning*, 18:4, 202-212.

[31] Bongaarts J. (1987) "Does Family Planning Reduce Infant Mortality Rates?" *Population and Development Review*, 13:2, 323-334.

[32] Pebley A. and Stupp P. (1987) "Reproductive Patterns and Child Mortality in Guatemala." *Demography*, 24:1, 43-60.

[33] LeGrand T. and Phillips J.F. (1996) "The Effect of Fertility Reductions on Infant and Child Mortality: Evidence from Matlab in Rural Bangladesh." *Population Studies* 50, 51-68.

infectious diseases amongst infants and young children in the home.[34] The effects of household composition and sibling rivalry do, however, depend somewhat on the gender and ages of the other siblings in the home. For instance, older siblings might contribute to infant survival by providing extra wages or sources of care.

Maternal age also affects infant mortality, although exactly how has been the subject of some debate. Hobcraft et. al.[35] found increased mortality only for younger mothers (under the age of 18), resulting from physical immaturity. Most authors, however, including Eberstein et. al.,[36] find that maternal age affects infant mortality in a J-shaped way: not only do younger mothers have suboptimal birth outcomes, as noted above, but older (over the age of 35) mothers also experience increased infant mortality resulting from chromosomal deficiencies, compared to younger women. These chromosomal problems can affect both the likelihood of miscarriage and the rate of congenital deformations in live births.

So, infants born to relatively younger or older mothers, infants born in rapid succession following a previous birth, and first-born or children of higher parity are at relatively higher risk compared to infants born after a greater interval, to mid-age-range (ages 18-35) mothers, and second- through approximately sixth-born children. I will examine whether the relative severity of the effect on infant survival of these different risk factors changes over time, as well as whether the composition of births changes over time to include relatively greater or fewer high-risk pregnancies.

I will also include as control variables two other factors that affect infant survival: the gender of the child and whether the child is part of a single or multiple birth. Females, in the absence of notable gender discrimination, have a slight natural survival advantage

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[34] Chidambam V.C., McDonald J. W., and Bracher M. D. (1985) "Infant and Child Mortality in the Developing World: Information from the World Fertility Survey*." International Family Planning Perspectives*, 11: 1, 17-25.

[35] Hobcraft J. N., McDonald J.W. and Rutstein S.O. (1985) "Demographic Determinants of Infant and Early Child Mortality: A Comparative Analysis." *Population Studies*, 39:3, 363-385.

[36] Eberstein I. W., Nam C. B., and Hummer R. A. (1990) "Infant Mortality by Cause of Death: Main and Interaction Effects." *Demography*, 27:3, 413-430.

over males, particularly in the first year of life.[37] Children from multiple births have a lower survivorship in the developing world due to lower birthweights, a tendency for premature birth and fewer resources available post-delivery for each infant.[38]

Data and Methods:

Data

The data for this paper come from two independent, cross-sectional surveys conducted in Kazakhstan, the 1995 and 1999 Demographic and Health Surveys (DHS).[39,40] These surveys are nationally representative of women of reproductive age (15-49), and include data on household and maternal characteristics, as well as retrospective birth histories that include information on all live births, including the month of birth and the month of death (when appropriate). The DHS uses the international definition of live birth in its questionnaires, and the birth data are provided by the mothers, rather than gathered from vital statistics sources.[41] The birth records in the DHS stretch back to a woman's first birth, over thirty years. Because I am interested in recent changes in infant mortality, however, I include in my sample only infants born after independence, which occurred on December 16th, 1991. Because I do not know the exact date of birth, only the month, I include infants born beginning in January 1992.

To increase sample size, the data from the two surveys are pooled together, resulting in over-representation from the earlier years of the study, 1992-94, since infants

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[37] Hobcraft J. N., McDonald J.W. and Rutstein S.O. (1985) "Demographic Determinants of Infant and Early Child Mortality: A Comparative Analysis." *Population Studies*, 39:3, 363-385.

[38] Guo G. and Grummer-Strawn L. M. (1993) "Child Mortality Among Twins in Lesser Developed Countries." *Population Studies*, 47:3, 495-510.

[39] National Institute of Nutrition, Academy of Preventative Medicine, and Macro International (1995) *Demographic and Health Survey Final Report, Kazakhstan 1995*. Almaty, Kazakhstan, and Calverton, MD: ORC Macro.

[40] Academy of Preventative Medicine and Macro International (1999) *Demographic and Health Survey Final Report, Kazakhstan 1999*. Almaty, Kazakhstan and Calverton, MD: ORC Macro.

[41] Having maternally-reported data may actually be better than vital statistics data, since the official definition of 'live birth' changed over time and, at least in Soviet times, official statistics were manipulated for political goals. The drawback of maternally-reported data is, of course, recall bias, as well as improper and inconsistent classification.

born in these years have the potential to be reported on in both the 1995 survey and the 1999 survey. In total, 3,551 live births from the years 1992 – 98, representing all children reported that were born after December 1991 but at least 12 months before the respective survey date, are included in the main analysis sample. Of these 3,551 live births, 186 infants subsequently died; children dying at or before one year were classified as infant deaths, to account for reported age of death 'heaping' at 12 months.

The surveys are constructed in two stages. Initially, regional areas are stratified by dividing the country into health blocks (in urban areas) and villages (in rural areas). In the first stage, primary sampling units are selected with probability proportional to population size; in the second stage, households are then randomly selected from within the primary sampling unit. To account for oversampling of certain regions, I have used the household weights included with the DHS survey as probability sampling weights.

Additionally, because of the way I have constructed my study sample, by appending the 1995 and 1999 datasets together, I have oversampled infants from the years in which the two studies overlap, from 1992-94. In order to control for this oversampling, I have multiplied the relevant probability sampling weights by .5 for these three years, since infants from these years are roughly twice as likely to be included in my study sample.

The standard errors of coefficient and sample mean estimates need to be corrected to account for clustering and sample stratification; to do so, I utilize design-effect adjusted models, incorporating the stratification and primary sampling unit variables that are included in the DHS. All logit models estimated utilized the svylogit command in STATA 8.0;[42] sample means are estimated using the svymean command. These commands apply appropriate weights and correct standard errors for sampling effects.

Additionally, there is differential recall bias for the years of overlap, 1992-94. Because the women from the 1999 survey are reporting on pregnancies that occurred up to seven years prior to survey, while women from the 1995 survey are reporting only on pregnancies during the previous three years, they are subject to differential recall periods. Other studies show that recall bias can affect the reporting of both births and deaths –

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[42] StataCorp. 2001. *Stata Statistical Software: Release 8.0*. College Station, TX: Stata Corporation.

mothers are more likely to omit births that end in deaths, and that this bias gets worse as more time elapses since the birth.[43] In order to assess whether there was substantial differential recall bias between the two samples, I determine the death and birth rates from the years of overlap, 1992-94, separately for the two surveys. Chi-squared tests of the birth and death rates for the two years of full overlap (1992 and 1993; 1994 contains fewer infants from the 1995 survey because the 1995 survey was conducted mid-year and only infants born at least a year prior to the survey were included in my sample) do not indicate significant differences between the two surveys.

In addition to differential recall bias for the period 1992-1994 arising from compiling data from two different surveys, there is also the danger of bias arising from the fact that the population of Kazakhstan has changed between the two surveys, and thus the 1995 survey population might be substantially different from the 1999 population in terms of ethnicity, urbanicity, region of residence, and other variables. Table 2 shows the means of key variables for the years of overlap, calculated separately for the two surveys. T-statistics on the differences between the two means do not indicate significant differences between the two surveys except for whether infants come from the southern or eastern/northern/central regions of the country or from an urban or rural area (the two are highly correlated), as well as significant differences in those who have not completed secondary education. These differences probably do not arise from differential recall bias between the 1995 and 1999 surveys, but from differential internal migration within the country, since women report where they are living not at the time of birth, but at the time of survey (e.g., an infant born in 1992 in the southern region, but living in Almaty at the time of survey, would be wrongly classified as having been born in Almaty). Fortunately, as Agadjanian and Qian[44] indicate, there is relatively little internal migration within Kazakhstan, so generally this question of using current residential data as a proxy for residence at time of birth should generally not be too problematic. A larger problem, one that cannot really be controlled for but should be kept in mind when interpreting results, is that women may have differentially migrated out of Kazakhstan between the 1995 and

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[43] Gubhaju B.B., Choe M. K., Retherford R.D., and Thapa S. (1987). "Infant Mortality Trends and Differentials in Nepal." *Studies in Family Planning* 18(1): 22-31.

[44]Agadjanian V. and Qian Z. (1997) "Ethnocultural Identity and Induced Abortion in Kazakhstan." *Studies in Family Planning*, 28:4, 317-329.

1999 surveys, especially women of different ethnicities, and thus may have given birth in 1992 but are no longer living in Kazakhstan (and thus have no probability of being included in the survey) by 1995 or 1999.

The largely insignificant t-statistics indicate that combining data from the two surveys should not lead to significant bias resulting from the differential recall period, only perhaps differential migration. Table 2 also includes means of key variables for the full 1995 and 1999 surveys (1992-95 and 1992-98, not just the years of overlap) as well as means for the full study sample (the appended 1995 and 1999 surveys, encompassing all infants born 1992-98 that appear in either survey).

Model

The goal of my empirical model is to use logistic regression to predict the odds of infant death. I run a linear, additive log-odds model, using as my dependent variable a dummy variable coded one if the infant died before or at 12 months and coded zero otherwise. In my model, I include 15 independent variables controlling for maternal characteristics, such as education, ethnicity, age, region and urbanicity of residence, economic status, as well as characteristics of the infant, including parity, gender, and how long ago a preceding birth took place. I have modeled the effect of time as a linear variable on a scale of zero to one, with zero representing the first year in the sample, 1992, and one representing the final year in the sample, 1998. In order to allow the coefficient effects to change over time, I have estimated two different models incorporating interaction effects between time and the other covariates. In the first, more parsimonious, model, I interact time only with the variables that I hypothesize contribute to changes in infant mortality over time: ethnicity, region of residence, maternal education, parity, preceding birth interval, and household economic status. As a control, I estimate a second model interacting time with all 15 independent variables.

I use dummy variables indicating whether the child is of Russian or 'other' ethnicity; Kazakh ethnicity is the reference group. Unfortunately, because the regional boundaries of administrative districts (oblasts) shifted over the time period studied, along with the definitions of the different 'regions' utilized by the survey administrators, the finest degree of regional variation that I can use are dummy variables indicating whether an infant was born in Almaty, the capital city until December 1998 and still the major urban area, where health care facilities and regional opportunity are concentrated, or in the south or west regions of the country, with infants born in the east/north/central regions of the country serving as the reference category. I have also included a dummy variable for whether an infant was born in a rural area.

I use a dummy variable indicating whether the preceding birth took place within the past two years, the length interval that Pebley and Millman,[45] among the others noted above, found was associated with a higher risk of poor birth outcomes. I use two dummy variables to control for parity; because of disagreement in the literature about the linear effect of more children, I have instead categorized parity, with the first child, children 2-6, and child 7 and higher being grouped together. These are not arbitrary cutoffs, but suggested by Millman and Cooksey as the groupings most relevant for predicting infant survival.[46] I use an indicator variable for being female, as well as an indicator variable for whether the infant is from a multiple birth. I model maternal age as a quadratic relationship, including both maternal age at birth (accurate to the month) as well as age squared.

Finally, I control for maternal education and household wealth by using a linear composite measure of household economic status (see below) and two dummy variables indicating whether mothers had some higher education or finished secondary school; the reference group is those who have not finished secondary school.

Household Wealth Index and Quintiles

As much information as the DHS includes, it does not include income and expenditure data, but does include information on asset ownership, housing quality, and sanitary conditions. In order to estimate the changes in economic status on the changes in infant survival, I have created an index using this information as an estimate of household

[45] Pebley A. and Millman S. (1986) "Birthspacing and Child Survival." *International Family Planning Prespectives*, 12:3, 71-79.

[46] Millman S. and Cooksey E. (1987) "Birthweight and the Effects of Birth Spacing and Breast Feeding on Infant Mortality." *Studies in Family Planning*, 18:4, 202-212.

wealth. The wealth index was derived using the methods of Filmer and Pritchett.[47] As argued by Filmer and Pritchett, although this method provides questionable data on current wealth, it is a good long-term approximation of household economic status, and relative differences between households. Furthermore, given the use of retrospective data on births, ultimately, a long-term measure of wealth is more appropriate here than information on current consumption.

The specific variables included in my wealth index are a series of dummy variables indicating whether a mother's household owns a telephone, radio, motorcycle, bicycle, car, refrigerator, and whether the home has running water, a flush toilet, and a non-dirt floor. In addition to representing a long-term measure of a household's wealth, some of these variables are also more proximal determinants of infant survival, such as access to running water, sanitation standards evident in the type of toilet facilities, and quality of food afforded by the ability to refrigerate items. Filmer and Pritchett, however, assert that these variables should be aggregated into a single index instead of being used individually in the regression, as these assets represent both direct and indirect effects. Their method allows for separation of variability of each asset relevant to household economic status from the variability resulting from the direct effect.

Filmer and Pritchett's index calculation method utilizes principle component analysis, which reduces a number of variables into a single index, detects structure in the relationship between variables, and utilizes this structure in determining household wealth. In order to test the veracity of the aggregation process, I tested whether the distribution of ownership of individual assets used to make the wealth index roughly correlated with the distribution of the aggregate index. I created quintiles of the wealth index using a cumulative distribution function (cdf) of the wealth index, then dividing this cdf into five equal parts. I then checked whether the percentage owning each asset increased from lowest to highest quintile; this result was the case for each component of the index except for motorcycle ownership, although in this case owning a motorcycle might be an inferior good relative to owning a car for transport, and thus ownership would be more correlated with the lower wealth quintiles (see Appendix A).

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[47] Filmer D. and Pritchett L. (2001) "Estimating Wealth Effects Without Expenditure Data – or Tears." *Demography*, 38:1, 115-132.

I also calculated the Cronbach alpha on the set of included variables, which measures how well a set of variables measures a single, underlying, unidimensional characteristic, in this case household wealth. The Cronbach alpha can also be thought of as a measure of inter-relatedness and internal consistency among the variables used for the aggregate index; when the correlation between the variables is high and positive, the Cronbach alpha will be larger. A Cronbach alpha of .7 is generally considered a 'good' score.[48] My Cronbach alpha value is .6614 (.68 when I leave out the dummy variable for motorcycle ownership, which is not as well correlated with higher quintiles as the other dummy variables).

It must be noted, however, that although my wealth index seems an internally consistent measure of household wealth, there is likely still some bias in estimating household wealth at the time of birth using current data. Generally speaking, assets accumulate as parents age, and thus an index constructed using current data will overestimate wealth from previous time periods, as assets will be reported in the survey that were not present in a household at the time of birth. In Kazakhstan, however, current data might underestimate household wealth from earlier time periods, as one method of economic survival widely employed by the Kazakh population in later years has been to sell off household assets, which might have been around at the time of birth but are gone by the time of survey. This bias must be considered when interpreting results, although, as Agadjanian and Qian[49] note, economic mobility in Kazakhstan is low, so infants born to poor families likely still belong to poor families, relatively speaking, at the time of survey.

Interpretation of Time Interaction Effects and Decomposition

Because I am interested in the factors that lead to the change in infant mortality rates over time, my main interest is in the interaction terms between time and the other

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[48] Peterson R. A. (1994) "A Meta-Analysis of Cronbach's Coefficient Alpha." *Journal of Consumer Research* 21 381-391.

[49] Agadjanian V. and Qian Z. (1997) "Ethnocultural Identity and Induced Abortion in Kazakhstan." *Studies in Family Planning*, 28:4, 317-329.

covariates. As Ai and Norton[50] point out, interaction terms in nonlinear models cannot be interpreted directly, as in a linear probability model, as the percentage point change in probability of infant death attributable to the interacting covariate associated with a one-unit change in time. Instead, the interaction terms must be incorporated into the link function (in this case, the logit function) in order to determine the change in predicted probability.

Furthermore, my main research question focuses not only on the factors that contribute to the change in infant mortality over time, but also on decomposing changes in the probability of infant mortality between 1992 and 1998 into component parts: changes due to composition effects and changes due to coefficient effects. This decomposition will allow me to determine whether the risk to various groups is changing over time, while the birth population structure stays relatively stable (a coefficient effect); whether the risk of infant death associated with various characteristics stays the same while the demographics of the birth pool change (a compositional effect); or both. Decomposing changes in the probability of infant mortality in this manner is crucial in determining why infant mortality rates are increasing, since this method allows me to isolate more precisely the often conflicting effects of rapid population changes concurrent with social and economic changes.

The decomposition method for linear models was pioneered by Oaxaca[51] and Blinder[52] as a means to explain differences in the dependent variable across two different groups. In linear probability models, the change in predicted probability, ΔP, equals $\Delta X\beta_1 + X_2\Delta\beta$, where ΔX is equal to the difference in the mean values of variable X between the first group and the second, and $\Delta\beta$ is equivalent to the coefficient on the interaction term between the X covariate of interest and a dummy variable indicting

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[50] Ai C. and Norton E. (2003) "Interaction Terms in Logit and Probit Models." *Economics Letters* 80, 123-129.

[51] Oaxaca R. (1973) "Male-Female Wage Differentials in Urban Labor Markets." *International Economic Review*, 14, 693-709.

[52] Blinder A. S. (1973) "Wage Discrimination: Reduced Form and Structural Variables." *Journal of Human Resources*, 8, 436-455.

which group a subject belongs in. This method can be applied to all the covariates and coefficients of interest.

In order to adapt the Oaxaca-Blinder decomposition method to my nonlinear model, I utilize the methods of Fairlie[53] and Yun,[54] who use a first-order Taylor series expansion to linearize the logit model around the β_{ij}'s (where the subscript i denotes the year, either 1992 or 1998[55], and the subscript j denotes the covariate that the coefficient corresponds to). Initially, we can calculate the predicted probability for each year (1992 and 1998), using the k covariates in the model, as:

$$P_{92} = p_{92} + \Sigma^{k}_{i=1} p_{92}(1-p_{92})\beta_{92}X(bar)_{92} + \varepsilon_{92}$$

and $$P_{98} = p_{98} + \Sigma^{k}_{i=1} p_{98}(1-p_{98})\beta_{98}X(bar)_{98} + \varepsilon_{98}$$

where I have linearized the equation by using the initial predicted probability for each year, p_{92} and $p_{98,}$ which I found using the logit regression results. We can then calculate the various component of the decomposition, $\Delta\beta$, ΔX, and ΔP.

1. $\Delta\beta$ for each covariate j is:

$$\Delta\beta_j = [(p_{98})(1-p_{98})\beta_{98j}] - [(p_{92})(1-p_{92})\beta_{92j}]$$

2. ΔX for each covariate j is:

$$\Delta X_j = X(bar)98_j - X(bar)92_j$$

3. The change in predicted probability, ΔP, between years 1992 and 1998 attributable to each covariate j is:

$$\Delta P_j = [(p_{98})(1-p_{98})\beta_{98j}X(bar)_{98j}] - [(p_{92})(1-p_{92})\beta_{92j}X(bar)_{92j}]$$

In order to determine the significance of the ΔX portion of the decomposition, I calculated a t-statistic on the difference between $(Xbar)_{92}$ and $(Xbar)_{98}$. To determine the significance of the $\Delta\beta$ portion, for now, I use the t-statistic on the interaction terms

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[53] Fairlie R. (2003) "An Extension of the Blinder-Oaxaca Decomposition Technique to Logit and Probit Models." Discussion Paper #873, Economic Growth Center, Yale University.

[54] Yun M. (2003) "Decomposing Differences in the First Moment." Discussion Paper #877, Institute for the Study of Labor, University of Bonn.

[55] I am using 1992 and 1998 in the decomposition equation in order to determine how population composition and coefficients have changed from 1992 to 1998. I confirmed that the effect of time was roughly linear using a modified Hosmer-Lemeshow test.

between the time covariate and other covariates (such as those for ethnicity, education, etc.) in the logit results.

Data Problems and Potential Biases

Chidambaram et. al. [56] established, through their analyses of infant mortality data from multiple countries, that in retrospective birth history surveys, such as the DHS, infants who have died are under-reported relative to those who have survived, due both to problems recalling children who have died, and due to psychological repressing of painful outcomes. Additionally, the old Soviet classifications of 'live birth' and 'stillborn' differed from the internationally-recognized WHO standard, in that the WHO definition defines as a live birth any infant that is expelled from the mother that subsequently breathes, has a heartbeat, a pulsation of the umbilical cord, or otherwise shows a sign of life. Conversely, the Soviet system defined a live birth only an infant that breathed; other infants were classified as stillbirths and thus omitted from infant mortality statistics.[57] Because women in Kazakhstan are likely more familiar with the old Soviet classifications, some births were likely classified by the survey respondents as stillbirths, rather than as live births that subsequently died. There is no way to know whether women have mis-reported differentially, which would bias the coefficients by biasing selection into the 'dead' sample, but there is no reason to believe that they would have done so based on the alternate definitions of live birth. Under-reporting dead infants due to recall or psychological bias would likely underestimate the effects of the covariates on infant death. Furthermore, as previously mentioned, bias could arise due to differential recall bias from compiling data for 1992-1994 from two different surveys conducted at different times. If this bias exists, it would bias the coefficients on the covariates. Fortunately, Table 2 indicates that there is little evidence for this differential recall bias.

[56] Chidambaram V.C., McDonald J.W., and Bracher M.D. (1985) "Infant and Child Mortality in the Developing World: Information from the World Fertility Survey*." International Family Planning Perspectives*, 11: 1, 17-25.

[57] Anderson B. A. and Silver B. D. (1986) "Infant Mortality in the Soviet Union: Regional Differences and Measurement Issues." *Population and Development Review*, 12:4, 705-738.

The time variable passes the Hosmer-Lemeshow test[58], which indicates whether a linear representation of a variable accurately represents the data. Therefore, modeling time linearly, rather than incorporating year effects with dummy variables, should not lead to bias of the coefficients on either the time variable or on the interaction effects. Forcing the other covariates to have linear effects, however, as the decomposition setup does, may bias the effects of these covariates. In order to check the linearity of these variables over time, I have regressed year dummy variables on each of the dummy covariates; the coefficients for these regressions are presented in Appendix B. Although the coefficients on the year dummies indicate that the trends for each covariate are not themselves perfectly linear, each of the individual dummy variables also pass the Hosmer-Lemeshow test of linearity, so bias from this source should not be much of a concern.

Results:

The first component of the decomposition analysis involves determining the compositional changes in the birth population between 1992 and 1998. The means of the independent variables in 1992 and 1998, along with the change between the two years, are presented in Table 3. In general, the population composition has not shifted significantly in terms of any explanatory variable, except for the ethnic composition of the population and a small (but significant) overall change in wealth. While the proportion of total births that are of Russian ethnicity has increased by almost ten percentage points between 1992 and 1998 (although this change is not quite significant), the proportion of the population that is of 'other' ethnicity has declined by almost ten percentage points; this change is significant.

The logit coefficients are presented in Table 4. In the initial regression, the results confirmed my prior beliefs regarding determinants of infant death. Infants born in Almaty had a survival advantage over those born in the eastern region, while infants born in the southern and western regions face a survival disadvantage relative to the eastern region, although only the disadvantage of being from the southern region is significant. Russian and other non-Kazakh ethnicities had an initially better chance of infant survival,

[58] Hosmer D. and Lemeshow S. (1989) *Applied Logisitic Regression*. New York: John Wiley.

although this initial advantage is only significant for those of ‘other’ ethnicity. Better economic status yielded a significant initial survival advantage. Maternal higher education yielded a survival advantage, as expected, but mothers completing secondary education, relative to those having less than a complete secondary education, actually is disadvantageous, although neither of the maternal educational effects are significant. As expected, females had significantly better survival rates than males, twins were at a significant survival disadvantage, and having children quickly after a previous birth proved significantly risky. The only surprising results in the regression was that children of very high parity (child 7 – 11) survived with significantly higher probability, although the number of children in this category is fairly low, and that being from a rural area actually conveyed a significant survival advantage. The coefficients and their significance do not change much once the other control time-covariate interactions are added (Model 2 in Table 4), indicating that the more parsimonious model (only including time-covariate interaction terms for key variables) is preferable.

To examine changes in the coefficient effects over time, I linearized the beta coefficients using the Taylor series expansion method described in the methods section. The main-effect coefficient (which can be found in Table 4) on each covariate serves as the non-linearized β_{92}; in order to linearize this coefficient, I use the predicted probability of infant death for all individuals born in 1992 to obtain $p_{92}(1-p_{92})\beta_{92}$, presented in Table 5. Next, I find the non-linearized β_{98} by adding the main-effect coefficient to the coefficient from the interaction effect between time and that covariate. I linearize this coefficient by using the predicted probability of infant death for all individuals born in 1998 to obtain $p_{98}(1-p_{98})\beta_{98}$, presented in Table 5, along with the difference between the two linearized coefficients, $\Delta\beta$.

Examining the beta shifts, the majority of the shifts in coefficient effects are not significant. The ones that are, however, indicate significant changes in the relative risk of infant mortality based on ethnicity. In 1992, non-ethnically Kazakh infants have a survival advantage relative to those of native Kazakh ethnicity, but this survival advantage erods over time. The risk of infant mortality for Russian infants increases by 9.7 percentage points, while the risk of infant mortality for infants of non-Kazakh, non-Russian ethnicity increases by 12.2 percentage points. These increases shrink to 7.3

percentage points (for Russians) and 11.2 percentage points (for other ethnic groups) when the other time interaction controls are added, but are still significant. These changes are of significant magnitude that, by 1998, non-ethnically Kazakh infants now have a net survival disadvantage relative to ethnically Kazakh infants.

When considering the changing coefficient effects of two control variables, female and twin, the risk of infant mortality for infants from multiple birth increases by 21.6 percentage points between 1992-98, significant at the 5% level, and the risk of infant death to female infants is also significantly increasing over time, by 3.3 percentage points. While infants from multiple births had a survival disadvantage to start with in 1992, relative to singlet infants, this survival disadvantage has increased over time. Females, conversely, had a survival advantage relative to male infants in 1992, and still have a net survival advantage in 1998, but the net survival advantage has decreased over time.

Turning to the decomposition of changes in predicted probability, ΔP, in Table 6, the increase in the probability of infant death over time for Russian and other non-Kazakh ethnicities, as well as the positive shift in the coefficient on the ethnic variables, indicate that the probability of infant death is increasing for non-ethnically Kazakh infants; the increase in overall births to Russian women over time also likely contributes to the overall population increase in infant mortality, especially because the initial survival advantage for Russian infants is eroding over time. There is a significant decrease in births to women of non-Russian, non-Kazakh ethnicity, although this decline does not lead to an overall negative contribution to overall probability of infant mortality; it is overridden by the significant increase in coefficient effects for those of these ethnicities. These results persist when control time interaction effects are added.

In terms of regional composition of births, there is a decline in births from Almaty, as well as the southern and western regions, but none of these compositional shifts are significant. The rural/urban mix of the birth pool also does not change significantly between 1992-98. The relative advantage of being born in Almaty, present in 1992, is increasing over time, while the traditional disadvantage of being born in the southern region of the country is declining, although not significantly so. The risk of death to infants born in the western region of the country, as well as in rural areas, is increasing, although neither of these effects are significant either. Similarly, the Almaty and southern region dummy covariates make a negative contribution to the overall change in probability of infant death, while the west and rural covariates make a positive contribution.

There is a decline in the number of women giving birth shortly following previous births, while the coefficient effect of having an infant within two years of previously giving birth is increasing, leading to a small decrease in the probability of infant death, although neither the coefficient effect nor the composition effect is significant. Although the percentage of infants who represent high-parity births (child number 7-11) is declining, the coefficient effect on probability of infant death is increasing, resulting in an overall positive contribution of this variable to the overall probability of infant death, although once again, neither effect is significant.

In terms of education, births to both higher and secondary-school educated women are declining, although these declines are not significant. The risk of death for infants born to women who have attained higher education is increasing over time, while the risk to infants born to secondary-schooled women is declining, although these changes are not significant.

While there is a significant shift in the mean value of the wealth score over time, this variable makes no net contribution to the overall change in the probability of infant death, indicating that changes in household wealth, at least as measured by this long-term aggregate measure, are not contributing to changes in infant mortality.

When control variables for the time interaction effects between the twin, female, and mother’s age covariates are added, the effects of the other variables do not change much. Interestingly, the risk of infant death among females and multiple births is significantly increasing, and both make a positive contribution to the overall probability of death.

Discussion:

In seeking to explain the contributions of different covariates to the overall increase in infant mortality between 1992 and 1998, changes in the ethnic composition of the population, as well as increased risk to infants of non-Kazakh ethnicity, yields the largest, most consistent, positive contribution to the increased infant mortality rates. Rather than a biological difference, this increased risk to Russian and other non-Kazakh ethnic populations is most likely attributable to factors such as differential out-migration of better-educated, better-employed members of these subpopulations and a shift in the balance of political power and social control to ethnically Kazakh citizens. As Buckley[59] notes, ethnically Kazakh citizens were long discriminated against in the Soviet period, receiving poorer housing, lower wages, and poorer employment prospects. She finds that with the shift in balance of power, the long-term resentment of the natives Kazakhs, especially towards ethnic Russians, could result in revenge discrimination that takes many forms, such as denying pension and wage payments, redistribution of housing, and preference for native Kazakhs in the more lucrative private sector jobs. Brubaker[60] also indicates that although, remarkably, ethnic tensions in Central Asia between indigenous populations and ethnic Russians have not yet led to violence, the specter of inter-ethnic conflict permeates everyday life.

Although the presence and effects of these types of discrimination and stress are hard to measure, it provides a plausible explanation for increased risk to non-Kazakh infants. As Marmot and Wilkinson[61] have noted, an inferior place in the social hierarchy can affect health status above and beyond the direct effects of fewer economic resources through increased stress. In Kazakhstan, the increased stress, for women of Russian ethnicity in particular, of living as a former ‘oppressor’ among the formerly ‘oppressed,’ could result in worsening birth outcomes, particularly as ethnic tensions rise. Lauderdale[62] has noted a similar, significant increase in infant mortality among Arab-American women in the United States that she attributes to increased psychological stress from increasing ethnic tensions and discrimination directed at Arab-Americans following

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[59] Buckley C. “Rural/Urban Differentials in Demographic Process: The Central Asian States.” *Population Research and Policy Review*, 17, 71-89.

[60] Brubaker R. (1998) “Migrations of Ethnic Unmixing in the ‘New Europe.’” *International Migration Review*, 32:4, 1047-1065.

[61] Marmot M. and Wilkinson R. G. (2001). “Psychosocial and Material Pathways in the Relation Between Income and Health.” *British Medical Journal*, 322, 1233-36.

[62] Lauderdale D. (2004) “Ethnicity-Related Stress and Low Birth Weight: Arab-American Women Before and After September 11th.” Demography Workshop Presentation Paper, 2.12.04, University of Chicago.

the September 11th, 2001, attacks. Conversely, the alleviation of institutionalized discrimination for native Kazakhs following independence from the Soviet Union could account for the relative improvement in native Kazakh birth outcomes.

An alternate explanation for the increased risk to non-Kazakh infants could be explained by the effects of differential selection into childbirth and differential migration of Russian and other non-Kazakh women. The non-native Kazakh minorities that emigrated during the 1990s were better educated and better off economically; although these factors are controlled for in the regression, those who remained behind might be less suited for raising infants in ways above and beyond education and income, such as having a higher likelihood of being an alcoholic, worsening maternal nutritional status, or a smaller extended family to draw upon for support and material assistance (a common survival mechanism amongst native Kazakhs throughout the economically tumultuous 1990s). Unfortunately, the survey data does not allow me to explore this potentiality. The role of the ethnic variables, therefore, is to alleviate some of this omitted variable bias, without offering a more specific explanation for the ethnic differences.

The shifts in regional and urban/rural patterns can be explained by a number of factors. Firstly, since independence, government spending on health care infrastructure and delivery has fallen, from 3.5% of GDP in 1990 to 2% in 1994 to 1.5% in 1998 (GDP itself is also falling over this time period); the funds that have been invested have not been invested equitably across the country but have instead been concentrated in urban areas.[63] The increase in risk to rural infants, and the continual decreasing risk to infants born in Almaty, could be attributed to differential depreciation of Soviet-era health infrastructure due to lack of investment outside of the main city. Secondly, the economic effects of the transition could result in different health risks for infants in different regions in ways that are not discernable from the long-term household welfare measure, such as public infrastructure deterioration, high levels of local unemployment, or shortages of food and medical supplies, especially drugs that treat infectious diseases

[63] McKee M., Healy J. and Falkingham J. (2002) Health Care in Central Asia. London: Open University Press.

common among infants. Although the regional compositional and coefficient effects are not yet significant, they could become so if differential deterioration or some other local factor increases infant deaths asymmetrically throughout the country in the future.

Although the lack of a significant change of the effect of household economic status over time is somewhat surprising, given the extent of economic changes throughout the 1990s, this lack of finding could be explained by the limitations of the wealthscore variable. The wealth score constructed using the Pritchett and Filmer method is, at best, a rough estimation of household wealth in general and is specific to the time of survey, rather than the time of birth. A better temporal measure of household wealth would yield more variation, as well as more accurate estimates of the effect of wealth on infant survival. Furthermore, there is very little community data available in the DHS survey about availability of health clinics, hospitals, and other medical resources; medical care cannot be purchased, no matter household wealth, if it is not available in the community. Better information on community resources, rather than just the individual's utilization of resources, would yield a more complete picture of how access to health care and infrastructure deterioration is affecting infant mortality. Examining differential mortality rates based on year of birth is a very rough proxy for these deteriorating social conditions. As others[64,65] have found that there is relatively little effect of household-level wealth on infant mortality, but rather focus on the link between infant mortality and community- and regional-level resources, these insignificant changes over time are consistent with these arguments, especially if local infrastructure is not changing much between 1992 and 1998. Furthermore, if the effects of economic changes are largely consistent across ethnic groups, the ethnicity variables may already be controlling for a significant degree of the effects of economic changes on infant mortality.

As the results show, neither the composition of births, in terms of maternal education, nor the effect of this education, has had a significant effect on changes in the infant mortality rate between 1992 and 1998. It could be that there will be an effect of

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[64] Farah A. A. and Preston S. H. (1982) "Child Mortality Differentials in Sudan." Population and Development Review 8(2): 365-383.

[65] Frenzen P. D. and Hogan D. P. (1982) "The Impact of Class, Education, and Health Care on Infant Mortality in a Developing Society: The Case of Rural Thailand." *Demography* 19(3): 391-408.

diminished quality and prevalence of state-sponsored education as time progresses, but that measuring changes in educational quality and effects 6-7 years post-independence does not allow for enough time to elapse to allow the educational system to deteriorate to the point where it begins to make a difference to infant mortality, since the women having children between 1992 and 1998 were largely educated under the Soviet system. Alternately, the efficacy of maternal education itself has been debated: Desai and Alva[66] have found that maternal education has, at most, a weak direct effect, and instead is better thought of as a proxy for household wealth. If this scenario is true, perhaps I am over-specifying the model by controlling for both household wealth and maternal educational status. In either instance, changes in maternal education do not yet play a significant role in the changes in infant mortality, although they could in the future as today's schoolgirls grow up to become tomorrow's mothers.

The significant decreased survival chances of infants from multiple births might point to either a household's or the health system's diminished capacity for caring for infants from high-risk pregnancies and with low birth weights. Finally, the significant increase of risk to female infants is somewhat worrisome, perhaps caused by increased gender discrimination in tough economic circumstance, or as a result of increased Islamization of the ethnic Kazakh population.

Future Directions

Although the consequences of economic transition, in terms of infant mortality, seem broadly to have been borne more by Russians, other non-Kazakh ethnic groups, those living in rural areas and those living outside of Almaty, the specific effects of transition on household economic status, and how these changes affect health status, including infant mortality, are still not well-documented. Although the Filmer/Pritchett wealth index controls somewhat for long-term economic status, since it is measured at the time of survey, rather than at the time at birth, it does not change over time for different households, and does not account for increased income inequality over time, and thus overestimates the relative effects of wealth in earlier years. Using World Bank Living

[66] Desai S. and Alva S. (1998) "Maternal Education and Child Health: Is There a Strong Causal Relationship?" *Demography*, 35:1, 71-81.

Standards survey data, I could get a much more accurate picture of how household wealth has changed over time, and how these changes affect infant mortality. I would also be interested to see how changes in household welfare affect access to health care, as the 'shadow price' of 'free' care increases with increasing expectations of informal payments to medical personnel over time.

Secondly, using time covariates and regional variables as proxies for deteriorating infrastructure is not as satisfactory a tool as direct measures of the availability and quality of health services that affect infant mortality. Although the prenatal care data present in the DHS are of low quality, being non-random and incomplete, and thus was not used for this study, I could use the DHS and Living Standards Survey data to look at other measures of primary care access and delivery that affect infant and child health, such as whether the percentage of children being fully vaccinated has fallen over time.

Finally, although I have answered questions about why infant mortality has increased in Kazakhstan, I still do not understand why infant mortality rates have risen in Kazakhstan while they have improved in other former Soviet countries in Central Asia, such as Kyrgyzstan. Further research can focus on factors that contributed to the decrease in infant mortality in these countries in order to analyze the reasons behind the different experiences of Kazakhstan and the rest of Central Asia.

Conclusion

Russian and non-Russian, non-Kazakh infants had a net survival advantage relative to ethnically Kazakh infants in 1992, an advantage that probably resulted from preferential treatment during the Soviet period in terms of education, housing, economic opportunities, and the indirect effects of social status on health that result from being in the socially and politically dominant group. The increasing probability of infant death between 1992 and 1998 to these non-ethnically Kazakh infants was of sufficient magnitude to yield a net survival disadvantage by 1998. This increase in infant mortality among non-Kazakh ethnic groups is perhaps only one of the health consequences of the sudden and swift change in the social, political, and economic status of non-Kazakh ethnic groups following independence from the Soviet Union, changes that exacerbated existing ethnic tensions and competition.

Table 1: Recent Mortality and Economic Conditions in the 15 Countries of the Former Soviet Union

Country	Infant Mortality 1980 [1,2]	Infant Mortality 2001 [1,2]	Life Expectancy 1980 [2,3]	Life Expectancy 2001 [2,3]	Average Annual % Growth GDP, 1990-2002 [4]
Eastern Europe:					
Belorus	21	17	71	68	-0.1
Estonia	17	11	69	71	1.0
Latvia	20	17	69	70	-1.0
Lithuania	20	8	71	73	-0.9
Moldova	41	27	66	67	-7.1
Russia	22	18	67	66	-2.7
Ukraine	22	17	69	68	-6.6
Caucasus:					
Armenia	48	31	73	74	0.4
Azerbaijan	76	77	68	65	1.2
Georgia	35	24	71	73	-4.3
Central Asia:					
Kazakhstan	50	81	67	63	-1.6
Kyrgyzstan	90	52	65	66	-2.2
Tajikistan	N/A	91	66	67	-6.8
Turkmenistan	67	69	64	65	-1.0
Uzbekistan	47	52	67	67	0.8

1. Per 1,000 live births
2. Data source: http://devdata.worldbank.org/hnpstats/files/Tab2_19.xls
3. Life expectancy at birth
4. Data source: http://www.worldbank.org/data/wdi2004/tables/table4-1.pdf

Table 2: Summary Statistics Associated With Combining Two Surveys[1]

variable	1995 survey mean[2]	1999 survey mean[2]	t-statistic[3]	significant?[4]	1999 survey mean[5]	total sample mean
Death Rate						
Infantdeath	0.049	0.060	0.57		0.056	0.055
Ethnicity						
Kazakh(ref)	0.673	0.649	-0.51		0.664	0.666
Russian	0.194	0.213	0.50		0.220	0.214
Other	0.134	0.139	0.14		0.117	0.120
Region						
East/North(ref)	0.340	0.480	3.53	***	0.476	0.446
Almaty	0.092	0.092	0.00		0.077	0.080
South	0.368	0.248	-3.28	***	0.261	0.284
West	0.200	0.180	-0.65		0.187	0.190
Urban(ref)	0.395	0.513	2.96	***	0.489	0.469
Rural	0.605	0.487	-2.96	***	0.511	0.531
Previous Birth Interval						
Prebirthint > 2 years(ref)	0.770	0.756	-0.41		0.791	0.786
Prebirthint < 2 year	0.230	0.244	0.42		0.209	0.214
Education						
Incomplete Secondary(ref)	0.182	0.103	-2.54	***	0.115	0.130
Higher	0.184	0.202	0.50		0.191	0.189
Completed Seconday	0.633	0.695	1.50		0.695	0.681
Economic Status						
Wealthscore	0.061	0.001	-0.75		-0.041	-0.019
N	1048	1321			2503	3551

1. data are weighted using probab ility sampling weights
2. mean for years of overlap in 1999 survey, 1992-1994
3. t-statistic on difference between two means
4. significance: + indicates $p < .1$, * $p < .05$, ** $p < .01$, *** $p <.001$
5. mean for total 1999 survey

Table 3: Shifts in Birth Population Composition, 1992-1998

Variable	1992 mean	1998 mean	ΔX	t-statistic[1]	significance[2]
Ethnicity					
Kazakh	0.633	0.621	-0.012	-0.07	
Russian	0.210	0.309	0.099	1.47	
Other	0.157	0.063	-0.094	-2.80	**
Region					
Almaty	0.126	0.081	-0.045	-1.17	
South	0.255	0.241	-0.014	-0.23	
West	0.210	0.190	-0.020	-0.35	
Rural	0.519	0.568	0.049	0.71	
Birth Characteristics					
Previous birth < 2 years	0.205	0.162	-0.043	-1.00	
Parity = 2-6	0.582	0.613	0.031	0.47	
Parity = 7-11	0.012	0.005	-0.008	-1.27	
Multiple birth	0.043	0.012	-0.031	-1.03	
Female	0.509	0.522	0.013	0.19	
Maternal Characteristics					
Mother's age	25.778	26.207	0.429	0.60	
Mother's age ^ 2	692.235	719.817	27.582	0.66	
Higher education	0.204	0.170	-0.034	-0.68	
Completed secondary	0.701	0.691	-0.010	-0.17	
Wealthscore	0.110	-0.167	-0.278	-1.90	+

1. t-statistic on the difference between two means
2. significance: + indicates p < .1, * p < .05, ** p < .01, *** p <.001

Table 4: Logit Coefficients and Standard Errors

	Without Controls		With Controls[3]	
Variable	Coefficient	Standard Error	Coefficient	Standard Error
Time	-1.32	1.09	0.31	8.10
Ethnicity				
Russian	-0.78	0.49	-0.49	0.43
Russian * time	1.86*	0.75	1.26+	0.70
Other	-1.40*	0.57	-1.19*	0.56
Other * time	2.35*	1.13	2.01+	1.14
Region				
Almaty	-0.14	0.70	-0.37	0.67
Almaty * time	-1.86	1.58	-1.36	1.55
South	.67+	0.36	0.60	0.36
South * time	-0.16	0.62	-0.09	0.62
West	0.10	0.42	0.02	0.41
West * time	0.64	0.74	0.79	0.76
Rural	-.77+	0.46	-.86+	0.46
Rural * time	1.02	0.90	1.27	0.91
Birth Characteristics				
Previous birth < 2 years	1.06**	0.38	1.03**	0.40
Previous birth < 2 years * time	0.00	0.76	0.02	0.81
Parity = 2-6	-1.04*	0.50	-1.02+	0.55
Parity = 2-6 * time	0.38	0.74	0.30	0.90
Parity = 7-11	-2.55+	1.54	-2.59	1.64
Parity = 7-11 * time	3.92	2.85	3.81	3.09
Multiple birth	2.64***	0.43	1.07	0.85
Multiple birth * time			3.19*	1.56
Female	-1.04***	0.21	-1.35***	0.31
Female * time			.77+	0.46
Maternal Characteristics				
Mother's age	0.38	0.26	0.45	0.35
Mother's age * time			-0.13	0.57
Mother's age ^2	-0.01	0.00	-0.01	0.01
Mother's age ^2 * time			0.00	0.01
Higher education	-0.21	0.56	-0.15	0.50
Higher education * time	0.41	1.05	0.23	0.98
Completed secondary	0.35	0.45	0.25	0.44
Completed secondary * time	-0.55	0.92	-0.39	0.91
Wealthscore	-0.58*	0.23	-.63**	0.23
Wealthscore * time	0.26	0.45	0.35	0.47

1. n=3,551
2. significance: + indicates p < .1, * p < .05, ** p < .01, *** p <.001
3. Includes time-covariate interaction terms for maternal age, female births, and multiple births

Table 5: Shifts in Coefficient Effects, 1992-1998

	Without Controls				With Controls[4]			
Variable	β_{92}[1]	β_{98}[1]	$\Delta\beta$[2]	significant?[3]	β_{92}[1]	β_{98}[1]	$\Delta\beta$[2]	significant?[3]
Ethnicity								
Russian	-0.040	0.057	0.097	*	-0.025	0.048	0.073	+
Other	-0.072	0.050	0.122	*	-0.061	0.051	0.112	+
Region								
Almaty	-0.007	-0.105	-0.097		-0.019	-0.108	-0.089	
South	0.035	0.027	-0.008		0.029	0.030	0.001	
West	0.005	0.039	0.034		0.005	0.055	0.050	
Rural	-0.040	0.013	0.053		-0.044	0.025	0.069	
Birth Characteristics								
Previous birth < 2 years	0.055	0.055	0.001		0.053	0.065	0.012	
Parity = 2-6	-0.054	-0.034	0.019		-0.052	-0.045	0.007	
Parity = 7-11	-0.131	0.072	0.203		-0.132	0.075	0.208	
Multiple birth					0.055	0.271	0.216	*
Female					-0.069	-0.036	0.033	+
Maternal Characteristics								
Mother's age					0.023	0.062	0.039	
Mother's age ^ 2					0.000	0.000	0.000	
Higher education	-0.011	0.011	0.022		-0.008	0.004	0.012	
Completed Secondary	0.018	-0.011	-0.029		0.013	-0.009	-0.021	
Wealthscore	-0.030	0.017	0.047		-0.032	-0.017	0.015	

1. linearized coefficient
2. $\Delta\beta$ standard errors are logit regression estimates of standard errors on interaction terms
3. significance: + indicates $p < .1$, * $p < .05$, ** $p < .01$, *** $p <.001$
4. Includes controls for female and multiple births, as well as mother's age and mother's age squared

Table 6: Decomposition Effects

Variable	ΔX	$\Delta\beta^3$	ΔP	ΔX	$\Delta\beta^3$	ΔP
Ethnicity						
Russian	0.099	0.097*	0.026	0.099	0.073+	0.020
Other	-0.094**	0.122*	0.014	-0.094**	0.112+	0.013
Region						
Almaty	-0.045	-0.097	-0.008	-0.045	-0.089	-0.006
South	-0.014	-0.008	-0.002	-0.014	0.001	0.000
West	-0.020	0.034	0.006	-0.020	0.050	0.009
Rural	0.049	0.053	0.028	0.049	0.069	0.037
Birth Characteristics						
Previous birth < 2 years	-0.043	0.001	-0.002	-0.043	0.012	0.000
Parity = 2-6	0.031	0.019	0.010	0.031	0.007	0.003
Parity = 7-11	-0.008	0.203	0.002	-0.008	0.208	0.002
Multiple births				-0.031	0.216*	0.001
Female				0.013	0.033+	0.016
Maternal Characteristics						
Mother's age				0.429	0.039	1.043
Mother's age ^ 2				27.582	0.000	0.017
Higher education	-0.034	0.022	0.004	-0.034	0.012	0.002
Completed secondary	-0.010	-0.029	-0.020	-0.010	-0.021	-0.015
Wealthscore	-0.278+	0.047	0.000	-0.278+	0.015	0.006

1. n=3551
2. significance: + indicates $p < .1$, * $p < .05$, ** $p < .01$, *** $p < .001$
3. $\Delta\beta$ standard errors are logit regression estimates of standard errors on interaction terms

Appendix A: Validity of Factor Component Analysis in Construction of Wealth Quintile

% in each quintile that report owning good

	Telephone	Toilet	Bicycle	Television	Water	Good Floor	Electricity	Radio	Refrigerator	Motorcycle
Poorest	0.6	0.8	6.6	60.2	1.9	89.5	84.6	16.8	8.7	12.3
Quintile 2	5.8	0.5	16.8	95.9	3.0	95.9	99.5	22.6	85.1	14.9
Quintile 3	34.2	6.9	24.9	96.7	34.3	99.5	99.4	49.2	88.1	12.3
Quintile 4	21.3	74.2	13.6	97.2	95.7	96.8	100.0	49.3	91.8	5.9
Wealthiest	95.4	100.0	23.1	100.0	100.0	99.2	100.0	65.0	100.0	4.7
p-value:	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000

1. p-value from chi-square analyses testing difference in ownership across quintiles

Appendix B: Testing the Linearity Assumption by Modeling Time Using Dummy Variables

Year	Russian	Other	Almaty	South	West	Rural	Higher Education	Completed Secondary Education	Short Previous Birth Interval	Female	Multiple Births	Parity = 2-6	Parity = 7/11
1993	-0.014	-0.034	-0.053	0.045	-0.020	0.030	-0.011	-0.078	0.050	0.020	-0.034	-0.001	0.027
1994	-0.002	-0.028	-0.054	0.104	-0.047	0.031	-0.021	-0.020	0.051	-0.027	-0.026	0.016	0.005
1995	0.014	-0.026	-0.032	0.029	-0.043	-0.010	0.024	-0.057	0.014	-0.003	-0.021	-0.007	0.003
1996	-0.040	-0.065	-0.070	0.027	0.005	0.031	-0.028	-0.020	0.022	-0.069	-0.010	0.064	-0.006
1997	0.015	-0.039	-0.079	0.002	-0.022	-0.038	-0.044	0.059	-0.065	-0.012	-0.034	-0.009	0.005
1998	0.099	-0.094	-0.045	-0.014	-0.020	0.049	-0.034	-0.011	-0.043	0.013	-0.031	0.031	-0.008
F-statistic p-value	0.483	0.051	0.164	0.243	0.830	0.808	0.848	0.135	0.030	0.798	0.614	0.884	0.257

n = 3,551