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Chronic Poverty in Kazakhstan

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Abstract

Given the lack of longitudinal data for Central Asia, research on poverty has largely ignored the time dimension. This study uses panel data constructed from the rotating cross-sectional Kazakhstan Household Budget Survey for the 2001-2009 period. The panel data provides an opportunity to measure chronic poverty levels and poverty transitions for the first time in Kazakhstan. We find that, despite the rapid and substantial reduction in poverty in the country since the turn of the century, and depending on the measure of chronic poverty employed, as much as a quarter of the population has experienced persistent poverty. However, the majority of chronically poor experience interrupted poverty spells. We apply the multiple-spell hazard model analysis to shed light on factors that impact on poverty exit and re-entry. The results of these estimates confirm that families with children under age six are experiencing higher probability of entry into poverty and lower probability of exit from poverty. Policy interventions are needed to improve the situation by providing an affordable state child care system in Kazakhstan.

Keywords: chronic poverty, longitudinal data, multiple-spell hazard model, Kazakhstan
JEL: I32 C23 C41

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1. Introduction

Poverty reduction is one of the major economic challenges in developing countries. Recent evidence illustrates a constant reduction in the incidence of absolute poverty in the developing world (Chen & Ravallion, 2012). The overall percentage of the population living below $1.25 a day in 2008 was 22 percent in developing countries, compared to 52 percent in 1981. Moreover, 1.3 billion people in 2008 lived below $1.25 a day, compared to 1.9 billion in 1981. The level of relative poverty also decreased from 63 percent in 1981 to 47 percent in 2008. However, the number of relatively poor increased by about 360 million over the 1981-2008 period (Chen & Ravallion, 2012). This evidence highlights the importance of poverty measurement in determining actual poverty levels. In particular, understanding the factors that lead to changes in poverty over time is essential for the effectiveness of poverty reduction policies. This understanding requires estimating alternative measures of poverty and their persistence over time. According to a study by the Chronic Poverty Research Center (2008),

> "Over the last five years, in an era of unprecedented global wealth creation, the number of people living in chronic poverty, understood as persistent poverty for over N years, has increased. Between 320 and 443 million people are now trapped in poverty that lasts for many years, often for their entire lifetime" (p. i).

Therefore, in addition to poverty reduction, chronic poverty reduction has also become an important issue for policy makers. As Hulme and Shepherd (2003) write:

"In a country where poverty is largely a transient phenomenon, with '*the poor*' at any particular time having a high probability of improving their position, then policies should focus predominantly on social safety nets that help people to manage their present deprivation, rapidly return to a non-poor status and reduce vulnerability. By contrast, in a country where a significant proportion of the poor are chronically poor, then policies to redistribute assets, direct investment toward basic physical infrastructure, reduce social exclusion (from employment, markets and public institutions) and provide long-term social security will be necessary" (p. 404).

From the policy perspective, therefore, it is important not only to identify the poor at one period of time, but also the chronically poor, i.e. those who have experienced poverty for extended periods or possibly all of their lives, and also the transient poor.

The study of conventional poverty measures alone, taken at a point in time, will not provide accurate information regarding the poverty level and number of poor (Biewen, 2006). Firstly, long periods of low income lead to greater welfare losses and damaging effects on self-perception and self-confidence for the affected people than a one-off poverty spell. Secondly, the presence of long poverty spells means that the burden of poverty is unequally distributed among the population, because it is only a small number of individuals who endure total poverty compared to a larger number of individuals who endure only short poverty spells. Thirdly, those in long-term poverty will consume a large part of the resources devoted to anti-poverty policies.

Until the late 1980s, the key techniques using the role of time in the study of poverty were developed in the form of poverty trends, seasonality, the timing of experiences, and historical accounts of poverty. Poverty trends usually compared headcounts of poverty across a population at two (or more) different times. However, contrasting poverty trends in this way does not describe whether individuals or households are persistently poor or if they typically move into and/or out of poverty over time (see Hulme & Shepherd, 2003; Carter & Barrett, 2006; Addison, Hulme & Kanbur, 2009). For example, Roberts (2001) studies chronic poverty in South Africa between 1993 and 1998 and finds that poverty incidence is 33.7 percent for 1993 and 41.5 percent for 1998. However, a richer picture is obtained by departing from conventional static poverty analysis and observing the dynamics of poverty, i.e. the real situation for individual households over time. In Robert's (2001) study, almost 22 percent of the sampled households qualified as poor in both periods and 11 percent were poor in 1993 but non-poor in 1998, meaning that they appear to have escaped poverty over this period; 19 percent were non-poor in 1993 but had fallen into poverty by 1998, and 47 percent were poor in neither period. Thus, in studying the dynamics of poverty, observing the same households over several years has gained importance in recent years.

Given the lack of panel data for transition countries, and specifically for Central Asia, very little analysis has been conducted on poverty dynamics and chronic poverty in this region. The few studies that have addressed the dynamics of poverty (Bierbaum & Gassmann, 2012; Commander, Tolstopiatenko, & Yemtsov, 1999) do so without considering the estimation of multiple-spell hazard models that focus on poverty transitions. A recent study on chronic and transient poverty in Russia reveals that the severity of poverty is found to occur largely from transient rather than chronic spells of economic hardship (Mills & Mykerezi, 2009). Further, Mills and Mykerezi (2009) find that a low level of post-secondary education is one of the major correlates of chronic poverty. A study by Okrasa (1999) finds that selling electronic durables is one of the coping strategies for households experiencing long-term poverty, and that savings accounts have a negative effect on being chronically poor in Poland. Bierbaum and Gassmann (2012) identify the main determinants of chronic poverty in Kyrgyzstan, such as location, low levels of human capital, and poor employment opportunities.

With respect to Kazakhstan, the majority of poverty studies have long been static, based on cross-sectional data (Anderson & Pomfret, 2002; Pomfret, 2006; Rhoe, Babu & Reidhead, 2008). Conventional static analysis in the literature mainly focuses on the poverty headcount ratio, indicating the proportion of the population that has fallen below a given income or expenditure threshold at a particular point in time. It compares the poverty trends at different times and defines the determinants of static poverty (Anderson & Pomfret, 2002; Pomfret, 2006; Rhoe, Babu & Reidhead, 2008). According to the World Bank, the poverty by headcount ratio in Kazakhstan by national standards has fallen since 2001, with 46.7 percent in 2001, dropping to 8.2 percent in 2009, and further reducing to 2.7 percent in 2015 (World Bank, 2016). Reviewing the literature on static poverty in Kazakhstan suggests that the following are significant correlates of poverty: geographic location (Anderson & Pomfret, 2002; Pomfret, 2006), composition of household (Jha & Dang, 2009), and education of the head of household (Anderson & Pomfret, 2002; Pomfret, 2006). One attempt was made to assess the vulnerability of households to future poverty based on cross-sectional data (Jha & Dang, 2009). Due to a lack of panel data, however, to our knowledge there are no studies on chronic poverty and poverty dynamics in Kazakhstan. Moreover, static poverty measures cannot provide a complete picture of poverty dynamics, e.g. whether individuals or households that were poor in previous years are the same this year, or whether new households or individuals have fallen into poverty. Static poverty measures also cannot reveal how long individuals remain in poverty. Thus, the aim of this study is to test the following hypotheses:

1. What is the chronic poverty level in Kazakhstan?

2. Do the chronically poor have more interrupted poverty spells[3]?

3.What are the triggers for poverty exit and re-entry?

In this study, we use panel data[4] for the period of 2001-2009 and, based on various measurements of chronic poverty, we find that almost a quarter of the Kazakh population is chronically poor. However, the majority of individuals are transient poor, due to transitions from poverty spells to non-poverty spells during the nine-year period. The following determinants positively influence poverty exit: head of household has a university degree, location in Almaty and Astana, and having some assets, such as a car or *dacha*[5]

The remainder of this paper is organised as follows. Section 2 defines chronic poverty measures applied in this study. Section 3 describes the multiple-spell hazard model. Section 4 presents the data description. Section 5 examines the empirical estimations of chronic poverty measures. Section 6 presents empirical results from the estimation of multiple-spell hazard regressions. Section 7 concludes.

2. Chronic Poverty Measures

Over the last two decades, research devoted to the measurement of poverty dynamics has been growing (Addison, Hulme, & Kanbur, 2009; Bossert, Chakravarty, & D'Ambrosio, 2012; Calvo & Dercon, 2007; Dercon & Porter, 2011; Foster, 2009; Hoy, Thompson, & Zheng, 2011; Hoy & Zheng, 2008; and Jalan & Ravallion, 2000, among others).

The definition of chronic poverty mainly depends on which of the different approaches is used to measure chronic poverty (Chakravarty, 2009; Hulme & Shepherd, 2003; McCulloch & Calandrino, 2003), such as the spells approach (based on duration of poverty spells), the components approach (which considers income or consumption shortfall), and vulnerability (probability of deficient future consumption) (Barrientos, 2007). Following Bane and Ellwood (1986), a poverty spell is identified as the set of consecutive periods during which income falls

[3] Spell of poverty means one or more consecutive periods below the poverty line.

[4] The analysis makes use of annual cross-section data extracted from the KHBS 2001–2009 and on a panel dataset constructed from these data. The KHBS is a cross-section survey, but the sampling frame remained largely unchanged during the period and a share of households was interviewed throughout the period. The panel dataset was constructed by matching observations across the annual data files (Kudebayeva, 2015).

[5] A small house other than the main dwelling, located near a city. Mainly used for rest and growing vegetables.

below the poverty line. In the analysis of chronic poverty, the important difference between the components and spells approaches is that the components approach assumes a compensation between low and high income periods and then the identification of who is poor during each period of time becomes unessential, while in the spells approach no compensation is allowed and one needs to identify who is poor in each period. According to McKay and Lawson (2003), the spells approach is more related to the concept of chronic poverty as persistent poverty, whereas the classification of the chronically poor in the components approach is prejudiced by the depth of poverty, although both offer important tools.

Foster (2009) develops the spells approach and presents a family of chronic poverty measures based on static Foster, Greer, and Thorbecke measures. Gradin et al. (2012) introduced a new family of aggregate intertemporal poverty indexes which take into consideration the incidence and the intensity of poverty dimensions in a dynamic framework. This measurement also includes sensitivity to the chain of poverty durations. However, the main contribution of Gradin et al. (2012) is the insertion of inequality poverty experiences, which cover the sensitivity of the measure to the equalisation of individual per-period poverty gaps. Moreover, this measure of chronic poverty is a generalised case of the chronic poverty measures of Foster (2009) and Bossert et al. (2010). The following expression is the aggregate intertemporal chronic poverty index proposed by Gradin et al. (2012):

$$P(Y,z) = \begin{cases} \frac{1}{N}\sum_{i=1}^{N}\left(\frac{1}{T}\sum_{t=1}^{T}\left(\frac{z-y_i^t}{z}\right)^{\gamma}\left(\frac{s_i^t}{T}\right)^{\beta}\right)^{\alpha} & if\ \alpha > 0 \\ \frac{q}{N} & if\ \ \alpha = 0 \end{cases}, \qquad (1)$$

where α is sensitivity of the aggregate intertemporal poverty measure to inequality among the intertemporal poor individuals; β is sensitivity of the individual intertemporal poverty indices to spells duration; γ is sensitivity of the individual intertemporal poverty measure to inequality among the intertemporal poor individuals; s_i^t is the duration in poverty of the individual *i*. One of the advantages of the framework given above is that it involves Foster's (2009) index when β=0 and α=1, which means normalised poverty gaps are not weighted by the poverty spell duration and the aggregate intertemporal poverty measure is simply the average of individual intertemporal indicators over the population, and hence insensitive to the indicators' distribution. When β=1 and α=1, we obtain Bossert et al.'s (2010) measure, which means that normalised poverty gaps are weighted proportionally to spell duration and the aggregation over the population average.

The components approach was proposed by Jalan and Ravallion (hereinafter "J-R") (2000). For identification of the chronically poor, an average or stable component of income is defined, and those individuals who lie below an appropriate poverty line are counted as chronically poor:

$$P_{JRi} = \frac{1}{N}\sum_{i=1}^{N}\left[1-\left(\frac{\bar{y}_i}{z}\right)^{\alpha}\right] \; where \; \bar{y}_i = \frac{1}{T}\sum_{i=1}^{T} y_i^t \,. \qquad (2)$$

However, it permits compensation when oscillating between prosperous and difficult periods and thus an individual with several severe poverty spells would be excluded from chronic poverty in the case of only one extremely good welfare measurement being received in the considered period.

The components approach for the measurement of chronic and transient poverty has been enlarged by using the equally distributed equivalent (EDE) poverty gaps and has developed a statistical correction for the biases that take place when the number of panel waves available is small (Duclos et al., 2010).[6] The approaches of J-R (2000) and Duclos et al. (2010) are distinct, both conceptually and methodologically. Firstly, J-R's (2000) measure of chronic poverty was connected to the average of welfare; however, Duclos et al.'s (2010) approach is linked to the average of ill-fare.[7] Secondly, the approaches of J-R (2000) and Duclos et al. (2010)s are distinct in their connection of out-of-poverty spells to chronic poverty. For instance, in J-R's (2000) approach someone in severe poverty over *t-1* periods may still be considered to be not chronically poor if his income over t^{th} period is sufficiently large to make the average income over *t* periods above the poverty line. Moreover, the EDE approach has a shortcoming: someone out of poverty during *t-1* periods will still contribute to total chronic poverty even if his average income over time exceeds the poverty threshold. Duclos et al. (2010) employed J-R's (2000) approach and EDE poverty gaps to estimate chronic and transient poverty for rural China, based on the Chinese Research Centre on Rural Economy (RCRE) surveys of 1986-2002. The main findings indicate a comparative scale of chronic and transient poverty, and therefore the policies that should be implemented depend on the measurement issues; e.g. the proportion of transient poor is 23 percent of the total poor by the EDE approach, which

[6] The level of individual ill-fare which, if assigned equally to all individuals and in all periods, would generate the same poverty measure as that produced by the distribution of normalised poverty gaps.

[7] Normalised poverty gaps are commonly used as measures of individual metrics of ill-fare.

substantially differs from that of the J-R (2000) estimation of transient poverty, which equals 75 percent.

Gradin et al.’s (2012) chronic poverty measure is a generalised case of Foster’s (2009) and Bossert et al.’s (2010) chronic poverty measures by the spells approach and is therefore implemented in our study. We apply the Jalan-Ravallion (2000) and Duclos et al. (2010) measures for the components approach.

3. Multiple-Spell Hazard Model

The amount of time households are in poverty within a period does not tell us how long poverty spells are or how long people stay out of poverty if they have been poor. A person who is poor in the first year of a nine-year period may have just entered poverty or may have already been poor for five years. Similarly, a person poor in the last year of the period may remain poor for another year or two, or leave poverty altogether. The differences in the time spent poor, or in the time spent non-poor, reflect differences in individuals’ poverty exit and entry rates. Therefore, a duration analysis based on hazard regressions is an important tool for the in-depth investigation of movements in and out of poverty.

Reducing poverty effectively requires context-specific knowledge about reasons for moving into and out of poverty. Since each type of poverty is likely to require a different policy response, there is an increasing demand to understand better the correlates of these transitions. This section constructs a model for evaluating the correlates of poverty exit and re-entry and observes the length of poverty spells for individuals who become poor. For those individuals who leave poverty, the model also examines the length of time between finishing one poverty-spell and starting another. Thus, the length of time at risk is a fundamental part of the analysis. In this section, the duration of dependence and individual heterogeneity in poverty exit and re-entry hazard rates are both considered using multivariate regression modelling approaches.

Initially, duration models by Bane and Ellwood (1986), and extended by Stevens (1999) and others, were applied for the estimation of poverty dynamics. Bane and Ellwood (1986) developed their technique because of their disappointment with the existing methodologies for investigating poverty dynamics, particularly variance components models and tabulations of the number of occurrences people were poor over a fixed time period. In variance components models, ‘*income*’ is the dependent variable and dynamics are introduced through the residual error structure, with assumptions about permanent and transient components. In studying poverty dynamics, Bane and Ellwood (1986) state that variance components models are better matched to modelling the dynamics of an individual's labour earnings than the dynamics of household income. Bane and Ellwood (1986) recognised that one could apply variance component modelling approaches with needs-adjusted family income as the dependent variable. Although they focused on single spells, the use of hazard models allowed them to gain new insights into the longitudinal structure of poverty. A significant result was that although most of those who fall into poverty will remain there for only a short time, the majority of the currently poor is, to a large extent, composed of individuals who are in the midst of a long spell.

Bane and Ellwood (1986) were criticised by Stevens (1999), who pointed out that focusing on single spells systematically underestimates poverty persistence, as the possibility of re-entry is ignored. A number of papers have pointed to the limitations of the implementation of single-spell methodologies in fitting the observed pattern of poverty persistence and have proposed a new approach that allows for the consideration of multiple poverty and non-poverty spells simultaneously. These approaches were first recommended by Stevens (1999) and then used by Devicienti (2002), Hansen and Wahlberg (2009), and Biewen (2006). In addition, Arranz and Canto (2012) studied the effect of spell recurrence on poverty dynamics.

Research on poverty persistence for developed countries has mainly implemented long panels and applied different types of hazard models, covariance structure models, and Markovian models (Cappellari & Jenkins, 2002; Canto, 2002; Biewen, 2006; Hansen & Wahlberg, 2009; Devicienti, 2011; Jenkins, 2011; Arranz & Canto, 2012). Among developing countries, few researchers have applied such methodologies (for China the discrete-time multi-spell duration model has been applied by You (2011) and Imai & You (2013); and for Ethiopia by Bigsten and Shimeles (2008)).

In order to shed light on how poverty duration and number of years spent poor are related to individual characteristics, we model poverty exit and re-entry rates conditionally on a number of characteristics and the duration already spent in the corresponding state. As all correlates determining exit and re-entry clearly cannot be included, it is essential to control for unobserved heterogeneity. It is also important to take into consideration a measurement error by conducting a sensitivity analysis by application of adjusted poverty lines. We conduct a stochastic dominance analysis that illustrates a decline in poverty levels during 2001-2009 in Kazakhstan regardless of the poverty lines used.[8] The results of the estimations for adjusted poverty lines indicate that evaluations of poverty exits and re-entries do not change substantially for small variations of poverty lines. Therefore, we avoid the measurement error in transitions of individuals from non-poverty to poverty status. Moreover, the difficulties of left-censoring data can be solved by excluding the first spell of poverty in the first wave of panel data.[9] Our panel data set contains nine waves, the first (non-) poverty spell starts from the second wave and the maximum duration is seven. The few studies that have attempted to include left-censored spells in the analysis (Stevens, 1999; Devicienti, 2011) have concluded that the left-censored biases are likely to be of second order in relatively long panels. However, right-censored data should also be taken into consideration because the individual could be at risk of exiting poverty; even in the last year of the panel he/she could be in poverty because there is no information about the state of the individual after this spell.[10] However, the empirical results illustrate that including the right-censored spells does not create problems in estimations (Devicienti, 2011).

Our approach is based on a joint discrete-time hazard model controlling for unobserved heterogeneity to estimate the determinants of transition into and out of poverty (Devicienti, 2002; Devicienti, 2011; Jenkins, 2011). This method allows us to implement the evaluations to forecast how long in total a poverty entrant will spend being poor, taking into consideration not only the initial poverty spell, but also possible later spells.

We consider two states between which individuals have moved: poverty and non-poverty. Survivor function$S(t_j)$ gives the probability that a person survives longer than some specified time t. The *survivor function* for poverty exit is defined for discrete time as follows:

$$S(t_j) = \prod_{(j|t_j \leq t)} \left(1 - \frac{d_j}{n_j}\right), \qquad (3)$$

where $t_1, t_2, \ldots t_j, \ldots t_k$ is survival time with equal intervals for simplicity, d_j is individuals or households end their poverty spells at t_j, n_j is individuals or households stay poor in at least j

[8] Stochastic dominance analysis is given in Appendix 2.

[9] Left-censoring means that the failure event (poverty exit or entry) occurs prior to the subject‘s entering the study.
[10] Right-censoring means that one runs a study for a pre-specified length of time, and by the end of that time, the failure event has not yet occurred for some subjects.

waves and are at '*risk*' of moving out of poverty at t_{j+1}. This is the probability of '*surviving*' past time t_j.

Hazard rates (hazard functions) of ending poverty at t_j:[11]

$$h(t_j) = \begin{cases} 1 - S(t_j) & if \quad j = 1 \\ \frac{S(t_j) - S(t_{j-1})}{S(t_j)} & if \quad j > 1 \end{cases} . \qquad (4)$$

For example, if we consider the waves as years, then h(7) gives the instantaneous potential per unit time for the event (in our case poverty exit) to occur, given that the individual has survived up to 7 periods (staying poor). The hazard rates of poverty re-entry are estimated similarly.

A multi-spell hazard model can be defined as below. Each individual could experience either a single type of spell (in-poverty or out-of-poverty) or both. For the latter case, an individual could have repeated spells of poverty and/or repeated spells of non-poverty. The probability that an individual *i* leaves the state at duration *d*, given that she/he has survived in the state to *d-1*, is assumed to be a standard *logit* hazard function (Devicienti, 2011; Jenkins, 2011):

$$E\left(\frac{d}{\theta_i^P}\right) = \frac{exp[\theta_i^P + \alpha_d^P + X_{it}^{P\prime}\beta^P]}{1 + exp[\theta_i^P + \alpha_d^P + X_{it}^{P\prime}\beta^P]}, \qquad (5)$$

where X_{it} is a set of covariates that differ across individuals and, potentially, also over calendar time, represented by *t*. These covariates can be fixed or time variant. The dependence of the hazard upon duration in the spell *d* is explicitly highlighted by (5), while dependence upon X_{it} and through X_{it} upon calendar time t is left implicit so as to simplify notation. Next, β is a vector of parameters to be estimated, and α_i^P represent interval-specific dummies aimed at capturing the shape of the baseline hazard function with fully flexible non-parametric specification.

The problem is unobserved heterogeneity, which can lead to underestimation of how hazard rates change with duration. Given that individuals differ in unobserved terms such as ability, effort, and tastes, and if these unobservables remain constant over the individual's lifetime, the assumption of uncorrelated spells might be inappropriate. Thus, people who have had long periods in poverty in the past might be more likely to have long spells again. Correspondingly, individuals that spent a long time out of poverty in the past might be less likely to experience

[11] The hazard function $h(t_j)$ gives the instantaneous potential per unit time for the event to occur, given that the individual has survived up to time t.

long spells of poverty in the future. In order to allow a correlation across a spell in the presence of unobserved heterogeneity, a joint maximum likelihood estimation of exit and re-entry rates is applied, in which hazard rates depend on spell-specific unobserved heterogeneity terms, and these terms are correlated across spells. In fact, the hazard rate in (5) is conditional on an individual-specific and state-specific effect θ_i^P, i.e. this random term is common across all poverty spells of individual *i*.

Correspondingly, for non-poverty spells a hazard of re-entry is specified:

$$R\left(\frac{d}{\theta_i^N}\right) = \frac{exp[\theta_i^N + \alpha_d^N + X_{it}^{N\prime}\beta^N]}{1+exp[\theta_i^N + \alpha_d^N + X_{it}^{N\prime}\beta^N]}, \qquad (6)$$

where θ_i^N is universal across all the non-poverty spells of person *i*, and *d* now shows duration in the present non-poverty spell. Duration dependence for out-of poverty spells is summarised by the interval-specific dummies α_d^N.

Allowing for temporal correlation across spells of the same type and correlation across spells of different types can be done by assuming that θ^P and θ^N are jointly distributed. For the model estimation, we use the Heckman and Singer (1984) estimator, where a joint distribution $G(\theta^P, \theta^N)$ is left unspecified, so as to minimise the misspecification biases, and is approximated by a bivariate discrete distribution with a number of support points to be determined by the data.

A potentially significant fact relates to the omission in the hazard-rate models of left-censored spells. If unobserved heterogeneity was not a determinant of the hazard rates in the factual model, then evaluation of exit and re-entry rates only from non-left censored spells would present reliable estimates of the parameters of covariates and duration terms. Conversely, when unobserved heterogeneity is a determinant of the true spell length, then the exclusion of a left-censored spell can result in potential biases, as one is likely to oversample individuals characterised by higher transition probabilities than the rest of the population. Stevens (1999) concluded that exclusion of left-censored spells is mainly without consequences in the results of regressions. In this context, selection into poverty or non-poverty at the beginning of the observation is not random but depends on individual characteristics, including the unobserved heterogeneity terms. This can be solved by adding a probability of being in poverty at the initial point observation (first non-left censored spell) of each individual as a function of the characteristics of unobserved heterogeneity term *q*, as follows:

$$P_{i0}(q) = \frac{exp\left[q+W_{i0}^{'}\delta\right]}{1+exp\left[q+W_{i0}^{'}\delta\right]}\,, \qquad (7)$$

where 0 is implemented to denote the time in which the first fresh spell started and W_{i0} is a vector of observables. This probability is then estimated with the in-poverty and out-of-poverty transition equations allowing for the unobservables to be correlated, by assuming that q, θ^P, and θ^N have a joint trivariate distribution.

The likelihood function for the extended model is given by the following expression:

$$L = \sum_{i=1}^{N} log \int_{H(\theta^P)} \int_{H(\theta^N)} \int_{H(q)} L_i(\theta^P, \theta^N, q) \cdot dG(\theta^P, \theta^N, q);$$

$$L_i(\theta^P, \theta^N, q) = P_{i0}(q)^{v_{i0}}(1 - P_{i0}(q))^{u_{i0}} \times \prod_{j=1}^{J_i-1} \left[\left(1 - E_{ij}(\theta^P)\right)^{1-y_i} E_{ij}(\theta^P)^{y_i}\right]^{v_i} \cdot$$

$$\left[\left(1 - R_{ij}\left(\theta^N\right)\right)^{1-y_i} \cdot R_{ij}\left(\theta^N\right)^{y_{it}}\right]^{u_i}, \quad (8)$$

where $L_i\left(\theta^P, \theta^N, q\right)$ denotes individual's *i* contribution to the likelihood function, conditional on the unobserved heterogeneity terms. H(q), H(θ^P), and H(θ^N) denote the support of *q*, $\theta^P and\ \theta^N$, respectively. J_i is the total number of interviews in which individual *i* has been observed in the panel, excluding years in left-censored spells. The subscript *j* indexes the sequence of poverty outcomes of each individual *i*, and *j=0* refers to the time in which his/her first non-left-censored spell started. The variable u_{ij} takes value 1 if the spell at *j* is a poverty spell, if not, it takes value 0. Similarly, v_{ij} is a dummy variable, which equals 1 if the spell at *j* is a non-poverty spell, otherwise it equals 0. The next variable is y_{ij} and is a dummy variable that is equal to 1 if the person exits from the current state in j, if not, it is equal to 0.

Based on Heckman and Singer (1984), the unobserved heterogeneity distribution $G(\theta^P, \theta^N, q)$ is left unspecified and is approximated by trivariate discrete distribution, with q, $\theta^P and\theta^N$ each assuming a finite number of support points. Finally, the estimates are obtained by maximizing the pseudo-likelihood function based on (8). Thus, the specified model helps to estimate the probabilities of exit and re-entry into poverty jointly, taking into consideration unobserved heterogeneity.

4. Data

The analysis in this paper relies on data from the Kazakhstan Household Budget Survey (KHBS) from 2001 to 2009 provided by the Agency of Statistics of the Republic of Kazakhstan (ASRK). The KHBS is a nationally-representative annual household survey that collects information on 12,000 households. The survey sample is representative down to the *oblast* (province) level, and it is stratified by rural and urban sectors and also by small, medium, and large cities. The questionnaires contain four modules: (i) daily expenditures on food and household necessities; (ii) quarterly expenditures for clothes, durables, utilities, education, healthcare, transportation, other expenditures and income of household members; (iii) housing conditions, livestock, equipment and machinery, education, and employment; and (iv) household composition and size. In 2002, 2003, and 2005, two supplementary modules surveyed health and education variables for household members.

The analysis below uses annual cross-sectional data extracted from the KHBS 2001-2009 and on a panel dataset constructed from these data. The panel dataset was constructed by matching observations across the annual data files (Kudebayeva, 2015). The KHBS is a cross-sectional survey (the survey also adopted a rotating sample, with 25 percent of households surveyed replaced every four quarters), but the sampling frame remained largely unchanged during the period, and a share of households was interviewed throughout the period. In total, 2,580 households were found to be present in all waves. Household and individual matching across the annual datasets was based on birth year, gender, and the first name of individuals in the household. Tests of robustness, representativeness, and attrition bias performed on the constructed panel dataset provide confidence regarding its properties. Appendix 1 provides more detailed information about the construction of the panel dataset.

Between 2001 and 2005, data from the full sample of households was collected quarterly. Beginning in 2006 until 2008, the survey methodology changed to surveying 3,000 households each quarter, with the information consolidated into annual datasets for the 12,000 households.

The poverty analysis will focus on per capita consumption expenditure as the welfare indicator. Asset holdings and borrowings usually make it possible to smooth consumption, with the implication that measures consumption is less volatile than income (Deaton, 1992). Focusing on consumption offers two advantages when analysing poverty dynamics. First, income-based measures may over-estimate variations in family economic well-being and the magnitude of poverty (Jorgenson, 1998). Second, expenditures appear to be less susceptible to systematic under-reporting than income, particularly among low-income families (Meyer & Sullivan, 2003). The focus on consumption expenditure better captures living standards among low-income groups. We focus on equivalised per capita consumption expenditures computed by dividing total household expenditures by the square root of the household size. Some researchers make a strong case for using adult equivalent expenditures to take account of household economies of scale and the different 'costs' of children (Deaton & Muellbauer, 1986; Deaton & Paxson, 1998; Lanjouw & Ravallion, 1995). Having explored this issue with the data, Kudebayeva (2015) found only marginal differences in poverty estimates using per capita household expenditure and alternative OECD and WHO equivalence scales.

The official poverty lines are set by tracking the value of a minimum subsistence consumption basket reflecting nutrition standards, as developed by the National Nutrition Institute. These nutrition standards also serve as the basis for cost of living calculations. Different baskets are constructed for the five regions, for nine age groups, and separately for females and males. This information is used to identify a mean national consumption basket. The cost of this consumption basket is calculated monthly, based on regional prices, separately for urban and rural areas. The costs of non-food goods and services are included as an adjustment to the food costs. Currently, non-food costs are estimated as an additional 40 percent of the food costs. Beginning in 2006, the Agency of Statistics applied a new methodology for the calculation of the subsistence minimum (SM) by expanding the range of goods included from 20 to 43 products, and setting a 2,175 kcal per day as the nutrition benchmark. The adjustment for non-food costs was raised from 30 percent to 40 percent. To enable comparison across regions and across years, gross per capita real consumption expenditures were adjusted with official regional poverty lines. The values of the minimum food basket were calculated by multiplying the norms of consumption for each food item and its average price at the regional (oblast) level in the middle of each month. The basic norms of food consumption did not vary among regions, and the difference in values of the subsistence minimum among regions reflected the difference in prices.

However, the stochastic dominance analysis shows a reduction in consumption poverty incidence regardless of which poverty lines and measures are used for the period 2001-2009 (see Appendix 2). Therefore, further estimates are based on 40 percent of equivalised per capita consumption expenditures taken as the relative poverty line.

5. Chronic Poverty Estimations

Table 1 illustrates the chronic and transient poverty measures of J-R (2000) and Duclos et al. (2010) for various values of α. In Table 1, the components approach, which defines the chronically poor as those individuals with average intertemporal equivalised consumption expenditures below the intertemporal poverty line (when α=0), shows the smaller share of transient poverty. This can be explained by the use of the relative poverty line as a poverty threshold. The reduction in chronic poverty measures due to the increase in α, illustrates less inequality among intertemporal poor individuals. The normalised poverty gaps are small for both poverty measures when α =1. The sensitivity of J-R's (2000) chronic poverty measure to the distribution of poverty gaps is low, whereas the sensitivity of Duclos et al.'s (2010) chronic poverty index to the equalised distribution of poverty gaps is larger. Moreover, the estimations from the Chinese Rural Household Survey yield larger transient poverty by J-R's (2000) approach than by Duclos et al.'s (2010) approach when α=2 (Duclos et al., 2010). However, Duclos et al. (2010) applied an absolute poverty line as a poverty threshold. Our estimations depict the same issue when transient poverty comprises about 63 percent of total poverty by J-R's (2000) approach and only 23 percent by Duclos et al.'s (2010) approach (when α=2). This result is explained by the fact that Duclos et al.'s (2010) measure assigns more weight to the poverty gap in each period and then aggregates it over the whole period of nine years for each individual, before then aggregating it over all individuals in the sample. However, J-R's (2000) measure weights the gap between average intertemporal consumption and the poverty line of each individual, and then aggregates it over all individuals in the sample. Thus, Duclos et al.'s measure (2010) indicates more inequality among the chronically poorest individuals.

The weakness of the component's approach in chronic poverty measurement (i.e. both J-R (2000) and Duclos et al. (2010)) is that an individual can experience a substantial decline in consumption in only one year out of nine, but can be counted as both poor and also as chronically poor in that year because his average intertemporal consumption is less than the intertemporal poverty threshold. However, the spells approach counts this individual as transient poor.

Table 1: Chronic and Transient Poverty by Components Approaches

Chronic Poverty Measures	Chronic Poverty	Transient Poverty	Total Poverty
J-R (2000), α=0	0.278	0.079	0.356
J-R (2000), α=1	0.045	0.036	0.080
J-R (2000), α=2	0.009	0.017	0.027
Duclos et al. (2010) α=1	0.080	0.000	0.080
Duclos et al. (2010) α=2	0.125	0.039	0.164

Source: Author's calculations based on KHBS 2001-2009.

The Table 2 illustrates Gradin et al.'s (2012) measure of chronic poverty, which is a more generalised version of the chronic poverty measures by the spells approach, i.e. Foster's (2009) and Bossert et al.'s (2010) poverty indexes.

Table 2: Chronic Poverty Measures by the Spells Approach

α	β=0			β=1		
	γ=0	γ =1	γ=2	γ=0	γ=1	γ=2
α=0	0.276	0.276	0.276	0.276	0.276	0.276
α=1	0.253	0.063	0.022	0.157	0.043	0.016
α=2	0.200	0.015	0.002	0.103	0.010	0.002

Source: Author's calculations based on KHBS 2001-2009.
Note: Gradin et al.'s (2012) chronic poverty measure, where α is sensitivity of aggregate intertemporal poverty measure to inequality among intertemporal poor individuals; β is sensitivity of individual intertemporal poverty indices to spells duration; γ is sensitivity of individual intertemporal poverty measure to inequality among intertemporal poor individuals. Gradin et al.'s (2012) measure yields Foster's (2007, 2009) chronic poverty measure when α=1 and β=0; when α=1 and β=1, it produces Bossert et al.'s (2010) chronic poverty index.

Table 2 reveals the sensitivity of intertemporal indices to variations in poverty gaps, and their intertemporal distribution for each individual, spell duration, and inequality in individual complete poverty practices over time. The analysis starts with the case in which $\beta=\gamma=0$ permits us to segregate the impact of changes in parameter α. The implication of progressively higher sensitivity to inequality of time spent in poverty across individuals (when $\alpha>0$) illustrates the decrease in chronic poverty. This means that the intertemporally poor are more equally distributed. Next, the analysis of the sensitivity of the aggregate intertemporal measure implies that larger weights on poverty spells of a long duration require the segregation of the effect on the choice for various values of β, when $\gamma=0$ and $\alpha=1$. The change in β from 0 to 1 shows a decline in chronic poverty measures of 38 percent. This fact confirms the larger experience of short-term periods of poverty among the intertemporally poor, because the penalisation of longer spells of poverty by higher weights caused the decline in indexes. There is further analysis on the effect of including sensitivity to inequality in the chronic poverty measure in a more comprehensive manner (when $\gamma=2$ and $\beta=1$), which takes into consideration poverty gaps and their intertemporal distribution for each individual along with poverty duration. The increase in α from 1 to 2 illustrates the larger decline in chronic poverty in percentage terms than when $\beta=\gamma=0$. The decrease is almost 88 percent. The results show all poverty-reducing features that are accumulated in the chronic poverty measure for Kazakhstan.

Thus, both estimates of chronic poverty by the components and spells approaches illustrate the robustness of the results. The percentage of chronically poor when a relative poverty line is applied is 27 percent. However, these measures of chronic poverty do not reflect transitions into and out of poverty.

6. Poverty Durations

This section analyses spells of poverty, durations of poverty, and poverty transitions. Table 3 below shows the duration of poverty for various categories of households.

Table 3: Proportions of times in poverty for individuals from different types of households

Time poor	Proportion for couples	Proportion for couple+adult+ children	Proportion for couples with children	Proportion of pensioner couples	Proportion of singles	Proportion of single with children
0	36.9	31.51	26.79	24.11	26.1	18.73
1	14.62	11.99	12.13	15.6	15.7	9.61
2	10.71	7.19	8.05	14.18	10.62	8.64
3	10.56	9.25	9.1	17.73	11.09	9.37
4	5.07	10.96	8.95	2.84	8.31	8.88
5	5.21	8.9	5.77	4.26	6.93	10.22
6	5.79	4.79	7.31	8.51	6.7	10.83
7	5.21	4.11	7.21	5.67	6	9
8	4.2	6.16	8.6	4.26	4.62	7.79
9	1.74	5.14	6.11	2.84	3.93	6.93

Source: Author's calculations based on KHBS 2001-2009.

Appendix 3 presents the results for all households and individuals with the application of absolute and relative poverty lines. It observes various household structures, such as couples without children, a couple plus one adult and children, couples with children, pensioner couples, singles, and singles with children. Non-poor individuals are mainly from households consisting of couples without children (i.e. 36.9 percent). The percentage of non-poor individuals from households with a couple with one adult and children is 31.51 percent, while for a single adult household with children, the percentage is only 18.73 percent.

With respect to persistently poor people, the proportion of poor individuals in the whole of the nine year period is one of the largest and is equal to 6.93 percent in households headed by a single parent with children, followed by individuals from households which include couples with children (6.11 percent); while for individuals from households consisting of couples without children, the percentage of always-poor is only 1.74 percent.

Transient poverty prevails among pensioner couples, for whom the percentage of poverty in periods of less than four years is one of the highest in comparison with other categories. For other categories of families, except for singles with children, the proportions of poor are higher for shorter periods of less than five years. As indicated in the last column of Table 3, only for singles with children is the percentage of poverty larger for longer periods (i.e. more than five years). Moreover, couples with children experience larger proportions of poverty for periods above five years in comparison with other categories of households. These results reveal the important issue of persistent child poverty in Kazakhstan and suggest that government policy should pay more attention to targeted social assistance programmes for poor families with children.

The data includes nine waves. We incorporate the right-censored data in the estimations and exclude the left-censored data from consideration. The exclusion of left-censored data means that the individuals who are poor and non-poor in all nine waves are excluded from our estimation. Therefore, now exits from the state of poverty can occur at any of the seven interviews following the one in which the individuals were found in poverty. First, we examine a wide outline of exits from and returns to poverty by using simple non-parametric estimates of the exit and re-entry rates, and look at how they differ with the duration of time people have been in and out of poverty. The exits rates relate to the group of individuals that are just falling into poverty and thus at risk of subsequent exit. The re-entry rates relate, as an alternative, to a group of individuals just starting a spell out of poverty, and thus are at risk of re-entering.

The estimated survivor and hazard functions in Table 4 indicate strong negative duration-dependence associated with the rates of poverty exit and re-entry. This implies a high probability for individuals to escape from poverty in the shorter term. This fact shows that, for a cohort of individuals just starting a poverty spell, the hazard of leaving after the first year is equal to 16.08 percent; after two years it is 8.1 percent, and drops further to 3.64 percent after four years. The probabilities of poverty exit then fall again after seven years, reaching 1.02 percent.

Table 4: Survivor and hazard function of spells in and out of poverty.

Time since the start of spell	Poverty exit				Poverty re-entry			
	Survivor function	(SE)	Hazard function	(SE)	Survivor function	(SE)	Hazard function	(SE)
1	0.8392	0.0040	0.1608	0.0044	0.8710	0.0035	0.1290	0.0038
2	0.7712	0.0047	0.0810	0.0036	0.8234	0.0041	0.0546	0.0027
3	0.7219	0.0052	0.0639	0.0036	0.7896	0.0045	0.0410	0.0026
4	0.6956	0.0055	0.0364	0.0033	0.7655	0.0048	0.0305	0.0026
5	0.6726	0.0058	0.0331	0.0036	0.7488	0.0051	0.0218	0.0026
6	0.6589	0.0061	0.0203	0.0033	0.7399	0.0052	0.0119	0.0022
7	0.6522	0.0064	0.0102	0.0030	0.7314	0.0055	0.0115	0.0025

Source: Author’s calculations based on KHBS 2001-2009.

The analysis of the survivor function for poverty exits illustrates that 83.92 percent of poverty entrants are still in the poverty pool after the first year; 77.12 percent are still in poverty after two years; 69.56 percent are poor after four years, after which the numbers drop further. After seven years, approximately 65.22 percent of the original pool of poverty entrants is still in poverty.

The hazard rates of re-entry are smaller than exit rates and indicate a significant risk for individuals out of poverty to fall back below the poverty threshold, particularly in the years just after an exit from poverty. Approximately 12.9 percent of the individuals ending a poverty spell will again have income below the poverty threshold after the first year; after two years the hazard of re-entry falls to 5.46 percent; and after four years the hazard of re-entry is only 3.05 percent (see Table 4). The survivor function for those who are out of a poverty spell indicates that almost 87.1 percent survive as non-poor after one year; 82.34 percent are non-poor after four years, and 76.55 percent are non-poor after seven years. The estimations illustrate that survival rates are higher for non-poor spells than for poverty spells.

Multivariate hazard regression models are a generalisation of life-table methods since they allow transition rate change, not only with time at risk, but also with personal characteristics. A special characteristic of the spell data in the given study is that spell lengths are observed only in units of a year; the data are grouped (‘*interval censored*’) rather than having exact dates for poverty-spell beginnings and endings. Thus, methods for discrete-time survival data are used rather than those for continuous time.

The data on spell lengths and censoring status summarise for each spell an '*event history*'; a sequence of years during which the individual was at risk of leaving poverty (in our case poverty exit is the event). Hence, for someone with a completed spell length of four years (i.e. the individual is not poor in the fifth year), there is a data sequence of four years with no exit event recorded for each of the first three years and one recorded for the final year. If, instead, the individual's spell is censored, the individual has been at risk of poverty exit for three years, but there is no exit event recorded for any of the years. Thus, the original data set is re-organised such that a person-year indicator of whether a transition occurred between that year and the next is embedded for the relevant individual.

The differences among individuals that are combined in hazard regression models mainly specify the differences in the structure of an individual's household and differences in measures of household labour market additions. For poverty transitions between some year *t-1* and *t,* the value of each explanatory variable used is the value in the base *t-1.*[12] The labour market characteristics are permitted to change by year within a spell. However, age variables are set to be equal to their values at the start of the spell.[13]

Household labour market variables are characterised by the employment status of the head of household. The inclusion of some individual specific variables, such as age and gender, does not show significant results. Therefore, we include dummy variables for the head of household, such as the individual is employed/unemployed, employed in the public sector, employed in the private sector, and self-employed. The age, gender, marital status, education level, and ethnicity of the head of household are also incorporated in the modelling. The demographic structure of the household is characterised by the quantity of adults, the elderly, and children under the age of six years. The variables that describe the demographic structure and the head of household's age and gender help to contrast the experience of single parents with married couples, large families with small families, and elderly people with younger people. The study

[12] According to Jenkins (2011: 299), '[*t]his is more satisfactory than using year t values because, in that case, there is a greater chance that the values are a consequence of the transition itself*'.

[13] Jenkins (2011: 299) argues that '[*t]his is done in order to avoid collinearities between age and duration dependence: spell length and age each increases by one year as time progresses*'.

of poverty duration suggests that individuals from single-parent households and couples with children have relatively long poverty spells.

The following significant assets of the household are also included as dummy variables: the household owns a *dacha* (a small house other than the main dwelling), owns a car, lives in a separate house or flat, and has access to water in the dwellings.

Thus, the estimation of the model is based on the pattern of poverty transitions for all individuals, although individuals vary in their characteristics. Some studies apply a sample of adults only, thus excluding children from the transition model (Biewen, 2006; Devicienti, 2011). The main argument for excluding children is that they are economically dependent on their parents, who can make the behavioural choice. However, as Jenkins (2011) points out, poverty transition models are descriptive rather than behavioural models, and therefore the above argument has become less significant. Further, the estimates of the model using only adult samples do not illustrate the substantive change as compared with the model that sample of all individuals, including children.

The data set is created as follows. An exit occurs in the next to the last wave in which the individual is poor (for entry, the same procedure is applied). However, the dummy variable for poverty exit allocates an exit to the same wave in which the individual was last in that state. Therefore, we do not need to create the lagged explanatory variables as we want to link the characteristics at *t-1* to exit in *t*. Due to the exclusion of left-censored observations, the individuals who are poor and non-poor in all nine waves are not observed in our estimations. The estimations of the joint model with unobserved heterogeneity did not fit well and created the problem of convergence. Therefore, we do not consider unobserved heterogeneity in our model. We also consider the '*initial conditions*' problem that appears when one permits unobservable differences in hazard rates. If the model is true, then an individual's poverty status when first observed is no longer random. For instance, individuals with a relatively high propensity to enter poverty have a higher likelihood of being found to be poor when they are first observed. Preliminary analysis for this paper followed Devicienti (2001) by modelling the probability that the first non-left-censored spell in the data is a poverty spell, estimating this probability jointly with the poverty exit and re-entry model specified in equations (5) to (6). The models did not fit well and the initial conditions parameters were statistically insignificant, as Devicienti (2001) also found. For the sake of brevity, the estimates are not provided here. Table 5 illustrates the results from the estimation of multivariate hazard rates of poverty exit and re-entry from joint multiple-spell regressions by using a *logit* model.

Table 5: Discrete time multiple-spell hazard models.

Variables	Poverty exit regressions		Poverty re-entry regressions	
	Coeff.	SE	Coeff.	SE
Duration dummies				
1	2.395*	0.092	2.624*	0.097
2	1.220*	0.096	1.374*	0.102
3	0.642*	0.101	0.891*	0.113
4	0.303**	0.124	0.321*	0.123
5	-0.052	0.138	0.138	0.148
6	-0.802*	0.183	-0.655*	0.203
7	-1.358*	0.300	-0.957*	0.237
Head of household:				
Age of head	-0.014*	0.001	-0.012*	0.002
Female head	-0.211*	0.050	-0.017*	0.053
Ethnicity is Kazakh	-0.083	0.061	-0.270*	0.065
Ethnicity is Russian	-0.170*	0.061	-0.038	0.066
(Omitted category –a head of the household is from an another ethnicity)				
Education of head:				
University	0.054	0.055	-0.392*	0.063
Vocational	-0.053	0.044	-0.179*	0.049
Not compl. second.	0.104	0.074	0.085	0.078
(Omitted category: head of the household has secondary education)				
Head is married	0.001	0.059	-0.238*	0.063
Head is widowed	-0.307*	0.071	-0.040	0.074
(Omitted category: head of the household is unmarried or divorced)				
Unemployed	0.067	0.079	0.136	0.086
Pensioner	0.108	0.083	0.247*	0.085
Public sector employee	0.012	0.048	-0.189*	0.052
Private sector employee	-0.153*	0.050	-0.068	0.055
Self-employed	-0.189*	0.051	-0.194*	0.055
(Omitted category: other category for the head of household, e.g. student, housewife, disabled or other				
Household demographics:				
Quantity of adults	-0.185*	0.018	-0.168*	0.021

Quantity of children 0-5 years aged	-0.172*	0.049	0.137**	0.053
Quantity of elderly	-0.215*	0.051	-0.370*	0.055
(Omitted category: school-age children)				
Assets of the household:				
Household has a separate house or flat	-0.457*	0.060	-0.348*	0.065
Household has a *dacha*	0.130	0.085	-0.169***	0.101
Household has water in the home	-0.211*	0.079	-0.007	0.084
Household has a car	0.038	0.061	-0.053	0.072
Location:				
Central	0.068	0.065	-0.612*	0.072
West	-0.104	0.069	-0.328*	0.070
North	-0.167**	0.074	-0.379*	0.075
East	-0.168**	0.078	-0.289*	0.084
Astana	0.062	0.224	-0.836*	0.268
Almaty	0.399*	0.135	-1.140*	0.181
(Omitted category is South)				
Urban	0.042	0.080	-0.144***	0.083
(Omitted category is rural)				
Log-likelihood	-7307.112			
Number of obs. (person-years)	17203			

Notes: Statistically significant at P<0.01, statistically significant at **P<0.05, statistically significant at ***P<0.1; SE are robust standard errors clustered by household.
Source: Author's calculations based on KHBS 2001-2009.

The results of the estimation of the multivariate multi-period joint model of the hazards of poverty exit and re-entry indicate that the negative poverty duration starts after four periods in poverty. The hazard rates of poverty re-entry become negative after five years in non-poverty. A one-year increase in the age of the head of household, other thing being equal, will reduce the hazard rate of poverty exit by 1.4 percent. Female headship compared to male headship will reduce the hazard rate of poverty exit by 0.8 times (exponent (-0.211) = 0.8). The head of household being Russian reduces the hazard rate of poverty exit by 0.84 times compared to other ethnicities; the other characteristics are identical. Only having a university degree positively effects the hazard of poverty exit. Widowed heads of the household decrease the hazard rate of poverty exit by 0.73 compared to single heads of the household, other thing being equal. Employment of the head of the household is not a significant factor for reducing the hazards of poverty exit, other things being equal. The larger the size of the household, the less the decrease in hazard rates of poverty exit, when other characteristics are equal. Living in a separate house or apartment also has a negative influence on the hazard rate of poverty exit because the majority of the households live in separate dwellings. Only households located in Almaty will increase the hazard rate of poverty exit for the individual.

The age of the head of the household, the head of the household being female, the head of the household being ethnic Kazakh, and the head of the household with vocational, university, or higher education negatively affect the hazard rate of poverty re-entry. The head of the household being married compared to being single, the quantity of adults and elderly, living in a separate dwelling, having a *dacha*, and living in locations except for rural areas and the south will also reduce the hazard rate of poverty re-entry. Only the head of the household being a pensioner, and having children under the age of six will increase the hazard rate of poverty re-entry.

The factor with significant positive influence on poverty exit is a location in Almaty. Many correlates of the model estimation have the same signs for the hazard rate of poverty exit and re-entries. This means that these factors are common for the transitory poor, who are moving in and out poverty in given periods of time. As defined previously, the existence of children under the age of six will increase the hazard rate of poverty re-entry.

7. Conclusion

By implementing the relative poverty lines and adult equivalence scales, we estimate and analyse the chronic poverty measures of Jalan-Ravallion (2000), Duclos et al. (2010), and Foster (2007,2009). We find that, despite the rapid and substantial reduction in poverty in Kazakhstan since the turn of the century, and depending on the measure of chronic poverty employed, as much as a quarter of the Kazakh population has experienced persistent poverty.

Moreover, the estimations of the intertemporal poverty measures of Gradin et al. (2012) assist in testing the sensitivity of individual intertemporal poverty indices to variability in per-period poverty and to spell duration. In addition, the measure of Gradin et al. (2012) incorporates the sensitivity inequality among intertemporally poor individuals. The estimation results indicate that the intertemporal poverty measure is less sensitive to all of the above issues, i.e. per period poverty and spell duration. Thus, the intertemporally poor are more equally distributed, which means that the shorter durations of poverty spells prevail and per-period poverty is less variable.

Our investigation of poverty duration among various household types indicates that the longest duration of poverty is experienced by single individuals with children and couples with children. The lowest duration of poverty is among adult couples without children and pensioner couples. Disabled people covered by the public health system and state social benefits are less vulnerable to poverty, and school-age children in the public education system are less likely to be chronically poor. However, the risk of poverty re-entry is considerable for individuals from households with children under the age of six, as parents have difficulty accessing the labour market due to a lack of public kindergartens and the necessity to care for children rather than earn money. Thus, with respect to policy implications, the construction of public kindergartens should be a priority for Kazakhstan.

In addition, we use multivariate hazard regression models to examine differences in individuals' experience of poverty over time. For individuals who enter poverty, the total time span in poverty consequently depends on both the chances of exit from poverty and the chances of re-entry into poverty. The results confirm the negative duration dependence of the hazard rates of exits from and re-entries into poverty. Factors that have a positive impact on the probability of poverty exit include location in Almaty, head of household with a university degree, and owning assets such as a car or *dacha*. Many correlates of the model estimation have the same signs for the hazard rate of poverty exit and re-entry. These factors are common for the transitory poor, who move in and out of poverty in a given period of time. This fact illustrates that the majority of the persistently poor, who were poor for more than a total of 5 years, experienced breaks between poverty spells. Thus, the majority experienced interrupted poverty spells during the whole period of the observation. Moreover, the existence of children under the age of six increases the hazard rate of poverty re-entry and decreases the probability of poverty exit.

This study of poverty transitions concludes that the majority of the population in Kazakhstan is transient poor or vulnerable to poverty. Hence, policies aimed at reducing vulnerability to poverty are required. In particular, an increase in the number of retraining centres is likely to result in increased pay as individuals obtain new professions. In addition, greater policy focus is needed on sectors of the economy with the lowest average wage, such as agriculture, education, and health care. Specifically, this problem may be solved by providing financial assistance for utility payments to workers in these sectors and to individuals with per capita income below the official poverty line, but not 40 percent of the official poverty threshold, which is currently applied to be eligible for targeted social assistance (TSA) in Kazakhstan.

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Appendix 1 - Panel dataset construction

The KHBS has a sample of 12,000 households, which are divided into four groups with one group visited each quarter. For the 2001-2005 period and for 2009, observations are matched on a quarterly basis and then added up to an annual dataset. The annual aggregate income and consumption expenditures of households were constructed based on data from one quarter for the period 2006-2008. The resulting number of households for each year is always less than 12,000 due to attrition between quarters and missing data on specific variables. The household attrition rate between quarters for the 2001-2005 period and for 2009 is less than 5 percent. From 2006 until 2008, the Statistical Office released an annual data set only, and therefore it is not possible to implement seasonal adjustments.

The attrition rates, reliability, and robustness at each stage of matching for various matching techniques were thoroughly evaluated. The total number of households remaining after matching all waves is 2,580 and the attrition rate is 70.82 percent. Attrition tests confirmed that high attrition rates do not bias the poverty measurement estimations. Table A1 shows descriptive statistics for cross-sections and panel data.

Table A1. Descriptive data of the KHBS 2001-2009

Variables	2001 Cross-section	2001 Panel	2009 Cross-section	2009 Panel
Poor by official poverty line	0.449	0.503	0.113	0.123
Poor by relative poverty line	0.290	0.291	0.301	0.269
Head of HH is male	0.533	0.562	0.440	0.443
Years of education of head of HH	10.940	10.950	12.140	11.670
Ethnicity of head of HH:				
1-Kazakh	0.461	0.507	0.526	0.503
2-Russian	0.382	0.340	0.342	0.343
3-Ukranian	0.054	0.061	0.045	0.056
4-Uzbek	0.011	0.011	0.012	0.016
5-Tatar	0.021	0.016	0.020	0.020
6-Uyghur	0.011	0.014	0.009	0.009
7-German	0.020	0.018	0.016	0.021
8-Other	0.040	0.033	0.030	0.033
Age of head of HH	50.550	49.990	49.680	54.050
Marital status of head of HH:				
1-married	0.657	0.678	0.638	0.586
2-never married	0.042	0.038	0.058	0.031
3-divorced	0.100	0.089	0.117	0.111
4-widow	0.201	0.196	0.187	0.273
The status of head of HH:				
0-employed	0.684	0.694	0.743	0.661
1-student	0.002	0.002	0.001	0.000
2-pensioner	0.235	0.218	0.205	0.283
3-housekeeper	0.026	0.023	0.023	0.020
4-disabled person	0.017	0.025	0.017	0.023
5-unemployed	0.030	0.029	0.010	0.013
6-other	0.007	0.008	0.001	0.001
Type of settlement:				
1-Astana city	0.020	0.022	0.020	0.019
2-rural settlement	0.370	0.514	0.446	0.527
3-large cities	0.320	0.306	0.327	0.288
4-meduim-size cities	0.073	0.057	0.075	0.059
5-small towns	0.130	0.053	0.045	0.065
6-Almaty city	0.088	0.048	0.087	0.041
Household size	3.800	4.210	3.447	3.443
Quantity of female in HH	2.017	3.500	1.900	3.710
Quantity of male in HH	1.790	3.720	1.590	3.170
Quantity of children in HH	1.240	1.480	0.923	0.821
Quantity of elderly in HH	0.430	0.380	0.365	0.455
Sample size	11679	2850	11782	2850

Source: Author's calculations based on KHBS. Note: Due to changes in administrative region divisions, some small towns belong to rural settlement in 2009. The descriptive statistics are available for all years. Note: HH means *household*.

Appendix 2 - Stochastic Dominance Analysis

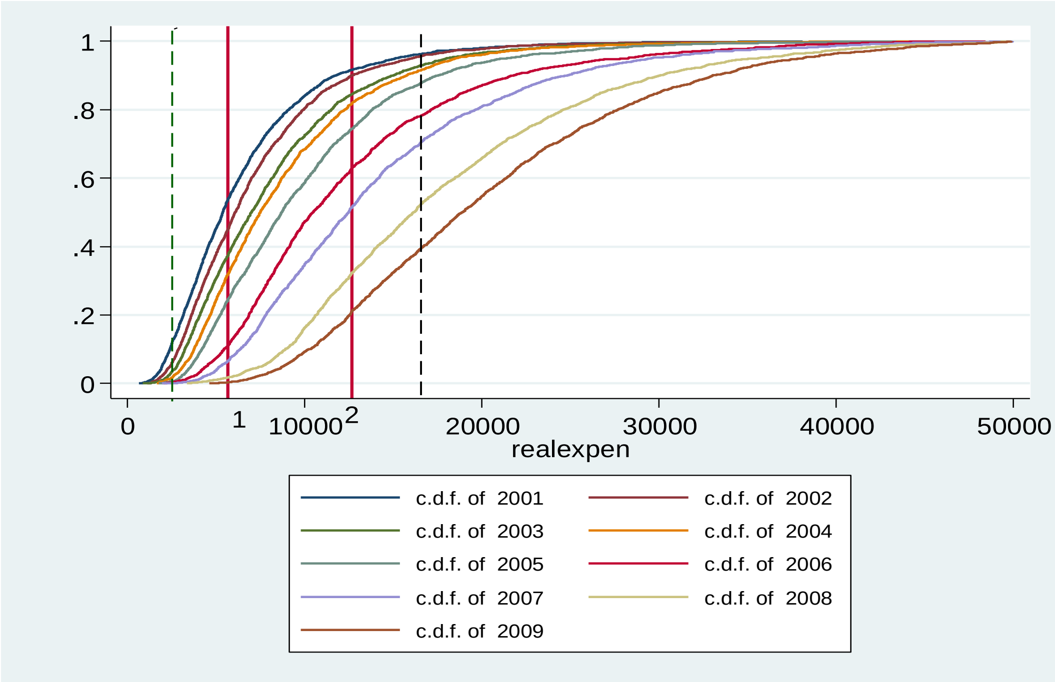
Stochastic dominance analysis may apply to the estimation of relations between a pair of distributions. Widespread implementation of stochastic dominance in the study of income distributions, poverty, and income inequality has been studied by various researchers (Atkinson, 1987; Davidson & Duclos, 2000; Foster & Shorrocks, 1988).

We analyse economies of scale and poverty lines for the panel data over the period 2001-2009 by application of stochastic dominance analysis. First, we consider real consumption expenditures per capita for absolute and relative poverty lines. Figures 1-2 illustrate the results graphically. Figure 1 draws the cumulative distribution functions (CDFs) of the real consumption expenditures per capita during 2001-2009. The graphs are restricted to expenditures below KZT 50,000 in order to see the behaviour of households that are close to poverty lines. Two bold vertical lines illustrate the subsistence minimums in 2001 and 2009 and two dashed lines set relative poverty lines based on 40 percent cut-off in real consumption expenditures per capita for 2001 and 2009, taken from the panel data of 2001-2009.

The CDF curves from Figure A1 are strictly below each other, in turn, between the Kazakh official poverty line in 2001 and the Kazakh official poverty line in 2009. This fact illustrates that the next year's CDFs first-order stochastically dominate the CDFs of previous years. This implies a reduction in consumption poverty incidence regardless of which poverty lines and measures are used in this domain. The same tendency is observed in the application of relative poverty lines.

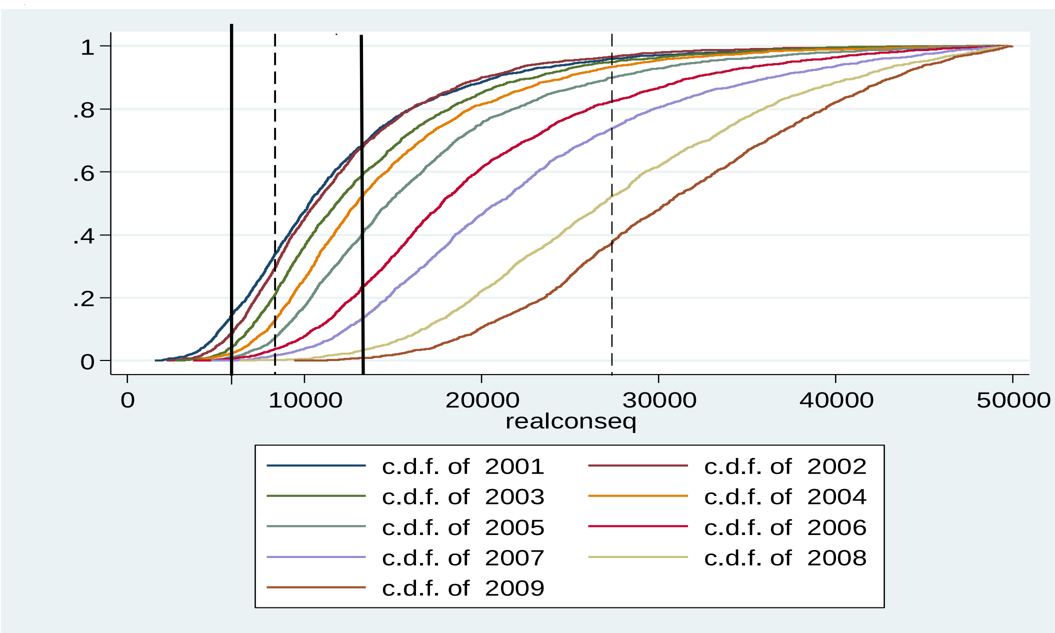
Figure A2 illustrates the graphs of CDFs for real consumption expenditures per adult equivalent from the panel data. It is also restricted to expenditures below KZT 50,000 to see the behaviour of households close to poverty lines. The square root of household size has been taken as the adult equivalent scale. Two bold vertical lines illustrate the subsistence minimums in 2001 and 2009, and two dashed lines set relative poverty lines based on 40 percent cut-off in 2001 and 2009 for real consumption expenditures per adult equivalent from the panel data of 2001-2009. Relative poverty lines for this case are larger than for the previous case (compare Figure A1 and Figure A2). The shapes of distribution for the later period indicate that the percentage of households below the official subsistence minimums is negligible.

Figure A1: Cumulative distribution of per capita real consumption expenditures of households in panel data for 2001-2009.



Source: Author's calculations based on KHBS.

Figure A2: Cumulative distribution of per adult equivalent real consumption expenditures of households in panel data for 2001-2009.



Source: Author's calculations based on KHBS.

The CDF curves from Figure A2 are almost below each other, in turn, between the Kazakh official poverty line in 2001 and the Kazakh official poverty line in 2009. The only exception is for 2001 and 2002, where the CDF curves intersect at one point. However, the point of intersection is above that of the poverty lines for 2001. This picture illustrates the restricted first-order stochastic dominance up to the absolute poverty line for 2009. This implies a reduction in consumption poverty incidence regardless of which poverty lines and measures are used for the period 2001-2009. The same tendency can be seen for the application of relative poverty lines.

Appendix 3

Table A3: Comparison of durations in poverty for households and individuals by different poverty lines and per adult equivalent consumption expenditures for 2001-2009.

Time poor	Households (per capita expenditures and absolute poverty line)		Individuals (per capita expenditures and absolute poverty line)		Individuals (per adult equivalent and relative poverty line[14])	
	Quantity	Proportion (%)	Quantity	Proportion (%)	Quantity	Proportion (%)
0	871	33.76	1565	27.62	1519	26.93
1	336	13.02	621	11.03	900	15.95
2	235	9.11	524	9.31	611	10.83
3	185	7.17	412	7.32	664	11.77
4	195	7.56	440	7.82	497	8.81
5	206	7.98	510	9.06	335	5.93
6	180	6.98	462	8.21	333	5.91
7	160	6.2	469	8.33	287	5.08
8	130	5.04	370	6.57	298	5.28
9	82	3.18	267	4.74	196	3.47

[14] As a relative poverty line, 40 percent of per adult consumption expenditures are taken.

Abstrakt

Tato studie využívá panelová data vytvořená z rotujících průřezových šetření Kazakhstan Household Budget Survey za období 2001-2009 k měření hladiny chronické chudoby a přechodu ze/do stavu chudoby v Kazachstánu. Zjišťujeme, že navzdory rychlému a podstatnému snížení chudoby v zemi od přelomu století, v závislosti na způsobu měření chronické chudoby, téměř čtvrtina populace čelí přetrvávající chudobě. Nicméně, většina chronicky chudých zažívá přerušovaná období chudoby. Studujeme faktory ovlivňující přechody ze/do stavu chudoby. Hlavním faktorem zvyšujícím pravděpodobnost propadu do stavu chudoby a snižujícím pravděpodobnost úniku ze stavu chudoby je přítomnost dětí do 6 let v rodině. Zavedení programu, který by poskytoval finančně dostupnou péči o děti, by změnilo tento stav.

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