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Longevity in Russia's Regions

Do Poverty and Low Public Health Spending Kill?

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Abstract

This paper examines the impact of changes in poverty and public health spending on inter-temporal variations in longevity using a unique regional-level dataset that covers 77 regions of Russia over the period 1994-2000. The dynamic panel data model is used as a tool for the empirical analysis. The model is estimated using the Arellano-Bond dynamic panel data estimator. The changes in regional levels of poverty and real per capita public health expenditure are identified to be significant determinants of the variations observed in longevity over time. The empirical results indicate that while male life expectancy responds more strongly than female life expectancy to economic circumstances, the latter appears to be more predisposed to the influence of public health spending. The results support the idea that the (positive) effect of public health spending on life expectancy is larger for those regions that experience higher incidences of poverty. The paper also finds that the financial crisis which hit Russia at the end of 1998 had a significant negative effect on longevity independently of the factors directly related to poverty and public health spending.

Keywords: life expectancy, poverty, public health spending, panel data

JEL classification: I12, I31

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1 Introduction

Poor health is a fundamental obstruction to human capital. It not only affects the longevity and quality of individual lives, but also undermines the future economic prospects of a country. These concerns make it very important from a policy perspective to understand the determinants of health outcomes. This paper explores a unique regional-level dataset on life expectancy, the incidence of poverty, and real per capita public health spending in Russia over the period 1994-2000 to explain the variations in longevity over time.[1]

It is by now a well-established fact that life expectancy, an aggregate measure of the population's health, declined dramatically in Russia during the first years of economic and social transition. It increased somewhat as the process of economic recovery started, but declined again in the aftermath of the 1998 financial crisis (see Figure 1).

The average life expectancy figures for Russia, however, disguise remarkable regional differences in the levels of, and changes in, life expectancy. The scope of these regional disparities is revealed in Figure 2, which compares all-Russia life expectancy during the period 1990-2000 to life expectancy in regions which are the least and most successful in terms of longevity.

Figure 2 provides clear indication that Russia's regions differ tremendously not only in the levels of longevity at any given period, but also in the changes in longevity over time. For instance, in 1990 life expectancy in the Dagestan Republic was 3.6 years higher than the all-Russia average, while in the Tyva Republic it was 7 years lower. Moreover, from 1990 to 1994 longevity in the former region declined by 2.4 years, as compared to 7.1 years in the latter. As one can see, even the *direction* of change in life expectancy at the regional level do not always correspond to the direction of change in the all-Russia life expectancy.

The magnitude of inter-regional and inter-temporal variations in longevity in Russia can be articulated further by the life expectancy maps of Russia for 1998 (before crisis) and 2000 (after crisis) presented later (see Figures 3 and 4).

In terms of inter-regional disparities, the life expectancy map suggests that the lowest levels of longevity are observed in the European North, Urals, Siberia and the Far East. The regions with the highest life expectancy are the Northern Caucasus and Volga-Vyatka (see Table A1, Annex). As mentioned in many studies, such regional patterns of mortality in Russia have been observed for many years (Shkolnikov *et al.* 1999).

In terms of the regional *trends* in longevity, it is worth noting that Northern Russia, East Siberia and the Far East have witnessed the greatest declines in life expectancy. Between 1990 and 1994, the period of the largest drop in longevity across all Russian regions, life expectancy declined by 7 years in the Northern region, 6.7 years in East Siberia, 5.8 years in Western Siberia, and 5.7 years in both Urals and the Far East. Notably, these regions were among those which experienced the greatest economic and

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1 In contrast to the crude death rates data, life expectancy data allow an abstraction from the age structure of the population and hence give a more accurate picture of differences in excess mortality across regions and over time (Nell and Stewart 1994). The quality of Russian demographic data is widely regarded to be comparable to that for the United States (e.g., Becker and Hemley 1998).

social disruption, as indicated by the rates of industrial output decline, job turnover, unemployment, divorce and fertility. As recognized in the literature, declines in life expectancy were the smallest in the agricultural regions of the south, and in the regions of Central Russia having the most developed medical services infrastructure. However, even there the longevity declines were quite substantial. So from 1990 to 1994, life expectancies in the North Caucasus and the Volga Basin regions declined by 3.1 and 4.1 years, respectively (see Table A1, Annex). Hence, dramatic changes in longevity during the 1990s were observed in *all* regions of Russia.

Figure 1
Life expectancy (LE) at birth in Russia, 1990-2000

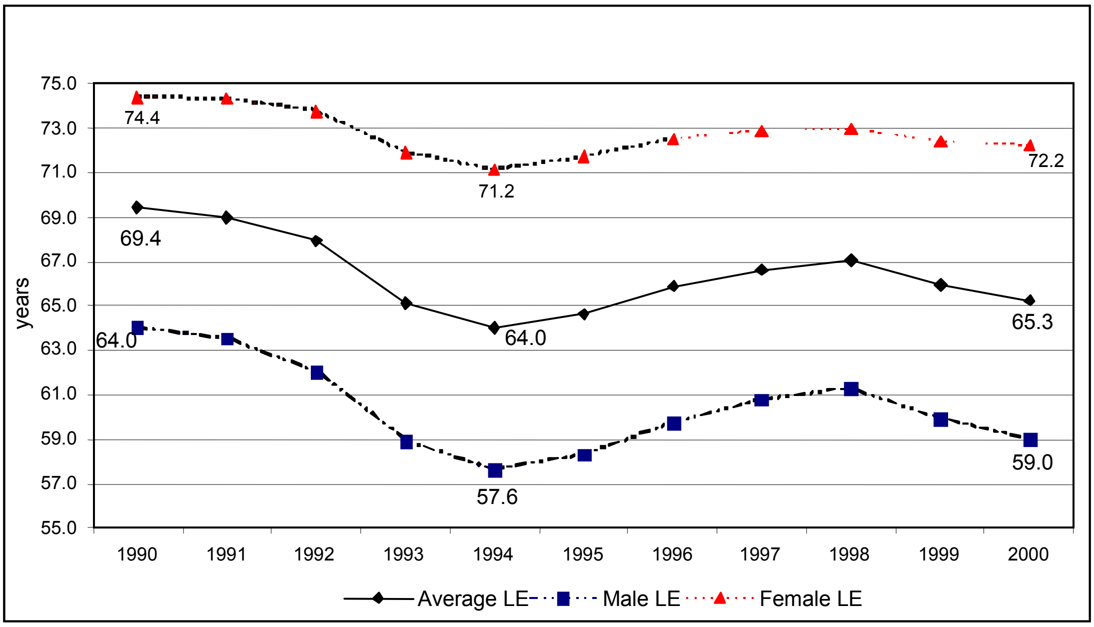
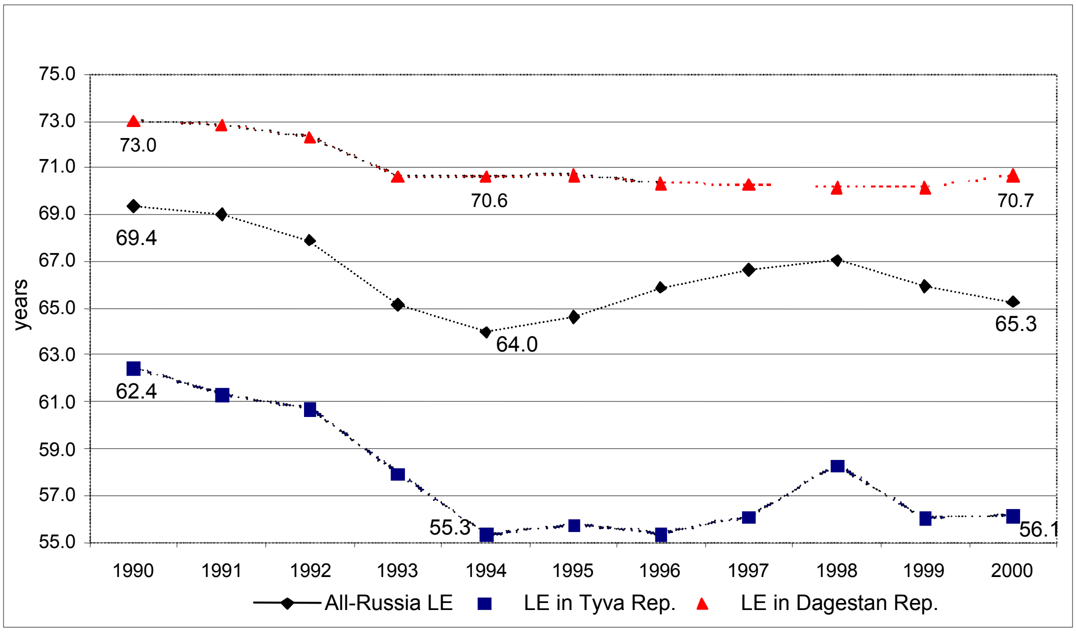


Figure 2
Life expectancy (LE) at birth in selected regions of Russia, 1990-2000



Notably, the regions with the largest declines in longevity during 1990-94 experienced the most noticeable improvements starting in 1995. For instance, life expectancy rose by an average 4.5 years in the North, 3.5 years in Siberia, and 3.4 years in the Far East from 1994 to 1998. However, longevity improvements after 1994 have not lasted very long. As Figures 3 and 4 suggest, even a relatively short-lived episode such as the financial crisis that erupted in the fall of 1998 has led to worsening longevity in all regions. In the North, Urals and Central regions life expectancy declined on average by more than 2 years in the period from 1998 to 2000. The smallest declines in longevity during the crisis were observed in the North Caucasus and Western Siberia.

Figure 3
Life expectancy in Russia's regions, 1998

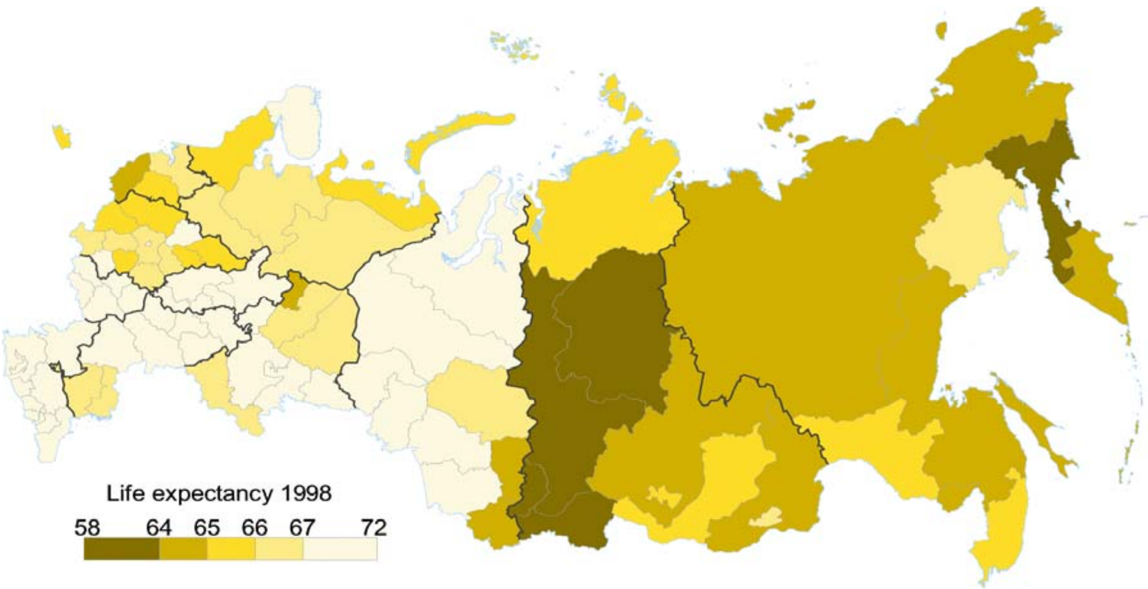
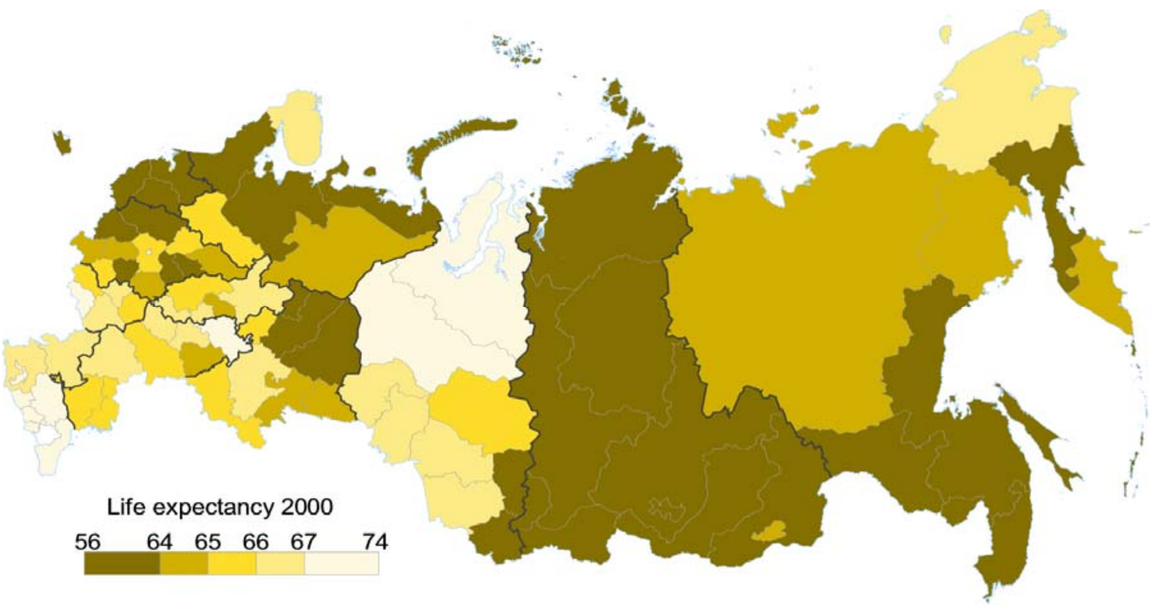


Figure 4
Life expectancy in Russia's regions, 2000



Although some region-specific time-invariant factors such as climate and environmental conditions are partly responsible for the observed differences in the levels of longevity across regions, they do not provide a comprehensive explanation for these differences, and clearly do not explain the changes in longevity over time. Hence, there should be some time-variant factors that help to explain the observed differences in life expectancy over time and across regions.

It should be noted here that research on the causes of the mortality crisis in Russia has been quite extensive.[2] However, previous research in this area has several shortcomings. First, most of the previous studies are of a highly descriptive nature and concentrate on the determinants of mortality by *immediate* causes of death (Shkolnikov *et al.* 1999; Shkolnikov and Mesle 1996; Walberg *et al.* 1998). Among the immediate causes identified to have contributed the most to the decline in life expectancy are cardiovascular diseases, injuries and violence, suicide, and alcohol-related disorders such as cirrhosis of the liver and accidental alcohol poisoning (Shkolnikov and Mesle 1996; Walberg *et al.* 1998). During the period 1989-94, circulatory diseases and accidents (trauma and poisoning) accounted for 42 and 33 per cent of the total *increase* in all deaths, respectively (Brainerd and Varavikova 2001).[3] Although undoubtedly important for gaining insight into the nature of the mortality crisis, these studies do not investigate the impact of the *underlying* factors affecting health outcomes.

Second, although many, if not all, studies recognize that changes in mortality are likely to have been effected by economic and social changes associated with transition, the empirical evidence on the *direct* impact of these changes on longevity (or mortality) in Russia remains scanty. A revision of the few those studies that attempt to quantify the impact of certain economic and social factors on mortality (see Becker and Hemley 1998; Brainerd and Varavikova 2001; Kennedy *et al*. 1998) makes it clear that they produce empirical results which are far from conclusive.

Third, while it has been speculated in many studies (Nell and Stewart 1994; Zohoori *et al.* 1998) that the health status of people in Russia is likely to be strongly affected by an increasing incidence of poverty and deteriorating public health infrastructure, the impacts of these factors on longevity (or other measures of health) have not been adequately explored empirically.[4]

Finally, the bulk of the studies have focused on explaining mortality during the initial period of economic transition—namely before 1995 (Shkolnikov *et al.* 1999; Walberg *et al.* 1998). The trends in health outcomes after 1994, particularly in the aftermath of the autumn 1998 financial crisis, have received very little attention. Since life expectancy showed a rapid decline from 1989 to 1994, the studies that cover this period effectively attempt to understand why the mortality rates increased. Although it is crucial to comprehend the causes of declining longevity, it is equally important to understand why longevity started to increase after 1994, and declined again in 1999.

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2 See, for example, the edited collections of papers by Cornia and Paniccia (2000a) and Cockerham (1999).

3 The same factors contributed the most to the decline in all deaths during 1994-98. The increase in mortality in the aftermath of 1998 was due primarily to increased deaths from circulatory diseases.

4 For two attempts in this direction, see Brainerd and Varavikova (2001) and Kennedy *et al.* (1998).

This paper is the first to examine the direct impact of poverty and public health expenditure on life expectancy in Russia using a compiled dataset that covers 77 regions over the 1994-2000 period. Although it may seem intuitively obvious that these factors are likely to be strongly associated with population health, econometric analyses that investigate their direct impact on health are scarce due to data limitations.

Considering that most health policy decisions are made at the country level, the empirical analysis of the relationship between health outcomes and socioeconomic conditions using regional data is expected to have much more policy relevance than cross-country studies. This study provides an attempt to evaluate the effects of regional poverty and public health spending on life expectancy in Russia's regions. The awareness of the magnitude of these effects is critical in the formation of policies aimed at improving health outcomes in the regions of Russia and the country as a whole.

The remainder of the paper is organized as follows. Section 2 discusses the reasons for believing that poverty and public health spending affect longevity in Russia. Section 3 reviews the body of earlier cross-country studies on the effects of poverty and public health expenditure on health outcomes. Section 4 first provides theoretical underpinnings of our empirical model, and then discusses the empirical specification and estimation methodology. Section 5 describes the data used in the empirical analysis. Section 6 discusses the estimation results. Section 7 tests the robustness of the results to alternative model specifications and estimation techniques. Section 8 concludes by highlighting the main findings and discussing their main policy implications.

2 Reasons to believe that poverty and public health spending affect longevity in Russia

Although empirical evidence on the effects of poverty and public health spending on longevity in Russia is almost non-existent, there are strong reasons to believe that the unprecedented regional disparities in these factors, as well as the variations in them over time, are the likely culprits for the observed regional and temporal patterns of longevity. These reasons are discussed below.

As emphasized in a number of studies, the increase in the death rates in Russia chiefly affected prime-age men.[5] Concurrently, the same population group experienced the highest loss in real and relative wages during the initial phase of the transition. For instance, in 1994 men with several decades of labour market experience earned on average lower wages than new labour market entrants (Brainerd 1998). It is argued that middle-aged men in Russia are among those hit the hardest by poverty and stress associated with loss of job security and the need to support a family in new economic circumstances (Zohoori *et al.* 1998).

[5] It has been estimated that from 1990 to 1995 Russia experienced between 1.3 and 1.6 million premature deaths, most of which were among working-age men. The mortality rate among men aged 35 to 44 was four times as high as that in Western Europe (Bennett *et al.* 1998). Deaths between the age of 35 and 64 made the largest contribution to the gap between regional *levels* of mortality (McKee 2001).

Many studies have documented a rapid increase in mortality due to causes likely to be strongly related to the lack of income. Several studies argue that rising mortality from cardiovascular disease, a dominant cause of increased mortality, is due to the high levels of stress experienced by Russia's population during the transition.[6] The very same studies recognize that rising stress levels should be regarded not as an addition to, but rather as the consequence of economic factors (Zohoori *et al.* 1998; Stegmayr *et al.* 2001; Shkolnikov *et al.* 1999). Among these factors are high labour force turnover, decline in real income and an increase in poverty. The increase in cardiovascular disease is also likely to be directly linked to poverty because of the insufficient consumption of fruits and vegetables, important sources of antioxidants.[7]

A number of studies indicate that the changes in life expectancy in Russia are related to changes in per capita income (Brainerd and Varavikova 2001) and income distribution (Walberg *et al*. 1998). Since at a given poverty line the prevalence of poverty in a region is a function of average per capita income and income inequality, these studies effectively provide implicit support for the hypothesis that poverty is fact that matters for longevity.[8]

A few studies find the crime rate to be a significant predictor of both male and female mortality (Brainerd and Varavikova 2001; Kennedy *et al.* 1998; Walberg *et al*. 1998). Notably, Kennedy *et al.* (1998) find that with the inclusion of the crime rate in the regression model, the effects of per capita income and poverty on life expectancy vanish, which indicates that the crime rate can be considered, *ceteris paribus*, as a function of economic circumstances. The same study identifies a statistically significant positive correlation between self-perceived economic hardship and the divorce rate, which is another variable often found to affect death rates, particularly those from accidental alcohol poisoning and suicide (Brainerd 2001; Brainerd and Varavikova 2001).

Increased alcohol consumption during the transition is another factor which many studies see as contributing significantly to the rise in mortality in Russia (Leon *et al.* 1997; Ryan 1995). Accidental alcohol poisoning and cirrhosis of the liver accounted for 9 per cent of the total increase in standardized death rates (SDR) from 1989 to 1994 (Brainerd and Varavikova 2001).[9] An additional impact of alcohol consumption on mortality is likely through its role in deaths from injuries and violence (McKee 2001). Unhealthy patterns of drinking are also found to increase the risk of sudden

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6 Increased stress may also result in the increased consumption of tobacco and alcohol, as well as other unhealthy behaviour.

7 It has been argued that malnutrition is not a major factor causing health problems in transition economies. Although it has been acknowledged that diets in many countries of the FSU have undoubtedly deteriorated, the *average* calorie intakes remain quite high. This does not rule out the possibility of substantial nutritional deficiencies among the *poor*, as is argued by Shkolnikov and Mesle (1996) to be the case for Russia. This would suggest that the poverty rate is a much better indicator of the possible incidence of malnutrition among the population than the *average* calorie intake.

8 Zohoori *et al.* (1998) using 1992-96 RLMS data find that changes in the poverty rates in Russia have closely followed the movements of household incomes.

9 It is worth mentioning that these causes of death made almost the same contribution to the *decline* in SDR during the 1994-98 period.

cardiovascular death (Britton *et al.* 1998). However, almost every study on the health impact of alcohol consumption recognizes the need to find the *underlying reasons* for the observed drinking patterns (Brainerd and Varavikova 2001; Shkolnikov *et al.* 1999; McKee 2001). Indeed, alcohol consumption is hardly considered an exogenous cause of mortality. There is some empirical evidence to support the notion that alcohol consumption and other deleterious health behaviour in Russia—and elsewhere—are functions of socioeconomic conditions. Using the 1992-96 micro-level data from the Russian Longitudinal Monitoring Survey (RLMS), Zohoori *et al.* (1998) find that per capita alcohol consumption rose with an increase in the prevalence of poverty (perhaps due to psychological stress related to poverty). The same study indicates that the largest amount of alcohol is consumed by men in the lowest income group.[10] Hence, drinking would simply reflect an important causal chain through which poverty exerts effects on health.

A number of studies have documented a rapid increase in mortality due to preventable diseases normally associated with deteriorating medical and sanitary services, such as infectious and parasitic diseases (Paniccia 2000; Shkolnikov and Cornia 2000; Shkolnikov and Mesle 1996). Among the increase in death rates, the rate for this group of diseases in the early 1990s was one of the largest. In the mortality toll due to infectious diseases, tuberculosis and diphtheria occupy a dominant place. Although far from being dominant, the relative contribution of infectious diseases to the total changes in mortality is not negligible. In 1999 mortality due to tuberculosis accounted for 6 per cent of the increase in male deaths for that year (Brainerd and Varavikova 2001). The incidence of tuberculosis increased from 34 cases per 100,000 people in 1991 to 91 cases in 2000 (Goskomstat 2001). At the same time, several studies recognize that even the rise in mortality from cardiovascular disease, a major contributor to changes in death rates in Russia, could have perhaps been smaller, had the health infrastructure not deteriorated as much as it did.

Facing economic hardship, many medical facilities were forced to start operating, legally or not, on a fee-for-service basis. The practice of medical staff demanding substantial under-the-table payments for the provision of services also became widespread (World Bank 2000). As a result, those unable to pay with private funds were practically abandoned by the health system.[11] Notably, while public spending on health as a share of GDP remained largely unchanged, the shrinking economy and strained government budget implied a substantial drop in *real* public health spending across regions, often to levels well below that necessary to cover the most basic expenditures.[12] While probably not the main factor in the short-term deterioration of

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10 The high level of alcohol consumption by those at the bottom of income distribution is possible, given that the relative price of alcohol declined noticeably during the transition.

11 It is worth noting that in those transition countries of Eastern Europe where life expectancy has on average increased during the 1990s (Czech Republic, Poland and Slovenia), real public health spending has increased despite the hardships of the transition period. In Russia, real public health spending has on average more than halved from 1989 to 1998.

12 Among government expenditures allocated to the health sector, salaries of the health sector employees accounted for the dominant share (although the average salary in the health sector continued to be below the country average), while resources for medicine and medical equipment were lacking. As a result, there have been numerous reports of cases of doctors performing surgery with razors and recycling disposable equipment (Davis 1993). A detailed account of the crisis in the Russian health care system is provided in Davis (1993) and Rosenfeld (1996).

health status in Russia, the weakness of the health care system is regarded as one of the dominant long-term determinants of health (Field 1995).[13]

Finally, previous attempts, using cross-country data, to identify the effects of poverty and public health spending on longevity generally found support that these factors matter for population health. We review some of these studies below as they have an immediate relevance to our work. The findings of these studies will provide a background comparison for our results. We also discuss the limitations of these studies.

3 Previous cross-country studies of the effects of poverty and public health spending on health outcomes

The main determinants of longevity (and of other health outcomes such as infant and child mortality) examined in cross-country studies include national incomes (per capita GDP), poverty, income inequality, private and public provision of health services. Although the literature on the determinants of longevity is voluminous, empirical evidence on the effects of poverty and public health spending is relatively scarce due to the lack of data comparable across countries and over time.

Using a sample of 22 developing countries, Anand and Ravallion (1993) investigate the effects on life expectancy of the poverty incidence and public health spending per capita. Their results suggest that a 1 per cent increase in the proportion of the population living below the poverty line is associated with a 0.21 per cent fall in life expectancy and that a 1 per cent increase in public health spending per capita leads to a 0.30 per cent rise in life expectancy. The authors argue that average income (per capita GDP) matters only through its close association with the real value of public health spending and the magnitude of absolute poverty.

In a later study using panel data on 35 developing countries, Bidani and Ravallion (1997) also find that the incidence of absolute poverty and per capita public health spending are significant determinants of longevity. Their results suggest that public health expenditure affects the health outcomes more for the poor than for non-poor.[14] The estimated elasticity of life expectancy for the poor with respect to public health spending indicates that a 1 per cent increase in public health spending per capita is expected to result in a 0.13 per cent rise in life expectancy.

Using a cross-section of 84 developing countries for 1990, Carrin and Politi (1995) explore the impacts of poverty incidence and public health expenditure on the population's health. They effectively revisit the results of Anand and Ravallion (1993) by applying the same model to an extended dataset. The only difference is that they use the share of public health expenditure in GDP rather than real per capita health spending as a measure of public investment in health. They argue that this variable reflects the willingness of the government to allocate public resources for health. As measures of

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[13] McKee (2001) argues that the weakness of the medical care system is likely to play an increasing role in the latter part of the 1990s, which is the period of our analysis.

[14] This is to be expected since the non-poor are in a better position to compensate with higher private health care spending for the lack of public provision of health services. Kakwani (1993) also finds that the income elasticity of life expectancy declines as income rises.

population health they use life expectancy at birth, and child (under 5) and infant mortalities. They find the incidence of poverty to be a significant factor determining health. The poverty elasticity of life expectancy estimated from a log-linear specification suggests that a 1 per cent increase in the incidence of poverty would be associated with a 0.04 per cent reduction in life expectancy. They also estimate a logarithmic reciprocal-type specification to account for the natural bounds of the dependent variable. In this case the estimated poverty elasticity of life expectancy is -0.02. Public health expenditure as a share of GDP is not found to affect health outcomes. However, this result is most likely explained by the fact that this variable does not measure the real value of health spending. An estimated instantaneous relationship also does not allow one to control for the probable endogenous nature of health spending.

Calfat (1996) investigates the impact of public health spending and rural poverty on life expectancy, infant mortality and child (under 5) mortality using a cross-section sample of 68 developing and developed countries for 1990. The share of health expenditure in total GDP is used as a measure of public health spending. The author finds that lower levels of rural poverty are associated with improved health for all health indicators used. The poverty elasticity of life expectancy estimated from the loglinear specification varies from -0.14 to -0.20 depending on which other explanatory variables are included in the model. This study, however, finds no evidence for the public health expenditure as a share of GDP to affect longevity or infant mortality. The use of cross-section data, however, does not allow the author to control for unobserved country effects and possible endogeneity of health spending.[15]

Filmer and Pritchett (1999) explore a cross-section of 100 developing countries to investigate the effects on health outcomes (infant and child (under 5) mortality) of socioeconomic conditions and public expenditures on health. They find that while a country's per capita income and inequality of income distribution explain a substantial portion of cross-national variation in mortality, public health spending does not appear to be a powerful determinant of mortality. The impact of public expenditure on under-5 mortality is found to be only marginally (at a 10 per cent level) significant in one specification. According to their estimates, doubling public health expenditure would decrease child mortality by 9 to 13 per cent.

Rajkumar and Swaroop (2002) study the impact of public health spending on child (under 5) and infant mortality using a cross-section of 98 developed and developing countries over two years (1990 and 1997). They look at the contemporaneous relationship between health spending and these health indicators. Their major finding is that increasing public resources for health lowers child and infant mortalities, although the efficacy of public health spending depends on the quality of governance. Countries that have less corrupt and more effective bureaucracies are shown to enjoy a higher elasticity of health outcomes with respect to public health spending. The elasticities estimated at the mean Quality of Bureaucracy Index suggest that a 1 per cent increase in the share of public health spending in GDP is linked with a 0.22 and 0.23 per cent reductions in child (under 5) and infant mortalities, respectively. They also find that an

[15] The author also argues that the share of health expenditure in total GDP may not capture the real impact of government health spending since health outcomes are likely to depend not only on the magnitude of expenditure, but also on the effectiveness of its use.

increase in per capita GDP results in an improvement in mortality, while an increase in income inequality (controlling for per capita GDP) causes mortality to deteriorate. The statistically significant parameter estimates on per capita GDP and income inequality indirectly suggest that an increase in poverty is associated with a deterioration in mortality.

Lichtenberg (2002) estimates the longevity model using annual US time-series data on life expectancy, health expenditure, and medical innovation for the period 1960-97. He finds strong support for the hypothesis that public expenditure on health contributed to a longevity increase during this period. The long-run elasticity of longevity with respect to total (public and private) health expenditure is estimated to be 0.09. The elasticity with respect to public health expenditure alone is found to be 0.04.

To sum up, the studies reviewed above seem to concur that the incidence of poverty is a significant determinant of health outcomes. The evidence on the health impact of public provision of health services seems to be much less conclusive. Nonetheless, earlier research is subject to several limitations which can affect the reported results. Perhaps the most important limitation is that all studies based on cross-sectional data risk identifying a merely associative rather than causative relationship between the dependent and explanatory variables of interest; this is because these studies do not address the possible endogeneity of the explanatory variables. In addition, the cross-sectional nature of the data does not permit one to account for the country or region-specific fixed effects, which presumably produce biased estimates. The cross-sectional studies are also by nature constrained to explore the cross-sectional variation in health outcomes. Hence, the critical variability in health indicators over time remains unexplained. Moreover, the empirical models used in the studies are static (for an exception, see Lichtenberg 2002), and thus fail to consider the dynamic nature of such health outcomes as life expectancy. Finally, the cross-country data explored in many studies are often of a quality that makes their comparability (and thus the very legitimacy of their use in the analysis) questionable.

Our study aims to address the above limitations, and establish the causal impact of poverty and public health spending on longevity using panel data.

4 Model

4.1 Theoretical grounds

We assume that *individual* health in year t is a function of the stock of health in year t-1 and the private and public real per capita health expenditures in year t, so that:

$$h_{it} = f(h_{it-1,} HE\_private_{it}, HE\_public_{it}, v_r) \quad (1)$$

where h_{it} is the health of individual i in year t, h_{it-1} is the lagged stock of health, $HE\_private_{it}$ is private investment in health by individual i in year t, HE_public_{it} is public health spending per individual i in year t, and v_r are time-invariant regional factors (e.g., climate, pollution) affecting health.

Private health expenditure can be best approximated by individual income, in which case equation (1) transforms into:

$$h_{it} = f_i\,(h_{it\text{-}1,}\ y_{it},\ HE\_public_{it},\ v_r) \qquad (2)$$

where y_{it} is income of individual i in year t.

There is considerable empirical evidence to suggest that the relationship between income and health is non-linear; that is, increasing income improves health, but after a certain point, exerts diminishing effects (Backlund *et al.* 1996; Deaton 2001). The assumed concave relationship between health and income implies that *population* health (which can be represented by a summary statistic such as life expectancy) depends not only on the mean income of the people living in a region, but also on the distribution of income around the mean.[16] At the same time, public health expenditure as a public good can be assumed to be equally available to each resident in a region. When aggregating to the regional (population) level, these assumptions allow us to express equation (2) as:

$$H_{rt} = f_r(H_{rt\text{-}1,}\ Y_{rt},\ I_{rt},\ HE\_public_{rt,}\ v_r) \qquad (3)$$

where H_{rt} is a measure of population health in region r in time t, H_{rt-1} is the lagged value of population health, Y_{rt} is the average income of people in a region r in time t, I_{rt} is the measure of income inequality in a region r in time t, HE_public_{rt} is real per capita public health expenditure in a region r in time t, and v_r are time-invariant regional factors (e.g., climate, pollution) affecting health.

At a given regional poverty line, the mean income in a region and the distribution of incomes around the mean would define the incidence of poverty. Hence, equation (3) can be alternatively presented as:

$$H_{rt} = f\,(H_{rt\text{-}1,}\ P_{rt},\ HE\_public_{rt,}\ v_r) \qquad (4)$$

where P_{rt} is the incidence of poverty (poverty headcount index) in region r at time $t,$ and all other variables are as defined above.[17]

Next we discuss the specification of equation (4) which is to be empirically estimated.

4.2 Empirical specification and estimation technique

As the rates of poverty and public health spending are likely to have diminishing marginal effects on longevity, in our empirical specification we follow many other studies (e.g., Anand and Ravallion 1993; Collins and Thomasson 2002; Lichtenberg 2002; Pritchett and Summers 1997) in assuming the log-linear relationship between life expectancy, an aggregate indicator of population health on the regional level, and the variables determining it. The baseline regression equation takes the following form:

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[16] Several hypotheses have been presented as to why average income and income distribution should affect the health of the population. A useful review of these is provided in Wagstaff and van Doorslaer (2000). In the same study, one can find formal proof that population health is a function of mean income and income inequality. Note that the authors also show that the assumption of the absolute income hypothesis (AIH) suggested by equation (2) is not critical in getting this relationship.

[17] It would be reasonable to expect that the impact of public health spending depends not only on its size, but also on the possibilities in the region for people to procure adequate health services through private means. In other words, public health expenditure is likely to be of more consequence in the poorer regions. We will test this hypothesis later in the paper.

$$\ln(LE_{rt}) = \alpha + \beta_1 * \ln(LE_{rt-1}) + \beta_2 * \ln(P_{rt}) + \beta_3 * \ln(HE\_public_{rt}) + v_r + \varepsilon_{rt} \quad (5)$$

where r indexes regions, t indexes time periods, v_r are time-invariant region-specific characteristics that affect life expectancy, and $\varepsilon_{rt} \sim$ iid $(0,\sigma_e^2)$ is a disturbance term.[18]

We estimate equation (5) separately for male and female life expectancies, thereby allowing the coefficients to be gender-specific.

There is substantial debate in the literature on whether one should use the natural logarithm of life expectancy or some other transformation of this variable in the regression equation (e.g., Anand and Ravallion 1993; Kakwani 1993). The former study suggests a nonlinear transformation of life expectancy of the form ln(*maximum achievable LE – actual LE*) as an alternative to ln(*actual LE*). Such a transformation would reflect the fact that life expectancy is bounded from above, and that it takes greater effort to increase life expectancy by the same number of years in a country where life expectancy is initially higher.[19] Therefore, to check the robustness of our estimates we will perform the regression analysis using this alternative definition of the dependent variable as well.

It can be argued that the amount of public health spending in a given period can be a reflection of the observed health situation in a region. In other words, one might expect that worsening health outcomes may induce regional (or federal) government to spend more on health. If that is indeed the case, then the estimated contemporaneous effect of public health spending on life expectancy will be biased. To deal with this issue, the lagged value of public health spending is used in the estimations. Another argument for using a lagged rather than contemporaneous value of public health spending is that the effect of this factor on longevity is likely to take time to be felt. There will also be a high correlation between simultaneous values of public health spending and poverty if higher levels of spending are geared towards poorer regions, or if poor regions spend less on public health.

The estimation of a dynamic regression model represented by equation (5) using OLS will result in biased and inconsistent estimates (Davidson and MacKinnon 1993: 330). To overcome this problem we use the dynamic panel data GMM estimator derived by Arellano and Bond (1991) as the main instrument of the empirical analysis.[20]

Expressing equation (5) in the first difference form, and using the *lagged* rather than contemporaneous public health spending, we get:

$$\Delta\ln(LE_{rt}) = \gamma_0 + \gamma_1 * \Delta\ln(LE_{rt-1}) + \gamma_2 * \Delta\ln(P_{rt}) + \gamma_3 * \Delta\ln(HE\_public_{rt-1}) + \Delta\varepsilon_{rt} \quad (6)$$

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[18] To check the robustness of the results, we also introduce the time-specific effects in the estimations to capture the unobserved factors influencing life expectancy in a given period.

[19] This may be seen from the fact that if life expectancy increases by the *same* number of years in countries A and B, the percentage decline in ln(*max* – LE_{rt}) index will be larger for a country with a higher initial life expectancy.

[20] The consistency of this estimator hinges on the assumption of no second-order autocorrelation in the first-differenced idiosyncratic errors (we test the validity of this assumption in section 6).

This transformation effectively removes region-specific fixed effects v_r which are present in equation (5). The second lag of the level, $\ln(LE_{rt-2})$, as well as subsequent lags, are used as instruments for $\Delta\ln(LE_{rt-1})$.[21] Instrumenting with the lagged levels gives us an advantage over instrumenting with the lagged differences in terms of gaining an additional time period (hence, additional 77 observations) in the estimation. Assuming that the poverty rate and lagged public health spending are exogenous variables, the first differences of these variables will serve as their own instruments. Nevertheless, if these variables are considered to be predetermined, then they must be treated similarly to the lagged dependent variable. We perform estimations that explore both of these assumptions. We also check the robustness of our findings to alternative model specifications. But before turning to the regression results, in the next section we discuss the data used in the empirical analysis.

5 The data

Any cross-country study that aims to establish the relationship between health outcomes and such factors as poverty and public health spending will inevitably face the problem of comparability of poverty and public health expenditure estimates across countries, and even over time for a given country.

The problem of the comparability of poverty data emerges because of differences in the survey instruments (e.g., Living Standards Measurement Survey (LSMS) versus Household Budget Survey (HBS)), sampling designs, definitions of variables, and richness of information used in the construction of the income or consumption aggregate. As argued by Deaton (2001), the problem of data consistency can be substantially diminished by using regional-level data for a country. This is because national surveys use a uniform survey instrument and the same methodology of estimating a welfare aggregate. Hence, the errors in the estimated poverty rates remain constant across administrative regions and over time within regions, and thus the spatial and inter-temporal comparisons are not affected. The problem that often emerges with the use of regional data is that household surveys, which are nationally-representative, are often not designed to be regionally-representative, which precludes the possibility of making cross-region comparisons.

For our analysis we compile a unique panel of comparable poverty estimates for the regions of Russia. The data come from the HBS conducted each year by the State Statistical Agency of Russia (Goskomstat).[22] HBS has a sample size of around 49,000 households, and is nationally and regionally representative.[23] Goskomstat poverty estimates are based on regional poverty lines (national poverty line expressed in regional prices), and use per capita money incomes (since 1992) and per capita

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[21] The full set of moment conditions is given by: $E[y_{i,t-s}(\Delta y_{i,t}-\alpha\Delta y_{i,t-1})]=0$, for t=3,…,T and s=2,…,(t-1). In our case y_i would denote life expectancy for region i, and T=7.

[22] Another survey instrument available for Russia, Russian Longitudinal Monitoring Survey (RLMS), has a small sample size and, in contrast to HBS, does *not* generate representative data at the regional level.

[23] The sample size and regional representativeness of the HBS make it very similar to the US Current Population Survey that covers 50,000 households and generates state-representative data.

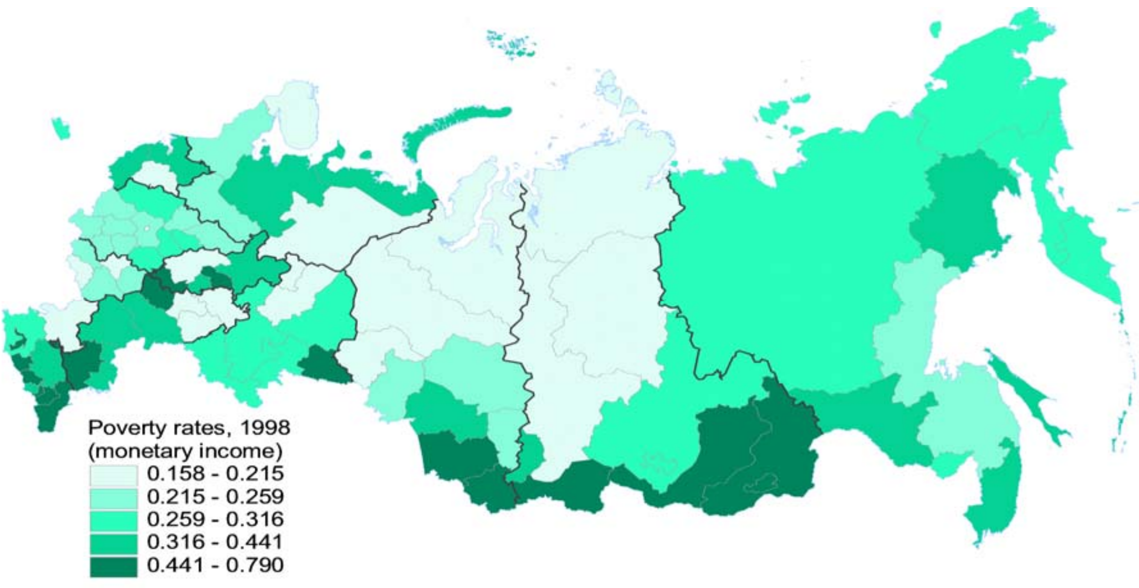
disposable resources (since 1997) as welfare measures. *Money income* is determined as total cash income received from formal sources (such as wages and salaries, social benefits, pensions and stipends), property income, plus estimates of income obtained outside officially registered economic activity (income from self-employment, sales of agricultural products, etc.). *Disposable resources* consist of cash expenditures, monetary assessment of in-kind consumption, plus withdrawn savings and borrowed funds during the survey period.[24] For the purpose of our study we use the money income poverty rates (poverty headcount indexes) because these data provide us with a longer time-series dimension. Nevertheless, the choice of the welfare indicator (money income versus disposable resources) is not expected to affect the results when the objective is to make comparisons across regions and over time using a given measure of

Table S1
Summary statistics of the poverty data, 1994-2000

Variable		Mean	Std dev.	Min	Max	Observations
hc	overall	31.998	14.538	11.500	98.595	N = 535
	between		13.002	15.230	87.399	n = 77
	within		7.859	-4.805	65.971	T = 7

Note: hc – poverty headcount index (incidence of poverty).

Figure 5
Money income poverty (headcount index) in Russia's regions, 1998



24 For a more detailed discussion of the welfare indicators used in Russia, see Yemtsov (2002).

of welfare.[25] Table S1 presents summary statistics of the regional incidence of money income poverty. [26]

It is worth noting that when we weight a sample by using the size of the regional population as a weight (in order to get nationally-representative summary statistics), we find that 28.6 per cent of the Russian population on average were poor during the period under consideration. This is slightly less than the sample mean of 32 per cent.[27] The poverty map of Russia clearly indicates the scope of regional disparities in the incidence of poverty (see Figure 5).

Figure 5 indicates that while in some regions less than 20 per cent of the population are located below the poverty line, in other regions the prevalence of poverty is almost universal. The picture of regional disparities appears to be similar when disposable resources are used as a welfare aggregate (see Figure 6).

For an indicator such as public health expenditure, any attempt in making cross-country comparisons can also introduce significant problems unless a researcher has detailed information on which entries make up the total public health spending in each country. In other words, the general term 'public spending on health' can, in fact, imply concepts that vary considerably across countries. Fortunately, we do not have this problem with the region-level data. The data on regional public health expenditures in current roubles are available from the Goskomstat. Regional public health spending is defined as the expenditure from the consolidated (federal plus local) budget on health care in a given region.[28] To obtain the per capita health spending in constant 2000 roubles, we deflate (in fact, inflate) total regional health expenditures expressed in current roubles with the non-food price index, and then divide them by the size of the regional population. Table S2 provides summary statistics of the per capita public health spending data.

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25 With regard to the sensitivity of a region's poverty ranking to the choice of welfare indicator, the analysis of the data suggests that the list of the richest and poorest regions is practically unaffected by our choice of measuring poverty in terms of per capita money income or in terms of per capita disposable resources.

26 The overall and within (over time) standard deviations are calculated for all 535 observations. The between (across regions) standard deviation is calculated over the means ($\overline{x}_i$) for 77 regions. Note that the *within* component can be negative since it is defined as ($x_{it} - \overline{x}_i + \overline{\overline{x}}$), where the global mean $\overline{\overline{x}}$ is added back to make results comparable.

27 In weighting a sample using regional population size, we take into account the fact that the *number* of people affected by poverty differs in each region.

28 In principle, public health spending also includes expenditure from the (non-budgetary) Health Insurance Fund. This expenditure (which constitutes on average about 18 per cent of the total) is not included in our measure of public health spending (which accounts for only budgetary expenses on health) due to the lack of region-disaggregated data.

Table S2
Summary statistics of the public health spending data, 1994-2000

Variable		Mean	Std dev.	Min	Max	Observations
hexp_pc	overall	880.322	538.388	238.707	4502.601	N = 539
	between		442.772	412.009	2934.507	n = 77
	within		309.841	-587.214	2448.415	T = 7

Note: hexp_pc – public health spending per capita (in 2000 prices).

Figure 6
Disposable resources poverty (headcount index) in Russia's regions, 1998

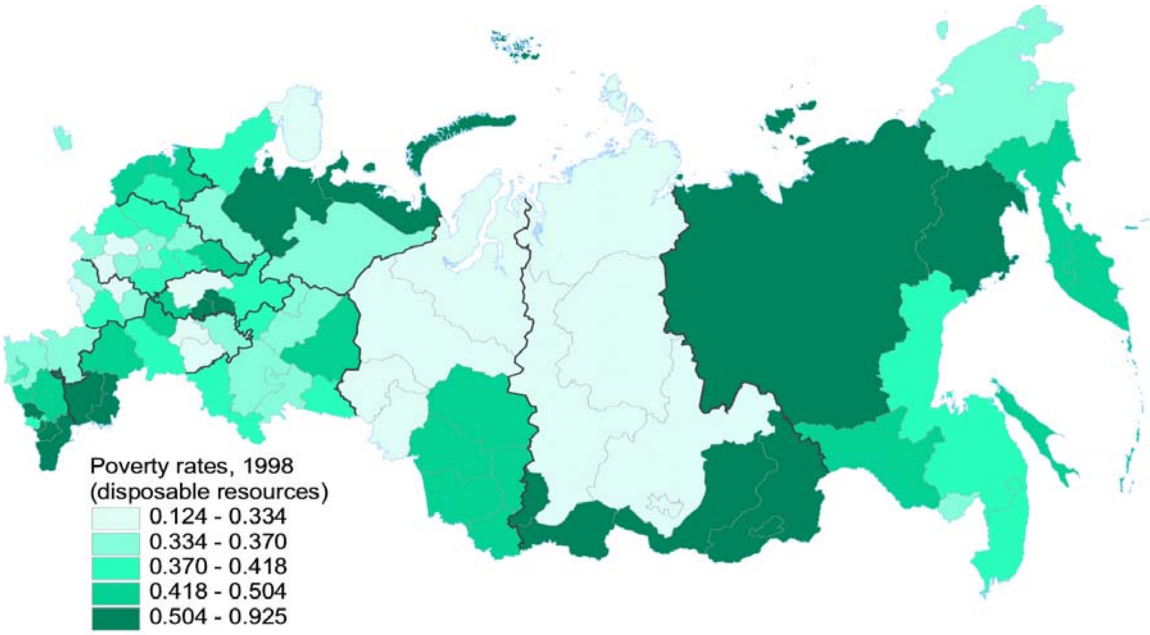

Figure 7
Public health spending per capita in Russia's regions, 1998

The table suggests that the real value of per capita public health expenditure varies considerably not only across regions, but also over time. The distribution of public health spending across Russia's regions is very unequal, as can be seen from the public health expenditure map in Figure 7.

The public health expenditure per capita in the region with the best publicly-financed health system exceeds the expenditure in the worst-funded region by as much as 19 times.

Finally, data on our dependent variable, life expectancy at birth, come from the *Regions of Russia 2001* publication of the Goskomstat.[29] This publication represents the most comprehensive source of socioeconomic information on the administrative regions of Russia. Table S3 shows summary statistics of the life expectancy data used in the empirical analysis.

The comparison of *within* (over time) and *between* (across regions) standard deviations indicates that, although life expectancy variation across regions dominates life expectancy variation over time, the latter is still quite substantial.

Table S3
Summary statistics of the life expectancy data, 1994-2000

Variable		Mean	Std dev.	Min	Max	Observations
le_all	overall	65.493	2.434	55.310	73.010	N = 537
	between		2.185	56.126	72.448	n = 77
	within		1.169	58.837	68.404	T = 7
le_m	overall	59.460	2.569	49.010	68.570	N = 537
	between		2.237	50.304	66.970	n = 77
	within		1.325	53.578	62.795	T = 7
le_f	overall	72.116	2.089	60.890	79.030	N = 537
	between		1.946	63.079	77.838	n = 77
	within		0.823	64.637	74.203	T = 7

Note: le_all – overall LE, le_m – male LE, le_f – female LE.

6 Estimation results

The results of estimating equation (6) using the Arellano and Bond (1991) GMM estimator are given below.[30] The reported results refer to the one-step heteroscedasticity-corrected GMM estimates.[31] Since consistency of the Arellano-Bond

[29] It is worth mentioning that even health outcome data cannot always be consistently compared across countries. For instance, until only recently the child mortality rate in Russia was measured differently than the standard western method.

[30] Note that a number of studies have investigated the small sample properties of the dynamic panel data estimators (e.g., Arellano and Bond 1991; Kiviet 1995; Judson and Owen 1999). While finding a (negative) bias on the autoregressive parameter, these studies indicate that GMM is virtually unbiased as far as the β vector of parameters is concerned.

[31] We have also used the two-step estimator, but since two-step standard errors tend to be downward-biased in small samples, we follow Arellano and Bond (1991) in using one-step results for inference on the coefficients.

GMM estimator is based on the assumption of no second-order autocorrelation in the first-differenced residuals, the results of the Arellano-Bond test for first and second-order autocorrelation are reported for each regression.[32] The estimation results for overall (male plus female) life expectancy are shown in column 1 of Table 1.

The estimation results suggest that current life expectancy is related positively to the past values of life expectancy and public health spending per capita, and negatively to the contemporaneous incidence of poverty. All parameter estimates are highly statistically significant. The estimated coefficients suggest that a 1 per cent increase in the incidence of poverty would be associated with a 0.1 month decline in life expectancy at the mean, while a 1 per cent increase in public health spending per capita would result in a 0.17 month rise in longevity. Interpreting the parameter estimates in an alternative way, we find that a 1 percentage point increase in poverty at the mean would lower longevity at the mean by 0.3 months, and that an additional 100 roubles (at year 2000 prices) in annual public health spending per capita would raise longevity at the mean by two months.[33]

Table 1
Estimates from the life expectancy equation

Dependent variable	$Ln(LE_{rt})$	$Ln(LE_{rt})$ (male)	$Ln(LE_{rt})$ (female)
Explanatory variable	Coefficient (std. error) 1	Coefficient (std. error) 2	Coefficient (std. error) 3
$Ln(LE_{rt-1})$	0.595***	0.671***	0.336***
	(0.063)	(0.048)	(0.082)
$Ln(P_{rt})$	-0.013***	-0.021***	-0.005*
	(0.005)	(0.005)	(0.005)
$Ln(HE\_public_{rt-1})$	0.022***	0.027***	0.013***
	(0.002)	(0.003)	(0.002)
Constant	-0.001	-0.001	0.001*
	(0.001)	(0.001)	(0.000)
No. of observations	382	382	382
No. of groups	77	77	77
Wald chi2(3)	447.81	640.00	156.05
AB test (ρ1=0):			
Prob. > z	0.003	0.000	0.000
AB test (ρ2=0):			
Prob. > z	0.638	0.933	0.520

Note: AB test (ρ1=0) [ρ2=0] refers to the Arellano-Bond test that average autocorrelation in residuals of order 1 is 0 [of order 2 is 0].

***, ** ,* indicate significance at 1, 5, 10% levels, respectively.

Source: Author's estimates.

32 Note that even if the residuals in the levels model (equation 5) are not autocorrelated, expressing the model in first differences is likely to induce AR(1) processes. We have also tested the validity of the over-identifying restrictions using the two-step Sargan test (Sargan 1958). The results of the Sargan test are not shown here since they lead to the same conclusions in all cases as the Arellano-Bond test of autocorrelation.

33 Note that 100 Russian roubles at year 2000 prices are equal to about USD 4.

We discuss next the results of estimating the male life expectancy regression (column 2 of Table 1). The parameter estimates suggest that a 1 percentage point increase in poverty at the mean would lower male longevity (at male-specific mean longevity) by 0.44 months, and that an additional 100 roubles in annual public health spending per capita (12 per cent increase at the mean) would raise longevity by 2.24 months.

The results of estimating female life expectancy regression are shown in column 3 of Table 1. As before, the parameter estimates are very significant statistically, but seem to be lower in magnitude than those for males. The coefficients suggest that for females a one-percentage point increase in poverty at the mean would lower longevity by 0.12 months, and that an additional 100 roubles in annual public health spending per capita would raise longevity by 1.36 months.

Next, we test the hypothesis that the effect of public expenditure on health in the region increases in the incidence of poverty. In other words, we expect that the impact of public spending is larger when private expenditure is not adequate for maintaining good health. We test this hypothesis by adding the interaction term between the lagged public health spending and lagged poverty into our baseline regression equation. The regression results are presented in Table 2.

Table 2
Estimates from the life expectancy equation (with interaction term)

Dependent variable	$Ln(LE_{rt})$	$Ln(LE_{rt})$ (male)	$Ln(LE_{rt})$ (female)
Explanatory variable	Coefficient (std. error) 1	Coefficient (std. error) 2	Coefficient (std. error) 3
$Ln(LE_{rt-1})$	0.625***	0.717***	0.345***
	(0.073)	(0.060)	(0.088)
$Ln(P_{rt})$	-0.012***	-0.020***	-0.004*
	(0.004)	(0.005)	(0.002)
$Ln(HE\_public_{rt-1})$	0.018***	0.022***	0.012***
	(0.002)	(0.004)	(0.002)
$Ln(HE\_public_{rt-1})*Ln(P_{rt-1})$	0.002***	0.002***	0.001**
	(0.000)	(0.000)	(0.000)
Constant	-0.001	-0.002**	0.001
	(0.001)	(0.001)	(0.001)
No. of observations	382	382	382
No. of groups	77	77	77
Wald chi2(4)	449.21	587.26	158.26
AB test (ρ1=0):			
Prob. > z	0.005	0.000	0.000
AB test (ρ2=0):			
Prob. > z	0.900	0.368	0.519

Note: AB test (ρ1=0) [ρ2=0] refers to the Arellano-Bond test that average autocorrelation in residuals of order 1 is 0 [of order 2 is 0].

***, **, * indicate significance at 1, 5, 10% levels, respectively.

Source: Author's estimates.

Table 3
Estimates from the life expectancy equation (with a time trend)

Dependent variable	$Ln(LE_{rt})$ (male)	$Ln(LE_{rt})$ (female)	$Ln(LE_{rt})$ (male)	$Ln(LE_{rt})$ (female)
Explanatory variable	Coefficient (std. error) 1	Coefficient (std. error) 2	Coefficient (std. error) 3	Coefficient (std. error) 4
$Ln(LE_{rt-1})$	0.462*** (0.058)	0.375*** (0.104)	0.457*** (0.065)	0.367*** (0.100)
$Ln(P_{rt})$	-0.012** (0.005)	-0.001 (0.004)	-0.016*** (0.005)	-0.002 (0.003)
$Ln(HE\_public_{rt-1})$	-0.007 (0.008)	0.012*** (0.002)	-0.012 (0.008)	0.011*** (0.002)
$Ln(HE\_public_{rt-1})*Ln(P_{rt-1})$	–	–	0.002** (0.001)	0.001 (0.001)
Pre-crisis year (dummy)	0.038*** (0.009)	0.019** (0.008)	0.038*** (0.009)	0.018*** (0.004)
Crisis year (dummy)	*ref.*	*ref.*	*ref.*	*ref.*
Post-crisis year (dummy)	-0.013*** (0.003)	-0.002 (0.003)	-0.016*** (0.003)	-0.003 (0.002)
Constant	0.008*** (0.002)	0.004** (0.002)	0.003** (0.001)	0.002** (0.001)
No. of observations	382	382	382	382
No. of groups	77	77	77	77
Wald chi2	1001.63	267.22	931.61	199.69
AB test (ρ1=0):				
Prob. > z	0.000	0.000	0.000	0.000
AB test (ρ2=0):				
Prob. > z	0.483	0.970	0.879	0.957

Note: AB test (ρ1=0) [ρ2=0] refers to the Arellano-Bond test that average autocorrelation in residuals of order 1 is 0 [of order 2 is 0].
***, **, * indicate significance at 1, 5, 10% levels, respectively.

Source: Author's estimates.

The results support the idea that the impact of public health spending on longevity is larger for those regions experiencing a higher incidence of poverty. Taking into consideration the interaction effect, the elasticity of longevity with respect to public health spending suggests that at the mean of poverty an additional 100 roubles in annual public health spending per capita would raise life expectancy by 2.7 months for men, and 1.5 months for women. For a given region, the interaction effects imply that the impact of public expenditure on health increases as more people in a region become poor.

As the next step of our empirical analysis we estimate the regression equation that includes time dummies among controls since we want to ensure that it is not simply the similar time trends in the dependent and explanatory variables driving the results. It is worth mentioning that at the very end of 1998, a financial crisis struck Russia. The consequences of the crisis, such as the declining real incomes due to rapidly growing prices, have been felt by the population mostly during 1999. Hence, the inclusion of the time dummies for the pre-crisis, crisis (1999) and post-crisis (2000) periods should allow us, in addition to simply controlling for a time trend, to see whether the financial crisis affected longevity through mechanisms other than the rising poverty and declining real value of public health spending.[34] The estimation results from the specifications which include the period dummies (for male and female life expectancies, with and without interaction terms) are given in Table 3.

The inclusion of a time trend into the male life expectancy regression only makes the coefficient on poverty remain statistically significant. It also reduces the magnitude of the parameter estimate on poverty. In the female life expectancy regression the parameter estimate on poverty becomes insignificant. Nevertheless, the coefficient on public health spending remains highly significant, and of the same size as in the specification without time dummies.

The parameter estimates reported in column 1 of Table 3 suggest a long-run elasticity of male longevity with respect to poverty of -0.022. This means that a *permanent* 1 per cent increase in the regional incidence of poverty would lead to a 0.022 per cent reduction in male life expectancy. Expressed alternatively, this elasticity suggests that a 1 *percentage point* increase in poverty at the mean would lower male longevity at the mean by half a month. The coefficient on health spending in column 2 (Table 3) indicates the *long-run* elasticity of female life expectancy with respect to health expenditure of 0.02. This elasticity translates into 1.94 months decline in longevity with a 100 roubles decline in public health spending.

We next estimate the specification of the equation that includes an interaction term between poverty and health spending along with period dummies. The estimates from these regressions are reported in columns 3 and 4 of Table 3. The levels of poverty and public health spending remain significant determinants of longevity for men and women, respectively. Nevertheless, the interaction term is found to be significant for men only.[35]

It is worth mentioning that the coefficients on the time dummies clearly suggest that the financial crisis in Russia had severe demographic costs beyond those associated with rising poverty and reduced public health spending triggered by the crisis. Such 'indirect' costs could be related, for instance, to an increased level of stress induced by the crisis. The period dummies indicate that, controlling for the poverty and health spending effects, life expectancy in 1999 was 3.9 per cent and 1.9 per cent lower than in the pre-crisis years for men and women, respectively. It continued to decline for the former group during 2000.[36]

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[34] We also estimated the model with time dummies for separate years, but since the coefficients on the dummies for the pre-1999 years are found *not* to be statistically different, we consider the year effects to be the same prior to 1999, and report the results in a more convenient form for interpretation.

[35] The insignificance of the interaction term for women could be due to its high correlation with the current period's poverty (ρ=0.61), and with the lagged value of public health spending (ρ=0.38), which enters the interaction term.

[36] Note that the results of the Arellano-Bond test of second-order autocorrelation in the first-differenced residuals for all specifications considered above indicate that the model is well-specified (i.e., the over-identifying restrictions are valid).

7 Robustness of results

This section investigates the robustness of our estimates to alternative specifications and the treatment of explanatory variables as predetermined rather than strictly exogenous. We start from the discussion of the choice of a functional form. A number of studies (Anand and Ravallion 1993; Kakwani 1993; Pritchett and Summers 1997) argue that the double log functional form for the life expectancy regression may not provide the best fit for the data since it imposes a constant elasticity. In fact, life expectancy is effectively bounded from above, which suggests a growing effort to achieve the same absolute gain in life expectancy for a country where life expectancy is initially higher. Clearly, if regions had fairly similar levels of life expectancy, the choice of the functional form would not be expected to significantly affect the results. However, the sheer diversity in longevity across the Russian regions is likely to render the functional form important.

To address a possible non-linearity in the relationship between longevity and its determinants, we alternatively use in the estimations a nonlinear transformation of life expectancy of the form proposed by Anand and Ravallion (1993). When one considers that in the sample the maximum life expectancy at birth equals 68.1 years for men and 78.6 years for women, the choice of a bound that equals respectively 70 and 80 years

Table 4
Estimates from the life expectancy equation (with a time trend and alternative dependent variable)

Dependent variable	*Ln(max-LE$_{rt}$)* (male, max=70)	*Ln(max-LE$_{rt}$)* (male, max=80)	*Ln(max-LE$_{rt}$)* (female, max=80)
Explanatory variable	Coefficient (std. error) 1	Coefficient (std. error) 2	Coefficient (std. error) 3
Ln(max-LE$_{rt-1}$)	0.666*** (0.102)	0.511*** (0.063)	0.362** (0.068)
Ln(P$_{rt}$)	0.071** (0.029)	0.035** (0.014)	0.020 (0.036)
Ln(HE_public$_{rt-1}$)	-0.004 (0.061)	0.013 (0.024)	-0.115*** (0.018)
Pre-crisis year (dummy)	-0.183*** (0.064)	-0.106*** (0.024)	-0.081 (0.078)
Crisis year (dummy)	*ref.*	*ref.*	*ref.*
Post-crisis year (dummy)	0.017 (0.030)	0.030*** (0.010)	-0.007 (0.020)
Constant	-0.024 (0.016)	-0.020*** (0.005)	-0.017 (0.014)
No. of observations	382	382	382
No. of groups	77	77	77
Wald chi2(6)	1067.46	1082.28	230.91
AB test (ρ1=0):			
Prob. > z	0.002	0.001	0.001
AB test (ρ2=0):			
Prob. > z	0.769	0.922	0.133

Note: AB test (ρ1=0) [ρ2=0] refers to the Arellano-Bond test that average autocorrelation in residuals of order 1 is 0 [of order 2 is 0].

***, **, * indicate significance at 1, 5, 10% levels, respectively.

Source: Author's estimates.

appears to be quite natural. The estimation results for the regressions that use ln*(bound-actual LE)* as a dependent variable are presented below. Note that since a dependent variable effectively measures a *shortfall of life expectancy* from the achievable maximum, a negative sign of the coefficient should be interpreted as a positive effect (life expectancy increases), while a positive sign should be considered as indicative of a negative effect (life expectancy declines). To make the exposition shorter, we discuss next only the estimates obtained from the specifications that include time dummies, since these specifications deliver more conservative estimates. The model which does not include an interaction term between poverty and public health spending is estimated first. The regression for men is estimated using 70 and 80 years as alternative values of the upper bound for the dependent variable in order to check the robustness of the results to the choice of the maximum level of longevity, while the regression for women uses only the dependent variable with 80 years as the upper bound. The parameter estimates from these alternative longevity (in fact, a shortfall from the maximum longevity) equations are shown in Table 4.

The regression results indicate that the coefficients on poverty for men, and health spending for women, are very significant statistically, and much larger in magnitude compared to their predecessors in the specification that uses ln(LE) as a dependent variable. The parameter estimate on poverty reported in column 1 of Table 4 implies a *long-run* poverty elasticity of male life expectancy of 0.21. This elasticity suggests that at the mean of male life expectancy, a permanent 10 per cent increase in the incidence of poverty in a region would lead to a 15.2 month decline in male longevity.

The results of the estimations that use ln(80-LE) instead of ln(70-LE) as a dependent variable for men are reported in column 2, Table 4. As before, poverty is found to have an impact on life expectancy. However, a change of the upper bound for longevity from 70 to 80 years halved the magnitude of the coefficient. The lower magnitude of the parameter estimate may reflect the fact that the choice of 80 years as the maximum achievable level of longevity for men is likely not very appropriate; male life expectancy in Russia did not exceed 70 years anywhere in the country during the period under consideration.

The estimation of the similar specification for women produces the results reported in column 3 of Table 4. The regression results for women indicate, as before, that only public health spending appears to affect longevity.[37] The parameter estimate suggests a *long-run* elasticity of female life expectancy with respect to health spending of -0.19. At the mean longevity for women, this elasticity indicates that a 10 per cent increase in per capita public health expenditure would lead to an almost 1.5 years increase in life expectancy. Noteworthy in these specifications is the fact that the ‘crisis’ effect is statistically significant only for men.

When the interaction term between the lagged value of the poverty headcount index and lagged public health spending is added to the specifications reported in Table 4, we find the interaction effect to be significant for men only.[38] As before, the sign of the

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[37] It is worth noting that the observed differential impact of poverty for men and women confirms the results of other studies for Russia (e.g., Brainerd and Varavikova 2001; Shkolnikov and Cornia 2000).

[38] These results are not shown here for the sake of brevity, but are available from the author on request.

interaction term indicates that public health spending has a greater (positive) effect on longevity at high levels of poverty.

We next estimate the model treating poverty and public health spending as predetermined rather than exogenous variables. This is a more relaxed assumption because the feedback effects from the lagged dependent variable (or lagged errors) to current and future values of the explanatory variables are not ruled out. Hence, with this assumption we effectively allow, for instance, that individual incentives for generating incomes (and related to these regional poverty rates) could be affected by past observations of longevity. The regression results for male and female longevity, with and without interaction terms, are shown in Table 5.

In the life expectancy equations for men the parameter estimates on poverty are highly significant, and notably larger than their counterparts from the specification that considers poverty as strictly exogenous rather than predetermined (see Table 4). The coefficient on poverty shown in column 1 of Table 5 implies a *long-run* poverty elasticity of 0.47. It is worth noting that treating poverty as predetermined makes it significant in also explaining female longevity (see column 3 of Table 5). The estimated long-term poverty elasticity for women indicates that a 1 percentage point increase

Table 5
Estimates from the life expectancy equation (treating explanatory variables as predetermined)

Dependent variable	$Ln(max\text{-}LE_{rt})$ (male, max=70)	$Ln(max\text{-}LE_{rt})$ (male, max=80)	$Ln(max\text{-}LE_{rt})$ (female, max=80)
Explanatory variable	Coefficient (std. error) 1	Coefficient (std. error) 2	Coefficient (std. error) 3
$Ln(max\text{-}LE_{rt\text{-}1})$	0.620***	0.493***	0.413***
	(0.097)	(0.058)	(0.083)
$Ln(P_{rt})$	0.180***	0.075***	0.128**
(*predetermined)*	(0.055)	(0.021)	(0.050)
$Ln(HE\_public_{rt\text{-}1})$	0.007	0.031	-0.103***
(predetermined)	(0.104)	(0.041)	(0.021)
Pre-crisis year (dummy)	-0.157	-0.108***	-0.087
	(0.109)	(0.041)	(0.160)
Crisis year (dummy)	*ref.*	*ref.*	*ref.*
Post-crisis year (dummy)	0.053***	0.042***	0.010
	(0.026)	(0.010)	(0.017)
Constant	-0.029	-0.023***	-0.007*
	(0.020)	(0.007)	(0.004)
No. of observations	382	382	382
No. of groups	77	77	77
Wald chi2	830.01	956.44	234.32
AB test (ρ1=0):			
Prob. > z	0.000	0.000	0.000
AB test (ρ2=0):			
Prob. > z	0.249	0.392	0.085

Note: AB test (ρ1=0) [ρ2=0] refers to the Arellano-Bond test that average autocorrelation in residuals of order 1 is 0 [of order 2 is 0].

***, **, * indicate significance at 1, 5, 10% levels, respectively.

Source: Author's estimates.

in poverty at the mean would reduce longevity at the mean by 2.9 months. Although still significant at a 1 per cent level, the coefficient on health spending for women is somewhat lower than its predecessor in the specification that treats this variable as strictly exogenous. Nevertheless, the estimation results for women reported in Table 5 have to be treated with some degree of caution since in this case a zero hypothesis of no second-order autocorrelation in the residuals cannot be rejected at a 10 per cent level.[39]

To sum up, the regression results reported in this section of the paper support the finding that the changes in poverty and public health spending have considerable impact on the changes in life expectancy. Treating longevity as bounded from above, and relaxing the assumption of strict exogeneity of the explanatory variables, increases the estimated impacts of poverty and inequality on life expectancy.

8 Conclusions

This paper aims to establish the causal impact of poverty and public health spending on life expectancy using a unique panel data covering 77 Russian regions over the period 1994-2000. The use of regional-level data from a single country overcomes the problem of data comparability which is often faced in studies that rely on cross-country data. The determination of longevity is modelled as a dynamic process. The estimation of the model is performed using the Arellano-Bond (1991) GMM estimator. We take advantage of the panel nature of the data in addressing the issue of endogeneity of the lagged dependent, as well as of the other explanatory variables, and controlling for the region-specific fixed effects which result in the omitted variable bias when the cross-section data are used. The robustness of the empirical results to several alternative model specifications and estimation assumptions is tested.

The results presented in this paper suggest that a reduction in the regional incidence of absolute poverty, and an increase in regional public investments in health generally have positive impacts on longevity. Nevertheless, the magnitudes of the estimated effects vary substantially depending on the model specification, and whether poverty and public health spending are treated as strictly exogenous or predetermined variables.

In accordance with the findings of several other studies (e.g., Brainerd and Varavikova 2001; Shkolnikov and Cornia 2000), the results of this paper indicate that the incidence of poverty in the region has a greater effect on the life expectancy of men. The coefficient on poverty is always significant (at a 1 per cent level) for men, but the estimated short-term poverty elasticity of life expectancy varies from -0.01 to -0.18

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[39] We have also estimated the static specification of the model using the random-effects GLS estimator, which represents a weighted average of the *between* and *within* estimators, and thus takes into account the variation in life expectancy across regions and over time (the estimation procedure also allowed the disturbance term to be first-order autoregressive). The estimation results (not shown here for brevity, but are available from the author on request) indicate that both poverty and public health spending are statistically significant (at a 1 per cent level) determinants of longevity for men, while only the latter factor is significant (also at a 1 per cent level) determinant of longevity for women. The interaction term between poverty and health spending is found to be significant (at a 1 per cent level), and have an expected sign, for both samples. The magnitude of the coefficients reported in the static specification cannot be compared directly to those reported in the main body of the paper since the lagged dependent variable in this case does not enter the regression equation.

depending on the model specification. The parameter estimate on poverty is insignificant in most of the specifications for women, and, when significant, varies from -0.01 (significant at a 10 per cent level) to -0.12 (significant at a 5 per cent level).[40] Therefore, the estimated poverty elasticity of life expectancy is generally in the range of estimates reported by other studies, which vary from -0.04 in Carrin and Politi (1995) to -0.21 in Anand and Ravallion (1993) and Calfat (1996).

In contrast to the effect of poverty, the impact of public health spending is found to be stronger for women. The parameter estimate on health spending in the estimations for men is insignificant in most cases, and, when significant, varies from 0.02 to 0.03 (both significant at a 10 per cent level). The parameter estimate on health spending is always significant (at a 1 per cent level) for women, and the estimated short-term elasticity of life expectancy with respect to public health spending varies from 0.01 to 0.12 depending on specification. This is generally within the range found in some other studies that use longitudinal data—from 0.04 in Lichtenberg (2002) to 0.13 in Bidani and Ravallion (1997), but lower than the 0.3 reported in cross-country studies by Anand and Ravallion (1993). That our estimates are lower than those based on cross-section data are likely due to the fact that we are able to deal better with the endogeneity of the explanatory variables, and that the estimator used in this study does not employ the cross-sectional variation in data.

The estimation results of this paper support the idea that the impact of public health spending on longevity is greater in those regions experiencing a higher incidence of poverty. In other words, the importance of publicly provided health care increases when the private resources that can be allocated to health care become more scarce. This confirms the finding of the study by Bidani and Ravallion (1997), which is based on cross-country data.

The finding of the substantial *long-run* effects of poverty and public health spending on longevity suggests that a permanent negative shock to the incidence of poverty and/or the amount of publicly-provided health care in a region results in enduring consequences for the health of the population. We also find a noticeable adverse impact of the financial crisis that erupted in Russia at the end of 1998 on the life expectancy of the country's population, even after controlling for the close association of the crisis with changes in poverty and the real value of public health spending.

The findings presented in this paper are significant from a policy perspective since they emphasize the need for stimulating regional economic development, and enhancing the health care provision in Russia. Importantly, the results indicate that measures aimed at the reduction in the incidence of regional poverty can be effective for the improvement of population health, especially under conditions when public provision of health services cannot be easily extended.

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40 The results of the estimates that consider life expectancy as a bounded variable should perhaps be considered as the most plausible.

Annex

Table A1
Life expectancy in Russian regions, 1990-2000

Region	Year	No. of observations (oblasts)	Mean	Std dev.	Min	Max
North	**1990**	**5**	**69.47**	**0.65**	**68.52**	**70.27**
	1991	5	69.14	0.92	68.30	70.60
	1992	5	66.77	1.01	65.73	68.21
	1993	5	63.85	1.19	62.51	65.22
	1994	5	62.27	1.16	61.07	63.80
	1995	5	62.87	1.32	61.24	64.26
	1996	5	64.78	0.77	63.78	65.64
	1997	5	66.32	0.95	65.59	67.89
	1998	5	66.74	1.23	65.70	68.85
	1999	5	65.30	1.52	63.78	67.47
	2000	5	64.70	1.28	63.31	66.13
North-West	**1990**	**4**	**68.80**	**0.95**	**67.87**	**70.12**
	1991	4	68.18	1.00	67.30	69.60
	1992	4	66.51	1.21	65.74	68.31
	1993	4	63.10	0.81	62.52	64.29
	1994	4	62.01	1.70	60.70	64.50
	1995	4	63.18	1.96	62.14	66.12
	1996	4	65.26	2.14	63.60	68.35
	1997	4	66.26	2.16	64.59	69.36
	1998	4	66.22	2.26	64.17	69.33
	1999	4	64.33	2.22	62.26	67.47
	2000	4	63.42	1.80	61.92	66.03
Central	**1990**	**13**	**69.50**	**0.42**	**68.66**	**70.18**
	1991	13	69.02	0.54	68.20	69.90
	1992	13	68.04	0.66	66.74	68.81
	1993	13	65.50	0.80	63.38	66.43
	1994	13	63.97	0.90	62.03	65.39
	1995	13	64.79	0.91	63.31	66.57
	1996	13	65.97	0.85	64.73	67.46
	1997	13	66.38	0.89	65.17	68.25
	1998	13	66.59	0.92	65.19	68.46
	1999	13	65.03	1.28	63.39	68.05
	2000	13	64.44	1.39	62.68	67.81
Volgo-Vyatka	**1990**	**5**	**70.08**	**0.64**	**69.25**	**70.86**
	1991	5	69.74	0.63	68.80	70.40
	1992	5	68.85	0.72	67.83	69.55
	1993	5	66.63	0.82	65.94	67.85
	1994	5	65.46	1.07	64.47	66.72
	1995	5	65.81	1.31	64.47	67.71
	1996	5	67.18	0.73	66.46	68.24
	1997	5	67.31	0.53	66.75	67.95
	1998	5	67.87	0.89	67.03	68.85
	1999	5	66.51	0.78	65.38	67.41
	2000	5	65.92	0.90	64.95	66.96

Table continues

Table A1 (con't)
Life expectancy in Russian regions, 1990-2000

Region	Year	No. of observations (oblasts)	Mean	Std dev.	Min	Max
Central-Chernozem	**1990**	**5**	**69.96**	**0.78**	**68.95**	**70.69**
	1991	5	69.18	0.77	68.30	69.90
	1992	5	68.81	0.56	68.20	69.54
	1993	5	66.96	0.53	66.55	67.55
	1994	5	66.09	0.94	65.26	67.39
	1995	5	66.83	0.89	66.02	68.16
	1996	5	67.48	0.96	66.34	68.87
	1997	5S	67.73	0.79	66.91	68.96
	1998	5	68.14	0.85	67.13	69.27
	1999	5	66.78	1.07	65.66	68.22
	2000	5	66.44	1.07	65.36	67.89
Volga	**1990**	**8**	**69.95**	**0.98**	**67.82**	**71.08**
	1991	8	69.76	0.83	68.10	70.70
	1992	8	69.14	0.66	68.01	69.99
	1993	8	66.89	0.77	65.85	68.13
	1994	8	65.86	0.69	64.85	66.87
	1995	8	66.36	0.65	65.42	67.19
	1996	8	67.18	0.73	66.14	68.20
	1997	8	67.42	0.49	66.66	68.18
	1998	8	67.57	0.89	66.35	68.77
	1999	8	66.53	0.81	65.93	68.37
	2000	8	65.85	0.93	64.48	67.54
North Caucasus	**1990**	**7**	**70.62**	**1.32**	**69.22**	**72.96**
	1991	9	70.14	1.58	68.30	72.80
	1992	9	69.85	1.45	67.87	72.25
	1993	8	68.07	1.74	65.65	70.60
	1994	8	67.53	1.88	65.16	70.61
	1995	9	68.04	2.15	65.51	71.34
	1996	9	68.44	1.63	66.66	71.51
	1997	9	68.83	1.73	67.14	72.57
	1998	9	69.01	1.37	67.58	71.91
	1999	9	68.81	1.95	66.75	73.35
	2000	9	68.76	2.32	66.35	74.01
Urals	**1990**	**7**	**69.75**	**0.62**	**68.92**	**70.60**
	1991	7	69.36	0.56	68.60	70.10
	1992	7	68.01	0.78	67.13	68.82
	1993	7	65.22	0.97	64.04	66.33
	1994	7	64.05	1.27	62.00	65.48
	1995	7	64.69	1.09	63.10	66.39
	1996	7	65.92	0.91	64.45	67.23
	1997	7	66.78	0.74	65.37	67.70
	1998	7	67.05	0.65	66.13	67.88
	1999	7	65.85	0.83	64.67	67.02
	2000	7	65.01	1.07	63.74	66.76

Table continues

Table A1 (con't)
Life expectancy in Russian regions, 1990-2000

Region	Year	No. of observations (oblasts)	Mean	Std dev.	Min	Max
Western Siberia	**1990**	**6**	**68.99**	**0.70**	**68.08**	**69.78**
	1991	7	68.21	1.35	65.50	69.60
	1992	7	67.03	1.79	63.79	69.06
	1993	7	64.03	1.71	61.45	66.47
	1994	7	63.19	2.01	59.95	65.67
	1995	7	63.90	1.94	61.09	66.51
	1996	7	64.75	1.78	62.39	67.12
	1997	7	65.49	1.82	62.58	67.11
	1998	7	66.66	1.66	64.14	68.17
	1999	7	65.75	1.81	63.15	67.69
	2000	6	65.73	1.38	63.19	66.64
East Siberia	**1990**	**5**	**66.62**	**2.36**	**62.43**	**67.90**
	1991	6	66.28	2.48	61.30	67.90
	1992	6	64.91	2.14	60.69	66.67
	1993	6	61.71	1.97	57.89	63.26
	1994	6	59.90	2.33	55.31	61.71
	1995	6	61.24	2.80	55.72	63.43
	1996	6	61.95	3.24	55.38	63.79
	1997	6	62.80	3.30	56.08	64.63
	1998	6	63.35	2.57	58.25	65.27
	1999	6	61.78	2.88	56.00	63.50
	2000	6	61.70	2.80	56.14	63.81
Far East	**1990**	**8**	**67.47**	**0.91**	**66.08**	**69.04**
	1991	8	67.23	0.72	66.40	68.70
	1992	8	65.63	0.96	64.16	67.31
	1993	8	62.76	1.20	60.38	64.41
	1994	8	61.81	1.08	60.34	62.98
	1995	8	61.96	2.93	55.34	64.81
	1996	8	63.51	1.18	61.96	65.85
	1997	8	64.72	0.79	63.40	65.85
	1998	8	65.22	0.76	64.20	66.63
	1999	8	64.48	0.53	63.94	65.38
	2000	8	63.99	0.61	63.12	65.00

Source: Author's calculations using data from the *Regions of Russia, 2001* publication.

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