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Research Article

Urban–rural disparities in adult mortality in sub-Saharan Africa

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Contents

Urban–rural disparities in adult mortality in sub-Saharan Africa

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Abstract

BACKGROUND

Empirical evidence showing higher survivorship in urban areas of sub-Saharan Africa (SSA) supports a theory of rural disadvantage. Yet this evidence mostly builds on infant or child mortality. There is practically no empirical evidence comparing adult mortality levels across urban and rural sectors.

OBJECTIVE

This study explores adult mortality differences by urban–rural residence across SSA for men and women. It considers whether existing differences across sectors vary over the course of development.

METHODS

The indirect orphanhood method is applied to 90 Demographic and Health Survey (DHS) datasets from 30 countries between 1991 and 2014. Conditional probabilities of dying between ages 15 and 60 (${}_{45}q_{15}$) are separately estimated for rural and urban populations.

RESULTS

Based on country averages over all time periods, the mean ${}_{45}q_{15}$ is 0.274 and 0.265 among adult women and 0.307 and 0.292 among adult men in urban and rural populations, respectively. The average urban to rural probability of dying ratio from the most recent data between 2000 and 2010 is 1.08 for females and 1.11 for males in SSA as a whole, indicating an urban penalty. Multiple checks highlight the robustness of our findings to methodological limitations inherent in the method. Multivariate regression models suggest that as countries develop, excess adult mortality is likely to shift from the urban to the rural sector.

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CONTRIBUTION
We provide evidence that, unlike child mortality, adult mortality remains higher in the urban sector for many countries and for SSA as a whole. This finding has policy implications regarding the spatial provision of health services. Urban poor should be given more attention.

1. Introduction

Empirical evidence from sub-Saharan Africa (SSA) consistently shows an urban mortality advantage. Mortality levels in cities appear substantially lower than those found in rural areas (Akoto and Tambashe 2002; Bocquier, Madise, and Zulu 2011; Cai and Chongsuvivatwong 2006; Fink, Günther, and Hill 2014; Gould 1998). This finding is unsurprising as it fits the general pattern predicted by the epidemiological transition where mortality rates in urban areas fall below those in rural areas after public health and sanitation systems expand and pandemics recede (Dye 2008; Omran 1971). This sectoral gradient also suggests that patterns of mortality across urban and rural SSA are broadly consistent with other regions of the world where urban mortality has been shown to be lower (Buckley 1998; Snyder 2016). Yet it turns out that the empirical foundations of the well-established rural disadvantage in mortality for SSA are primarily built on evidence from infant or child survivorship, and more occasionally on maternal mortality differences (Akoto and Tambashe 2002; Bocquier, Madise, and Zulu 2011; Fink, Günther, and Hill 2014; Günther and Harttgen 2012; Harpham 2009). For example, 25% higher rural mortality for all ages has been assumed in estimating rural to urban migration using census survival (United Nations 2001). To the best of our knowledge there is little support for this assumption when applied to adults. In fact, general adult mortality is sometimes extrapolated from estimates on child mortality due to limitations of reliable data, particularly for model life tables (Bradshaw and Timæus 2006; de Walque and Filmer 2013). Despite recognition of the value of lowering and disaggregating mortality in SSA objectives of the Sustainable Development Goals (SDG) (United Nations 2015a), empirical evidence comparing adult mortality levels across urban and rural sectors is sorely lacking (see Günther and Harttgen (2012) as an exception).

On the one hand, in spite of a dire shortage of empirical evidence, there are strong reasons to expect adult mortality in rural SSA to exceed that found in the urban sector. Adults in cities should be expected to enjoy lower mortality particularly given longstanding urban bias (Lipton 1977), expanded health service provision in cities (The Lancet 2015), and better educational and economic opportunities in urban areas (Lipton

1977; Sahn and Stifel 2003). Furthermore, any direct urban health advantage should be reinforced if rural to urban migrants are positively selected for health (Lu 2008; Marmot, Adelstein, and Bulusu 1984; Nauman et al. 2015) or if urban migrants return to the rural sector due to old age or poor health (Arenas et al. 2015; Clark et al. 2007).

Yet, on the other hand, there are reasons to expect higher urban mortality with countries in SSA still undergoing a mortality transition. An ‘urban penalty’ in mortality has been identified in historic transitions (Reher 2001; Woods 2003). In fact, in countries where the transition is complete, the urban penalty has not really disappeared; rather the consequences are no longer fatal. Urban living takes its toll on people’s health due to the risks and vulnerability associated with city life (Dye 2008; Gould 1998; Reher 2001). In developing countries, urban populations, especially the poor, suffer from a ‘double burden’ of disease – both noncommunicable or chronic illnesses associated with later stages in the epidemiological transition and infectious diseases (Agyei-Mensah and Aikins 2010; Mberu et al. 2015; Soura, Lankoande, and Millogo 2014). In this article we evaluate whether SSA urban–rural adult mortality differences for men and women support the case for an urban advantage or penalty, providing a series of empirical robustness tests to further substantiate our results, given the challenges involved in measuring subnational adult mortality with limited data. To shed more light into future anticipated trends, we then consider whether mortality differences across sectors vary by region and over the course of development and urbanisation.

Our analysis builds on both the expanded availability of survey data across SSA and improvement in the methodological tools available to estimate adult mortality. In recent decades advances in indirect methods for estimating mortality make it easier to evaluate adult mortality in low income settings where vital registration is often incomplete or inaccurate (Feehan, Mahy, and Salganik 2017; Gakidou, Hogan, and Lopez 2004; Timæus 1991). These advances make it possible to assess the extent of spatial variation in adult mortality within a country in SSA. However, these methods rely on various assumptions, making the point estimates of mortality difficult to ascertain with confidence. For this reason, our main findings are buttressed by multiple tests of robustness to help determine the direction of potential biases.

Our focus is on sectoral differences in adult mortality. We show that many factors produce similar mortality biases across the urban and rural settings, generating relatively little impact on our estimated differences in adult mortality across sectors. Ultimately, our findings, which appear robust, indicate that urban conditional probabilities of dying between ages 15 and 60 (${}_{45}q_{15}$) – our measure of adult mortality – either exceed or are equal to rural probabilities for most countries in SSA.

1.1 Urbanisation, development, and adult mortality in sub-Saharan Africa

In spite of indisputable advantages associated with urban life that include superior infrastructure (Sahn and Stifel 2003), greater availability of health and educational professionals (Dussault and Franceschini 2006; McEwan 1999), political favouritism (Konadu-Agyemang and Shabaya 2005), and higher income (Lipton 1977), historical experiences of the mortality transition suggest an urban penalty in mortality (Fox 2012; Reher 2001; Woods 2003). This paradox leads us to explore rural and urban mortality in relation to urbanisation and the mortality transition. It is well recognized that both declining fertility and mortality across urban and rural sectors play a key role in urbanisation (Dyson 2011; de Vries 1990). Before the onset of the demographic transition, urban growth is highly constrained due to high mortality levels and is typically sustained by rural to urban migration. Once the conditions for mortality transition emerge, urban mortality declines first with reduced death rates from infectious diseases (Omran 1971; Woods 2003), stimulating urban population growth. The initiation of rural mortality decline increases rural cohort sizes, producing a larger pool of potential migrants, leading to increased rural to urban migration and further boosting urban populations (Zelinsky 1971). However, transition theory suggests that only after declines in fertility in both rural and urban sectors will urban mortality be lower than rural mortality (Dyson 2011; de Vries 1990). When this sequence occurs, urban rates of natural increase may exceed rural rates, potentially leading to 'autonomous urbanisation' (de Vries 1990).

Alongside this inherent link between the demographic transition and urbanisation, the demographic transition plays a key role in development (Dyson 2001). For instance, declines in fertility lead to decreased gender differentiation as women gain more independence and pursue activities unrelated to childrearing. Sustained mortality decline and urbanisation drive development as society becomes more complex, with increased division of labour and occupational specialisation, expanded transport and communication, and democratic advancement (Davis 1965; Gibbs and Martin 1962; Wilson and Dyson 2016).

Whether this relationship among the demographic transition, urbanisation, and development holds true for SSA is not clear. National mortality rates began declining in SSA at least from the 1950s (Moser, Shkolnikov, and Leon 2005), despite setbacks from the spread of HIV in the 1980s and 1990s (McMichael et al. 2004). From the 1990s fertility rates in SSA began to decline too, though following a unique pattern (Bongaarts 2016). These declines are consistent with transition theory, suggesting that urbanisation too should be well underway. Yet SSA urbanisation levels remain substantially lower than in other world regions (United Nations 2014). Furthermore, subregional differences in SSA in urbanisation levels (United Nations 2015b) and in adult mortality (McMichael et al. 2004; Wang et al. 2012) allude to regional differences in mortality transition and subsequently in urban–rural mortality differentials.

In addition, there appears to be some disassociation between transition, urbanisation, and development in SSA. Since the 1980s, urbanisation in SSA has been distinguished as virtually disconnected from economic growth (Fay and Opal 2000). The relatively weak state of industrialization common to many SSA states and the continued economic stagnation across much of the continent would seem to offer little or no basis for predicting trends in adult mortality and their differences across the urban and rural sectors. Despite this disconnect between economic development and urbanisation, other dimensions of development, such as health and living standards, have been associated with urbanisation in SSA (Njoh 2003). This association should mean lower urban mortality in countries that are more developed.

1.2 Importance of understanding urban–rural inequalities in adult mortality

Sub-Saharan Africa offers a compelling opportunity to explore whether an urban advantage in adult mortality emerges as countries urbanise and develop. Furthermore, there are extensive implications for understanding the spatial distribution of urban–rural gaps in adult mortality in developing countries. Since the most socially and economically active population is concentrated among adults between ages 15 and 60, adult deaths often impose heavy burdens on families, communities, and states in developing countries (Beegle, de Weerdt, and Dercon 2008; Dixon, Mcdonald, and Roberts 2002) – burdens that may differ by sector. Evidence indicates that insufficient resources in developing countries have focused on avoiding premature adult deaths, leading to higher adult mortality (Murray and Feachem 1990; Rajaratnam et al. 2010). In fact, the strong emphasis of development and health projects on under-five mortality may have unintentionally contributed to a relative neglect of adults. Hence, policies overcoming these tendencies by targeting premature deaths by urban–rural sector are imperative (The Lancet 2015; United Nations 2015a). Our analysis of spatial inequalities in mortality is opportune to facilitate the SDG framework of equity (United Nations 2015a), especially in SSA where urbanisation is still underway.

2. Methodology and data

Most countries in SSA lack sufficient population registration and mortality records necessary for direct estimation of adult mortality (Timæus 1991). Surveys can provide a means for estimating adult mortality in developing countries (e.g., Bendavid, Seligman, and Kubo 2011). Yet surveys that incorporate questions on recent adult deaths are typically limited by sample size (Timæus 1991), making them less useful for calculating regional or sectoral estimates. The development and refinement of indirect methods to estimate mortality, including the sibling and orphanhood methods, has proven to be a critical step (Bicego 1997; Brass 1975; Brass et al. 1981; Gakidou, Hogan, and Lopez 2004; Timæus and Jasseh 2004; Zaba, Timæus, and Ali 2001). These methods use a very limited set of questions collected from surviving relatives – children or siblings – to estimate past mortality. Both methods have been shown to be effective at capturing broad trends and levels of adult mortality used for population forecasting and resource allocation (Timæus 1991; Timæus and Jasseh 2004).

One critical advantage of the orphanhood method employed in our analysis, over the sibling survivorship method, concerns the implicit assumption of coresidency. When estimating adult mortality within a country by using the orphanhood survivorship method, mortality of parents of young children will have typically occurred while children and parents were coresident and thus sharing their urban–rural residential status. Of course, it is possible that child residence is not a good proxy for parental residence (Lankoande 2016), as children are fostered or parental mortality may lead to orphan migration – concerns we consider below. Yet despite widespread fostering, evidence shows that in many SSA countries the majority of children under 15 live with both parents or with either mother or father if orphaned (Beegle et al. 2010; Hosegood et al. 2007; Lloyd and Desai 1992). The necessary assumption of shared urban–rural status for adult siblings is more difficult to make, reducing the value of the sibling approach for estimating separate urban–rural adult mortality levels.

The orphanhood method builds on respondents' ages and their responses to questions regarding parental survival to calculate life table measures of conditional survivorship in adulthood (Brass et al. 1981; Timæus 1986, 1992, 2013). The method produces estimates of past adult mortality that are most recent when data is obtained from children aged 5 to 14. Although obtaining mortality estimates from older individuals yields more historic estimates, its main advantage is to lessen bias from the 'adoption effect' (Blacker 1984; Timæus 1991) – a bias we discuss below in detail. Bias may also occur if the parental probability of survival is related to the number of surviving children, although such biases have appeared to be small (Palloni, Massagli, and Marcotte 1984). Biases may also arise due to coverage error, since childless adults are not included. Parents necessarily show better survival, since deaths later in the reproductive life span are less likely to affect childless individuals (Luy 2012). Higher survival of parents has been found to only slightly underestimate adult mortality (Festy 1995). The bias could be magnified when estimating separate urban and rural adult mortality rates within countries. Nulliparous adults have been found to be more common in urban populations that follow the trends observed in Western societies

(Bloom and Pebley 1982; Rowland 2007; Tanturri and Mencarini 2008; Veevers 1971). In this case, lower estimated urban adult mortality may lead to a smaller urban–rural differential. However, SSA is generally characterised by relatively low levels of childlessness. On average 6.3% of women ages 25–49 have never given birth (ICF International 2014) compared to both contemporary and historical childlessness in European countries, which average 14% and between 15%–25%, respectively (Kohler, Billari, and Ortega 2002; Rowland 2007; Tanturri and Mencarini 2008). Thus, we expect minimal bias due to exclusion of childless adults, although it may become a bigger factor in the future.

Our analyses are based on nationally representative data from the Demographic and Health Surveys (DHS), which include questions on parental survivorship from household members up to age 17 in the household portion of the survey, customarily posed to household heads.[3] Applying the orphanhood method (Timæus 2013), we estimate separate national, urban, and rural adult mortality levels from 90 surveys conducted between 1991 and 2014 that cover a total of 30 SSA countries. These survey estimates are not substitutable for mortality measured through vital registration records; however, they have been shown to be effective in capturing adult mortality levels and trends in SSA, and we argue they can be used to identify urban–rural differentials (Chisumpa and Dorrington 2011; Timæus 1991).

The surveys included in our analyses were selected based on availability of data on parental survival and age of birth, for mainland SSA countries. Accordingly, 71% of all available DHS surveys between 1988 and 2014 were included. Within selected surveys, cases of unknown and nonresponses to questions on parental survival were excluded, assuming a similar distribution to respondents who answered or knew about their parental survival. Together, missing and unknown responses were slightly higher in the urban sector – 0.32% of all responses on maternal survivorship and 0.76% of responses on paternal survivorship (compared to rural proportions of 0.24% and 0.70% respectively). These differences were not statistically significant.

Since respondents' mothers were alive when respondents were born (and fathers alive at time of conception), exposure of living parents to risk of dying is the age of the respondents. Using information on the mean age at which mothers give birth in the rural and urban sectors, it is possible to predict life table survivorship from age 25 to 35, or l_{25+n}/l_{25} in standard demographic notation, where l_x represents survival probability to age x. The mean age at which men have children is estimated by adding an index (calculated from the sex difference in median ages of those currently married) to women's mean age of childbearing (Timæus 2013). Although this estimate is a rough indicator of age at childbearing, which often differs according to marital status, direct

[3] Unfortunately, the DHS does not collect parental survivorship information from adults, which would have allowed us to examine urban–rural adult mortality with reduced bias from the 'adoption effect.'

measures of this mean age by urban–rural sector are lacking. While this factor should be considered, it is not of great concern as mortality estimates are not sensitive to errors in this measure (Timæus 2013). Life table survivorship for men is from age 35 to 45, given that men are typically older than their partners. Survivorship ratios are then converted to a common index of mortality, the probability of dying between exact ages 15 and 60, conditional on survival to age 15, ${}_{45}q_{15}$. We convert the survivorship ratios to allow for comparison of estimates across age ranges. The survivorship ratios are converted by fitting a one-parameter relational logit model life table. We use the Princeton South Standard Life Table with a life expectancy of 60 as its standard.[4] The parameters for estimating survivorship are based on proportions of respondents with living mothers (Timæus 1992, 2013).The male mortality estimate requires two proportions of male survivorship, obtained from both the 5–9 and 10–14-year-olds. The female mortality estimate requires only one proportion of female survivorship, allowing for its estimation from the two age groups of orphans (5–9 and 10–14-year-olds). We use an average of the two ${}_{45}q_{15}$ estimated for women to obtain a single point estimate for each survey. Similarly, the reference date is a midperiod date for women.[5] We aggregate female estimates within individual survey rounds since time trends in mortality using the orphanhood method have been shown to poorly capture short-term mortality changes (Timæus 2013).

2.1 Potential concerns with the orphanhood method

We consider three potential concerns associated with use of the orphanhood method that may lead to bias when estimating urban–rural differences in adult mortality: a) orphan migration, b) the 'adoption effect,' and c) HIV/AIDS. Our main findings on adult mortality differences across the urban–rural sectors are supported by a series of robustness tests addressing these concerns, presented in full below. Here, we briefly present the rationale of these tests and their implications for our ability to estimate the adult mortality gap across the urban–rural sectors.

[4] We tested sensitivity to the choice of model and found that estimates using the UN general life table produced almost identical urban–rural mortality ratio estimates (with an R squared of .99). Using the UN general life table raises the level of mortality in both rural and urban settings, but the ratio between them remains constant. In addition, we tested whether using the UN general life table for the urban sector and Princeton South standard life table for the rural sector affected the urban–rural mortality ratios, and vice versa. Here too the ratios are almost identical with an R squared of .997 in both cases. Similarly, we tested sensitivity using patterns of mortality based on African surveillance sites (Clark 2002). Results indicated that these urban–rural mortality ratios are also very similar to our results (with an R squared of .98), and using different standards for the urban–rural sectors correspond to our results (R squared of .93 is obtained).

[5] A 'time location' estimates how many years prior to survey the cohort survivorship ratio equaled the period survivorship ratio.

2.1.1 Orphans may not share the urban–rural residence of their parents

Families may become separated through migration of either parents or orphans. A particular concern arises if households adjust their composition following a parental death with some members, including orphans, moving from rural to urban areas or vice versa. This mobility could affect estimates of adult mortality based on orphanhood if orphaned children are systematically shifted across sectors, as found by Lankoande (2016) in Burkina Faso. However, we propose three separate tests for upward bias of urban estimates arising from rural to urban migration of orphans and show how the direction of the bias is unlikely to lead to overestimation of urban adult mortality.

2.1.2 The 'adoption effect'

In parts of Africa high rates of fostering and adoption may lead to underestimation of adult mortality. If children are too young to remember their true parents, or cultural norms lead children to call their foster parents 'mother' or 'father' (Blacker 1984), adoption may distort mortality estimates (Blacker and Mukiza-Gapere 1988). Furthermore, the misreporting of true parental death is most pronounced for young children, often leading to an overestimation of mortality decline (Timæus 2013). It is worth noting that misreported fostering in the DHS surveys is the outcome of an adult misreporting on child household members.[6] In some circumstances this misreporting may weaken the likelihood of an adoption effect, though it may also similarly cause bias when adoptive parents want to give the impression of biological parentage or when the interviewer does not probe and makes assumptions regarding the relationship. Solving this issue directly is very challenging, but a comparison of fostering and adoption prevalence in rural and urban population in Africa can offer some insight into the stability of the urban–rural mortality gap. Ultimately, if the adoption effect functions similarly across sectors, then differences between rural and urban mortality should be robust.

2.1.3 An HIV/AIDS correction

Mortality estimates for SSA based on the orphanhood method may be biased downward in countries with widespread HIV/AIDS and limited access to treatment (Timæus and Nunn 1997). The bias derives from two main sources: HIV positive women can

[6] Children who belong to the household are included in the survey household roster whether they are related to the household head or not, and details on the children are collected.

transmit the virus to their children who are less likely to survive to report on parental death, and HIV positive women typically have lower fertility than uninfected women (Chen and Walker 2010; Fabiani et al. 2006). Thus, too few children in the population report on parental death, exaggerating the proportion of mothers alive.

While a correction for this bias has been developed for countries with moderately severe HIV epidemics (Timæus and Nunn 1997), two considerations lead us to refrain from applying the correction in our main analyses. First, the necessary data for application of the corrections is typically not available by urban–rural sector for most countries in SSA. Urban–rural data does exist for HIV prevalence, with evidence across SSA countries showing higher urban prevalence rates (Barongo et al. 1992; Dyson 2003; Killewo et al. 1990). This data suggests a particularly strong impact on urban survivorship. In effect, applying the correction in our case in these higher HIV/AIDS prevalence countries would lead to a larger urban–rural adult mortality gap. A second reason not to correct for HIV bias is the increased spread of HIV antiretroviral treatments (ART) in recent years, changing the relationship between HIV and mortality (Bor et al. 2013) and possibly rendering a correction obsolete for estimates from more recent surveys. All the same, we do verify the robustness of our findings by applying the correction (Timæus 2013) for East African countries with HIV prevalence of between 5% and 10% during the peak HIV epidemic period between 1995 and around 2005 (before the onset of widespread ART access).[7]

2.2 Urbanisation, development, and adult mortality: Multivariate analysis

We use the adult mortality estimates to probe urban–rural inequalities as countries urbanise and develop. We capture the inequality between sectors with the urban to rural ratio of the conditional probability of dying between ages 15 and 60, ${}_{45}q_{15}$.[8] Our

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[7] The HIV correction adjusts the reported proportions of mothers alive downwards by using a correction factor, revised coefficients for estimating life table survivorship from proportions of respondents with living mothers, and an adjusted model life table for moderately severe HIV/AIDS settings (with life expectancy of 50 for men and 52.5 for women as its standard) – all derived by Timæus and Nunn (1997). The correction factor is based on estimates of HIV prevalence, mother-to-child (vertical) HIV transmission rate, and the relative fertility of HIV positive women to HIV negative women. The proportion of men with infected partners is also required when adjusting male mortality. HIV/AIDS prevalence estimates by rural and urban sector were taken from the DHS and the World Health Organization. The vertical transmission rate and the relative fertility of women are considered the same in both urban and rural populations: –0.33 (Timæus 2013) and 0.6 (Chen and Walker 2010) respectively. The proportion of men with infected partners was estimated by combining data on sero-discordant couples data from five SSA countries (de Walque 2007) with data from Kenya by sector (Kaiser et al. 2011), providing average rural (0.04) and urban (0.05) estimates.

[8] An alternative approach to expressing the urban–rural inequalities in mortality is to use odds ratios. Our main results are qualitatively unchanged when using odds ratios. For all 30 SSA countries, the mean odds ratio for women is 1.07 and for men 1.14, both indicating higher urban odds of adult mortality.

multivariate analyses explore the relationship between the urban to rural ratio of adult mortality and various predictors including sex, time period, proportion of country that is urban, national development, total fertility rate, and HIV prevalence. Our analyses are based on a series of fixed effects models that account for sources of heterogeneity between countries. Both observed differences and differences that are fixed but unobserved are included. The estimates within countries further help to reduce the impact of variation in definitions of 'urban' across countries – an important limitation (Bocquier 2004). Indeed, it is important to note that each country uses its own definition of urban, and the DHS does not standardise the urban definitions. Thus, for example, the meaning of urban mortality in Liberia, where an 'urban locality' is defined as a settlement with over 2,000 inhabitants, is different to urban mortality in Senegal, where 'urban' is defined as a place with over 10,000 inhabitants. In addition, national definitions of 'urban' may change over time (although over the 25 years we examine, we expect this change is not common).[9]

3. Results

The standard orphanhood approach is used to estimate national, rural, and urban adult mortality from 1986 until 2011. Overall, the estimated adult mortality by residence indicates that the SSA mean conditional probability of dying between ages 15 and 60 for adult women is 0.274 in urban settings compared to 0.265 in rural. The mean ratio of urban to rural mortality is 1.05 for women, indicating that urban conditional probability of death is 5% greater. For adult men, the urban average conditional probability of dying is 0.307 compared to the average rural probability of dying of 0.292. The overall mean urban to rural ratio for men is 1.08. The average SSA estimates for adult mortality largely indicate higher ${}_{45}q_{15}$ in urban areas for both sexes. When considering only recent data, from 2000 to 2010, the mean SSA urban to rural mortality ratio is 1.08 for women and 1.11 for men. Weighted by national population size, the urban to rural mortality ratios remain above one – 1.04 and 1.19 for women and men respectively.[10] A detailed version of our main findings, showing the estimated conditional probabilities of dying between ages 15 and 60 and the urban to rural mortality ratio by sex, for 30 SSA countries and each round of data collection, is found in the Appendix (Table A-1).

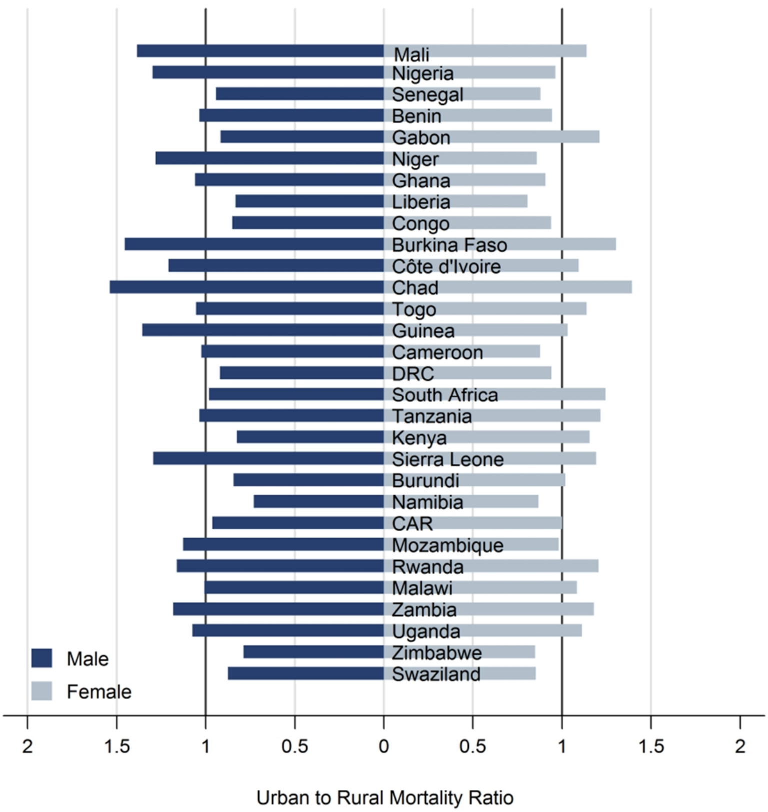
[9] We base this expectation on a comparison of the 'urban' definitions for each country gathered by the UN for their Demographic Year Books of 2005 and 2010. Of countries that feature in both books, 8% of the countries changed their definitions of 'urban' over this period.

[10] Weighting the mortality estimates leads to a larger increase in the urban to rural ratio among men than among women. This difference may be due to larger countries having greater gender inequality.

Unsurprisingly, the overall SSA average conceals considerable heterogeneity across countries. Figure 1 shows average urban to rural ${}_{45}q_{15}$ ratios for each country over all periods by sex and is sorted by national adult mortality levels from top to bottom (also estimated using the orphanhood method with DHS survey data). West African countries (which mostly feature in the upper section of the plot) have lower adult mortality, in part because they are less affected by HIV/AIDS, which has had a large impact in Southern Africa (Timæus and Jasseh 2004). According to the figure, there is an urban adult mortality disadvantage (urban to rural mortality ratios in excess of one) in more than half (60%) of the countries. Moreover, while some countries show an urban mortality advantage (urban to rural mortality ratios below unity), the estimates are more likely to lie far above, rather than below, one. In our data, there are eleven countries where male urban mortality is 10% greater than rural but only seven where rural is 10% greater than urban among men. Countries with higher rural mortality average 13% greater mortality, and countries with higher urban mortality average 15.5% greater mortality.

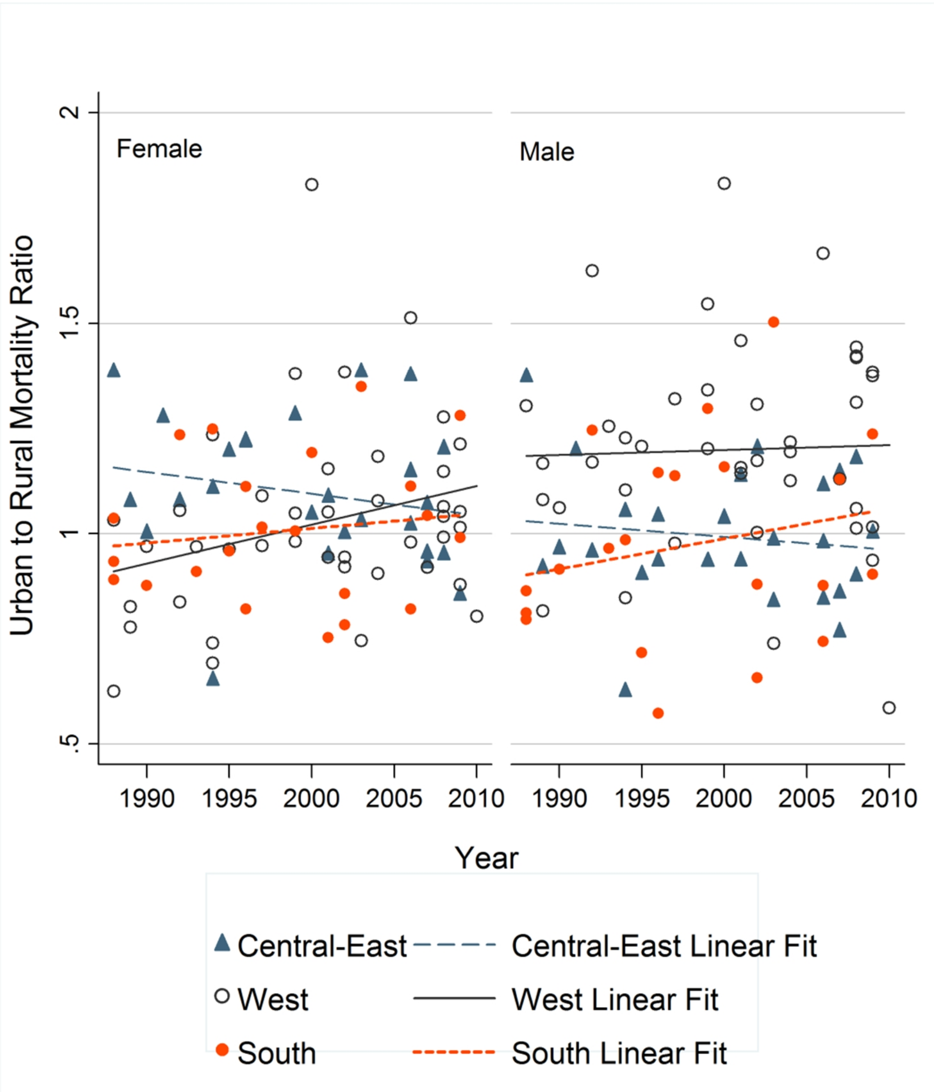
While most of our work explores SSA as a whole, regional differences have been shown to be marked. National levels of adult mortality tend to vary across regions within SSA (McMichael et al. 2004; Wang et al. 2012). In light of these differences, we further explore whether there are distinct subregional differences in urban–rural mortality within sub-Saharan Africa. In Figure 2 we examine how the subregional urban to rural mortality ratio shifts over time by sex. The figure indicates that urban female mortality is higher than rural for all subregions in later years. An urban disadvantage evolves over time in West African countries among females. In contrast, men have considerably higher urban mortality in West Africa, with a mean urban to rural ratio of 1.199 over the whole period (Appendix). In Central East Africa an opposite trend is implied, with a rural disadvantage occurring in many countries in more recent years. This trend is also evident by the male urban to rural mortality ratio over all periods in Central East Africa, 0.991. In Southern Africa the linear trend line suggests an increase in urban to rural ratios over time, shifting to higher urban mortality. These differences between regions of SSA in the urban to rural adult mortality ratio probably reflect important regional characteristics. In particular, differences exist between regions in leading causes of mortality – malaria in West Africa, HIV/AIDS in Southern Africa – and lifestyle-related morbidity – higher liver cancer rates in West Africa and higher lung cancer rates in Southern Africa – (Lozano et al. 2012). Additionally, Southern African trends are probably related to higher HIV prevalence in the region, especially in the urban sector (Barongo et al. 1992; Dyson 2003; Killewo et al. 1990; Timæus and Jasseh 2004). Countries in this region – namely Namibia, Zimbabwe, and Swaziland – have low urban to rural mortality ratios on average, as seen in Figure 1. This rural disadvantage in adult mortality is surprising in these countries considering higher urban HIV prevalence, as seen in Table 1. Only in Namibia is HIV prevalence higher in the rural sector, among females. It is possible that these countries may be experiencing higher rural adult mortality due to the combined effect of HIV with drought and food insecurity in the rural sector (de Waal and Whiteside 2003). South Africa too has high HIV prevalence, and an urban to rural ratio just below one (for men) and above one (for women). Considering South Africa is a more developed country, we would expect higher rural mortality, yet HIV seems to be slowing down the transition away from an urban adult mortality disadvantage.

Figure 1: Mean urban to rural adult conditional mortality ratio by country and sex in sub-Saharan Africa, sorted by national level of mortality



Note: Our analysis is based on the country mean of all surveys presented in the Appendix. A ratio above one means higher urban mortality.

Figure 2: Urban to rural adult mortality ratio by sex over time in regions of sub-Saharan Africa



Note: Our analysis is based on the urban to rural ratios of adult mortality estimated using the orphanhood method. A ratio above one means higher urban mortality.

Table 1: HIV/AIDS prevalence by urban–rural sector for selected Southern Africa countries

Country	Survey Date	Female Urban	Female Rural	Male Urban	Male Rural
Zimbabwe	2010–2011	19.6	16.8	13.1	12
Namibia	2013	15	19.3	11.5	10.1
Swaziland	2006–2007	36.8	29.1	25.6	17.4

Source: Measure DHS Statcompiler (ICF International 2014).

Given uncertainty regarding the time location of the estimates of adult mortality, we concentrate primarily on overall average urban–rural differences rather than temporal trends in these differences. The results presented so far offer no evidence to suggest that rural adult mortality in SSA exceed those levels found in cities. A one-sided t-test ($p<0.01$) suggests that the urban to rural mortality ratio is greater than one – indicating an urban disadvantage. The difference in the mean rural and urban mortality probabilities is about one-tenth the magnitude of the standard deviations of the mortality probabilities.

In contrast to the urban adult mortality disadvantage we identify in our SSA data, the empirical literature has been mostly consistent in showing that child mortality is lower in urban areas across much of SSA (Akoto and Tambashe 2002; Bocquier, Madise, and Zulu 2011; Fink, Günther, and Hill 2014). Similarly, when compared to the urban to rural mortality ratio of under 5-year-olds, based on data from the DHS (ICF International 2014), we find that urban child mortality exceeds rural in only 4% of our sample. The lack of consistency is apparent in running a simple correlation analysis between the child mortality urban to rural ratios and the adult mortality urban to rural ratios. We find that the R-squared values are less than 1% for males and 7% for females, raising concern that our understanding of spatial differences for child mortality do not have direct implications for adult mortality differentials. This inconsistency may be particularly relevant to methods that rely on sectoral gaps in child mortality to proxy those among adults. While acknowledging the greater accuracy in child mortality estimates and the severe constraints on adult urban–rural mortality data for SSA, our findings suggest that such approaches need to be carefully considered.

Notwithstanding the broad consistency in our findings, it is important to note that our results are not population measures of mortality. We turn now to our robustness analysis, which explores in more detail those factors presented above that are likely to bias either urban or rural estimates and examines whether these factors lead us to alter our main conclusions.

3.1 Robustness tests

3.1.1 Rural–urban migration of orphans

In theory, rural to urban migration of orphans may artificially inflate urban adult mortality estimates and deflate rural estimates, producing upwards bias in the urban to rural adult mortality ratio. Here, we offer three arguments for why the direction of any bias is unlikely to reverse our overall conclusion that urban adult mortality levels exceed those in rural areas.

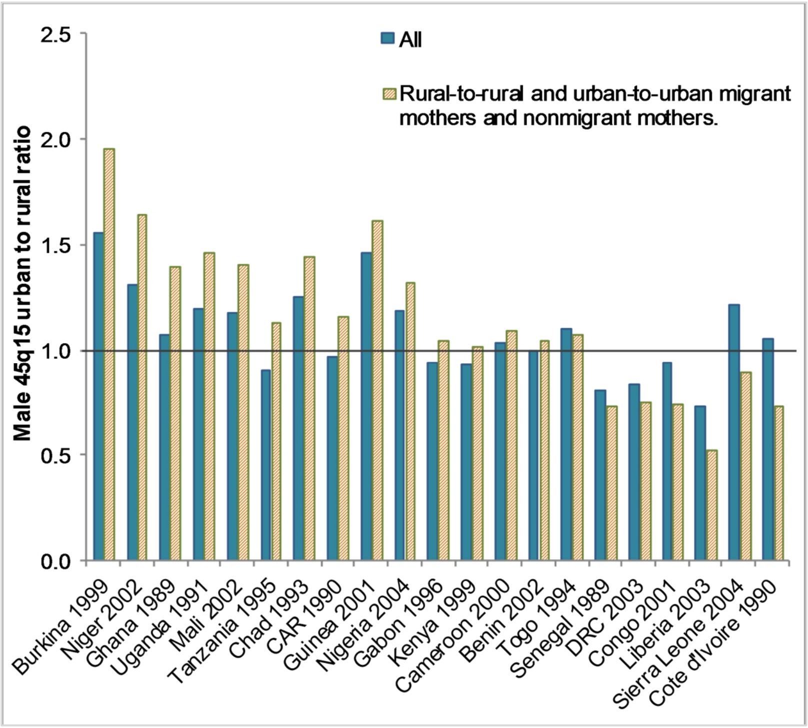
3.1.1.1 Existing empirical evidence on orphan migration

Studies show that children in the urban sector are more likely to migrate to the rural sector after being orphaned in comparison with the probability of orphaned children in the rural sector migrating to the urban sector (van Blerk and Ansell 2006; Isiugo-Abanihe 1985; Monasch and Boerma 2004; Sahn and Catalina 2013). In effect, this evidence suggests that the method based on orphanhood leads to systematic underestimation of the urban to rural adult mortality ratio.

3.1.1.2 Testing paternal orphan differences

We test the impact of orphan migration on paternal mortality estimates using DHS data for 21 West and Central East African countries that includes questions on maternal residence. Since the residence and migration status of the surviving parent (the mother) is known in these cases we can overcome a possible threat to the validity of our findings. Adult male mortality is estimated from single orphans whose mothers have not migrated, or whose mothers migrated only between rural areas or between urban areas. Under the assumption that paternal orphans remain with their mothers after paternal deaths, estimation of male adult mortality from this subsample offers a useful comparison group. Results shown in Figure 3 indicate that urban mortality tends to exceed rural when the estimates are restricted to a subgroup with relatively little migration. Instead of the limitation reducing the urban to rural ratio, we find that in most cases the limitation raises the ratio. Furthermore, whereas two countries (Sierra Leone and Côte d'Ivoire) show the ratio falling under 1.0 with the migration restriction, at least four show the opposite effect.

Figure 3: Male urban to rural mortality ratios excluding rural-to-urban and urban-to-rural migrants for 21 countries



Note: Our analysis includes one survey from West and Central East African countries where data on maternal residence is available. A ratio above one means higher urban mortality. Countries are sorted from largest negative gap to largest positive gap.

3.1.1.3 Testing duration of exposure and migration

An alternative evaluation test builds on the fact that older children have longer exposure to the possibility of migration following parental death. Thus, if the migration of orphaned children leads to an overestimate of adult urban mortality, the urban to rural mortality ratio estimated from older children (aged 10–14) should be higher than the urban to rural mortality ratio estimated from younger children (aged 5–9). We compare young and older children from data on 11 West African countries chosen to reduce the potential influence of HIV. The results show that the urban to rural female adult mortality ratio for younger children is higher in 26 of the 34 cases examined, offering further support for our main findings.

3.1.2 The adoption effect

Although we cannot directly estimate how many respondents refer to a foster mother rather than a deceased mother, the odds that this uncertainty affects our results is substantially diminished if orphanhood and fostering rates are similar across sectors. We find that the adoption effect is negligible based on DHS data on fostering and orphanhood prevalence by rural and urban sector (ICF International 2014). Average SSA orphanhood prevalence (for countries included in our analysis covering all dates) is 9.6% in rural populations and 10.3% in urban populations. Similarly, fostering prevalence averages 26.1% and 25.3% in rural and urban populations respectively in SSA, with a mean difference not statistically significantly different from zero.[11] This evidence is in line with literature that found that the proportion of households that foster out young children is roughly similar across the rural and urban sectors (Vandermeersch 2002).

In addition, if our mortality estimates were differentially affected by urban–rural fostering, urban to rural mortality and fostering ratios should be correlated. However, we find weak correlations between the urban to rural fostering ratio and the urban to rural mortality ratio, with an R^2 of 0.27 among males, and 0.06 among females. Thus, any bias resulting from fostering in the estimated level of mortality differs little across rural and urban estimates. Moreover, while overall estimates of mortality may be biased downwards, this bias should be minimal, considering that the respondents are adults reporting on household child members.

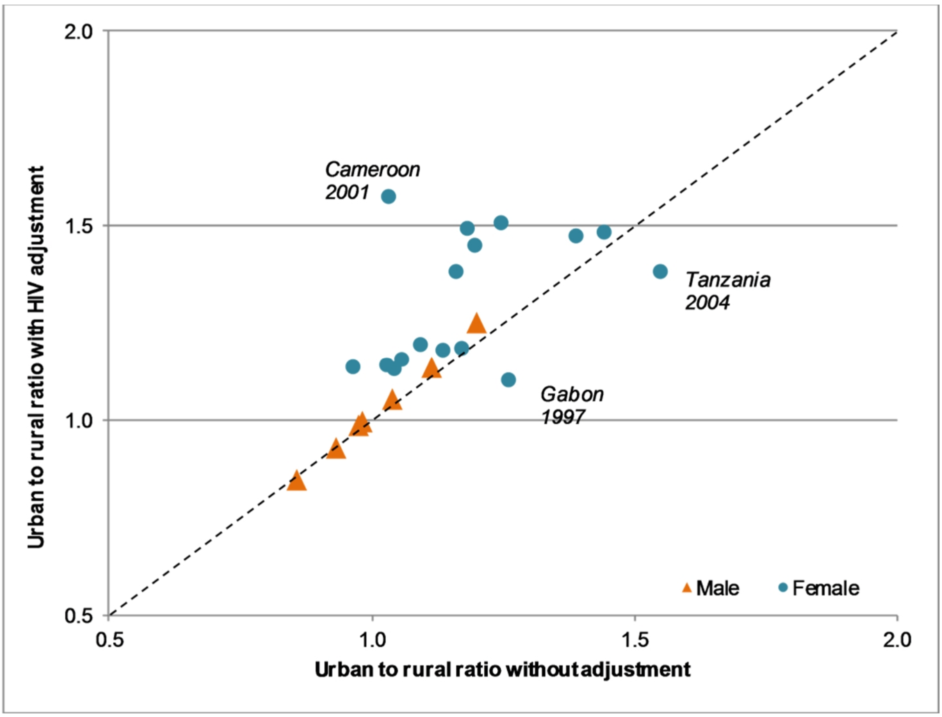
3.1.3 An HIV/AIDS correction

Our final robustness test involves the use of an adjustment to the orphanhood method to correct bias in estimates due to widespread HIV/AIDS-related mortality (Timæus and Nunn 1997). We compare our results with and without the adjustment in countries experiencing moderately severe HIV epidemics where the adjustment is relevant (Timæus and Nunn 1997). We also focus on peak HIV prevalence periods when the availability of ART was limited (1995–2005). Six countries from Central East Africa fit these criteria. The results in Figure 4 clearly show that the urban to rural mortality ratio

[11] Fostering prevalence is measured as the percentage of households with foster children who are under age 18 living with neither mother nor father.

is mostly higher when the adjustment is applied. In two of twelve cases, we find female mortality ratios higher without the adjustment. In effect, using the HIV adjustment would increase urban to rural adult mortality ratios, primarily because of higher prevalence of HIV in the urban sector, which supports an urban adult mortality disadvantage.

Figure 4: Urban to rural mortality ratios for six Central East SSA countries with moderately severe HIV prevalence (1995–2005): Comparison using the HIV adjustment



Note: Adult mortality calculated from DHS samples using the orphanhood method. From each survey one mortality estimate for males is calculated, and two estimates are calculated for females. A ratio above one indicates that urban mortality exceeds rural.

3.2 A prolonged urban disadvantage?

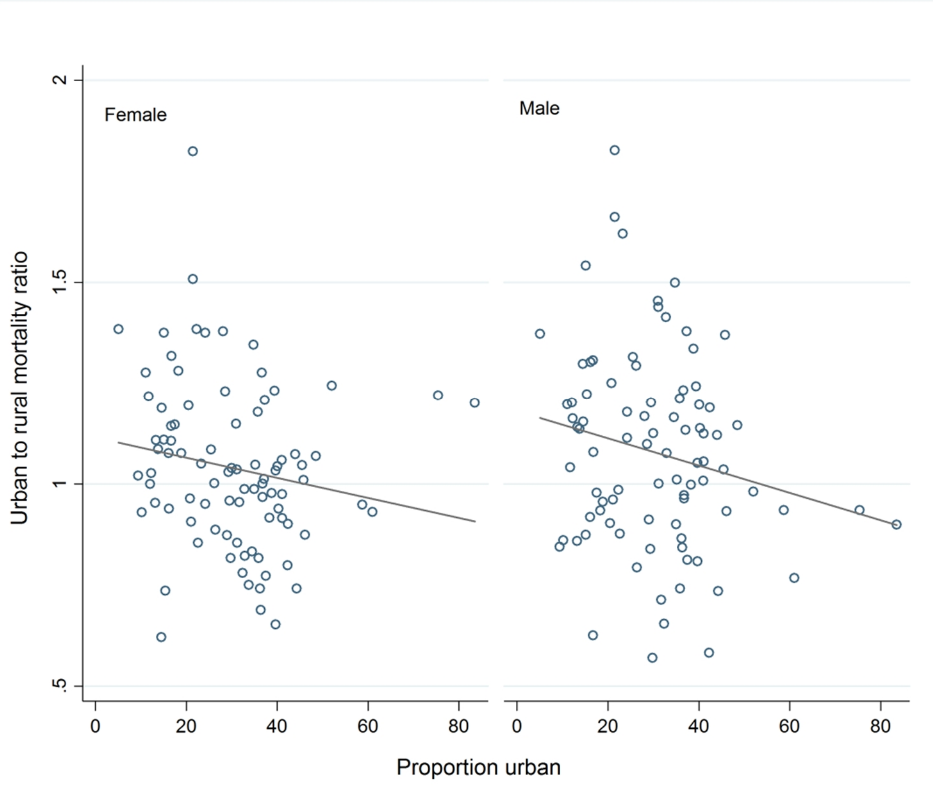
Our findings offer compelling evidence that rural adult mortality across much of sub-Saharan Africa is mostly comparable to and often lower than levels found in the urban sector. Although no firm conclusion can be drawn from the orphanhood method alone due to its limitations, our analysis has allowed a systematic review of urban and rural adult mortality in the majority of SSA countries based on nationally representative surveys. Though the results are surprising when compared to urban–rural child mortality and broader understandings of urban–rural mortality differences, they are consistent with Günther and Harttgen's (2012) findings of an urban disadvantage based on the sibling method of mortality estimation in the 2000s using a more limited sample of SSA countries.

Underlying our quest to uncover patterns of adult mortality across urban and rural settings is a broader aim of learning whether SSA is heading towards an urban mortality advantage – one that has been delayed – or whether urbanization in SSA simply has little relationship with adult mortality. An 'urban penalty' (especially visible among adult males) has been shown to change over the course of the demographic transition in Europe, with mortality decline being much faster in urban areas (Reher 2001). Our subsequent analysis examines the relationship between adult mortality and national level indicators of development status to determine whether a shift in the urban penalty can also be identified in SSA.

An initial perspective is gained by considering the shift in the urban to rural mortality ratio as countries urbanise, shown in Figure 5. There is no clear visual association from the data, although the bivariate regression fit to the data does offer some weak indication that urban mortality is higher than rural for both sexes when countries have low urban proportions. Although this association covers all reference dates and the mean proportion urban is merely 31% for the SSA countries in our sample, there is some indication that the rural disadvantage increases at higher levels of urbanisation. This modest relationship between urbanisation and the urban to rural ratio reflects the distinctiveness of the urban transition in SSA – one that is relatively slow and only weakly related to development (Cohen 2004; Potts 2016).

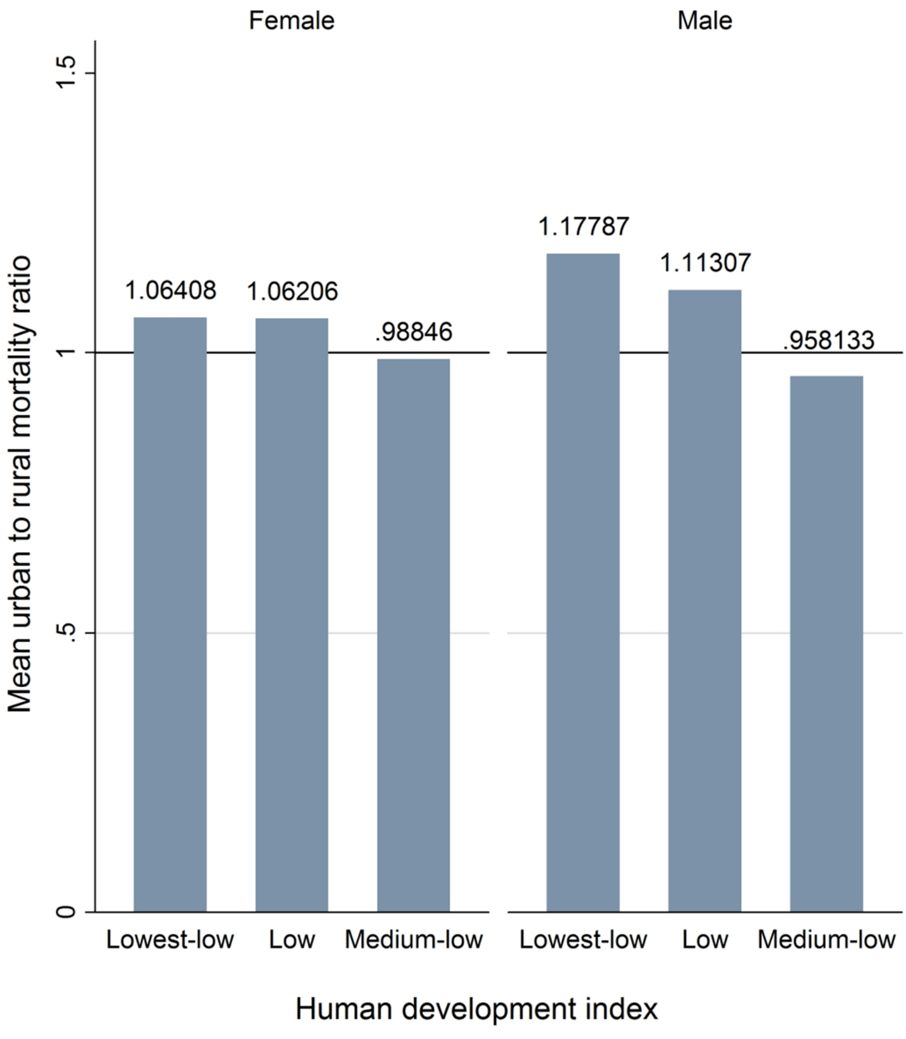
A second perspective is gained by considering the shift over the course of development. The Human Development Index (HDI) is a powerful indictor of national development based on life expectancy at birth, mean years of schooling, and gross national income per capita (UNDP 2015). We examine whether urban–rural adult mortality differentials vary systematically with levels of HDI. Rather than considering how countries progress over time, we see how they progress here by level of HDI. Therefore, countries may maintain the same HDI level for all reference dates, or they may change levels. Our findings, shown in Figure 6, indicate that urban to rural mortality ratios decline as countries develop. With very low HDI of below 0.37, urban mortality is higher than rural. In SSA countries with higher HDIs of above 0.45 the urban to rural ratio falls to below one (though close to equity among females), indicating a greater rural disadvantage in adult mortality.

Figure 5: Urban to rural adult mortality ratio in sub-Saharan Africa by proportion urban



Note: Our analysis is based on the urban to rural ratios of adult mortality estimated using the orphanhood method for 30 SSA countries, by proportion urban. A ratio above one means higher urban mortality.

Figure 6: Mean urban to rural adult mortality ratio in sub-Saharan Africa by HDI level and sex



Note: Our analysis includes all $_{45}q_{15}$ estimates presented in the Appendix, by HDI terciles. A ratio above one means higher urban mortality.

In Table 2, we move beyond the bivariate impressions and pool both within and between country estimates to model the variance in urban to rural adult mortality ratios. In these models, each case is an estimate of the male or female urban to rural adult mortality ratio at a given time for a given country. Our models control for national adult mortality ${}_{45}q_{15}$, sex, temporal patterns in mortality, proportion of country urban (as a measure of urbanisation), HIV prevalence, HDI, and total fertility rate (TFR). Models 1 and 2 are OLS, and Model 3 is linear country-level fixed effects (FE). The FE model explores changes in the urban to rural mortality ratio within countries. It also overcomes differences in urban definitions across countries, assuming definitions within the countries are unchanging over recent decades.

These models of urban to rural mortality ratios support our findings in Figures 5 and 6. Proportion urban is negatively associated with the mortality ratio, though not significant in all models. Additionally, countries with higher HDI experience substantial declines in the mortality ratio, leading to a rural disadvantage. An increase from a low HDI of 0.3 to a medium HDI of 0.6 is associated with a decline of 0.57 in the urban to rural ${}_{45}q_{15}$ ratio. This weak relationship between the urban to rural mortality ratio and urbanisation on one hand and strong relationship with HDI on the other may be explained by the unique urbanisation process in SSA that has slowed down and been disconnected from economic growth (Cohen 2004; Fay and Opal 2000; Potts 2016).

Although HDI is composed in part of life expectancy, it only partly reflects national levels of adult mortality. In our sample, the correlation coefficient between the two is 0.046. In all models, higher national ${}_{45}q_{15}$ reflects a lower ratio. A mean national adult mortality probability within countries indicates a decline of 0.32 in the urban to rural ratio. Where there is higher adult mortality at the national level, a rural disadvantage is expected. Total fertility rate is included in models to allow us to identify whether the adult mortality gap varies as fertility declines. The coefficient on TFR in Model 2 indicates that higher fertility is associated with a higher mortality ratio. That is, moving to lower fertility shifts mortality from an urban to a rural disadvantage.

Our findings (in Appendix Table A-1) demonstrate that adult women's mortality is lower, consistent with previous national level findings (Nathanson 1984; Rajaratnam et al. 2010). Although these results are encouraging, we suggest caution in comparison by sex as it may be possible that some bias between sexes is introduced once the mortality estimates are disaggregated by sector. In our models, female mortality is associated with an average decline in the urban to rural ${}_{45}q_{15}$ ratio by 0.06 or by about 6% of the average ${}_{45}q_{15}$ ratio. A lower urban to rural mortality ratio among women may thus be explained by less bias in women's estimates or by lower exposure to mortality risks such as smoking or hazardous occupations among urban women. Yet evidence suggests that urban women actually have lower mortality than rural men in SSA (de Walque and Filmer 2013).

The temporal pattern in the models indicates that mortality ratios have actually been rising over time. Adult mortality ratios peak between 1996 and 2000, then start to decline, though remaining higher than in the late 1980s (the reference category). In Model 3, the ratio rises again in recent years. Overall, over time these findings indicate a shift to an urban disadvantage, as seen in Figure 2 in southern and western Africa. These findings are also consistent with studies on mortality trends that indicate increases in mortality since the 1970s in SSA, followed by more recent mortality declines (McMichael et al. 2004; Rajaratnam et al. 2010; Wang et al. 2012). The role of HIV is found to be positive but insignificant, indicating that variation in overall HIV prevalence has little relation to urban to rural adult mortality ratios.

Table 2: Linear country fixed effects models showing determinants of adult mortality ${}_{45}q_{15}$ in sub-Saharan Africa

	Model 1	**Model 2: OLS**	**Model 3: FE**
	b/se	**b/se**	**b/se**
National ${}_{45}q_{15}$	–0.586**	–0.510*	–1.141**
	(0.16)	(0.227)	(0.268)
Female	–0.053^	–0.050^	–0.069**
	(0.03)	(0.03)	(0.026)
1991–1995	0.07	0.087^	0.057
	(0.049)	(0.05)	(0.05)
1996–2000	0.198**	0.220**	0.264**
	(0.05)	(0.053)	(0.061)
2001–2005	0.145**	0.173**	0.260**
	(0.048)	(0.053)	(0.066)
2006–2011	0.113*	0.158**	0.330**
	(0.046)	(0.052)	(0.09)
Proportion urban	–0.004**	–0.002	–0.005
	(0.001)	(0.001)	(0.008)
HIV prevalence		0.003	0.006
		(0.004)	(0.018)
HDI		–0.13	–1.897**
		(0.254)	(0.501)
TFR		0.046^	0.001
		(0.024)	(0.053)
Constant	1.277**	0.922**	2.119**
	(0.07)	(0.249)	(0.595)
R-squared	0.112	0.133	0.478
No. of cases	272	272	272

Note: The models predict the urban to rural ratio of the probability of dying between ages 15 and 60. The omitted category of time is 1986–1990. ^ p<.10, * p<.05, ** p<.01 [two-tailed tests]. Standard errors are in parenthesis.

4. Discussion

In contrast to established findings that show that infant mortality is higher in rural areas, we find adult mortality to be higher in urban areas across many SSA countries. Historically, before the onset of the demographic transition, Europe experienced higher mortality in urban settings (Reher 2001; Woods 2003). Our results suggest that contemporary SSA may be facing similar challenges with high urban population densities and insufficient health and sanitation infrastructure (Konteh 2009; Szreter 1997). Urban populations may face higher mortality in part due to factors related to natural resistance to disease (Johansson and Mosk 1987) – including the high proportion of poverty in urban areas (Gould 1998; Ravallion, Chen, and Sangraula 2007) and growing urban slums (Fink, Günther, and Hill 2014; Rice and Rice 2009). Indeed, a narrowing gap in urban–rural child mortality differentials is attributed to the conditions in urban slums (Kimani-murage et al. 2014). A ‘double burden’ of disease is common in SSA cities. Poor populations living in slums suffer mostly from infectious diseases (Mberu et al. 2015; Soura, Lankoande, and Millogo 2014), though urban populations are also vulnerable to noncommunicable diseases as they begin to age and the middle class expands (Agyei-Mensah and Aikins 2010). Furthermore, while rural to urban migrants may be healthier than their rural counterparts initially, by adopting an urban lifestyle, migrants may have worse health, raising urban mortality levels (Ebrahim et al. 2010; Hernández et al. 2012).

As most countries in SSA are still undergoing the urban transition, urban mortality is expected to remain higher than rural. Urbanisation may not be very important for reducing urban mortality, even though higher urban proportions are associated with lower urban to rural mortality ratios. In general, the more people living in cities, the greater the access of the population to education, health services, and economic opportunities. Considering the uncoupling of urbanisation from economic growth experienced in SSA in recent decades (Fay and Opal 2000), we turned to a composite measure of development, HDI, to assess shifts in urban–rural mortality. The results indicate that countries with lower HDI have higher urban adult mortality among both sexes. Higher levels of development do appear to be associated with conditions of higher rural relative to urban ${}_{45}q_{15}$.

The declines in life expectancy seen in SSA, depicted by McMichael et al. (2004) as reversals of trends, highlight the importance of understanding mortality dynamics within countries in the region, particularly between urban and rural populations. Our findings identify a delayed adult mortality decline in urban populations in SSA, possibly due to low (and slow) urbanisation. What is equally worrisome is if and how the urban adult mortality disadvantage may affect the course of progress of African states. Unfortunately, the implications may be exacerbated by the economic costs associated with a prolonged urban mortality disadvantage occurring in the midst of a demographic transition. Excess adult mortality could reduce productivity in peak labour productivity ages, reducing the potential demographic dividend within the urban sector. This possibility suggests that more effective policies are needed to address excess adult mortality in urban areas where education and employment opportunities are greater and the productivity of young adults is relatively high. SDG-defined programmes should invest in decreasing urban mortality to ensure both equality within countries and harnessing economic growth.

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Appendix

Table A-1: Estimates of adult mortality in sub-Saharan Africa between ages 15 and 60 by rural and urban residence and ratio of urban–rural mortality, by sex based on DHS survey data

Country	Survey date	Reference date	Female			Male		
			Urban $_{45}q_{15}$	Rural $_{45}q_{15}$	Urban to rural ratio	Urban $_{45}q_{15}$	Rural $_{45}q_{15}$	Urban to rural ratio
Western Africa								
Benin	1996	1992	0.202	0.243	0.83	0.231	0.198	1.17
	2001	1997	0.172	0.177	0.97	0.223	0.229	0.97
	2006	2002	0.195	0.213	0.93	0.226	0.226	1.00
	2011	2008	0.153	0.145	1.06	0.187	0.186	1.01
Burkina Faso	1993	1989	0.286	0.278	1.03	0.311	0.267	1.16
	2003	1999	0.304	0.221	1.37	0.361	0.234	1.54
	2010	2006	0.199	0.132	1.51	0.257	0.155	1.66
Chad	1997	1993	0.239	0.248	0.97	0.328	0.262	1.25
	2004	2000	0.363	0.199	1.83	0.366	0.200	1.83
Côte d'Ivoire	1994	1990	0.189	0.196	0.99	0.227	0.215	1.06
	2005	2001	0.219	0.210	1.08	0.291	0.252	1.15
	2011	2008	0.289	0.227	1.29	0.308	0.217	1.42
Ghana	1993	1989	0.203	0.247	0.83	0.264	0.245	1.08
	1998	1994	0.157	0.228	0.70	0.187	0.222	0.84
	2003	1999	0.181	0.174	1.06	0.254	0.212	1.20
	2008	2004	0.184	0.171	1.07	0.249	0.222	1.12
Guinea	1999	1995	0.247	0.258	0.96	0.313	0.260	1.20
	2005	2001	0.259	0.225	1.15	0.322	0.222	1.45
	2012	2008	0.259	0.263	1.03	0.322	0.227	1.41
Liberia	2007	2003	0.186	0.251	0.75	0.189	0.257	0.74
	2013	2009	0.155	0.178	0.88	0.230	0.247	0.93
Mali	1996	1992	0.204	0.194	1.06	0.270	0.166	1.62
	2001	1997	0.207	0.191	1.11	0.245	0.186	1.32
	2006	2002	0.247	0.179	1.41	0.216	0.185	1.17
	2012	2008	0.152	0.147	1.03	0.155	0.108	1.44
Niger	1992	1988	0.183	0.295	0.62	0.255	0.196	1.30
	1998	1994	0.174	0.236	0.73	0.219	0.179	1.22
	2006	2002	0.222	0.236	0.94	0.235	0.180	1.30
	2012	2008	0.221	0.194	1.14	0.191	0.146	1.31
Nigeria	2003	1999	0.242	0.248	0.98	0.257	0.192	1.34
	2008	2004	0.149	0.165	0.91	0.208	0.175	1.19
	2013	2009	0.146	0.144	1.01	0.217	0.159	1.37
Senegal	1992	1989	0.158	0.204	0.78	0.190	0.234	0.81
	2005	2001	0.208	0.222	0.95	0.275	0.241	1.14
	2010	2007	0.171	0.176	0.98	0.204	0.181	1.13
	2012	2008	0.145	0.158	0.91	0.201	0.190	1.06
	2014	2010	0.121	0.151	0.81	0.122	0.210	0.58
Sierra Leone	2008	2004	0.344	0.292	1.18	0.394	0.325	1.21
	2013	2009	0.309	0.255	1.23	0.405	0.293	1.38
Togo	1998	1994	0.280	0.228	1.23	0.303	0.276	1.10
	2013	2010	0.229	0.218	1.04	0.277	0.274	1.01
Average			0.213	0.210	1.032	0.256	0.216	1.199
Weighted average			0.197	0.198	1.002	0.244	0.196	1.255

Table A-1: (Continued)

Country	Survey date	Reference date	Female			Male		
			Urban $_{45}q_{15}$	Rural $_{45}q_{15}$	Urban to rural ratio	Urban $_{45}q_{15}$	Rural $_{45}q_{15}$	Urban to rural ratio
Central East Africa								
Burundi	2010	2006	0.316	0.310	1.02	0.306	0.363	0.84
Cameroon	1991	1987	0.175	0.237	0.75	0.214	0.247	0.87
	1998	1994	0.194	0.298	0.65	0.317	0.301	1.05
	2004	2000	0.276	0.264	1.05	0.326	0.315	1.04
	2011	2007	0.287	0.268	1.07	0.302	0.264	1.15
CAR	1995	1990	0.337	0.337	1.01	0.353	0.367	0.96
Congo	2005	2001	0.276	0.291	0.95	0.236	0.253	0.93
	2011	2007	0.189	0.203	0.93	0.149	0.195	0.77
DRC	2007	2003	0.287	0.279	1.03	0.239	0.285	0.84
	2013	2009	0.252	0.295	0.85	0.267	0.266	1.00
Gabon	2001	1996	0.263	0.215	1.22	0.175	0.187	0.94
	2012	2008	0.200	0.167	1.22	0.196	0.218	0.90
Kenya	1993	1989	0.170	0.158	1.08	0.252	0.274	0.92
	1998	1994	0.260	0.235	1.10	0.230	0.368	0.63
	2003	1999	0.400	0.313	1.29	0.417	0.447	0.93
Rwanda	1992	1988	0.327	0.237	1.38	0.458	0.333	1.37
	2005	2001	0.449	0.414	1.08	0.611	0.538	1.14
	2010	2006	0.294	0.256	1.14	0.330	0.337	0.98
Tanzania	1991	1987	0.290	0.220	1.32	0.283	0.262	1.08
	1996	1992	0.283	0.263	1.08	0.292	0.306	0.96
	1999	1995	0.352	0.294	1.20	0.286	0.317	0.90
	2007	2003	0.351	0.253	1.40	0.329	0.334	0.99
	2010	2006	0.321	0.233	1.40	0.317	0.284	1.11
	2012	2008	0.221	0.232	0.95	0.349	0.296	1.18
Uganda	1995	1991	0.509	0.398	1.28	0.528	0.441	1.20
	2001	1996	0.490	0.402	1.22	0.447	0.429	1.04
	2006	2002	0.417	0.417	1.00	0.533	0.443	1.20
	2011	2007	0.301	0.315	0.95	0.314	0.366	0.86
Average			0.303	0.279	1.093	0.323	0.323	0.991
Weighted average			0.302	0.284	1.069	0.315	0.326	0.964

Table A-1: (Continued)

Country	Survey date	Reference date	Female			Male		
			Urban $_{45}q_{15}$	Rural $_{45}q_{15}$	Urban to rural ratio	Urban $_{45}q_{15}$	Rural $_{45}q_{15}$	Urban to rural ratio
Southern Africa								
Malawi	1992	1988	0.316	0.340	0.92	0.243	0.283	0.86
	2000	1996	0.439	0.396	1.11	0.438	0.384	1.14
	2005	2000	0.489	0.411	1.19	0.521	0.452	1.15
	2010	2006	0.379	0.342	1.10	0.319	0.366	0.87
Mozambique	1997	1993	0.381	0.420	0.91	0.314	0.327	0.96
	2003	1999	0.334	0.333	1.00	0.408	0.316	1.29
	2011	2007	0.343	0.331	1.04	0.399	0.355	1.13
Namibia	1992	1988	0.165	0.186	0.90	0.222	0.280	0.79
	2000	1996	0.247	0.303	0.81	0.257	0.451	0.57
	2007	2003	0.354	0.454	0.78	0.330	0.505	0.65
	2013	2009	0.332	0.336	1.02	0.334	0.371	0.90
South Africa	1998	1994	0.212	0.170	1.25	0.358	0.365	0.98
Swaziland	2007	2002	0.464	0.543	0.85	0.558	0.637	0.88
Zambia	1992	1988	0.268	0.259	1.03	0.242	0.300	0.81
	1996	1992	0.418	0.340	1.23	0.464	0.374	1.24
	2002	1998	0.456	0.450	1.01	0.548	0.483	1.13
	2007	2003	0.504	0.375	1.35	0.555	0.370	1.50
	2014	2009	0.341	0.267	1.28	0.405	0.329	1.23
Zimbabwe	1994	1990	0.209	0.239	0.87	0.302	0.331	0.91
	1999	1995	0.361	0.378	0.95	0.379	0.530	0.71
	2011	2006	0.436	0.533	0.81	0.451	0.608	0.74
Average			0.355	0.353	1.020	0.383	0.401	0.974
Weighted Average			0.300	0.281	1.115	0.376	0.378	1.006
SSA Average			0.274	0.265	1.048	0.307	0.292	1.082
SSA Weighted Average			0.254	0.244	1.052	0.294	0.275	1.111

Note: From each survey, one mortality estimate for males is calculated and two estimates for females. For females, the average from these two estimates is used. Regional and population-weighted averages are included.