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Three Essays On Family, Gender, And Educational Change Across Low- And Middle-Income Countries

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Abstract
Over the past half century, low- and middle-income countries (LMICs) have undergone profound transformations in the realm of the family, accompanied by shifts in gender norms and practices, and dramatic increases in schooling. Rising educational attainment has in turn been a by-product of micro-level behavioral changes on the part of families, alongside macro-level socio-structural factors such as industrialization, urbanization, and targeted educational policies. This dissertation advances the field of social demography by exploring the interrelations between family, gender, and educational dynamics across LMICs. Although the three essays represent self-contained articles, they all trace linkages between these three dimensions with a focus on LMICs, thus contributing new empirical knowledge on policy-relevant population processes in contexts that have to date received less scholarly attention. The first chapter provides a macro-level overview on the changing nature of families across multiple domains with advances in socio-economic development. Its focus is on family change, yet gender features in the type of indicators considered, some of which are computed separately for men and women – showing vastly divergent patterns – while others capture men and women's bargaining power within the couple. Educational expansion features throughout the discussion as one key driver of family change and one component of the Human Development Index (HDI) proxying for socio-economic development. The second chapter provides an overview on trends, variation, and implications of educational assortative mating for inequality in sub-Saharan Africa (SSA). Mating patterns are vital to understanding the demographic makeup of households, such as family formation, composition, and breakdown. The focus on education and gender is inherent in the type of question raised (educational homogamy/heterogamy) and perspective adopted (couple). The third chapter explores the effect of a cash-transfer intervention given to parents on children's schooling and unpaid work in rural Morocco. As such, the family focus is tied to a parental investment perspective, while the educational focus comes from the policy considered – a cash transfer promoted by the government – and the outcomes analyzed – school dropout and grade progression. Lastly, gender features throughout the discussion as analyses consider heterogeneity by gender, and unpaid care dynamics show striking gender differences.

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THREE ESSAYS ON FAMILY, GENDER, AND EDUCATIONAL CHANGE ACROSS LOW- AND MIDDLE-INCOME COUNTRIES

Luca Maria Pesando

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To my family and friends, and to whatever will come next in my journey of knowledge.

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ABSTRACT

THREE ESSAYS ON FAMILY, GENDER, AND EDUCATIONAL CHANGE ACROSS LOW- AND MIDDLE-INCOME COUNTRIES

Luca Maria Pesando

Hans-Peter Kohler

Over the past half century, low- and middle-income countries (LMICs) have undergone profound transformations in the realm of the family, accompanied by shifts in gender norms and practices, and dramatic increases in schooling. Rising educational attainment has in turn been a by-product of micro-level behavioral changes on the part of families, alongside macro-level socio-structural factors such as industrialization, urbanization, and targeted educational policies. This dissertation advances the field of social demography by exploring the interrelations between family, gender, and educational dynamics across LMICs. Although the three essays represent self-contained articles, they all trace linkages between these three dimensions with a focus on LMICs, thus contributing new empirical knowledge on policy-relevant population processes in contexts that have to date received less scholarly attention. The first chapter provides a macro-level overview on the changing nature of families across multiple domains with advances in socio-economic development. Its focus is on family change, yet gender features in the type of indicators considered, some of which are computed separately for men and women – showing vastly divergent patterns – while others capture men and women's bargaining power within the couple. Educational expansion features throughout the discussion as one key driver of family change and one component of the Human Development Index (HDI) proxying for socio-economic development. The second chapter provides an overview on trends, variation, and implications of educational assortative mating for inequality in sub-Saharan Africa (SSA). Mating patterns are vital to understanding the demographic makeup of households, such as family formation, composition, and breakdown. The focus on education and gender is inherent in the type of question raised (educational

homogamy/heterogamy) and perspective adopted (couple). The third chapter explores the effect of a cash-transfer intervention given to parents on children's schooling and unpaid work in rural Morocco. As such, the family focus is tied to a parental investment perspective, while the educational focus comes from the policy considered – a cash transfer promoted by the government – and the outcomes analyzed – school dropout and grade progression. Lastly, gender features throughout the discussion as analyses consider heterogeneity by gender, and unpaid care dynamics show striking gender differences.

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PREFACE

Over the past half century, low- and middle-income countries (LMICs) have undergone profound transformations in the realm of the family, partly driven – or at least accompanied – by shifts in gender norms and practices, and dramatic increases in schooling. Rising educational attainment has in turn been a by-product of micro-level behavioral changes on the part of families, alongside macro-level socio-structural factors such as industrialization, urbanization, and targeted educational policies. This dissertation seeks to advance the field of social demography by exploring the complex interrelations between family, gender, and educational dynamics across LMICs. The dissertation follows a three-chapter format. Although the essays represent self-contained, independent articles, they share the commonality that they all – more or less directly – trace linkages between these three dimensions with a specific focus on LMICs, thus contributing new empirical knowledge on policy-relevant population processes in contexts that have to date received less scholarly attention.

Chapters 1 and 2 adopt a cross-country macro-level perspective, while Chapter 3 adopts a country-specific micro-level approach. Chapter 1 provides a macro-level overview on the changing nature of families across multiple domains with advances in socio-economic development. Its focus is primarily on family change, yet gender features in the type of indicators considered, some of which are computed separately for men and women – showing vastly divergent patterns – while others capture men and women's relative position and bargaining power within the couple. Educational expansion features throughout the discussion as one key driver of family change and one of the three components of the Human Development Index (HDI) used as proxy measure for socio-economic development. Chapter 2 provides a macro-level overview on trends, variation, and implications of educational assortative mating for inequality in sub-Saharan Africa (SSA) over the past three decades. Assortative mating patterns are vital to understanding a whole set of dynamics in the demographic makeup of households, such as family formation, composition, and breakdown. The focus on education and gender is inherent in the type of research question raised (partners' educational homogamy/heterogamy) and perspective adopted (couple-level). Chapter 3 explores the effect of a cash-transfer intervention given to parents on children's schooling and unpaid care work outcomes in rural Morocco. As such, the family focus is tied to a parental investment perspective, while the education focus comes from the policy considered – a cash transfer promoted by the government – and the type of outcomes analyzed – school dropout and timely grade progression. Lastly, gender features throughout the discussion as all analyses consider heterogeneity by gender and unpaid care dynamics – the latter showing striking patterns by gender.

Moving to the specifics of each chapter, Chapter 1 builds on the idea that tracking family transformations in LMICs constitutes a first step towards providing a quantum leap in the study of the shifts in family forms and dynamics that are unfolding globally. Three are the guiding research questions: (i) How have families across LMICs changed with advances in development over the past 30 years? (ii) Which family domains have been changing the most? (iii) Is there evidence of cross-country convergence in family domains over socio-economic development? The paper uses pooled micro-level data from 293 Demographic and Health Surveys (DHS) across 84 low- and middle-income countries to provide cross-national evidence on the association between the HDI and a series of indicators spanning multiple family domains such as fertility, timing of life-course events, union formation (marriage and cohabitation), household structure (inter-generational relationships), and within-couple decision-making dynamics (intra-generational relationships).

The motivation behind the study relies on two fundamental premises. First, family change in LMICs is unlikely to be a simple extension of patterns observed in high-income countries (Furstenberg 2013). Second, the shift in focus onto understanding global social change requires endorsement of the idea that changes in families matter independently of changes in fertility (van de Walle 1993), despite family formation remains a precondition for fertility in some contexts. The analysis proceeds by means of descriptive graphical associations between 18 family indicators and HDI – both globally and regionally – complemented by more formal statistical tests of convergence (beta- and sigma-convergence) borrowed from the growth convergence literature (Barro and Sala-i-Martin 1992; Sala-i-Martin 1996).

Findings support the aforementioned premises. First and foremost, while advances in development are associated with declines in fertility across all low- and middle-income regions, other domains such as marriage and inter-generational relationships move at a far slower pace, showing widespread cross-regional heterogeneity. While fertility, intra-couple decision-making, and women's life-course timing indicators are strongly associated with HDI, cross-country convergence is limited to the latter domain. Marriage, cohabitation, household structure, and men's life-course timing indicators show a weaker association with HDI and span a broad spectrum of convergence dynamics ranging from divergence to modest convergence. We describe this scenario as "persistent diversity with development." The remarkable persistence of the family in several domains suggests that its centrality is not going to fade anytime soon and urges to move beyond the common narrative that ties fertility and family change together.

The nature of these findings enriches the literature on global family change by engaging with previous attempts at conceptualizing family change such as Goode's "convergence hypothesis" (Goode 1963), Therborn's "five family systems" (Therborn 2004), and Lesthaeghe and van de Kaa's Second Demographic Transition (SDT) theory (Lesthaeghe 2010, 2014). Our analysis suggests a more nuanced and contextualized understanding of family change that helps account for patterns of convergence, divergence, and persistent differences. Lastly, the paper devotes ample space to discussing regional idiosyncrasies, and highlights the role of sub-Saharan Africa (SSA) as the region most lagging behind in dimensions of development, hence most likely to experience swift transformations in family domains in the near future with socio-economic progress.

Chapter 2 picks up on this SSA "uniqueness" – not backed by enough comparative research within the region – to explore how educational change interacts with union/marriage formation dynamics in sub-Saharan Africa. For this study, I use micro-level data from 126 DHS surveys across 39 SSA countries to provide a closer look at changing patterns of educational assortative mating in the region. The study seeks to: (i) highlight patterns of mating by marriage cohort, sub-region of SSA (Western, Central. Eastern, and Southern), and household location of residence (rural/urban); (ii) compare observed patterns of mating with those that would prevail under random mating and investigate the extent to which the observed trends are driven by changes in compositional factors versus partners' actual preference for spousal educational resemblance; (iii) assess implications of mating for between-household wealth inequality – measured through the International Wealth Index (IWI) provided by the Global Data Lab. To address these questions, I rely on graphical analyses of associational measures, contingency tables, log-linear models, and counterfactual simulations. To assess implications for wealth inequality I ask the following counterfactual questions: what would have happened to the wealth distribution if mating within each cohort was random instead of assortative? What would have happened to wealth inequality if couples in the latest marriage cohort matched as those in the earliest one?

The paper builds on the premise that never has the assortative mating literature focused exclusively on patterns of mating within SSA, while research is more extensive in other low- and middle-income contexts such as Latin America (Esteve and McCaa 2007; Esteve, McCaa, and López 2013; Ganguli, Hausmann, and Viarengo 2014; Gullickson and Torche 2014; Torche 2010) and South-East Asia (Borkotoky and Gupta 2016; Hu and Qian 2015; Park and Smits 2005; Smits and Park 2009) – not to mention high-income societies such as Europe and the US. Also, existing studies on declining hypergamy throughout the world (Esteve, Garcia, and Permanyer 2012; Esteve et al. 2016) might not suffice to assess whether we observe increasing educational resemblance of spouses net of shifts in marginal distributions, and how trends vary by sub-region and location of residence within SSA – a vastly heterogeneous region that has followed different urbanization trajectories and retains context-specific socio-cultural repertoires. As documented in Chapter 1, SSA is undergoing rapid changes in the realm of union formation – such as delays in mean ages at life-course events – together with increases in female educational attainment and labor force participation rates (Bongaarts, Mensch, and Blanc 2017). Yet the gender gap in education has not reversed yet, and the region remains the only one where the share of households in which the husband is the sole decision-maker reaches up to 40 per cent (Chapter 1). Therefore, there is reason to believe that trends towards increasing assortative mating documented globally, often unfolding along with reversals in gender gaps in education and increases in women's empowerment, might occur differently in SSA. Also, no study has yet assessed the validity of the competing theoretical hypotheses (general openness hypothesis, status attainment hypothesis, inverted U-curve hypothesis) relating socio-economic development and educational homogamy in the SSA context.

My results show that mating in SSA has followed rather different trajectories by sub-region and location of residence. While there is evidence of positive educational assortative mating throughout SSA – i.e., men and women with the same level of education marrying more frequently than what would be expected under a marriage pattern that is random with respect to education – mating has increased over subsequent cohorts in Western, Central, and Eastern Africa, yet it has flattened out and somewhat decreased in Southern Africa. Heterogeneity is also evident in levels and relative growth, as mating was lower in Western Africa for early cohorts, yet the sub-region has witnessed the steepest increase in the marital sorting parameter. Additionally, findings show that increases in mating have been largely driven by rural areas – where the trend for SSA as a whole is consistent with the status attainment hypothesis – while mating in urban areas has shown a mild increase followed by an incipient decline – consistent with the inverted U-curve framework and the increasing applicability of the general openness hypothesis. Overall, the documented heterogeneity – and, foremost, the diverging trends between Western and Southern Africa – is consistent with the economic (e.g., urbanization), socio-demographic (e.g., changes in families), and cultural specificities (e.g., patriarchal norms) of each sub-region. As for the inequality analysis, I find that mating accounts for a non-negligible share (3 to 12 percent, varying by sub-region) of the cohort-specific inequality in household assets, yet – in line with the literature from high-income societies – changes in mating over time hardly move the time trends in wealth inequality.

Chapter 3 focuses on one single country in Northern Africa, Morocco, and builds on experimental data from a randomized policy evaluation – a cash-transfer intervention called "Tayssir"

– implemented throughout the country in rural areas between 2008 and 2010. The paper aims to shed new light on the interplay between household inequality – as driven by differential parental investments, gender, and unpaid care work dynamics – and children's schooling. Specifically, I assess the impact of the cash transfer on school progression outcomes, defined as progressing through grades in a timely fashion, allowing for heterogeneity of the treatment along socio-demographic lines such as the gender of the child and the amount of time spent by children on unpaid care work prior to intervention implementation. As unpaid care work emerges as a negative predictor of school progression, the analysis concludes with an examination of whether the cash transfer had any effect on lessening the care burden itself. The data are publicly available from the online platform of the World Bank and the Abdul Latif Jameel Poverty Action (J-PAL) Lab.

The study builds on the premise that girls in low-income contexts – such as rural settings in Middle-East and North-African countries – are often at higher risk of not completing primary education due to rooted traditions and moral and religious beliefs that perpetuate gender inequalities since young ages. These inequalities affect the role girls play within the household, the distribution of activities, and the amount of time spent on them, leading to unequal allocations of care tasks. In this work I bring a gender lens to the analysis of the effects of a cash transfer to investigate whether these household-level dynamics shape the effectiveness of the policy implemented. By assessing whether the impacts of the transfer on school progression differ by gender of the child and burden of unpaid care work prior to intervention implementation, I evaluate whether the disproportionate care burden affecting girls may – at least partly – explain differential educational outcomes among heterogeneous groups of children.

My results show that a cash transfer explicitly tied to an educational purpose affected school progression on top of enrolment and attendance by stemming dropout (extensive margin of school participation) and increasing the likelihood of timely grade progression (intensive margin of school participation). While the former effect operated similarly for boys and girls, the latter was null for boys. In other words, the cash transfer was effective in reducing boys' dropout, though it did not alter their grade progression path. Furthermore, while the effect of the intervention on dropout operated equally for children who performed unpaid care work versus those who did not, the beneficial effect of the treatment on timely grade progression was cut by a half to a third for girls overburdened by household chores. Therefore, as a result of the treatment girls engaged in unpaid care tasks were staying in school more but were less likely to progress on time. This claim lends itself to two implications. First, it points to the need to design policy interventions with the potential to help children attend and progress through school in a timely manner, as a monetary "nudge" might not be sufficient to simultaneously achieve both goals. Second, it stresses the importance of tackling rooted gender inequalities within the household by directing policies to vulnerable children that may need ad hoc targeting.

Besides dealing with family, gender, and educational dynamics, the proposed chapters share some other commonalities. First of all, they all center on LMICs, with a key focus on Africa. Chapters 1 and 2 are very much related in other respects, as they are comparative in nature and aimed at understanding cross-country and/or cross-regional and inter-temporal variation in family dimensions in contexts that are undergoing rapid socio-economic change. Also, they both rely on micro-level DHS data to get at population-level processes. Not least, the two papers devote particular attention to the demographic nature of the phenomena under investigation by disentangling purely compositional factors from broader socio-demographic and behavioral considerations underlying the changing nature of families. For instance, in Chapter 1 we age-standardize age-sensitive indicators to ensure that comparisons identify the extent to which basic demographic trends – such as changing age structures – account for differences within/across countries versus more fundamental social processes that are arguably driving family changes. In a similar spirit, in Chapter 2 I am interested in determining whether mating patterns are driven by "mechanical" changes that result from proportionally faster increases in women's education, versus behavioral responses related to the shifting value of education and spousal preferences for educational resemblance. In Chapter 1 we also compute indicators for fertility, marriage, and cohabitation that incorporate information on mortality (Net Reproduction Rate and life expectancies in the married and cohabiting states using the Sullivan method), claiming that bringing mortality into the picture is key in contexts that are experiencing rapid increases in life expectancy. These "demographically savvy" considerations interacting family domains with basic demographic forces such as fertility and mortality are essential when conducting comparative analyses, yet mostly overlooked in existing research.

Indeed, there are numerous differences between the chapters. First is the scope of the research questions raised. Chapter 1 sets the stage for a whole research agenda on global family change by assessing changes across multiple family domains, while Chapter 2 focuses on a specific facet of family change, i.e., educational assortative mating, and Chapter 3 focuses on parental investments, educational policies, and children's outcomes. Second is the geographical scope. While Chapter 1 is "global," Chapter 2 is regional as it focuses on SSA, and Chapter 3 is country-specific as it focuses on Morocco only. Third are the data sources. While for Chapter 1 we use DHS combined with HDI data from the Human Development Report Office of the United Nations Development Programme (UNDP), and official life tables from the World Population Prospects: The 2015 Revision, for Chapter 2 I combine DHS with data on the International Wealth Index (IWI) from the Global Data Lab. For Chapter 3 I use publicly available (World Bank/J-PAL) data from an existing randomized experiment. Lastly, throughout the dissertation I combine a variety of theoretical and methodological approaches from demography, sociology, and economics. Methodologies range from simple correlation coefficients and data visualization techniques to ordinary least squares – adapted to fit the growth convergence framework – contingency tables, log-linear models, counterfactual simulations, and more causally-oriented approaches such as policy evaluation and quasi-experimental techniques (e.g., difference-in-differences).

Chapter 1. Global Family Change: Persistent Diversity with Development

1.0 Abstract

This paper provides a broad empirical overview of the relationship between family change and socio-economic development drawing on 30+ years of Demographic and Health Survey data from 3.5 million respondents across 84 low- and middle-income countries (LMICs). We conduct two sets of analyses. First, we document global and regional-level associations between the Human Development Index (HDI) and novel indicators reflecting multi-dimensional family change. Second, we use methods from the growth convergence literature to examine whether – and in which domains – there is evidence of cross-country convergence in family indicators over levels of development. We show that families in LMICs have transformed in multiple ways, changing differently across domains, world regions, and genders. Fertility, intra-couple decision-making, and women's life-course timing indicators are strongly associated with HDI, yet cross-country convergence is limited to the latter domain. Marriage, cohabitation, household structure, and men's life-course timing indicators are more weakly associated with HDI, and span a broad spectrum of convergence dynamics ranging from divergence to modest convergence. We describe this scenario as "persistent diversity with development," and shed light on the underlying regional heterogeneity – driven primarily by sub-Saharan Africa.

1.1 Introduction

The family remains a fundamental building block of human societies, affecting health, reproduction, and well-being of both present and future generations. Decades of sweeping demographic, economic and social change have radically transformed forms, gender roles, power relations, and intergenerational bonds of families worldwide (Bianchi 2014; Furstenberg 2014) – initially in high-income countries (HICs), and more recently in low- and middle-income countries (LMICs). At the global level, however, the process of change in families and family domains is inadequately understood (Ruggles 2012; Therborn 2014). This gap in knowledge about Global Family Change (GFC) is striking when compared to discrete socio-demographic events such as fertility, mortality, migration, and education, for which high-quality data exist across most world regions (e.g., UN World Population Prospects, Global Bilateral Migration Database, and Global Human Capital Project). No equivalent data resource exists for GFC. This paucity of comparable data capturing variation in family patterns over time and across space has limited scholars' capacity to evaluate theories of GFC and its driving forces, and assess the interactions between GFC and broader social and economic development.

Transformations in families are a late-comer in social and economic changes occurring during the demographic transition, as declines in fertility and mortality are often preconditions for substantial systemic change that is associated to alterations in the life course of family members. While a lack of focus on family change might have been acceptable during the initial stages of the demographic transition, this is no longer the case. The preconditions for fundamental transformations of the families exist globally, and GFC has emerged as a central aspect of global social change. An improved understanding of GFC therefore constitutes a critical step in social scientists' research on global change.

Several potential drivers of GFC have been identified during recent decades, all of which are particularly relevant in LMICs: the largest-ever cohort of youth currently entering adulthood; dramatic technological change; rising economic uncertainty; longer lives and lower fertility; narrowing gender gaps in schooling and the labor market; globalization forces affecting the flow of information, goods and people across the globe. Families have adjusted in diverse and sometimes surprising ways to these forces (Therborn 2004). Arguably, the transformation of the family that has occurred across high-income countries since the 1960s is currently entering its peak in LMICs. But GFC in LMICs is unlikely to be a simple extension of patterns observed in high-income countries (Furstenberg 2013). Heterogeneity in social, institutional, cultural and legal contexts, and differences in roles and functions of families may result in a diversity of GFC patterns that far exceeds the divergences in family trajectories that have been documented to date (Breen and Buchmann 2002).

This paper provides a broad empirical overview of the relationship between family change and socio-economic development drawing on 30 years of Demographic and Health Survey (DHS) data from 3.5 million respondents across 84 LMICs. We seek to advance the understanding of the changes in families that are occurring in many economically less developed nations by focusing on a set of agreed-upon domains that are key to the notion of the "family," namely fertility, timing of life-course events, union formation (marriage and cohabitation), household structure, and intra-couple decision-making.[1] To this end, we first document global and regional-level associations between the Human Development Index (HDI) and novel indicators reflecting multi-dimensional family change. Second, we examine whether there is evidence of cross-country convergence in these indicators over levels of

[1] We acknowledge that family change stems from the combination of complex, multiple, and concurrent processes that operate at different levels of analysis and cannot be fully captured through DHS data (e.g., coresidence with elderly). Yet we aim in this paper to measure change in some core dimensions that are well-understood and agreed-upon as being central to the functioning of families. By focusing on these family domains, we by no means claim that these provide an exhaustive picture of the phenomenon – family change – under investigation, yet we claim that they provide a good starting point.

development and, if so, in which family domains. Our analysis draws on a growing literature on whether fertility is converging across contexts (Casterline 2001; Dorius 2008; Wilson 2001, 2011), and extends this literature to broader domains, such as timing of life-course events, union formation (marriage and cohabitation), household structure, and intra-couple decision-making.

Our focus on GFC in this study takes the perspective of women (and men, partly) in young- and primary-adult (i.e., reproductive) ages, rather than old ages. The motivation is threefold. First, family change at these ages is tied to socio-economic considerations relating to household production and investments in children that are of particular relevance for LMICs. Second, existing paradigms of family change, such as the Second Demographic Transition (SDT) theory, have generally focused on young- and primary-adult ages. Third, our aim to consider multiple dimensions of GFC in the largest possible universe of countries requires the use of DHS data, which are generally restricted to ages 15–49. We also focus on country-level analyses, as no similar comparative GFC studies exist to-date, and sub-national analyses will be addressed in subsequent GFC research.

Our analysis makes three important contributions. First, we extend the literature to LMICs, where comparatively less is known about cross-country patterns in family transformations. Second, we rely on a wealth of micro-level data to compute innovative indicators accounting for key sources of demographic variability such as changing age distributions and increasing life spans. Third, we extend the fertility convergence literature to look at convergence trends in multiple family domains. In so doing, we move beyond most of the attempts at conceptualizing convergence in families, which have so far been embedded into fertility-related discussions (Casterline 2001).

Our framework of analysis is summarized in the stylized graph below, along with a general overview of the findings (Figure 1.1). The plane is comprised of four quadrants defined by different combinations of weak/strong associations with HDI (x-axis, from left to right) and divergence/convergence patterns over HDI (y-axis, from bottom to top). Our diagrammatic representation points to each quadrant being occupied by at least one family domain, suggesting that families in LMICs are distinct in many possible ways, and changes in families with development occur differently across domains. A strong association with HDI is observed for fertility, intra-couple decision-making, and women's life-course timing indicators, yet cross-country convergence over HDI is limited to the latter domain (top-right quadrant). The remaining domains are more weakly associated with HDI, and cover a broad spectrum of convergence dynamics ranging from divergence (marriage) to modest convergence over HDI (men's life-course timing, cohabitation, and household structure indicators). We refer to this heterogeneity as "persistent diversity with development." In what follows we further describe and categorize this diversity, shedding better light on the underlying regional heterogeneity and the key role of sub-Saharan Africa in departing from the overall trends.

1.2 Background

Global families, ***quo vadis?***[2]

More than fifty years ago, Goode's World Revolution and Family Patterns predicted that, as a consequence of industrialization, family patterns would globally converge to a prevalence of the "conjugal family form" of the West (Goode 1963). He concluded that individuals had become less dependent on extended family groups during the industrial revolution in the West, and hypothesized that other societies would go through the same family changes as they, too, went through the industrialization and urbanization processes. According to a 50-year analysis of global family change, Goode's prediction about the convergence of family systems in high-income countries and the rising prevalence of conjugal families has not been realized (Cherlin 2012). Over the second half of the twentieth century the Western family became complex in ways that Goode did not anticipate (e.g., the rise in cohabitation, single-parent families, and stepfamilies). Furthermore, Goode's implicit assumption that all developing countries would follow the same path to industrialization did not come to fruition (Ruggles and Heggeness 2008). Sub-Saharan Africa (SSA) has been largely left behind due to economic crises and the AIDS epidemic that stressed the SSA family; the Middle-East and South Asia have struggled; progress in Latina America and Southeast Asia has been uneven, and only in East Asia has consistent economic growth occurred, along with changes in families consistent with Goode's hypothesis of declining family control (Cherlin 2012).[3]

More recently, scholars argued that changes in families are driven by a diffusion of new ideas about family forms via social networks, language and culture-based networks, and global communication networks (Johnson-Hanks et al. 2011). The idea of change in families being driven by the interaction of socioeconomic development and ideational change has manifested itself in the theory of "developmental idealism," i.e., the notion that in many societies around the world the Western family is associated with a higher form of development (Thornton 2001, 2005). Thornton demonstrates that the developmental paradigm included the belief that Western Europe had transitioned from a predominantly extended-family system to a predominantly nuclear family system during its progression through the stages of development (Cherlin 2012). A central tenet of this framework posits that

[2] "Quo vadis" is a Latin expression meaning "Where are you heading?".
[3] Goode did not really focus on Latin America, where signs of a Second Demographic Transition (SDT) can be seen even before the first transition is complete (Cherlin 2012).

modern social structures and modern family behaviors have reciprocal causal influences. Convergence in family systems is an implication of developmental idealism, in that one of the central values of the framework is the desirability of a modern family – compatible with the needs of an industrialized society – along the lines of the nuclear Western family form.

However, the past half century has shown that there is not a singular pattern of family behaviors that constitutes "the modern family," and thus the potential endpoint of converging family systems. Rather than convergence, scholars are increasingly emphasizing the continuity of long-standing differences in family patterns, functions and behaviors across the world. At least in the area of fertility, divergent demographic trajectories have started to characterize high-income societies (Billari and Wilson 2001; Rindfuss, Choe, and Brauner-Otto 2016). Such persistent heterogeneity is also predicted by Therborn (2004, 2014), who postulates that two aspects of family change are certain: 'First, the family pattern will look different in different parts of the world, and the future will offer a world stage of varying family plays. Second, the future will not be like the past' (2014, p. 3). Therborn argues that three social phenomena drive family change differently in different parts of the world – namely the decline of patriarchy, changes in marriage forms and prevalence, and fertility decline – thereby producing seven family systems that differ by different combinations of these factors.[4]

The emerging hypothesis of a "convergence to divergence" in global family systems is not adequately reflected in other conceptual frameworks that are often invoked for guiding analyses of family change, including the two dominating lines of family research starting from work on the West: Becker's New Home Economics (Becker 1981; Becker, Landes, and Michael 1977) and Lesthaeghe and van de Kaa's SDT theory (Lesthaeghe and van de Kaa 1986; Lesthaeghe 2010, 2014). Both frameworks predict a weakening of the family due to women's pursuit of career and ideational changes valuing individual autonomy and self-fulfillment. According to the SDT, ideational change triggers declines in fertility, which set in motion other life-course transformations such as less and later marriage, a multitude of living arrangements, a disconnection between marriage and procreation, and increased women's independence inside and outside of unions. The SDT theory correctly anticipated the unfolding of different patterns of partnership formation, the shift in value orientations that emerged as driving forces in childbearing decisions – such as attitudes about politics, sex, religion, and education – and the emergence of sub-replacement fertility as a lasting feature of advanced societies. Albeit starting as a Western theory, it has been recently expanded to other regions such as East Asia

[4] These are the *Christian-European* family, the *Islamic West Asian/North African* family, the *South Asian* family, the *Confucian East Asian* family, the *sub-Saharan African* family, the *Southeast Asian* family, and the *Creole* family.

(Lesthaeghe 2010) and Latin America (Esteve, Lesthaeghe, and López-Gay 2012), and broadened to account for the role of path dependency and geo-historical legacies (Esteve and Lesthaeghe 2016). Whether the theory applies to other regions that still lag behind in the demographic transition (e.g., SSA), or where religion protects and reinforces patriarchal kinship organizations (e.g., Islam) is still an open question (Cherlin 2012).

Lastly, a very recent line of family research (Esping-Andersen and Billari 2015; Goldscheider, Bernhardt, and Lappegård 2015; McDonald 2000) has theorized a new phase of family life characterized by a profound "gender revolution" that leads men and women to not only participate in the public sphere in equal terms, but also to share household and childrearing tasks more equitably than in the past. This theoretical foundation posits a long-run return to "more family" as gender egalitarianism gains increasingly normative status, and the predictions of fewer marriages, children, and greater couple instability show signs of reversal. Again, while the theory has proved applicable to some European countries, it is unclear whether it extends beyond European borders. If this theory has any applicability to LMICs, we will likely observe changes over the next decades. As of now, most LMICs might still be undergoing the "first half" of the gender revolution, characterized by dramatic growth in female labor force participation rates which bring challenges to family formation and union stability (Goldscheider, Bernhardt, and Lappegård 2015). As trends are still underway, it might thus be premature to theorize a new "rosy" future for family life, especially at the global level (Cherlin 2016; England 2010).

Convergence in a demographic perspective

As the foundations of the convergence concept were initially developed within a demographic transition framework (Chesnais 1992), scholarly research to date has mostly dealt with mortality (Goesling and Firebaugh 2004; Janssen et al. 2016; McMichael et al. 2004; Montero-Granados, de Dios Jiménez, and Martín 2007; Neumayer 2003, 2004) and, to a lesser extent, fertility studies of convergence (Casterline 2001; Coleman 2002; Crenshaw, Christenson, and Oakey 2000; Dorius 2008; Wilson 2001, 2011).[5] Combined with modernization theory, this framework postulates that fertility and mortality rates vary overtime in a predictable and uniform manner, and less developed countries would follow a

[5] The formal assessment of convergence trends in demographic behaviors is a fairly recent development, with some of the most influential contributions dating back to the early 2000s. For studies of convergence in health and life expectancy, see Goesling and Firebaugh (2004), Janssen et al. (2016), McMichael et al. (2004), Montero-Granados, de Dios Jiménez, and Martín (2007), and Neumayer (2003, 2004). For studies of convergence in fertility, see Casterline (2001), Coleman (2002), Crenshaw, Christenson, and Oakey (2000) Dorius (2008), and Wilson (2001, 2011). For studies of convergence in broader living standard indicators such as wealth, educational enrolment, literacy, and television availability, see Jordá and Sarabia (2015), Kenny (2005), and Neumayer (2003).

path of economic and social progress similar to the one observed in more developed countries, thereby eventually converging in their fertility and mortality rates. The gradual transition from a high fertility and mortality scenario to one characterized by low vital rates would hence constitute the clearest example of demographic convergence (Salvini et al. 2015), pushing the world towards a new "demographic equilibrium" (Wilson 2001).

Despite clear theoretical predictions, the most puzzling aspect of existing studies of global demographic convergence is the ambiguous nature of their findings. In the area of mortality, global convergence has been modest throughout the past half century, and has been replaced by divergence since the late 1980s (Goesling and Firebaugh 2004; Moser et al. 2005), due in large part to declining male life expectancy in Eastern Europe, and the spread of HIV/AIDS in sub-Saharan Africa. In the area of fertility, research points towards high levels of inter-country and intra-regional variation in the pace of fertility decline (Casterline 2001). The only conclusive statistical evidence for convergence in fertility is found after the mid-1990s, about two decades after the onset of the decline in world mean fertility (Dorius 2008).

Most important from the perspective of this paper, the demographic literature on convergence – with its heavy focus on the drivers of the demographic transition – shows evident gaps when it comes to embedding additional family dimensions in the picture.[6] Most of the attempts at conceptualizing family change to date have been incorporated into broader discussions of fertility dynamics (Casterline 2001; Ram 2012; Skinner 2014). For instance, it is undisputable that domains such as union formation and fertility are closely tied, and union formation remains a precondition for fertility in several contexts. Yet the shift in focus onto understanding global social change requires endorsement of the idea that changes in multiple domains of the family matter independently of changes in fertility (van de Walle 1993). For instance, the changing frequency of the types of unions that occur in a society – customary or civil marriages with full social recognition versus informal or temporary unions – may influence the prevalence of female-headed households and the economic environment of children. Similarly, delays in the age at which men and women marry has implications for the organization of family life and for gender relations within society (Mensch, Singh, and Casterline 2005).

Scholars have recently attempted to assess convergence in family dimensions such as marriage, cohabitation, divorce, though rarely has the focus been comparative and LMICs-oriented. For instance, Lundberg, Pollak, and Stearns (2016) documented diverging patterns in marriage, cohabitation, and childbearing over the last 60 years, but their focus is on comparing within-country trends among

[6] A similar claim could be made for convergence studies in the area of migration.

population subgroups – more versus less-educated in the United States – rather than countries or world regions. Conversely, Billari and Wilson (2001) and Billari and Borgoni (2002) carried out cross-country comparative analyses of convergence in family dimensions and transition to adulthood markers such as early home leaving, timing of first union, total first marriage rate, and total divorce rate, but their focus is on European countries exclusively. To our knowledge, the present paper is the first comparative study to formally assess convergence in family indicators across multiple domains in LMICs. Our motivation is not to provide a further rejection of Goode's model of family convergence, but to identify which domains have been converging and which have not, in order to build a more comprehensive and theoretically robust framework for understanding GFC in LMICs.

1.3 Data and measures

Data

This paper uses repeated cross-sectional DHS data from 84 LMICs across five world regions (Figure 1.2), namely Americas, Asia, Former Union of Soviet Socialist Republics (USSR), Middle East and North Africa (MENA), and sub-Saharan Africa (SSA). For the countries included, their regional classification, and the number (and year) of survey waves per country, see Appendix Table A1.1.[7]

Our analyses include micro-level information from 293 DHS survey waves – on average five waves per country – collected between 1985 and 2016. We combine socio-demographic information from these surveys with HDI time series provided by the Human Development Report Office of the United Nations Development Programme (UNDP 2015). The HDI is a summary measure of average achievement in three key dimensions of human development: living a long and healthy life ("health and life expectancy"), being knowledgeable ("human capital"), and having a decent standard of living ("wellbeing").[8] The HDI is the geometric mean of normalized indices for each of the three dimensions, and it ranks countries into four tiers of development, namely "very high" (>0.8), "high" (0.7- 0.799),

[7] DHS samples are nationally representative cross-sections of the population selected via a two-stage procedure in which the primary sampling units most often correspond with census enumeration areas. Households are then randomly selected within each primary sampling unit, and all women aged 15-49 residing in the selected households are invited to participate. In some countries, a sample of men between the ages of 15 and 59 are also interviewed. DHS are collected by ICF Macro in collaboration with host country governments, and collect detailed information on fertility and marriage histories, population health, family planning, and anthropometrics. Standardization of survey questionnaires allows for comparability across countries and survey waves (ICF Macro 2009).

[8] There is by now consensus that development is a multidimensional concept, which, in addition to income, also should consider social indicators. This line of argumentation has gained prominence among academics over the last decades, thus resulting in many attempts to synthesize different aspects of well-being, in a composite index which offers a more comprehensive perspective of such a process than per capita income alone (Jordá and Sarabia 2015).

"medium" (0.550-0.699), and "low" (<0.550), thus providing a standardized measure of well-being across diverse contexts.[9] We acknowledge that the composite nature of the HDI makes it less obvious to appreciate which socio-economic factors are more closely associated with family indicators, yet we rely on the HDI as the primary index used by the UNDP to monitor broadly-defined development goals (UNDP 2015). In doing so, our work aligns with a long stream of scholarly tradition in sociology, demography, and economics (Bijwaard and van Doeselaar 2014; Bongaarts and Watkins 1996; Bystrov 2014; Harttgen and Vollmer 2014; Kreidl and Hubatkova 2014; Jordá and Sarabia 2015; Myrskylä, Kohler, and Billari 2009; Tanaka and Johnson 2016). In the Appendix, we provide ancillary analyses using the three HDI-components separately, most of which support the conclusions reached using an HDI-based approach.

In sum, our work encompasses all available DHS surveys 1985-2016, drawing information from about 3.5 million respondents, and representing the most comprehensive dataset for which these analyses are possible. By relying on high-quality surveys that provide comparable measures for a well-defined universe of countries – such as the DHS – we face an obvious trade-off between country coverage and data quality. This resulted in the exclusion of important LMICs for which no DHS is available, such as China. We acknowledge that this is a limitation of our study, and we aim to extend GFC-analyses to China and other excluded LMICs in subsequent research. Yet we show in Appendix Figure A1.1 that the sample of DHS countries included in the analysis well covers the complete range of the development spectrum of LMICs.

GFC-indicators

We focus on five family domains: fertility, timing of life-course events, union formation (marriage and cohabitation), household structure (vertical, or inter-generational relationships), and intra-couple decision-making (horizontal, or intra-generational relationships). This multifaceted conceptualization reflects the complexities and interrelatedness of global family change, with 20 indicators classified along three conceptual axes of analysis (Figure 1.3): Family Events and Behaviors (FEB), Linked Lives (LL), and Life-Course Patterns (LCP). Best documented to date have been changes in indicators of family events and behaviors, such as increases in unmarried cohabitation prevalence, and delays in the timing of marriage and/or onset of sexual intercourse (Bongaarts, Mensch, and Blanc 2017; Hayford, Guzzo, and Smock 2014; Manning, Brown, and Payne 2014). Indicators of linked lives (Elder 2001) illustrate the extent to which GFC transforms social relations, both within and across generations. Life-course pattern indicators combine measures of fertility, marriage, and cohabitation with information on

[9] We imputed missing values on the HDI for specific years using linear interpolation.

mortality conditions to capture the average number of years spent in different family "constellations" across the adult life-course. The rationale behind this latter group lies in the premise that many LMICs have experienced increases in life expectancy during the last three decades (UN-DESA Population Division 2017), driven in part by declines in young- and adult-age mortality. These mortality changes are important for understanding family change, for instance, because the person-years spent married by individuals can increase despite a delay in entering marriage and/or increased rates of marriage dissolution. Similarly, the number of surviving children decreases less than fertility, or might even increase, if development is also associated with declines in infant mortality. Prior studies of family change have not extensively featured the role of mortality on life-course family patterns (Bianchi 2014; Hagestad 1988). Overall, this multidimensional framework permits us to enrich the GFC literature by looking at global changes in domains other than fertility, and constructing indicators that have seldom been emphasized in previous scholarship.

Indicators of family events and behaviors include the Total Fertility Rate (TFR), marriage and cohabitation prevalence, and a set of sex-specific timing indicators measuring mean ages at three critical life-course events – first sex, first marriage, and first birth – commonly used in prior analyses (Bongaarts, Mensch, and Blanc 2017). Following prior studies (Bongaarts and Blanc 2015; Clark and Brauner-Otto 2015), timing indicators are estimated using singulate mean ages (SMAFS: Singulate Mean Age at First Sex; SMAFM: Singulate Mean Age at First Marriage; SMAFB: Singulate Mean Age at First Birth), a methodology first developed for mean ages at first marriage (Hajnal 1953), and then adapted to the estimation of mean ages at first birth in contexts that lack accurate vital statistics (Bongaarts and Blanc 2015; Casterline and Trussel 1980; Booth 2001). The SMAFB relies on age-specific proportions childless, and is defined as the average length of life with no children among those who have children before age 50. Analogous definitions and data requirements apply for the SMAFS and SMAFM.

Indicators of linked lives are constructed combining information from the women's file with information from the household roster. Intra-generational indicators are obtained from the women's file and include the share of households in which the husband is the sole decision maker on women's health, household purchases, and women's visits to family and friends. Conversely, inter-generational indicators include the share of children living with both parents, and a set of three indicators measuring the share of women living in a nuclear (a household where only one couple resides, either with or without kids), three-generation (a household where at least one member of the household roster reports that his/her relationship to the household head is one of grandparent or grandchild), and complex household (a household where at least one member of the household roster reports a lateral relationship with the household head). While the former indicator is obtained from the household roster only, the latter three are computed following a multi-step process that draws information on women's living arrangements from both a woman and a household-level perspective. Essentially, in instances in which the household roster does not provide conflicting or additional information with respect to the categorization based on the woman's perspective, the latter categorization is kept. Otherwise, both perspectives are combined to come up with a more precise categorization.[10] By computing these indicators, we do acknowledge that our aim in this paper is not to capture coresidence patterns in society, but rather to capture the household context of women in reproductive ages – which may significantly differ from broader coresidence patterns in society described, for instance, in Ruggles (2009, 2010) and Ruggles and Heggeness (2008). For instance, we acknowledge that our indicators are a poor measure of coresidence with the elderly, which would require data on individuals aged 65 or older – which the DHS does not provide by design.[11]

Life-course pattern indicators – comprising the Net Reproduction Rate (NRR), marital expectancy at age 15, cohabitation expectancy at age 15, and marital and cohabitation expectancy at age 15 – are constructed using mortality information from the UN Population Division's World Population Prospects: The 2015 Revision. We combine every DHS survey wave with available life tables from the year closest to the survey. We compute marital and cohabitation expectancies at age 15 using the Sullivan method (Imai and Soneji 2007), widely used to estimate healthy life expectancy, yet rarely applied to the family realm. State-specific life-expectancies (or person-years spent in marriage, cohabitation, etc.) are obtained via $e_x^i = \frac{1}{l_x}\widehat{\sum_x}\ {}_n\hat{\pi}_x^i\ {}_nL_x$ where ${}_n\hat{\pi}_x^i$ are age-specific proportions in a certain state i computed from a survey (e.g., proportion of married women between ages 25 and 30),

[10] In terms of household classifications, if the household includes more than one couple, then it is not classified as *nuclear* but *complex*. Households classified as complex are essentially those that deviate from the nuclear case. Note that a three-generation household can also be complex, and vice versa. The multi-step process followed to reach the above categorization combines information on women's living arrangements obtained from both a woman and a household-level perspective. It proceeds as follows: (i) We first classify women in categories based on the information provided in the women's file; (ii) We identify the household context using information on household members from the household-level file; (iii) We combine the woman and household-level perspectives in such a way that the categorization based on the woman's perspective is kept if there is no additional or conflicting information provided in the household roster (otherwise the two are combined); (iv) We merge back to the woman-level to obtain estimates at the individual level (e.g., share of *women* living in nuclear households), rather than at the household level (e.g., share of nuclear *households*).

[11] It is not clear that the patterns that pertain to the 65+ population extend to the younger population, which is the focus of our paper. Actually, almost certainly the patterns are quite distinct. Therefore, the likely divergence of our findings to those in Ruggles is due to a focus on different parts of the life-course (broadly reproductive ages in our case, ages 65+ in Ruggles' case), and the different patterns are not necessarily contradictory. In short, we acknowledge that we are likely to miss a considerable amount of co-residence by relying on DHS data, yet we can get some other relevant information from the perspective of women in reproductive ages.

and l_x and ${}_nL_x$ are period life-table quantities. Expectancies are computed at age 15 as the DHS provide no data below 15.

Our indicators are age-standardized to eliminate influences resulting from age-structure differences across countries.[12] Age-structure differences can affect many indicators, such as the proportion married or the share of women living in a specific type of household, because age-groups are differentially weighted in country-averages. Age-standardization is thus critical – yet rarely adopted in comparative family studies – to ensure that comparisons identify the extent to which observed differences in indicators across countries are due to social processes underlying the changing nature of families, rather than driven by demographic considerations such as changing age structures. Age-standardized indicators are computed by combining age-specific proportions from DHS micro-data with national age-structure data provided by the UN Population Division's World Population Prospects: The 2015 Revision. All indicators are standardized to the 2000 age distribution for less developed countries excluding the least developed (2000 is the average survey year in the sample). Supplemental analyses show that age-standardization significantly shifts the observed distribution of age-sensitive indicators. For instance, cross-regional differences in marriage prevalence are narrower after age-standardization (Figure A1.2 in the Appendix), suggesting that findings based on crude indicators could lead to overstate (understate) the role of behavioral (compositional) factors underlying family changes. Descriptive statistics of GFC-indicators are shown in Table 1.1.[13] All indicators and analyses will be made publicly-available.

1.4 Associations between GFC-indicators and HDI

Analytical strategy

Our methodological approach proceeds in two stages. First, we conduct a series of descriptive analyses of family change indicators over levels of socio-economic development – as measured by HDI – in a spirit similar to Myrskylä, Kohler, and Billari (2009) and Anderson and Kohler (2015). These exploratory investigations – purely descriptive and associational – are crucial to highlight comparative

[12] Some indicators are already age-standardized, such as the TFR. Others were age-standardized by us. Age-standardization was the main reason why we delved into the micro-data to compute our estimates, rather than simply borrowing aggregate information from online platforms such as the DHS StatCompiler.

[13] We would like to caution the reader that the number of observations (country-years) varies by indicator (as also mentioned in the note to the table), hence there might be some country-year observations included in the computation of, for instance, the SMAFB but not in the SMAFM. It follows that the comparison of the averages between indicators needs to be interpreted with caution. For instance, the Table shows that in MENA the SMAFM (23.49) is much higher than the SMAF (19.85). However, once the same set of country-year observations are retained the SMAFB increases to an average that is almost the same as the SMAFM.

macro-trends in family patterns, such as the clustering of countries or regions in specific domains. We plot each indicator against HDI, and assess whether a linear approximation summarizes the association reasonably well. To ease visualization and enable comparability across measures, we summarize the 20 scatter plots in a single graph that reports standardized associations (slopes) from a linear regression of each indicator on HDI. Indicators and the HDI are standardized on the pooled sample so that coefficients reflect changes in indicators measured in standard deviation (SD), per one SD change in HDI. Standard errors are clustered at the country level, and estimates are weighted for the number of survey waves by country to account for the fact that some countries have repeated observations (e.g., 11 waves in Peru) while others do not (e.g., one wave in Myanmar).

Although avoiding any claims of causality, we test the robustness of the associations using both contemporaneous and lagged values of HDI (the latter reported in the Appendix) to – at the very least – assuage endogeneity concerns due to reverse causation. Also note that throughout the study we provide both global and regional evidence. Whenever we provide regional evidence, we remove one region at a time rather than running separate analyses by region, as for some regional groupings the number of country-years would be too limited to warrant an adequate sample size. By excluding one region at a time we are able to preserve sample variability and appraise the contribution that each excluded region provides to the overall association/coefficient. This approach is similar to Dorius (2008).

Results

Figure 1.4 summarizes the association between HDI and the 20 GFC-indicators. Indicators are grouped by color and shape, following the three-way conceptual classification in Figure 1.3. Corresponding scatter plots are provided in Appendix Figure A1.3. Note that the indicators in this graph – reported on the vertical axis – have been rephrased in terms of "trends," as we are interested in comparing the strength of the positive association of the indicators with HDI. For instance, as the TFR is negatively associated with HDI, we rephrased the indicator "TFR" as "Reduction in TFR". Each marker corresponds to the coefficient of a regression of the respective GFC indicator on HDI. Filled markers refer to statistically significant estimates ($p\text{-value}<0.05$), and larger markers indicate more precisely estimated associations. The detailed regression estimates are provided in Appendix Table A1.2 (panel [a]), along with robustness checks using lagged HDI values (panels [b] and [c]).

Our analyses corroborate the well-established finding that increased socio-economic development is associated with lower fertility (Bryant 2007; Myrskylä, Kohler, and Billari 2009). A one SD increase in HDI – corresponding to approximately a 10-point increase in HDI on a 0-100 scale –

is associated with a 0.65 SD reduction in TFR. Accounting for mortality as reflected in the Net Reproduction Rate (NRR) weakens the association by about 0.1 SD, suggesting that reductions in infant mortality make the number of surviving children to decline less than overall fertility levels. The association of fertility with HDI is the strongest among those considered in this study, followed by decision-making indicators (horizontal linked-lives indicators) – in areas as varied as women's health, freedom of movement, and purchases for the household – and timing indicators measured by the SMAFS, SMAFM, and SMAFB. For the latter, important gender differences emerge: statistically significant associations between timing indicators and HDI are found for women, while the relationship is weak and statistically insignificant for men. SMAFM is associated most strongly with development, followed in turn by SMAFB and SMAFS. This divergence is likely due to (mostly first) births increasingly occurring outside of marriage.

Coefficients on indicators of family events and behaviors are aligned with SDT predictions of lower fertility, delayed markers of adulthood, and increased women's autonomy associated with increases in HDI. As a complement to the SDT, which is mostly silent on gender convergence in the transition to adulthood, we also provide evidence of reduced gender differences: as women tend to have earlier ages at first sex, first marriage and first birth than men, the stronger association with HDI for women as compared to men is suggestive of a trend towards converging gender patterns in the transition to adulthood. This diminishing sex discrepancy is consistent with global trends towards women's increasing commitment to education and labor force participation (Esteve et al. 2016).

Vertical linked-lives indicators are more weakly associated with HDI than the above family events and behaviors and horizontal linked-lives indicators. While a one-SD increase in HDI is associated with a 0.2 SD gain in the share of children living with both parents, there is no significant association between HDI and the share of women living in nuclear and three-generation households. This finding somewhat departs from theories postulating an inverse association between household complexity and socioeconomic development (Goode 1963; Le Play 1884), yet it is consistent with the stability in traditional family forms found in analyses of intergenerational coresidence across 15 developing countries by Ruggles and Heggeness (2008). We caution, though, that results are not fully comparable with census data from IPUMS, as with DHS we can only capture the living arrangements of women in reproductive ages.

Marital status indicators show that higher HDI is associated with a decline in marriage prevalence (0.2 SD reduction per 1 SD increase in HDI), while there is no statistically significant association with prevalence of cohabitation. Yet, once we account for period mortality conditions by computing person-years (PY) lived married, marriage shows remarkable persistence, as the association with HDI turns insignificant and its magnitude is reduced by more than 50 per cent (0.08 SD). Conversely, differences between prevalence of cohabitation and the average number of years spent cohabiting are minimal, due to the fact that increases in cohabitation are driven by coresidence at relatively young ages, where the mortality effect is weakest. Summing the number of PY spent in marriage and cohabitation delivers an even clearer finding: increases in HDI are not associated with declines in the number of years adults spend in unions. This finding is due to reduced adult mortality and the increased person-years lived as adults that have occurred in LMICs in recent decades. When seen through the lens of life-course pattern indicators (rather than family events and behaviors) such as person-years in marriage or cohabitation, our descriptive analysis therefore suggests that during the development process, families in LMICs have not been characterized by dramatic shifts in patterns of union formation. Overall, the associations are robust to replacing contemporaneous HDI with 2-year and 5-year lagged HDI values (Appendix Table A1.2, panels [b] and [c]), and to substituting HDI with its components (Appendix Table A1.4 and Figure A1.4).[14]

Figure 1.5 expands our analyses in Figure 1.4 to consider heterogeneity by region. Following the same approach, each line in Figure 1.5 reports five (rather than one) markers that correspond to the coefficients of a linear regression of the respective indicator on HDI, excluding one world region at a time (e.g., the red circle in the NRR line captures the association between the NRR and HDI on the pooled sample excluding the Americas). The cross (x) locates the global association provided in Figure 1.4. Filled markers indicate statistically significant estimates ($p<0.05$). Detailed regression estimates are provided in Appendix Table A1.3. For each line, markers that are clustered near each other indicate cross-regional homogeneity in the estimated association between the indicator and HDI. Close markers suggest that removing regions does not affect the association to a significant extent, i.e., no particular region is driving the association in either direction. This is the case for the TFR, with the association with HDI robust to the exclusion of each region, and stable around 0.6-0.7 SD. The case of the NRR is similar, although excluding SSA here results in a stronger association with HDI (0.63 SD versus the 0.54 SD shown in Figure 1.4). This is reasonable, as SSA has experienced substantial mortality declines along with increases in HDI in recent decades. Figure 1.5 reveals varying degrees of cross-regional homogeneity in women's timing and decision-making indicators. Homogeneity is particularly pronounced for women's SMAFM and intra-household decision-making

[14] Note that in panel [b] of Appendix Table A2 the HDI is lagged by two years – rather than one – because some demographic estimates are obtained from DHS surveys which were collected over two years. Therefore, by taking the HDI value two years before we make sure we are not taking any contemporaneous value.

on women's visits to family and friends, while for decision-making on women's health, excluding SSA would result in a stronger positive association with HDI (0.72 SD versus 0.55 SD shown in Figure 1.4).

The regional picture for vertical linked-lives indicators is complex, as the Americas contribute to making the positive association between HDI and the share of children living with both parents smaller than it would be, while the null or weak associations between HDI and the share of women living in nuclear and three-generation households is mostly driven by SSA and Asia. The case of SSA is also interesting in that the share of women living in a three-generation household is positively correlated with HDI Although further research (beyond the scope of this paper) is needed to untangle why this is the case, we suspect this is due to rapid demographic changes in the region, such as fastest mortality declines which increase young generations' opportunities to reside with parents or have a living grandparent (Ruggles and Heggeness 2008). The influence of the HIV epidemic and differential patterns of migration might also play an explanatory role. Lastly, consistent with prior findings on heterogeneous trends in men's ages at reproductive transitions (Bongaarts, Mensch, and Blanc 2017), the highest cross-regional heterogeneity in the associations is observed for men's timing indicators.

Overall, it is worth noting that coefficients excluding SSA are most often at the lower or upper extremes of each line (with Former USSR at the opposite end, yet somewhat less distal from the average), suggesting that SSA is the region that contributes the most to the observed heterogeneity. Excluding SSA, regional trends would depart less from global trends in the associations between HDI and family domains. For instance, in the absence of SSA we would observe stronger associations between HDI and NRR reduction, delay in mean ages at first birth, women's empowerment, decline in the share of women living in complex households, and decline in years spent in marriage or cohabitation.

Figures 1.4 and 1.5 combined suggest strong associations between women's family events and behaviors and horizontal linked-lives indicators and human development, with little cross-regional variability. Conversely, men's family events and behaviors, vertical linked-lives, and life-course pattern indicators are more weakly associated with HDI, and the associations show widespread heterogeneity. However, associational evidence of this kind provides little guidance to understand how change unfolds over advances in development. Specifically, it does not tell us whether LMICs are becoming more "similar" as HDI improves, and whether the extent of intercountry variability in family indicators narrows as countries move along the development path.[15] In what follows we carry out formal assessments of the convergence hypotheses outlined in the background section.

1.5 Convergence in GFC-indicators over HDI

Analytical strategy

To delve into dynamics of change, we complement the previous cross-sectional investigation with formal statistical analyses of whether there has been convergence in family indicators over HDI using approaches pioneered by Barro and Sala-i-Martin (1992), Sala-i-Martin (1996), and Dorius (2008) – see also Gächter and Theurl (2011), Jordá and Sarabia (2015), Salvini et al. (2015), and Janssen et al. (2016) for more recent applications. Differently from previous related scholarship (e.g., Dorius 2008), we explore whether convergence has occurred over levels of socio-economic development rather than time. Specifically, we test for convergence over HDI, and report parallel analyses for convergence over HDI per unit of time ("pace of development") in the Appendix (Table A1.6).

Our analysis tests for beta-convergence (β), that is, the catching-up of countries "lagging behind" in specific indicators. In line with Dorius (2008), we estimate β-convergence following equation 1, where ln is the natural log, subscript j refers to the jth country, $Y_{t_{i+n}}$ is the value of the demographic indicator observed in survey year i+n, Y_{t_i} is the value of the same demographic indicator observed in survey year i, $(HDI_{t_{i+n},j} - HDI_{t_i,j})$ is the difference in the value of the HDI between two repeated cross-sections (t_i and t_{i+n}) for the same country j, β is the convergence coefficient, α is the constant, and e_j is the error term for the jth country.[16]

$$\frac{ln\left(Y_{t_{i+n},j}\right) - ln\left(Y_{t_i,j}\right)}{\left(HDI_{t_{i+n},j} - HDI_{t_i,j}\right)} = \alpha + \beta\left(Y_{t_i,j}\right) + e_j \quad (1)$$

For every country, each previous cross-sectional survey forms the base measurement for the calculated growth rate. Hence, if a country has three repeated cross-sections, two growth rates are calculated over the corresponding periods. It follows that the set of country-years included in this second stage of the analysis is reduced (henceforth, "convergence sample"), as countries with only one survey are automatically excluded (i.e., 26 countries – and 26 survey waves – are excluded, resulting in

[15] We acknowledge that this study does not focus on within-country inequalities, which may contribute to explaining some of the outlined trends, yet further GFC research will pick up on this point.
[16] When we test for convergence over HDI per unit of time ("pace of development"), the denominator $(HDI_{t_{i+n},j} - HDI_{t_i,j})$ is replaced with $(HDI_{t_{i+n},j} - HDI_{t_i,j})/n$, where *n* is the number of years between the two repeated cross-sections.

a sample of N=267 country-year combinations and 58 countries). For an overview of the countries with only one DHS survey, see Appendix Table A1.1. In Appendix Table A1.5 we show that averages of the GFC-indicators between the overall sample (N=293) and the convergence sample (N=267) are quite aligned. As the DHS countries with one survey are primarily from the Former USSR region, some differences emerge for the TFR and NRR (higher in the convergence sample) and the timing indicators (lower in the convergence sample). Differences are, however, small, and unlikely to invalidate our findings. If anything, the convergence sample is more representative of the low-income (rather than low and middle-income) world.

A negative sign on the β-convergence coefficient indicates that lagging countries are catching up with leading countries, i.e., they are converging; a coefficient not significantly different from zero indicates that differences between countries are maintained, while a positive coefficient indicates that lagging countries are falling farther behind, i.e., they are diverging. Applied to fertility, for instance, β-convergence occurs when the rate of decline among countries with high fertility is greater than the rate of decline among countries with low fertility. In line with the literature, we complement tests for β-convergence with analyses of sigma-convergence (σ), the reduction of between-country variability – as measured by the coefficient of variation (CV) – in indicators over HDI levels.[17] Sigma-convergence analyses are reported in the Appendix.

There is much debate in the literature on the appropriateness of using population weights in these types of analyses. Early studies of inter-country dynamics treated each country equally as the principal units of interest were economies. Subsequent scholarship suggested that whenever the research focus is on individuals, then countries should be weighted by population size/shares (Firebaugh 1999; Korzeniewicz and Moran 1997). In this study country-years are the main units of analysis, hence we opt for the former approach. Doing so ensures that a change in Y for a large country like India does not disproportionately affect the estimates as compared to a similar change for a smaller country like Malawi.[18]

[17] We assess trends in the repeated cross-sectional coefficient of variation (CV) for each family change indicator. The CV is the ratio between the standard deviation and the mean of each indicator at the country level. While there is no "cut-off point" for the CV to define high or low variability, this approach permits to comparatively assess which indicator displays more heterogeneity. A negative trend in the CV implies a decline in the variability relative to the mean, i.e., convergence; a flat trend implies differences are maintained; and a positive trend implies increasing heterogeneity, i.e., divergence. S*igma*-convergence analyses are a natural complement for β-convergence analyses in that the latter is a necessary but not sufficient condition for the former to hold (Barro and Sala-i-Martin 1992; Young, Higgins, and Levy 2008).

[18] In line with Neumayer (2004), Kenny (2005), and Dorius (2008), weighted and unweighted estimates differ significantly. As our focus is on country-years, we do not report the former (results available upon request).

Results

Table 1.2 reports results from a β-convergence model where the growth rate of each indicator over HDI is regressed onto its initial level.[19] Panel [a] provides global estimates, while panel [b] provides estimates on the sample excluding one region at a time, in an effort to isolate the contribution that each region provides to the overall global convergence coefficient. The left column of panel [a] reports coefficients from an unconditional β-convergence specification, while the right column controls for region-specific dummies to account for within-region heterogeneity, thereby providing conditional β-convergence estimates (Lall and Yilmaz 2001).[20] Conditional β-convergence estimates imply that pathways of convergence hinge upon the structural specificities of each region.

Despite the majority of coefficients are negative in sign – thereby pointing to convergence – statistical evidence of global convergence (panel [a]) is limited to a narrow subset of indicators, namely women's timing indicators (SMAFS, SMAFM, and SMAFB) and the number of person-years spent in unions (marriage plus cohabitation). Unconditional β-convergence estimates offer no evidence of a catching-up process in fertility (TFR and NRR), thereby aligning with Dorius's (2008) analysis of global convergence in fertility over time, which shows that knowing the initial value of the TFR for the average country tells little about subsequent fertility decline over a 50-year time span (except for limited periods such as the 1995-2005 decade). Controlling for regional dummies provides some evidence of within-region convergence for the NRR and, in line with Dorius (2008), our estimates align with the idea that sub-Saharan African countries exert a braking effect on the global β-convergence coefficient. The regional counterfactuals (panel [b]) indicate that, without these countries, the sign of the β-coefficient would turn from null/positive to negative, leading to a 1.3 per cent decline in the growth rate of the NRR over HDI in response to a one-unit increase in initial NRR. Conversely, excluding any other region would leave the global convergence coefficient virtually unchanged.

The idea that family change trajectories may follow different patterns by gender – suggested in Figure 1.3 – is confirmed by strong cross-country convergence in women's (but not men's) postponement of first sex (SMAFS), first marriage (SMAFM), and first birth (SMAFB). For instance, a one-year increase in initial SMAFM for women reduces the average growth rate over HDI by about 0.1 per cent. This gender discrepancy is likely to lead to growing similarities in transition to adulthood

[19] The computation of growth rates over HDI yields extreme values due to small changes in the HDI between survey waves (ΔHDI, i.e. the denominator of the rate). In our preferred specification, we exclude outliers that fall outside the first quartile (Q1) minus three times the interquartile range (IQR), and the third quartile (Q3) plus three times the IQR.

[20] With the addition of region fixed-effects we are assuming different intercepts for each region, but common speed of convergence (i.e. same slope).

patterns between the sexes, thereby affecting couple-formation strategies and patterns of assortative mating by age and education (Mensch, Singh, and Casterline 2005). Again, panel [b] provides some evidence for the unique role of SSA yet, differently from above, excluding SSA would result in weaker – rather than stronger – convergence coefficients. SSA countries are therefore seemingly speeding up the convergence process in women's timing indicators.

Horizontal and vertical linked-lives indicators show negative yet non-significant coefficients, hinting at persistent differences with development. In a way that parallels fertility indicators, LMICs would be more strongly converging with development in intra-household decision-making in the absence of SSA, suggesting that countries in SSA are lagging behind other regions when it comes to improvements in intra-household bargaining. Similarly, coefficients on vertical linked-lives indicators indicate that heterogeneity within MENA and Former USSR drives results away from convergence in the share of women living in nuclear and three-generation households, while heterogeneity within SSA drives results away from divergence in the share of children living with both parents.

Lastly, marriage prevalence is the only domain in which clear evidence of divergence over HDI is observed, implying that countries where marriage prevalence is high have experienced a relatively slower decline in marriage as compared to countries where marriage prevalence is low. Note that these divergence trends are driven by the Americas, and findings excluding these countries would be consistent with a scenario of persistent differences with development. Combined with the observation that women across the world are converging in their mean age at first marriage – and in cohabitation practices, at least within regions – these findings suggest that in the realm of union formation we are likely to observe the emergence of heterogeneous clusters of countries varying by different combinations of marriage prevalence and timing.

Beta-convergence estimates over HDI per unit of time – reported in Appendix Table A1.6 – fully confirm our evidence of (i) global convergence in women's (but not men's) timing indicators and person-years spent in unions, (ii) global divergence in marriage prevalence, (iii) within-region convergence in fertility and cohabitation, and (iv) the peculiar role of regions (mostly, SSA) in slowing down or speeding up global convergence patterns.

Analyses of σ-convergence over HDI – reported in Appendix Figure A1.5 – show a good degree of consistency with β-convergence coefficients except for timing indicators, which display too little variability to detect meaningful trends. The linked-lives decision-making indicators follow an inverted-U shape, confirming that sub-Saharan African countries – with mean HDI around 0.45 – contribute the most to divergence patterns in these family domains. These findings align with the idea that in the absence of countries in the lowest HDI tiers, LMICs would converge towards more equal intra-household dynamics. Moreover, while variability is increasing in the share of married women, it is unambiguously decreasing for cohabitation indicators, confirming trends towards convergence in cohabitation practices.

1.6 Conclusions and discussion

This paper has provided a comprehensive empirical assessment of the relationship between family change and socio-economic development drawing on 30 years of survey data from 3.5 million respondents across 84 LMICs. We conducted two sets of analyses. First, we documented – in a descriptive and cross-sectional way – global and regional-level associations between HDI and novel indicators reflecting multi-dimensional family change. Second, focusing on this same set of indicators we explored whether – and in which domains – families in LMICs have converged over levels of development. In this second effort, our analysis built on Dorius (2008), and extended his focus to domains other than fertility, thereby becoming the first study to expand the GFC literature to cross-country analyses of convergence across multiple family domains in LMICs.

We documented strong associations between HDI and women's indicators of family events and behaviors and horizontal linked lives, with little cross-regional variability. Conversely, HDI is more weakly associated with men's family events and behaviors, vertical linked-lives, and life-course pattern indicators, exhibiting regional idiosyncrasies that call for a more contextualized understanding of GFC. Although the emerging picture is complex and opens several research avenues, this simple analysis has value in that it questions the narrative that fertility and family change are closely synchronized during the demographic transition. Moreover, the conceptualization of family change adopted emphasizes the importance of interacting multiple axes of analyses – such as measures of both prevalence and timing, horizontal and vertical dynamics, and person-years in diverse constellations – with underlying demographic considerations such as increasing life expectancy and changing age structures. While in contexts like Europe or the United States neglecting mortality trends in studies of family change is likely to distort analyses to a minimal extent, our life-course pattern indicators point to a significant and oft-neglected role of mortality in shaping family trends across LMICs, particularly in SSA where recent mortality declines have been fastest. The "mortality effect" is most pronounced for union formation, where the combination of delays in marriage timing and increases in life expectancy suggest that the overall number of years spent in unions has remained unchanged across wide ranges of development.

Looking at the findings from a general theory of convergence, we provided evidence of global convergence in some dimensions of families over HDI, such as the timing of women's life-course events and person-years spent in unions (mainly driven by cohabitation), accompanied by persistent differences and/or divergence in other domains, such as marriage prevalence. With reference to the timing indicators, we identified clear gender differences whereby countries are converging towards delayed first sexual intercourse, first marriage, and first birth for women, but less so for men. Whether these changes are purely structural, because more women live in cities and have gained schooling – factors that tend to delay the passage into adulthood – or whether they represent a profound transformation in the patterns of early and universal marriage that affect the entire population is a question that cannot be convincingly settled through this initial analysis (though we expect to pursue this and related questions in future GFC papers). The changes are certainly linked with deep family transformations and are accompanied by, or perhaps in part caused by, increasing female independence inside and outside of unions.

Overall, our results combined suggest that families in LMICs have transformed in multiple ways, with changes occurring differently across domains, world regions, and sexes. The picture that best conforms to our findings is one of persistent diversity with development. Also, a point that is clear from the analysis is that development is not a powerful driver of convergence for all outcomes, suggesting the need to take into account additional factors that might contribute to explaining the observed heterogeneity. Among these, GFC scholars need to better consider geo-historical legacies and long-standing differences in social and economic institutions (Esteve and Lesthaeghe 2016), which play a role in shaping how globalization impacts upon life-course patterns. The main institutional considerations that might enlighten the understanding of the demography of adult life in LMICs are those which pertain to the education system, and the housing and labor markets (Furstenberg 2014; Grant and Furstenberg 2007). Some of these factors are embedded in indicators of human development, yet these macro-measures miss deeper elements underlying global social change, such as the path-dependent nature of institutions, institutional and cultural constraints, social norms, and the intangible role of diffusion processes in promoting or hindering change.

Our study has implications for theorization on global family change. First, our results engage with the convergence hypothesis advanced by Goode (1963) by suggesting a more nuanced categorization of convergence dynamics whereby convergence is "partial" and limited to specific domains. The idea that convergence towards the nuclear family type of the West has not occurred is not new and has been widely accepted by scholars over the past decades (Cherlin 2012, 2017). This paper enriches existing knowledge by providing further evidence of which domains are experiencing cross-country convergence. Second, our findings of a positive association between HDI and women's family events and behaviors and horizontal linked-lives indicators align with the SDT theory, reflecting the so-called "postponement" (i.e., the upward shifts in ages at marriage and first births) and "non-conformist" (i.e., the growth of alternative family formation strategies and the rise of individualization) transitions (Lesthaeghe and López-Gay 2013). However, the SDT proponents claim that the weakening of the institution of marriage is one of the main features of the SDT (Zaidi and Morgan 2017). Our focus on life-course pattern indicators accounting for mortality (PY) suggests that this conclusion is less likely to hold in the context of LMICs.[21] Third, our findings provide an empirical assessment of Therborn's postulated diversity of global "family systems" (2004). Although in this paper we do not deal with family systems – and we are hence not able to conclude that there are seven persistent family systems worldwide – our findings align with Therborn's idea that families are 'on the whole not converging and in some respects rather diverging; they will also characterize the world in the foreseeable future' (Therborn 2014, p. 3). Lastly, our findings on reduced imbalances in intra-couple decision-making and sex-discrepancies in timing of life-course events engage with the "gender revolution" framework predicting an increasing role of men in sharing household and childrearing tasks more equitably – which in turn translates into more gender-equal marriages and partnerships (Esping-Andersen and Billari 2015; Goldscheider, Bernhardt, and Lappegård 2015).

The emerging picture of persistent diversity with development – along with the GFC patterns that stem from the computation of innovative life-course pattern indicators – have important implications for understanding the social and economic consequences of global development and globalization, and should be considered in the policy for sustainable development and for increasing individual and family well-being. For instance, although development is associated with declines in fertility and delays in marriage, which in turn unfold along with more gender-equal dynamics within couples, unresolved remains the question of whether, globally, higher empowerment within the household translates into increasing female independence outside of unions, and how existing institutional support enhances female agency by reconciling family life and labor market opportunities. Given the heterogeneity of institutional, cultural, and policy contexts across LMICs, further research is required to investigate the extent to which gender equity in individual-oriented institutions combines with gender-equity in family-oriented institutions to sustain or hinder these trends (McDonald 2000). Other considerations stem from implications of family change for child wellbeing during development.

[21] Also, the SDT is rather silent on dimensions of household composition, such as our vertical linked-lives indicators.

Closely related to these policy considerations is the role of sub-Saharan Africa that emerges from this study. Countries in SSA – the region with the lowest levels of HDI – turn out to be those that contribute the most to the observed global heterogeneity, both in terms of associations of GFC-indicators with HDI and of β-convergence patterns. Figure 1.6 reproduces Figure 1.1 showing how the framework modifies if we exclude SSA from the convergence analysis. After removing SSA countries we move from covering the whole plane (i.e., four quadrants) to two polarized quadrants (bottom-left and top-right), thereby suggesting a significant decrease in "persistent diversity with development." The peculiar nature of SSA is apparent in the area of fertility (Bongaarts and Casterline 2013; Shapiro and Hinde 2017) and intra-couple decision-making, where detailed estimates show that the world would converge towards lower fertility and increased women's empowerment in the absence of SSA. We take this evidence as suggesting that further advances in dimensions of development such as education, literacy, income, and health within SSA might contribute to reducing disparities across regions and put the developing world on a more defined convergence trajectory. Although a convergence trajectory in these domains is not desirable per se, it is deemed beneficial to the extent that it is also conducive to longer human capital investments, higher female labor force participation rates, and increased compatibility between family life and economic success.

This study has important limitations that lay the ground for subsequent research. First, we acknowledge that in various world regions differentiating factors other than those captured in the UN HDI are at work. We recognize that the HDI may be subject to criticism for its "narrowness" and inadequacy in capturing all possible dimensions of development. The HDI has also been criticized on the grounds of construction (Kelley 1991), selection of variables (Srinivasan 1994; Alkire 2002), arbitrary weighting scheme (McGillivray and White 1993), and redundancy with its components (Cahill 2005; McGillivray 1991; Ravallion 1997). Yet our study is cross-country and comparative in scope and, as such, it encompasses multiple family dimensions across a wide range of countries. This inevitably requires reliance on a set of summary measures that are well-known and broadly understandable. Second, the synthetic-cohort nature of some of our indicators relies on the stationarity assumption, which may not hold when different cohorts undergo changes in family domains at different times and under different conditions. Although using age-specific proportions may be an alternative option, we believe there is no better measure a priori.[22] Third, it is likely that the quality of data pertaining to cohabitation and household composition might not be fully reliable, and the variables used might not be measured in exactly the same way and/or attributed the exact same meaning across

[22] Actually, if measurement error is not correlated with age, these all-age (15-49) measures might be more reliable than age-specific proportions.

time and space (Ruggles 2012; van de Walle 1993). DHS data were collected in a comparable manner in all countries, yet the differing cultural norms and practices regarding formation and dissolution of unions can affect the way in which respondents report their marital status (Shapiro and Gebreselassie 2014; Westoff, Blanc, and Nyblade 1994). This is particularly so given the well-documented multiplicity of conceptualizations and definitions of marriage (which has a direct effect on the definition of living together or cohabiting) in SSA (Hertrich 2002; Mokomane 2006). Similarly, some studies – mostly from the African context – have shown that survey data may be problematic when measuring household structure due to the country-to-country variability in the definition of households (Randall and Coast 2015; Randall et al. 2015). These concerns pose a threat to the validity of the estimates and suggest that findings in these areas need to be handled with care. Nevertheless, we embrace the view that for data quality to be improved in the future, presently available information ought to be used to produce comparative research, despite its flaws. Lastly, this study does not adequately incorporate the idea that countries are not independent entities but are part of an international system or network that extends across international borders which, by means of peer influence and concerted efforts (e.g., family planning programs), is likely to shape some family domains more than others (Cherlin 2012). Future GFC analyses will pick up on these important points.

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1.8 Tables

Table 1.1: Summary statistics on HDI and GFC-indicators, by region

	Type	Overall (N=293)		Americas (N=54)		Asia (N=43)		Former USSR (N=14)		MENA (N=24)		SSA (N=158)	
		Mean	(SD)	Mean	(SD)	Mean	(SD)	Mean	(SD)	Mean	(SD)	Mean	(SD)
HDI		0.52	(0.117)	0.60	(0.071)	0.55	(0.071)	0.66	(0.058)	0.54	(0.126)	0.44	(0.094)
GFC Indicators													
Total Fertility Rate (TFR)	FEB	4.14	(1.439)	3.53	(0.897)	3.23	(1.106)	2.29	(0.908)	4.06	(1.280)	5.14	(1.016)
Net Reproduction Rate (NRR)	LCP	1.72	(0.521)	1.59	(0.369)	1.43	(0.464)	1.05	(0.424)	1.76	(0.518)	2.02	(0.374)
SMAFS - women	FEB	19.24	(2.253)	19.97	(0.846)	19.48	(3.742)	21.58	(1.306)			18.41	(1.599)
SMAFS - men	FEB	19.46	(2.397)	18.23	(1.218)	19.86	(4.145)	20.38	(1.487)			19.51	(2.270)
SMAFM - women	FEB	21.82	(1.945)	22.03	(1.171)	22.23	(1.551)	22.35	(1.318)	23.49	(1.907)	21.40	(2.260)
SMAFM - men	FEB	25.90	(1.883)	25.19	(1.389)	24.72	(0.922)	26.40	(1.297)			26.15	(2.093)
SMAFB - women	FEB	21.48	(2.029)	22.54	(1.033)	20.80	(2.464)	23.92	(1.337)	19.85	(2.596)	21.04	(1.457)
SMAFB - men	FEB	26.48	(2.131)	26.40	(1.493)	24.31	(2.486)	28.13	(1.182)			26.79	(1.850)
Husband decides about women's health (prop. HH)	LL	0.30	(0.198)	0.13	(0.075)	0.22	(0.127)	0.12	(0.108)	0.24	(0.133)	0.40	(0.193)
Husband decides about household purchases (prop. HH)	LL	0.32	(0.178)	0.17	(0.065)	0.23	(0.111)	0.16	(0.124)	0.33	(0.117)	0.42	(0.166)
Husband decides about women's visits (prop. HH)	LL	0.24	(0.162)	0.10	(0.036)	0.16	(0.107)	0.12	(0.105)	0.19	(0.042)	0.34	(0.153)
Children living with both parents (prop.)	LL	0.68	(0.160)	0.65	(0.096)	0.82	(0.085)	0.82	(0.085)	0.91	(0.039)	0.58	(0.127)
Women living in a *nuclear* household (prop.)	LL	0.40	(0.114)	0.45	(0.093)	0.38	(0.096)	0.46	(0.066)	0.47	(0.046)	0.36	(0.124)
Women living in a *three-gen*. household (prop.)	LL	0.29	(0.090)	0.27	(0.048)	0.32	(0.066)	0.35	(0.085)	0.22	(0.092)	0.29	(0.097)
Women living in a *complex* household (prop.)	LL	0.36	(0.169)	0.29	(0.106)	0.31	(0.115)	0.19	(0.091)	0.20	(0.087)	0.45	(0.164)
Prevalence of marriage (prop.)	FEB	0.54	(0.186)	0.38	(0.117)	0.67	(0.084)	0.63	(0.063)	0.64	(0.054)	0.52	(0.213)
Marital expectancy at age 15 (yrs)	LCP	18.55	(6.427)	13.63	(4.027)	23.82	(2.540)	22.82	(2.153)	23.22	(1.689)	16.86	(6.962)
Prevalence of cohabitation (prop.)	FEB	0.14	(0.152)	0.23	(0.105)	0.02	(0.037)	0.02	(0.019)			0.14	(0.168)
Cohabitation expectancy at age 15 (yrs)	LCP	4.43	(4.999)	7.83	(3.586)	0.70	(1.236)	0.64	(0.645)			4.57	(5.410)
Marital and cohabitation expectancy at age 15 (yrs)	LCP	22.15	(3.113)	21.46	(1.346)	24.11	(2.225)	23.46	(1.854)	23.22	(1.689)	21.24	(3.756)

Notes: Estimates are weighted by the number of survey waves available per country. N refers to the number of country-year combinations. As the actual number of observations is indicator-specific (i.e., not the same set of country-years is included in the computation of each indicator), the reported N refers to the maximum number of observations per group across all indicators. FEB: "Family events and behaviors", LL: "Linked lives"; LCP: "Life-course patterns". Information on cohabitation, sexual intercourse, and timing indicators for men is not available for MENA countries.
Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

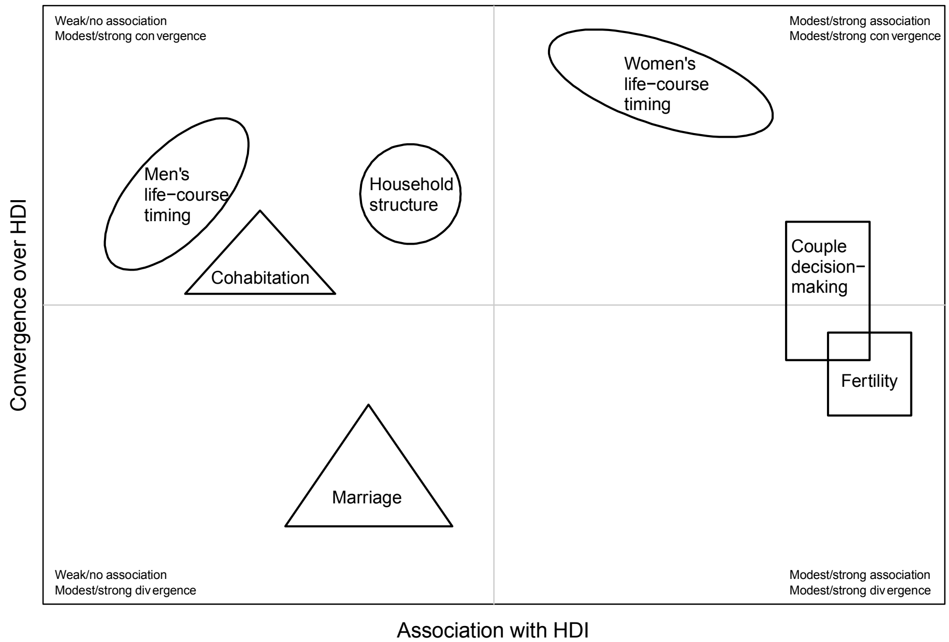
Table 1.2: Beta-convergence over HDI, global (panel a) and regional (panel b) analysis. Beta-convergence coefficients reported

Indicator	*a. Global*		*b. Region excluded*				
	No controls	**Regional dummies**	**Americas**	**Asia**	**Former USSR**	**MENA**	**SSA**
Total Fertility Rate (TFR)	0.002	-0.005	0.002	0.002	0.001	0.003	-0.006
	(0.003)	(0.003)	(0.004)	(0.004)	(0.002)	(0.003)	(0.005)
Net Reproduction Rate (NRR)	0.003	-0.021*	0.001	0.002	-0.001	0.008	-0.013+
	(0.011)	(0.009)	(0.013)	(0.014)	(0.008)	(0.013)	(0.008)
SMAFS, women	-0.001*	-0.001***	-0.000	-0.001*	-0.002***	-0.001*	-0.000
	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)
SMAFS, men	-0.000	-0.000	-0.000	-0.000	-0.000	-0.000	0.000
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
SMAFM, women	-0.001*	-0.001*	-0.001*	-0.001*	-0.001*	-0.001*	-0.001
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)
SMAFM, men	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
SMAFB, women	-0.001*	-0.002**	-0.001*	-0.001*	-0.002**	-0.002**	-0.001
	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
SMAFB, men	-0.001	-0.002*	-0.001	-0.001	-0.001	-0.001	-0.000
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
Husband decides about women's health (prop. HH)	-0.000	-0.063	0.039	0.086	-0.046	-0.010	-0.442+
	(0.091)	(0.092)	(0.088)	(0.102)	(0.086)	(0.095)	(0.255)
Husband decides about household purchases (prop. HH)	-0.040	-0.131	-0.002	-0.030	-0.047	-0.037	-0.541*
	(0.075)	(0.118)	(0.061)	(0.103)	(0.081)	(0.074)	(0.202)
Husband decides about women's visits (prop. HH)	0.064	-0.103	0.036	0.095	0.033	0.068	-0.321+
	(0.088)	(0.124)	(0.099)	(0.101)	(0.089)	(0.087)	(0.179)
Children living with both parents (prop.)	-0.012	-0.010	-0.016	-0.012	-0.008	-0.024+	0.038**
	(0.011)	(0.012)	(0.011)	(0.013)	(0.011)	(0.013)	(0.011)
Women living in a *nuclear* household (prop.)	-0.073	-0.064	-0.069	-0.065	-0.082+	-0.071	-0.054
	(0.044)	(0.045)	(0.049)	(0.046)	(0.044)	(0.045)	(0.051)
Women living in a *three-gen*. household (prop.)	-0.098	-0.110	-0.097	-0.093	-0.108	-0.132+	-0.013
	(0.064)	(0.073)	(0.067)	(0.074)	(0.072)	(0.066)	(0.084)
Women living in a *complex* household (prop.)	-0.054	-0.061	-0.053	-0.040	-0.062	-0.066	0.009
	(0.040)	(0.074)	(0.046)	(0.045)	(0.042)	(0.046)	(0.065)
Prevalence of marriage (prop.)	0.042*	0.017	0.025	0.036*	0.043*	0.041*	0.068**
	(0.016)	(0.020)	(0.016)	(0.018)	(0.016)	(0.016)	(0.023)
Marital expectancy at age 15 (yrs)	0.001	-0.000	-0.000	0.000	0.001	0.001	0.002*
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Prevalence of cohabitation (prop.)	-0.086	-0.168+	-0.097	-0.081	-0.153+	-0.086	0.136
	(0.091)	(0.097)	(0.111)	(0.097)	(0.086)	(0.091)	(0.195)
Cohabitation expectancy at age 15 (yrs)	-0.001	-0.005+	-0.001	-0.001	-0.005+	-0.001	0.010
	(0.003)	(0.003)	(0.004)	(0.003)	(0.003)	(0.003)	(0.008)
Marital and cohabitation expectancy at age 15 (yrs)	-0.001**	-0.002**	-0.001**	-0.002**	-0.001**	-0.001**	-0.001
	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)	(0.001)

Notes: Estimates are weighted by the number of survey waves available per country. Note that MENA countries report no indicators for SMAFS (women and men), SMAFM (men), SMAFB (men), prevalence of cohabitation, average number of years spent cohabiting, and average number of years spent married and cohabiting. Therefore, the regional coefficient corresponding to MENA for those indicators is equivalent to the pooled 'global' one. Standard errors clustered at the country level. Estimates weighted by the number of survey waves per country. Contemporaneous values of HDI used. Regional dummies not included in panel b.
Sig: *** $p<0.001$, ** $p<0.01$, * $p<0.05$, +$p<0.1$
Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

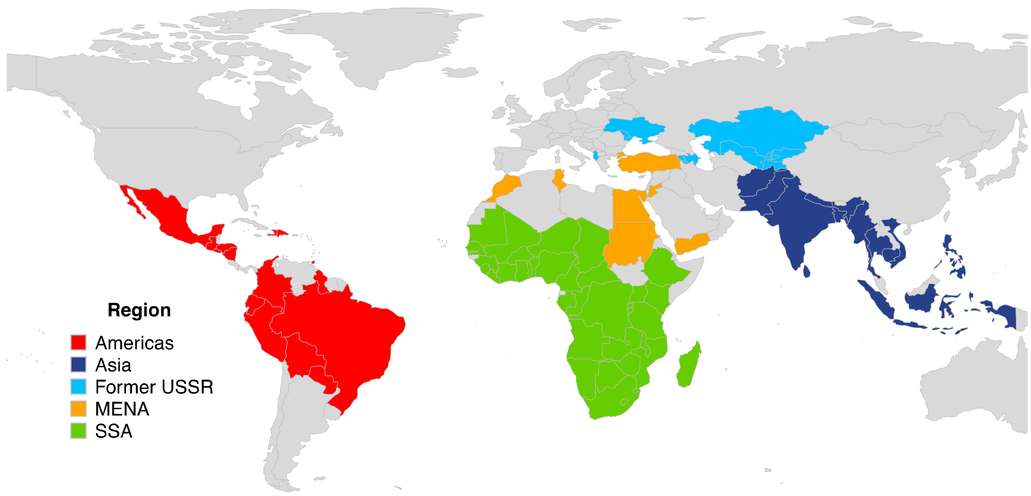
1.9 Figures

Figure 1.1: Framework of analysis and general overview of findings



Notes: This is a stylized diagram that builds on pooled (i.e., all LMICs combined) associations and coefficients presented throughout the paper. The horizontal axis measures the association of indicators of family change with HDI. Note that the association of family change indicators with HDI can be negative (e.g., fertility is negatively associated with HDI), yet the above graph summarizes the strength of association, i.e., it abstracts from the signs of the coefficients. The vertical axis measures convergence over HDI (specifically, *beta*-convergence, as defined later in the paper). The gray line that cuts the plane horizontally corresponds to a null *beta*-convergence coefficient, pointing to persistent differences, i.e., neither convergence (above the gray line) nor divergence (below the gray line).

Figure 1.2: Map of countries included in the analysis. 293 DHS survey waves available for 84 LMICs, grouped into five regions: Americas, Asia, Former Union of Soviet Socialist Republics (USSR), Middle-East and North Africa (MENA), and sub-Saharan Africa (SSA)



Sources: Demographic and Health Surveys (DHS)

Figure 1.3: Conceptual axes of analysis and indicators of Global Family Change (GFC)

Family Events and Behaviors (FEB)

(Total Fertility Rate, % women married, % women living in a cohabiting union, singulate mean age at first sexual intercourse by gender, singulate mean age at first marriage by gender, and singulate mean age at first birth by gender)

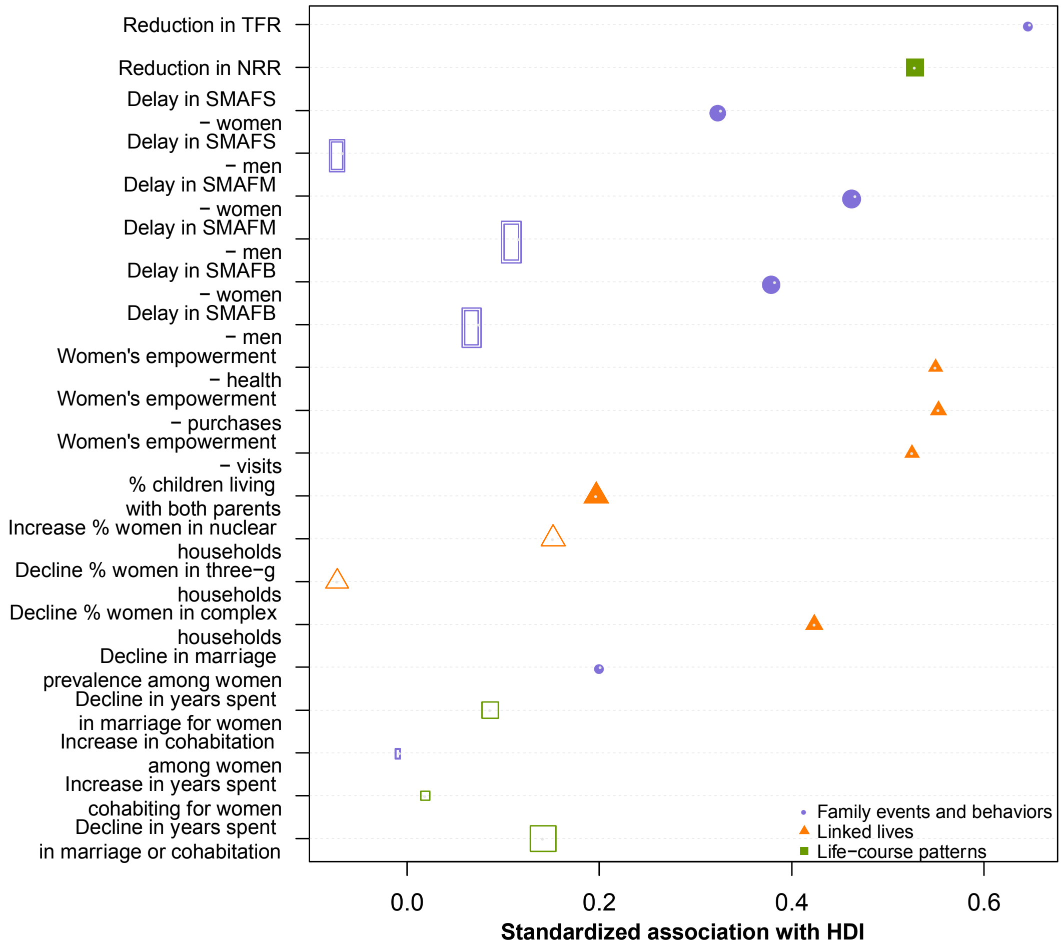
Linked Lives (LL)

(*Vertical*: Children living with both parents (%), share of women living in nuclear, three-generation, and complex households
Horizontal: Intra-household decision making on women's health, household purchases, and visits)

Life-Course Patterns (LCP)

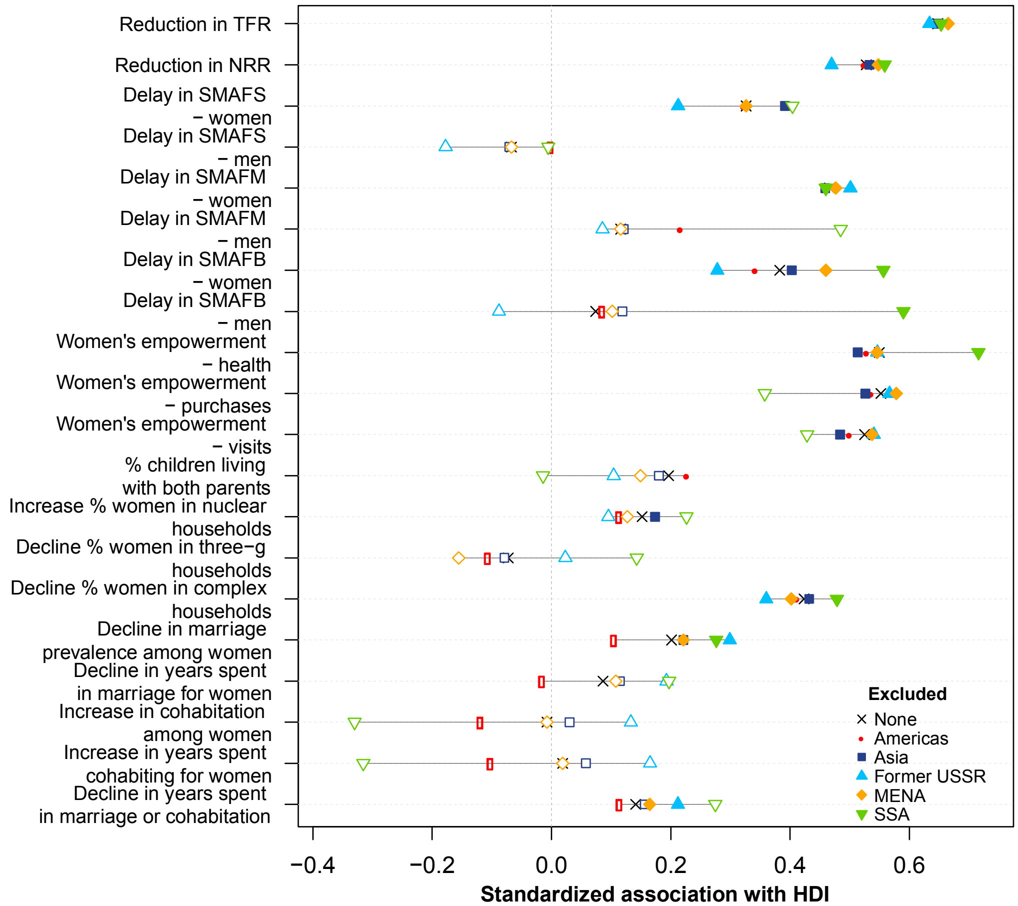
(Net Reproduction Rate, marital expectancy at age 15, cohabitation expectancy at age 15, marital and cohabitation expectancy at age 15)

Figure 1.4: Associations between HDI and GFC-indicators, global analysis



Notes: Standardized beta coefficients reported. The central point corresponds to the estimated slope of the relationship. The area of each marker is inversely proportional to the spread of the distribution of each indicator. Filled markers identify statistically significant estimates (p-value<0.05). Standard errors clustered at the country level. Estimates weighted by the number of survey waves per country. Contemporaneous values of HDI used.
Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

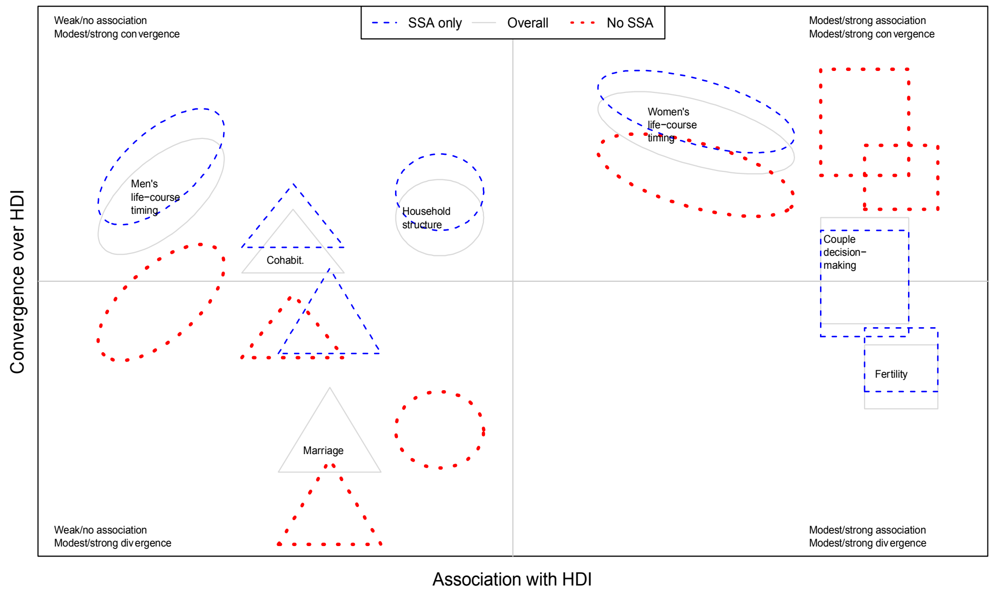
Figure 1.5: Associations between HDI and GFC-indicators, regional analysis



Notes: Standardized beta coefficients reported. Each point corresponds to the estimated slope of the relationship for the overall sample, excluding the region that corresponds to the respective color. Filled markers refer to statistically significant estimates (p-value<0.05). Note that MENA countries report no indicators for SMAFS (women and men), SMAFM (men), SMAFB (men), prevalence of cohabitation, average number of years spent cohabiting, and average number of years spent married and cohabiting. Therefore, the regional coefficient corresponding to MENA for those indicators is equivalent to the pooled 'global' one. Standard errors clustered at the country level. Estimates weighted by the number of survey waves per country. Contemporaneous values of HDI used.

Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

Figure 1.6: Framework of analysis and general overview of findings, isolating the role of sub-Saharan Africa



Notes: This is a stylized diagram. The horizontal axis measures the association of indicators of family change with HDI. Note that the association of family change indicators with HDI can be negative (e.g., fertility is negatively associated with HDI), yet the above graph summarizes the strength of association, i.e., it abstracts from the signs of the coefficients. The vertical axis measures convergence over HDI (specifically, *beta*-convergence). The gray line that cuts the plane horizontally corresponds to a null *beta*-convergence coefficient, pointing to persistent differences, i.e., neither convergence (above the gray line) nor divergence (below the gray line).

1.10 Appendix

Table A1.1: Classification of countries (N=84) by region (N=5) and number of surveys waves (N=293)

Americas (15 countries - 54 waves)	**Former USSR (9 countries - 14 waves)**	**SSA (39 countries - 158 waves)**
Bolivia (5 - 1989, 1994, 1998, 2003, 2008)	Albania (1 - 2008)	Angola (3 - 2006, 2011, 2015)
Brazil (3 - 1986, 1991, 1996)	Armenia (4 - 2000, 2005, 2010, 2015)	Benin (4 - 1996, 2001, 2006, 2011)
Colombia (7 - 1986, 1990, 1995, 2000, 2005, 2010, 2015)	Azerbaijan (1 - 2006)	Botswana (1 - 1988)
Dominican Republic (7 - 1986, 1991, 1996, 1999, 2002, 2007, 2013)	Kazakhstan (2 - 1995, 1999)	Burkina Faso (4 - 1998, 2003, 2010, 2014)
Ecuador (1 - 1987)	Kyrgyzstan (2 - 1997, 2012)	Burundi (3 - 1987, 2010, 2012)
El Salvador (1 - 1985)	Moldova (1 - 2005)	Cameroon (4 - 1991, 1998, 2004, 2011)
Guatemala (4 - 1987, 1995, 1998, 2014)	Tajikistan (1 - 2012)	Central African Republic (1 - 1994)
Guyana (2 - 2005, 2009)	Ukraine (1 - 2007)	Chad (3 - 1996, 2004, 2014)
Haiti (4 - 1994, 2000, 2005, 2012)	Uzbekistan (1 - 1996)	Comoros (2 - 1996, 2012)
Honduras (2 - 2005, 2011)		Congo (3 - 2005, 2009, 2011)
Mexico (1 - 1987)		Congo, DR (2 - 2007, 2013)
Nicaragua (2 - 1998, 2001)		Cote d'Ivoire (4 - 1994, 1998, 2005, 2011)
Paraguay (1 - 1990)		Ethiopia (4 - 2000, 2005, 2011, 2016)
Peru (13 - 1986, 1991, 1996, 2000, 2004, 2005, 2006, 2007, 2008, 2009, 2010, 2011, 2012)		Gabon (2 - 2000, 2012)
Trinidad and Tobago (1 - 1987)		Gambia (1 - 2013)
		Ghana (7 - 1988, 1993, 1998, 2003, 2008, 2014, 2016)
		Guinea (3 - 1999, 2005, 2012)
Asia (14 countries - 43 waves)	**MENA (7 countries - 24 waves)**	Kenya (7 - 1989, 1993, 1998, 2003, 2008, 2014, 2015)
Afghanistan (1 - 2015)	Egypt (8 - 1988, 1992, 1995, 2000, 2003, 2005, 2008, 2014)	Lesotho (3 - 2004, 2009, 2014)
Bangladesh (7 - 1993, 1996, 1999, 2004, 2007, 2011, 2014)	Jordan (6 - 1990, 1997, 2002, 2007, 2009, 2012)	Liberia (4 - 2007, 2009, 2011, 2013)
Cambodia (4 - 2000, 2005, 2010, 2014)	Morocco (3 - 1987, 1992, 2003)	Madagascar (7 - 1992, 1997, 2003, 2008, 2011, 2013, 2016)
India (3 - 1992, 1998, 2005)	Sudan (1 - 1989)	Malawi (7 - 1992, 2000, 2004, 2010, 2012, 2014, 2015)
Indonesia (7 - 1987, 1991, 1994, 1997, 2002, 2007, 2012)	Tunisia (1 - 1988)	Mali (7 - 1987, 1995, 2001, 2006, 2010, 2012, 2015)
Maldives (1 - 2009)	Turkey (3 - 1993, 1998, 2003)	Mauritania (1 - 2000)
Myanmar (1 - 2015)	Yemen (2 - 1991, 2013)	Mozambique (4 - 1997, 2003, 2009, 2011)
Nepal (5 - 1996, 2001, 2006, 2011, 2016)		Namibia (4 - 1992, 2000, 2006, 2013)
Pakistan (3 - 1990, 2006, 2012)		Niger (4 - 1992, 1998, 2006, 2012)
Philippines (5 - 1993, 1998, 2003, 2008, 2013)		Nigeria (5 - 2003, 2008, 2010, 2013, 2015)
Sri Lanka (1 - 1987)		Rwanda (7 - 1992, 2000, 2005, 2007, 2010, 2013, 2014)
Thailand (1 - 1987)		Sao Tome and Principe (1 - 2008)
Timor-Leste (1 - 2009)		Senegal (11 - 1986, 1992, 1997, 2005, 2006, 2008, 2010, 2012, 2014, 2015, 2016)
Vietnam (3 - 1997, 2002, 2005)		Sierra Leone (3 - 2008, 2013, 2016)
		South Africa (1 - 1998)
		Swaziland (1 - 2006)
		Tanzania (9 - 1991, 1996, 1999, 2003, 2004, 2007, 2010, 2011, 2015)
		Togo (3 - 1988, 1998, 2013)
		Uganda (7 - 1988, 1995, 2000, 2006, 2009, 2011, 2014)
		Zambia (5 - 1992, 1996, 2001, 2007, 2013)
		Zimbabwe (6 - 1988, 1994, 1999, 2005, 2010, 2015)

Table A1.2: Associations between HDI and GFC-indicators, global analysis. Contemporaneous HDI (panel a), two-year lagged HDI (panel b), and five-year lagged HDI (panel c)

	a. Contemporaneous HDI			b. HDI lagged 2 years			c. HDI lagged 5 years		
	Stdz. Beta	**P-value**	**Sig.**	**Stdz. Beta**	**P-value**	**Sig.**	**Stdz. Beta**	**P-value**	**Sig.**
Reduction in TFR	0.650	p<0.001	***	0.641	p<0.001	***	0.654	p<0.001	***
Reduction in NRR	0.531	p<0.001	***	0.523	p<0.001	***	0.552	p<0.001	***
Delay in SMAFS - women	0.328	p<0.001	***	0.331	p<0.001	***	0.308	p<0.001	***
Delay in SMAFS - men	-0.067	p>0.1		-0.068	p>0.1		-0.084	p>0.1	
Delay in SMAFM - women	0.468	p<0.001	***	0.476	p<0.001	***	0.460	p<0.001	***
Delay in SMAFM - men	0.117	p>0.1		0.139	p<0.1	+	0.147	p<0.1	+
Delay in SMAFB - women	0.384	p<0.001	***	0.391	p<0.001	***	0.363	p<0.001	***
Delay in SMAFB - men	0.075	p>0.1		0.098	p>0.1		0.099	p>0.1	
Women's empowerment - health	0.550	p<0.001	***	0.535	p<0.001	***	0.515	p<0.001	***
Women's empowerment - purchases	0.557	p<0.001	***	0.540	p<0.001	***	0.523	p<0.001	***
Women's empowerment - visits	0.526	p<0.001	***	0.514	p<0.001	***	0.495	p<0.001	***
% children living with both parents	0.200	p<0.05	*	0.181	p<0.05	*	0.156	p<0.1	+
Increase % women in *nuclear* HH	0.150	p<0.1	+	0.141	p<0.1	+	0.114	p>0.1	
Decline % women in *three-gen.* HH	-0.067	p>0.1		-0.067	p>0.1		-0.089	p>0.1	
Decline % women in *complex* HH	0.425	p<0.001	***	0.425	P<0.001	***	0.395	p<0.001	***
Decline in marriage prevalence among women	0.204	p<0.05	*	0.193	p<0.05	*	0.198	p<0.05	*
Decline in years spent in marriage for women	0.087	p>0.1		0.079	p>0.1		0.091	p>0.1	
Increase in cohabitation among women	-0.007	p>0.1		-0.020	p>0.1		-0.007	p>0.1	
Increase in years spent cohabiting among women	0.019	p>0.1		0.007	p>0.1		0.018	p>0.1	
Decline in years spent in marriage or cohabitation	0.142	p<0.1	+	0.145	p<0.1	+	0.173	p<0.05	*

Notes: Standardized beta coefficients reported. Standard errors clustered at the country level. Estimates weighted by the number of survey waves per country. Estimates reported in panel a correspond to those shown in Figure 1.3 in the main text.
Sig: *** p<0.001, ** p<0.01, * p<0.05, +p<0.1
Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

Table A1.3: Associations between HDI and GFC-indicators, regional analysis. Contemporaneous HDI.

	Region excluded														
	Americas			Asia			Former USSR			MENA			SSA		
	Beta	P-value	Sig.	Beta	P-value	Sig.	Beta	P-value	Sig.	Beta	P-value	Sig.	Beta	P-value	Sig.
Reduction in TFR	0.643	p<0.001	***	0.652	p<0.001	***	0.637	p<0.001	***	0.668	p<0.001	***	0.658	p<0.001	***
Reduction in NRR	0.528	p<0.001	***	0.536	p<0.001	***	0.472	p<0.001	***	0.550	p<0.001	***	0.563	p<0.001	***
Delay in SMAFS - women	0.325	p<0.001	***	0.394	p<0.001	***	0.214	p<0.05	*	0.328	p<0.001	***	0.408	p<0.1	+
Delay in SMAFS - men	0.001	p>0.1		-0.071	p>0.1		-0.179	p<0.1	+	-0.067	p>0.1		-0.005	p>0.1	
Delay in SMAFM - women	0.481	p<0.001	***	0.461	p<0.001	***	0.504	p<0.001	***	0.479	p<0.001	***	0.464	p<0.05	*
Delay in SMAFM - men	0.218	p<0.05	*	0.122	p>0.1		0.086	p>0.1		0.117	p>0.1		0.490	p>0.1	
Delay in SMAFB - women	0.344	p<0.001	***	0.406	p<0.001	***	0.280	p<0.001	***	0.462	p<0.001	***	0.560	p<0.001	***
Delay in SMAFB - men	0.088	p>0.1		0.120	p>0.1		-0.088	p>0.1		0.103	p>0.1		0.595	p<0.05	*
Women's empowerment - health	0.533	p<0.001	***	0.517	p<0.001	***	0.550	p<0.001	***	0.546	p<0.001	***	0.719	p<0.001	***
Women's empowerment - purchases	0.539	p<0.001	***	0.529	p<0.001	***	0.568	p<0.001	***	0.580	p<0.001	***	0.364	p>0.1	
Women's empowerment - visits	0.502	p<0.001	***	0.487	p<0.001	***	0.545	p<0.001	***	0.540	p<0.001	***	0.436	p>0.1	
% children living with both parents	0.231	p<0.05	*	0.184	p<0.1	+	0.107	p>0.1		0.151	p<0.1	+	-0.016	p>0.1	
Increase % women in *nuclear* HH	0.113	p>0.1		0.172	p<0.05	*	0.095	p>0.1		0.130	p>0.1		0.222	p<0.1	+
Decline % women in *three-gen*. HH	-0.105	p>0.1		-0.075	p>0.1		0.023	p>0.1		-0.159	p<0.05	*	0.160	p>0.1	
Decline % women in *complex* HH	0.415	p<0.001	***	0.436	p<0.001	***	0.362	p<0.001	***	0.404	p<0.001	***	0.481	p<0.001	***
Decline in marriage prevalence among women	0.106	p>0.1		0.220	p<0.05	*	0.300	p<0.01	**	0.224	p<0.05	*	0.279	p<0.05	*
Decline in years spent in marriage for women	-0.013	p>0.1		0.115	p>0.1		0.194	p<0.1	+	0.108	p>0.1		0.199	p>0.1	
Increase in cohabitation among women	-0.117	p>0.1		0.032	p>0.1		0.135	p>0.1		-0.007	p>0.1		-0.330	p>0.1	
Increase in years spent cohabiting among women	-0.100	p>0.1		0.058	p>0.1		0.167	p>0.1		0.019	p>0.1		-0.319	p>0.1	
Decline in years spent in marriage or cohabitation	0.117	p>0.1		0.157	p<0.1	+	0.213	p<0.05	*	0.166	p<0.05	*	0.277	p>0.1	

Notes: Standardized beta coefficients reported. Standard errors clustered at the country level. Estimates weighted by the number of survey waves per country. Estimates reported correspond to those shown in Figure 1.4 in the main text.
Sig: *** $p<0.001$, ** $p<0.01$, * $p<0.05$, +$p<0.1$
Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

Table A1.4: Associations between HDI-components and GFC-indicators, global analysis. GNI (panel a), education (panel b), and life expectancy (panel c)

	a. GNI Index			b. Education Index			c. Life Expectancy Index		
	Stdz. Beta	**P-value**	**Sig.**	**Stdz. Beta**	**P-value**	**Sig.**	**Stdz. Beta**	**P-value**	**Sig.**
Reduction in TFR	0.493	p<0.001	***	0.470	p<0.001	***	0.543	p<0.001	***
Reduction in NRR	0.423	p<0.001	***	0.406	p<0.001	***	0.384	p<0.001	***
Delay in SMAFS - women	0.197	p<0.05	*	0.249	p<0.001	***	0.325	p<0.001	***
Delay in SMAFS - men	-0.096	p>0.1		-0.088	p>0.1		0.079	p>0.1	
Delay in SMAFM - women	0.445	p<0.001	***	0.358	p<0.001	***	0.282	p<0.001	***
Delay in SMAFM - men	0.209	p<0.05	*	0.083	p>0.1		-0.005	p>0.1	
Delay in SMAFB - women	0.259	p<0.001	***	0.325	p<0.001	***	0.276	p<0.001	***
Delay in SMAFB - men	0.066	p>0.1		0.084	p>0.1		0.030	p>0.1	
Women's empowerment - health	0.358	p<0.001	***	0.480	p<0.001	***	0.389	p<0.001	***
Women's empowerment - purchases	0.366	p<0.001	***	0.490	p<0.001	***	0.383	p<0.001	***
Women's empowerment - visits	0.334	p<0.001	***	0.439	p<0.001	***	0.415	p<0.001	***
% children living with both parents	0.094	p>0.1		0.062	p>0.1		0.362	p<0.001	***
Increase % women in *nuclear* HH	0.044	p>0.1		0.132	p<0.05	*	0.158	p<0.05	*
Decline % women in *three-gen.* HH	-0.067	p>0.1		-0.044	p>0.1		-0.078	p>0.1	
Decline % women in *complex* HH	0.260	p<0.01	**	0.331	p<0.001	***	0.378	p<0.001	***
Decline in marriage prevalence among women	0.215	p<0.05	*	0.172	p<0.05	*	0.032	p>0.1	
Decline in years spent in marriage for women	0.136	p>0.1		0.108	p>0.1		-0.112	p>0.1	
Increase in cohabitation among women	0.046	p>0.1		-0.059	p>0.1		-0.013	p>0.1	
Increase in years spent cohabiting among women	0.068	p>0.1		-0.042	p>0.1		0.016	p>0.1	
Decline in years spent in marriage or cohabitation	0.233	p<0.01	**	0.184	p<0.01	**	-0.167	p<0.05	*

Notes: Standardized beta coefficients reported. Standard errors clustered at the country level. Estimates weighted by the number of survey waves per country. Estimates reported correspond to those shown in Figure A1.4 in the Appendix.
Sig: *** p<0.001, ** p<0.01, * p<0.05, +p<0.1
Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

Table A1.5: Differences in sample averages between the overall and the 'convergence' sample

	Type	All (N=293)	W>1 (N=267)	Diff.
HDI		0.52	0.50	0.017
GFC Indicators				
Total Fertility Rate (TFR)	FEB	4.14	4.38	-0.238
Net Reproduction Rate (NRR)	LCP	1.72	1.79	-0.071
SMAFS - women	FEB	19.24	18.99	0.249
SMAFS - men	FEB	19.46	19.19	0.269
SMAFM - women	FEB	21.82	21.53	0.283
SMAFM - men	FEB	25.90	25.66	0.246
SMAFB - women	FEB	21.48	21.28	0.196
SMAFB - men	FEB	26.48	26.28	0.205
Husband decides about women's health (prop. HH)	LL	0.30	0.31	-0.019
Husband decides about household purchases (prop. HH)	LL	0.32	0.33	-0.009
Husband decides about women's visits (prop. HH)	LL	0.24	0.25	-0.011
Children living with both parents (prop.)	LL	0.68	0.68	0.004
Women living in a *nuclear* household (prop.)	LL	0.40	0.40	-0.003
Women living in a *three-gen.* household (prop.)	LL	0.29	0.28	0.010
Women living in a *complex* household (prop.)	LL	0.36	0.36	-0.002
Prevalence of marriage (prop.)	FEB	0.54	0.55	-0.009
Marital expectancy at age 15 (yrs)	LCP	18.55	18.63	-0.079
Prevalence of cohabitation (prop.)	FEB	0.14	0.13	0.003
Cohabitation expectancy at age 15 (yrs)	LCP	4.43	4.32	0.113
Marital and cohabitation expectancy at age 15 (yrs)	LCP	22.15	22.30	-0.154

Notes: Estimates weighted by the number of survey waves available per country. N refers to the number of country-year combinations. Specifically, as the N is indicator-specific, it refers to the maximum number of observations per group. FEB: "Family events and behaviors", LL: "Linked lives"; LCP: "Life-course patterns". Information on cohabitation, sexual intercourse, and timing indicators for men is not available for MENA countries.
Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

Table A1.6: Beta-convergence over HDI per unit of time, global (panel a) and regional (panel b) analysis. Beta-convergence coefficients reported

Indicator	*a. Global*		*b. Region excluded*				
	No controls	**Regional dummies**	**Americas**	**Asia**	**Former USSR**	**MENA**	**SSA**
Total Fertility Rate (TFR)	0.009	-0.045*	0.007	0.007	0.008	0.010	-0.065*
	(0.016)	(0.021)	(0.019)	(0.019)	(0.013)	(0.017)	(0.028)
Net Reproduction Rate (NRR)	-0.008	-0.182**	-0.015	-0.024	-0.020	-0.002	-0.141+
	(0.055)	(0.062)	(0.064)	(0.068)	(0.047)	(0.059)	(0.072)
SMAFS, women	-0.007*	-0.006**	-0.003	-0.008*	-0.009***	-0.007*	-0.002
	(0.003)	(0.002)	(0.003)	(0.004)	(0.002)	(0.003)	(0.002)
SMAFS, men	-0.006	-0.005	-0.005	-0.005	-0.006	-0.006	-0.006
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.010)
SMAFM, women	-0.010**	-0.012**	-0.009**	-0.010**	-0.009**	-0.011***	-0.012
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.008)
SMAFM, men	-0.005	-0.006	-0.004	-0.005	-0.006	-0.005	-0.001
	(0.006)	(0.007)	(0.006)	(0.007)	(0.006)	(0.006)	(0.010)
SMAFB, women	-0.007*	-0.011**	-0.008*	-0.006+	-0.009**	-0.010**	-0.003
	(0.003)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
SMAFB, men	-0.006	-0.009*	-0.005	-0.006	-0.007	-0.006	-0.003
	(0.005)	(0.004)	(0.005)	(0.005)	(0.005)	(0.005)	(0.009)
Husband decides about women's health (prop. HH)	0.101	-0.250	0.185	0.535	-0.118	0.075	-1.766+
	(0.446)	(0.446)	(0.450)	(0.497)	(0.415)	(0.462)	(1.070)
Husband decides about household purchases (prop. HH)	0.104	-0.294	0.135	0.288	0.102	0.119	-2.307*
	(0.309)	(0.518)	(0.265)	(0.418)	(0.334)	(0.304)	(0.854)
Husband decides about women's visits (prop. HH)	0.108	-0.808+	-0.088	0.240	-0.054	0.126	-1.287
	(0.382)	(0.462)	(0.401)	(0.472)	(0.374)	(0.384)	(1.091)
Children living with both parents (prop.)	-0.093	-0.049	-0.120+	-0.100	-0.052	-0.162+	0.162*
	(0.070)	(0.070)	(0.070)	(0.078)	(0.060)	(0.084)	(0.070)
Women living in a *nuclear* household (prop.)	-0.463	-0.388	-0.429	-0.435	-0.497+	-0.463	-0.319
	(0.282)	(0.280)	(0.307)	(0.299)	(0.285)	(0.288)	(0.318)
Women living in a *three-gen.* household (prop.)	-0.411	-0.504	-0.420	-0.422	-0.474	-0.685*	0.368
	(0.347)	(0.375)	(0.369)	(0.396)	(0.387)	(0.322)	(0.595)
Women living in a *complex* household (prop.)	-0.338	-0.406	-0.341	-0.275	-0.393+	-0.379+	-0.009
	(0.210)	(0.270)	(0.227)	(0.221)	(0.214)	(0.222)	(0.324)
Prevalence of marriage (prop.)	0.220*	0.170	0.240	0.187	0.223*	0.229*	0.218+
	(0.100)	(0.151)	(0.147)	(0.112)	(0.100)	(0.100)	(0.111)
Marital expectancy at age 15 (yrs)	0.005+	0.004	0.005	0.005	0.006+	0.006+	0.006+
	(0.003)	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)
Prevalence of cohabitation (prop.)	-0.879	-1.409**	-1.068+	-0.861	-1.301*	-0.879	0.870
	(0.571)	(0.491)	(0.627)	(0.609)	(0.497)	(0.571)	(1.206)
Cohabitation expectancy at age 15 (yrs)	-0.032+	-0.048**	-0.039+	-0.033+	-0.045**	-0.032+	0.020
	(0.018)	(0.017)	(0.021)	(0.019)	(0.016)	(0.018)	(0.034)
Marital and cohabitation expectancy at age 15 (yrs)	-0.008**	-0.009**	-0.007*	-0.011**	-0.007*	-0.008*	-0.006
	(0.003)	(0.003)	(0.003)	(0.004)	(0.003)	(0.003)	(0.006)

Notes: Estimates are weighted by the number of survey waves available per country. Note that MENA countries report no indicators for SMAFS (women and men), SMAFM (men), SMAFB (men), prevalence of cohabitation, average number of years spent cohabiting, and average number of years spent married and cohabiting. Therefore, the regional coefficient corresponding to MENA for those indicators is equivalent to the pooled 'global' one. Standard errors clustered at the country level. Estimates weighted by the number of survey waves per country. Contemporaneous values of HDI used. Regional dummies not included in panel b.
Sig: *** p<0.001, ** p<0.01, * p<0.05, +p<0.1
Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

Figure A1.1: Kernel density of countries over Human Development Index (HDI), by group of countries

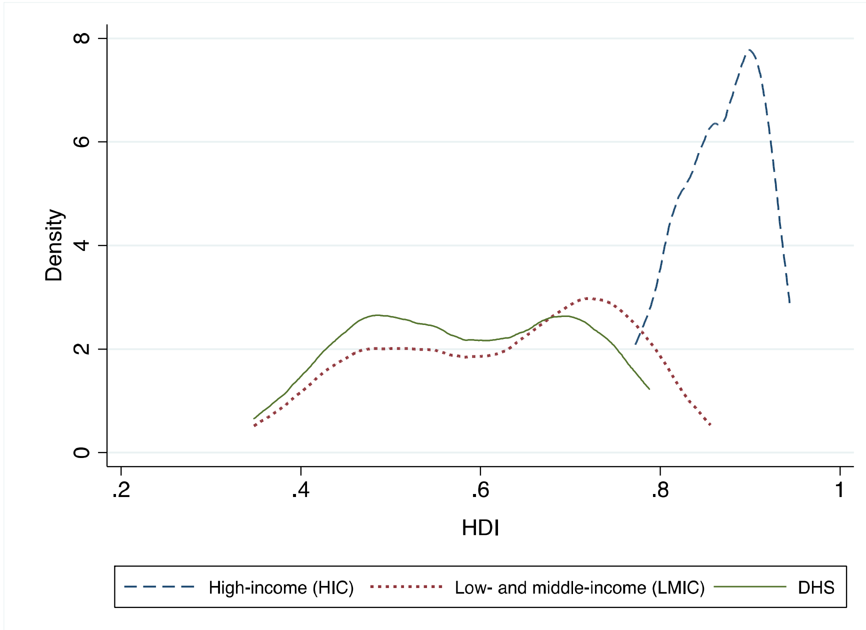


Figure A1.2: Age-standardized vs crude family indicators. Example shown: age-standardized vs crude marriage prevalence

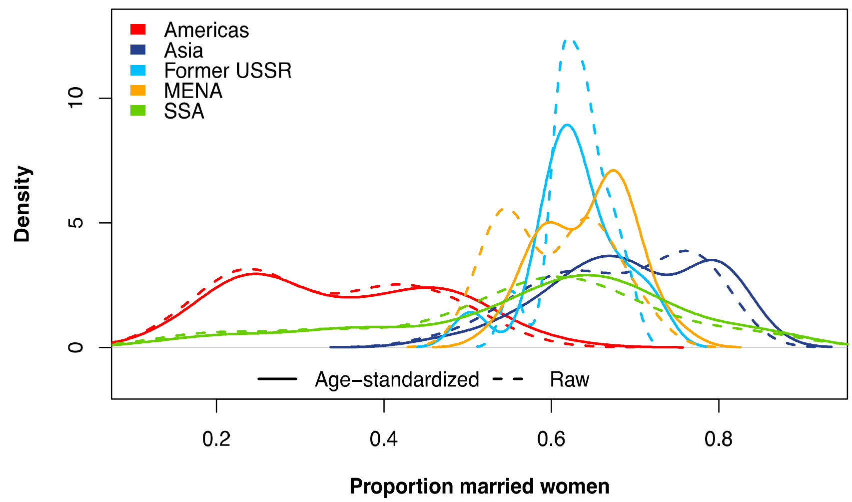
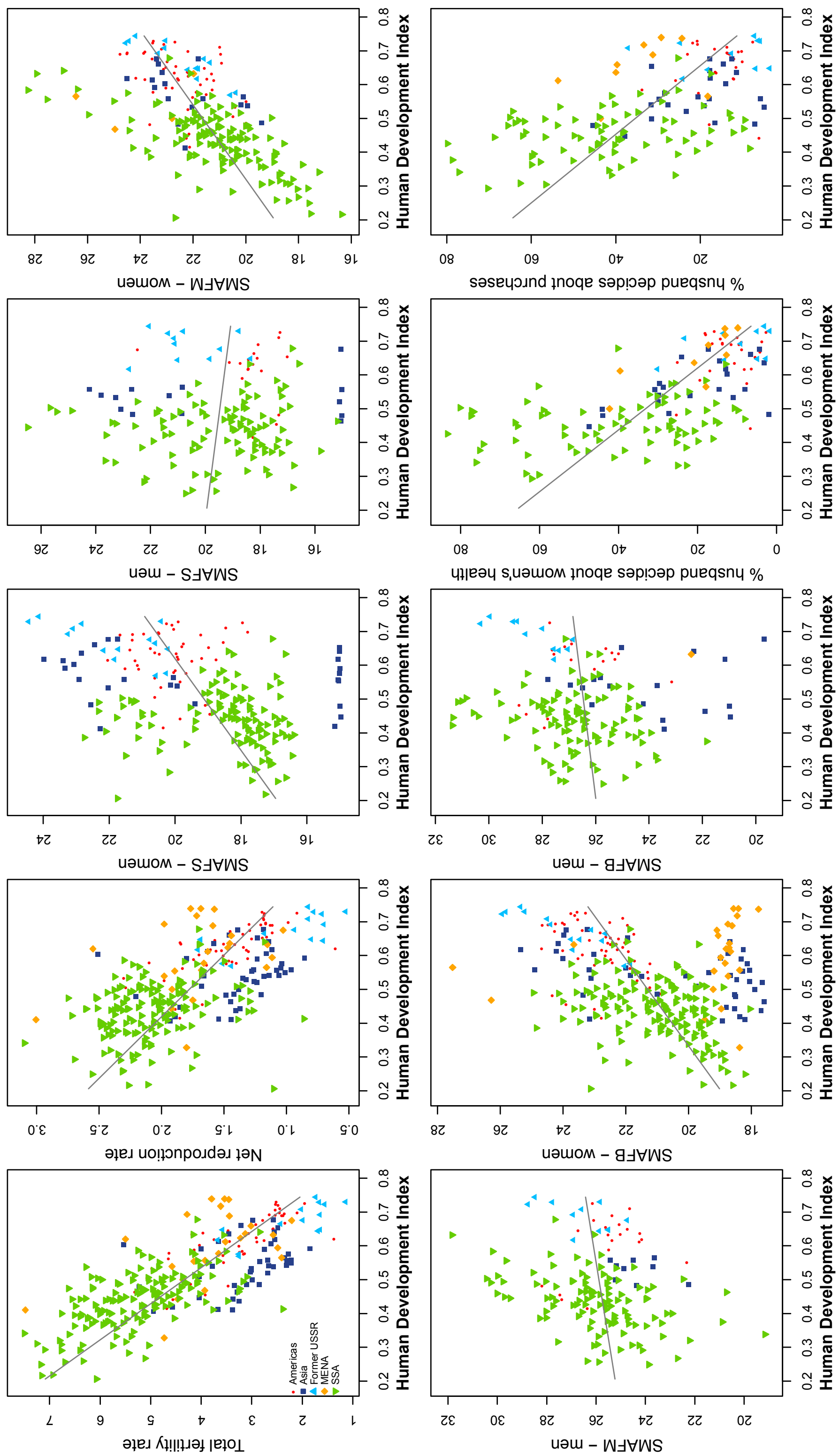
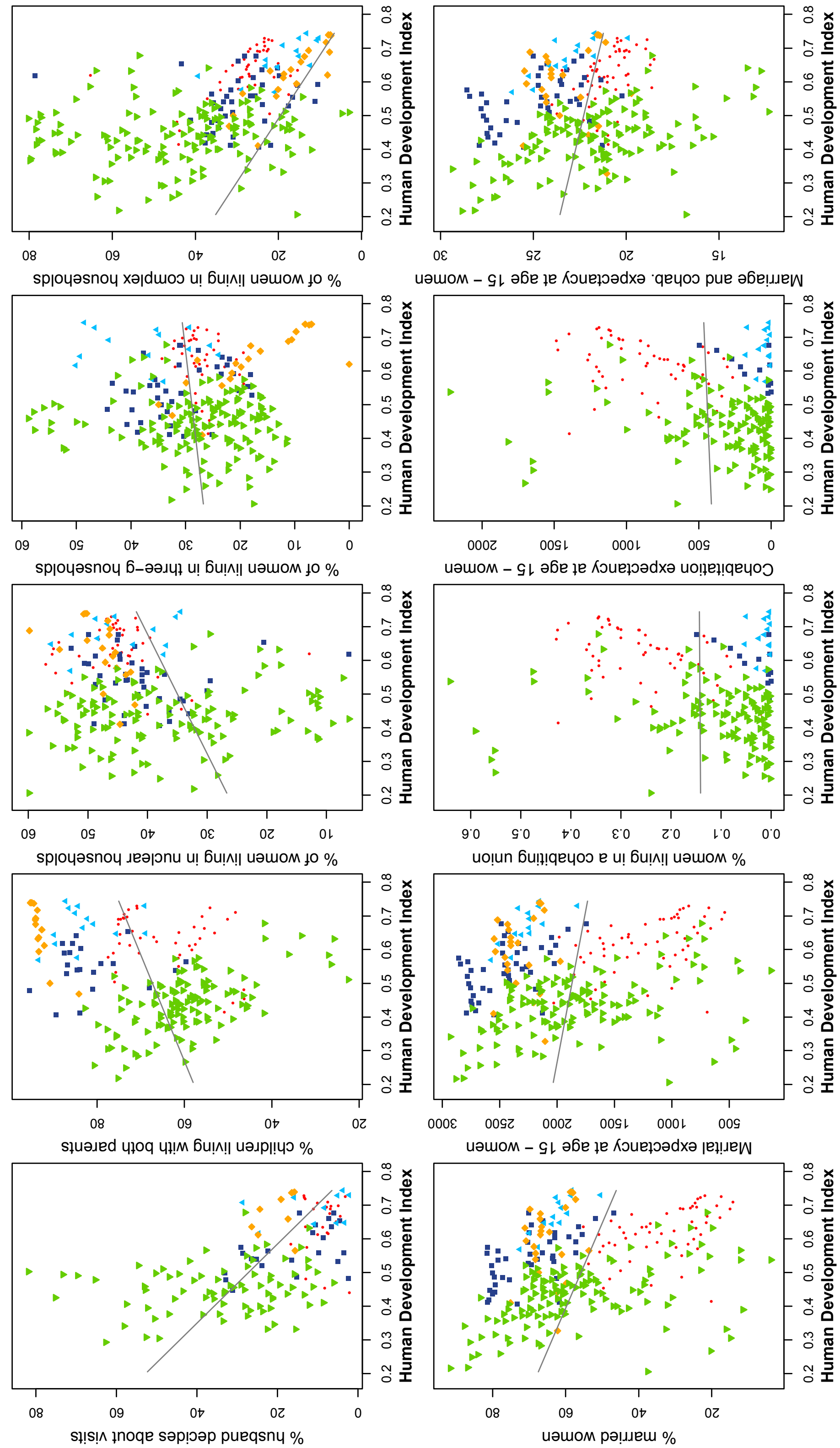


Figure A1.3: Scatter plots of the association between the 20 GFC-indicators and HDI





% husband decides about visits
% married women
% children living with both parents
Marital expectancy at age 15 − women
% of women living in nuclear households
% women living in a cohabiting union
% of women living in three−g households
Cohabitation expectancy at age 15 − women
% of women living in complex households
Marriage and cohab. expectancy at age 15 − women
Human Development Index

Figure A1.4: Associations between HDI-components and GFC-indicators, global analysis. GNI (left), education (center), and life expectancy (right)

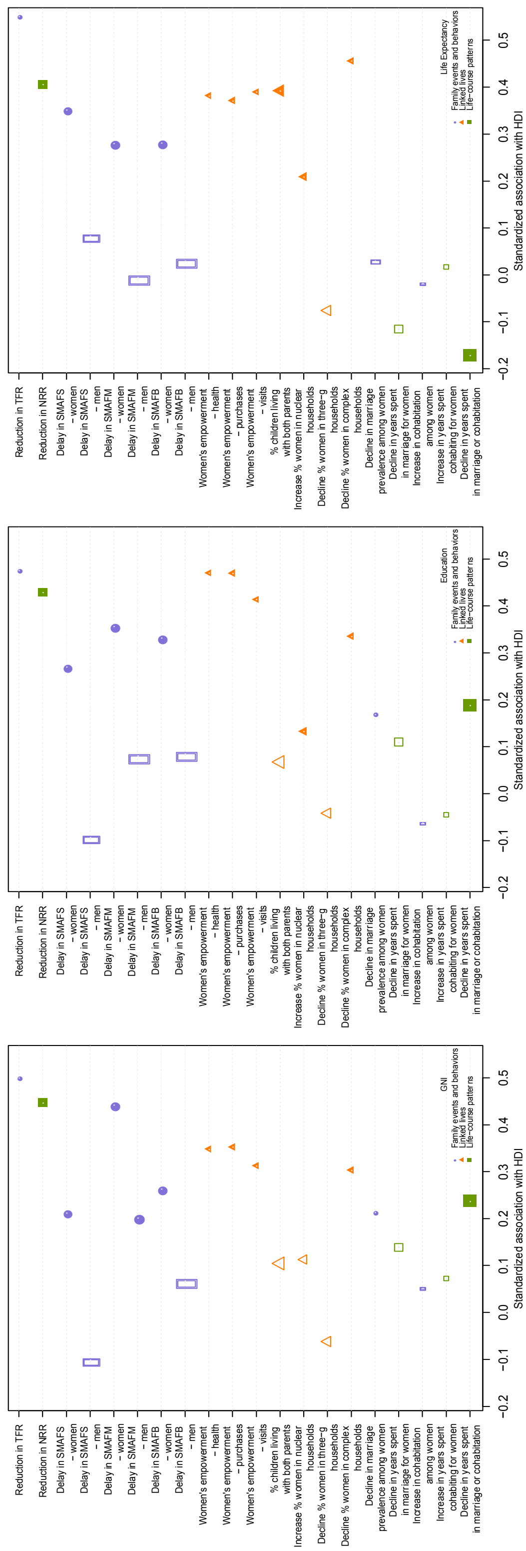
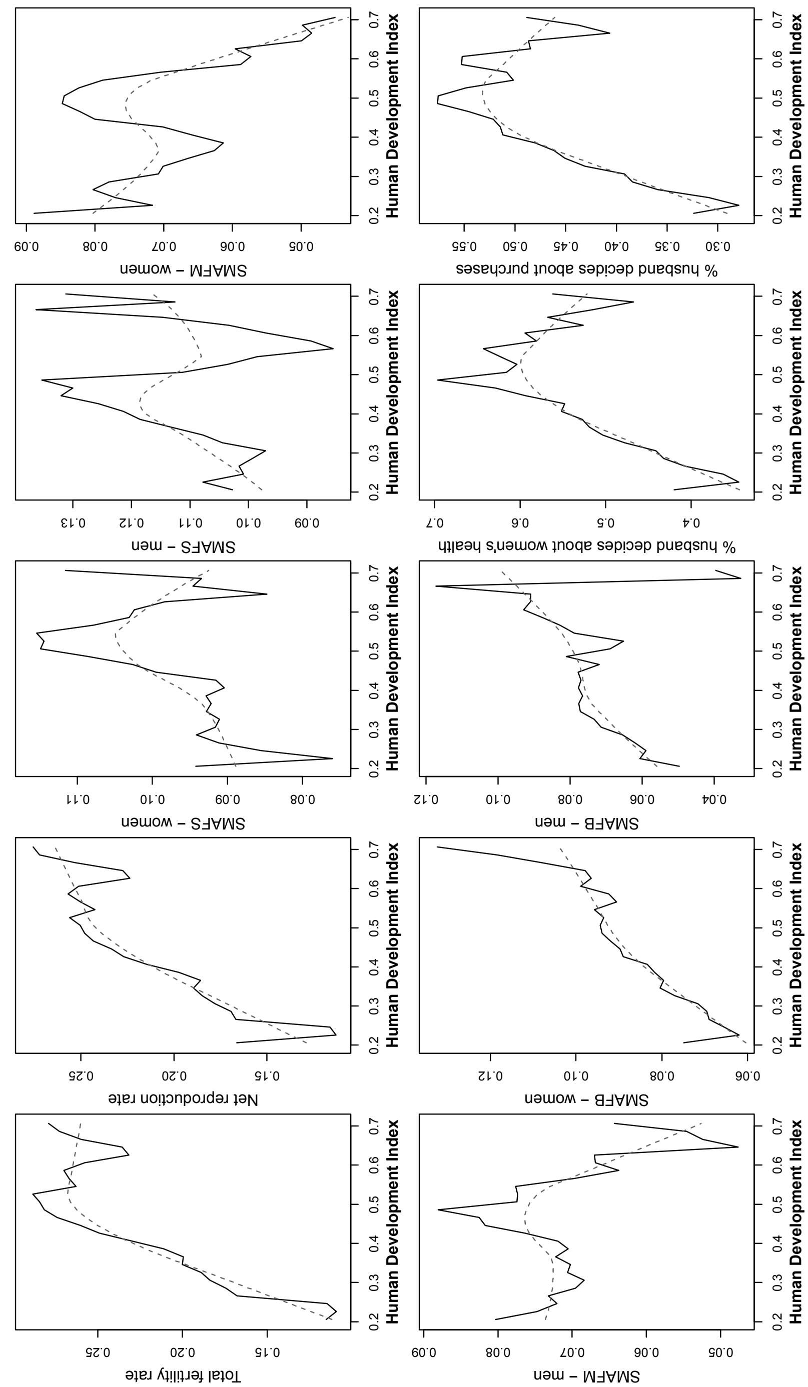
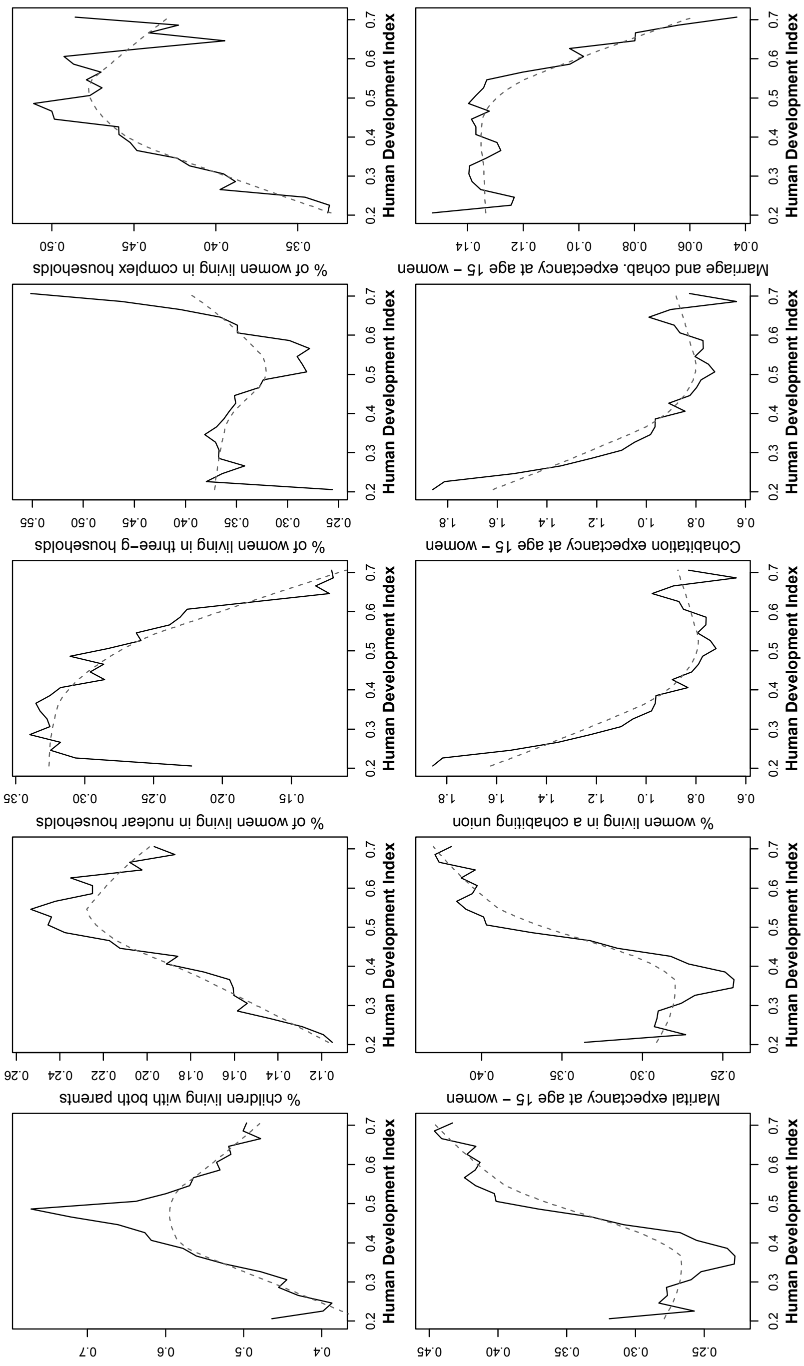


Figure A1.5: Sigma-convergence over HDI





Notes: This figure presents trends in the coefficient of variation (CV) over HDI. To boost the number of observations we use a moving-average approach We use two-percentage-point steps to compare the moving measures. At each step we include countries over a bandwidth of ten percentage points for the assessment over HDI. The dotted line is estimated using local polynomial regression smoothing techniques. A CV of 0.1 indicates that the standard deviation is at most 10 per cent of the mean. A negative trend in the CV implies a decline in the variability relative to the mean, i.e., convergence; a flat trend implies differences are maintained; and a positive trend implies increasing heterogeneity, i.e., divergence.
Sources: Demographic and Health Surveys (DHS), UNDP, and UN-DESA Population Division.

Chapter 2. Educational Assortative Mating in sub-Saharan Africa: Compositional Changes and Implications for Household Wealth Inequality

2.0 Abstract

Over the past decade, there has been increasing interest in patterns of educational assortative mating across the developing world. However, research on the topic in sub-Saharan Africa (SSA) is still scant. Using Demographic and Health Survey (DHS) data from 39 countries between 1986 and 2016, this paper is the first to offer a comparative overview of changes in educational assortative mating in SSA. The study has three aims. First, I examine whether educational assortative mating has increased over time, by sub-region and location of residence. Second, by comparing observed patterns of mating with the ones that would prevail under random mating, I investigate the extent to which trends are driven by compositional – shifting educational distributions – versus residual factors. Third, I adopt an accounting-based methodology to assess implications of educational assortative mating for household wealth inequality. Results show that mating has increased over marriage cohort in all sub-regions except for Southern Africa, with increases driven mostly by rural areas. Moreover, I find that educational assortative mating accounts for a non-negligible share (3 to 12 percent, varying by sub-region) of the cohort-specific inequality in household wealth, yet changes in mating over time hardly move the time trends in wealth inequality.

2.1 Introduction

Over the past decades, there has been growing interest in patterns of educational assortative mating around the world. Assortative mating is a powerful driver of societal change, as it shapes the way people organize within families, affecting in turn individuals' access to resources and their distribution across families (Schwartz 2013). In light of the role that education plays for economic inequality and its persistence from generation to generation, most of the focus on educational assortative mating to date – at least in high-income societies – has been tied to its implications for the study of inequality and the social stratification system (Mare 1991, 2016; Rosenfeld 2008; Schwartz and Mare 2005; etc.). Yet patterns of mating with regard to couples' socio-economic characteristics are vital to understanding a whole set of dynamics in the demographic makeup of households, such as family formation, composition, and breakdown (Schwartz and Han 2014). They also have consequences for outcomes that are directly or indirectly linked to the family, such as longevity, health, fertility preferences, and fertility behavior (Huber and Fieder 2011; Ntoimo and Mutanda 2017; Trimarchi, Schnor, and Van

Bavel forthcoming). A proper understanding of mating patterns thus ultimately permits to shed light on fundamental changes underlying the demography of the population. This study focuses on trends, variation, and implications of educational assortative mating for inequality in sub-Saharan Africa (SSA), a region of the world that has been experiencing rapid socio-economic and demographic change yet has been largely neglected in the assortative mating literature.

The paper has three aims. First, using 126 Demographic and Health Surveys (DHS) collected between 1986 and 2016, I provide an overview of educational assortative mating patterns across 39 countries in SSA. Despite a series of global and comparative studies documenting declining hypergamy around the world (Esteve et al. 2016; Esteve, Garcia-Roman, and Permanyer 2012), never has the assortative mating literature focused exclusively and comparatively on patterns of change within SSA.[23] To put it simply, evidence is lacking on questions as simple as whether educational assortative mating has increased or decreased overtime. This is surprising, as scholars have devoted ample space to trends, determinants, and consequences of educational assortative mating in other parts of the world, such as the United States (Breen and Salazar 2011; Eika, Mogstad, and Zafar 2017; Greenwood et al. 2014; Mare 1991, 2016; Qian 1998; Schwartz and Mare 2005), Latin America (Esteve and McCaa 2007; Esteve, McCaa, and López 2013; Ganguli, Hausmann, and Viarengo 2014; Gullickson and Torche 2014; Torche 2010), South-East Asia (Borkotoky and Gupta 2016; Hu and Qian 2015; Park and Smits 2005; Smits and Park 2009), and Europe (Boertien and Permanyer 2017; Breen and Andersen 2012; Breen and Salazar 2010; De Hauw, Grow, and Van Bavel 2017; Esteve and Cortina 2006; Grow and Van Bavel 2015). Ntoimo and Mutanda (2017) is a notable exception, yet they examine patterns of homogamy and heterogamy in Ghana, Nigeria, and Zambia only.[24]

A more comprehensive study of mating in SSA is critical for several reasons. First and foremost, SSA countries are undergoing swift transformations in the realm of union formation, such as delays in mean ages at first union (Bongaarts, Mensch, and Blanc 2017; Koski, Clark, and Nandi 2017; Shapiro and Gebreselassie 2014), along with increasing educational attainment particularly for women, and expanding female labor force participation rates (Lloyd et al. 2005). Underlying these

[23] Educational *homogamy* is defined as union formation between individuals who are similar in terms of education. The alternative is educational *heterogamy*, defined as union formation between individuals with different levels of education. Heterogamous couples can in turn be educationally *hypergamous* – if the female partner/wife has lower education than the male partner/husband – and educationally *hypogamous* – if the female partner/wife has higher education than the male partner/husband.

[24] Another notable exception is Behrman (2016), who focuses on patterns of educational assortative mating across five Eastern and South-Eastern African countries. Yet her study is aimed at examining the implications of mating for intra-couple gender dynamics (e.g., decision-making, intimate partner violence, unmet need for family planning, etc.), rather than providing a comparative overview of trends and demographic determinants of mating.

changes has been a massive growth in urbanization, spreading "modern" ideals stressing the value of education, encouraging later marriage, and reducing the influence of kin who control the timing of marriage and choice of spouse (Cherlin 2012; Singh and Samara 1996). As these factors are key drivers of mating, there is reason to believe that important changes in educational assortative mating might have occurred in SSA over the past half century. Yet, demographic change and urbanization have followed uneven trajectories within SSA, partly as a function of the various cultural repertoires, diversified economies, political systems, but also on the crises – such as conflicts, civil wars, food shortages, and the HIV/AIDS epidemic – that countries or even entire sub-regions have experienced (Cherlin 2012; Tabutin and Schoumaker 2004). Hence, a closer look at within-region dynamics is likely to deliver a more nuanced picture of the phenomenon, highlighting sub-regional heterogeneity and diverging patterns of change that are masked in "global" studies of mating such as Esteve et al. (2016).

Alongside the dramatic changes outlined above, there is yet evidence that SSA countries still lag behind other low- and middle-income countries (LMICs) in areas such as gender and couple-related dynamics. For instance, gender gap reversals in education are occurring more slowly in SSA than in other world regions (Esteve et al. 2016) – the gender gap in education has narrowed but not reversed yet. Similarly, previous research has found stark gender imbalances in intra-household bargaining dynamics, to the extent that SSA remains the only region where the share of households in which the husband is the sole decision-maker reaches up to 40 percent (Pesando et al. 2018).[25] Accordingly, there is ground to hypothesize that trends towards increasing assortative mating documented globally, often unfolding along with reversals in gender gaps in education and increases in women's empowerment (Esteve, Garcia-Roman, and Permanyer 2012; Grow and Van Bavel 2015), might have occurred differently in SSA.

One challenge in studies of mating is to determine whether increases in educational homogamy arise due to secular changes in educational attainment of women versus men, or because of shifts in educational assortative mating itself. For instance, the narrowing of the gender gap in education may increase the chance that someone with secondary education is married to someone else with secondary education even in the absence of changes in the assortativeness of marriage (Liu and Lu 2006). Therefore, as a second contribution I compare the observed patterns of mating to those predicted under random mating and investigate the extent to which trends in educational assortative mating are driven by compositional – i.e., changes in educational distributions – versus residual changes up and

[25] Averages reach up to 70 percent in countries (and selected years) such as Burkina Faso, Chad, Mali, Niger, and Senegal.

beyond changes in educational distributions. In other words, I explore the extent to which shifts towards homogamy can be accounted for by "mechanical" changes that result from proportionally faster increases in women's education, as compared to other responses related to the shifting value of education and spouses' preferences for educational resemblance. To address this question I use contingency tables, sorting indicators, and log-linear models.

As a third contribution, I rely on this same accounting-based methodology to assess implications of educational assortative mating for household wealth inequality – defined as inequality between households in asset possession.[26] As a measure of wealth I adopt the International Wealth Index (IWI), the first comparable asset-based wealth index covering the complete developing world, based on data for over 2.1 million households (Smits and Steendijk 2015). I am interested in asking the following counterfactual questions relating mating and inequality: what would happen to the wealth distribution if in every marriage cohort mating was random instead of assortative? Similarly, what would happen to wealth inequality if couples from the latest marriage cohort matched as those in the earliest ones? These analyses address the broader puzzle of whether marital sorting affects household wealth inequality, another yet unexplored question in the SSA context, and rarely addressed in the assortative mating literature in general. More broadly, a more comprehensive understanding of determinants, trends, and implications of educational assortative mating in SSA has the potential to shed better light on the reciprocal linkages between demographic change, family change, and the social stratification system.

2.2 Background

Theoretical perspectives on educational assortative mating across Western and non-Western societies

The assortative mating literature documents extensive variation in educational homogamy among countries, time periods, and even among studies focusing on the same country. Starting around the 1960s, a close focus on temporal and spatial variation in the patterns of association between spouses' educational attainment originated from studies on high-income Western societies, very much driven by the idea that industrialization brings progress, and differences in countries' level of socio-economic development may explain variation in homogamy. The underlying logic – embedded in theoretical perspectives such as modernization theory (Blau and Duncan 1967; Parsons 1971), industrialization theory (Kerr 1983), and individualization theory (Beck 1986; Giddens 1991) – builds on the premise

[26] Throughout the paper I refer to the terms "wealth inequality" and "asset inequality" interchangeably.

that industrialization and social modernization unfold in tandem with trends towards social openness and meritocratization, thus weakening societies' social structures and social boundaries.

Within this modernization macro-perspective, scholars have formulated and tested three competing hypotheses relating socio-economic development and educational homogamy both across countries and within a country overtime (Smits, Ultee, and Lammers 1998). First is the general openness hypothesis, which postulates that industrialization leads to less educational homogamy because of the decrease in parents' control over the marriage process and the increase in the number of contacts between individuals from different classes and status groups, occurring through greater geographical mobility, more education, and the spread of mass communication (Blossfeld 2009; Smits, Ultee, and Lammers 1998). The second one, called the status attainment hypothesis, postulates instead a positive relationship between economic development and educational homogamy due to the increased importance of education as a marker of social status – which in turn pushes high-educated individuals to increasingly select their partners based on educational considerations (Blossfeld and Timm 2003; Kalmijn 1998; Treiman 1970). The third hypothesis (inverted U-curve hypothesis) combines the previous two and predicts an increase in educational homogamy in the first phase of the industrialization process, where status considerations and parental authority still play an important role in partner choice. Conversely, the decrease in educational homogamy takes place in a second phase, where rising wages and more binding laws loosen the parental bond and give individuals more freedom to marry whom they like, hence following more closely the logic of 'romantic love'.[27]

Gradually, research examining trends and variation in educational assortative mating has expanded to other societies across Latin America (Esteve and McCaa 2007; Esteve, López, and McCaa 2013; Ganguli, Hausmann, and Viarengo 2014; Gullickson and Torche 2014; Torche 2010), East Asia (Hu and Qian 2015; Park and Smits 2005; Smits and Park 2009), South Asia (Borkotoky and Gupta 2016; Prakash and Singh 2014), adopting a more large-scale comparative approach (Esteve et al. 2016; Esteve, Garcia-Roman, and Permanyer 2012; Raymo and Xie 2000; Smits 2003; Smits, Ultee, and Lammers 1998, 2000; etc.). Studies evaluating the applicability of the aforementioned hypotheses to different contexts has delivered quite mixed – and often conflicting – findings.[28] Evidence in favor of

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[27] Smits et al. (2000) and Smits (2003) elaborated a fourth related hypothesis called the *saturation hypothesis*. This postulates a decrease in homogamy in modernizing societies which slows down and eventually stops in societies that have reached a high level of openness. As this hypothesis is more applicable to highly industrialized societies that are far from the sub-Saharan countries included in this study, I leave it aside.

[28] As claimed by Blossfeld (2009), several analytical choices contribute to explaining this mixed empirical evidence, such as the analytical focus (e.g., first marriages versus stock of marriages), the study population (e.g., marital matches versus all individuals at "risk" of entering a union), the type of data (cross-sectional versus panel), methodology (e.g., log-linear models versus event history models), etc.

a trend towards more educational homogamy has been found for several highly developed Western societies – mostly the United States (US) and some European countries – by scholars as diverse as Blossfeld and Timm (2003), Kalmijn (1991), Mare (1991), Qian and Preston (1993), Schwartz and Mare (2005), and Smits, Ultee, and Lammers (2000).

However, research including LMICs suggests a more complex picture. Using data from 65 countries, Smits, Ultee, and Lammers (1998) found a cross-sectional inverted U-shaped relationship between level of development and educational homogamy. The status attainment hypothesis (higher development, higher homogamy) was supported only when comparing the least-developed countries with countries at intermediate levels of development, while the general openness hypothesis (higher development, lower homogamy) was supported when comparing countries at intermediate levels with the most developed ones. Consistent with this finding, in a follow-up study covering 55 countries, Smits (2003) found declining educational homogamy and more openness in more rapidly developing countries. Although evidence from long-term trend studies remains scarce, less educational homogamy in higher developed countries – mostly Asian – has also been confirmed in trend studies such as Raymo and Xie (2000), Smits, Ultee, and Lammers (2000), and Smits and Park (2009). Focusing on ten East-Asian societies, the latter authors documented a trend towards less homogamy over time and claimed that educational homogamy tends to be lower in societies that are more modern, have higher female labor force participation rates, and experience less Confucian influence. Research from India (Borkotoky and Gupta 2016; Prakash and Singh 2014) and Latin America (Esteve and McCaa 2007; Ganguli, Hausmann, and Viarengo 2014) suggests instead an increase in educational assortative mating over time.

Although comparative studies such as Smits, Ultee, and Lammers (1998; 2000) included a few SSA countries, none of the above hypotheses has been wholly evaluated in the African context. This study attempts to do so by adopting a time-trend perspective. It is challenging to generalize claims on patterns of educational assortative mating in a region of the world as diverse and heterogeneous as SSA, yet documenting trends by sub-region and location of residence (urban/rural) is a first step towards a better understanding. As – comparatively – SSA countries rank lowest on development indices such as the Human Development Index (HDI), the above theories would suggest an increase in educational assortative mating over time in line with the status attainment hypothesis, with considerable differences by sub-region of SSA and location of residence, consistent with different rates of modernization and urbanization. Specifically, I hypothesize a more marked increase in mating in rural areas paralleled by a less marked increase (or even an incipient decline) in urban areas, where the general openness hypothesis is more likely to take hold – as driven by greater geographical mobility, educational expansion, cross-cultural exchange, and mass communication.

Educational expansion, urbanization, family change, and mating in sub-Saharan Africa

Over the last few decades, there has been an increase in mean grades of schooling attained among young women in all regions of the developing world (Mensch, Singh, and Casterline 2006). Yet in their recent global study of declining educational hypergamy, Esteve et al. (2016) claim that African countries have the lowest proportions of the population with college education and the lowest levels of women's education compared to men's. According to their study, time trends indicate little progress in expanding college education in Africa, but substantial progress in women's education that has contributed to narrowing the gender gap, which still favors men. A key factor underlying the expansion of education has been the massive growth in the share of the population living in cities, which started from very different levels across sub-regions, as Southern Africa was already far more urbanized than the other sub-regions in the 1950s. Heterogeneity in the degree of urbanization between sub-regions has lessened since the 1950s, as the least urbanized regions 50 years ago (Eastern Africa, followed by Western and Central Africa) have experienced the highest urban growth, with the urban population multiplied by roughly 20 between 1950 and 2000 (Tabutin and Schoumaker 2004).

In tandem with these macro-structural transformations, African families have also changed in domains that are likely to relate to mating patterns. First and foremost, age at marriage has risen throughout the continent (Koski, Clark, and Nandi 2017; Tabutin and Schoumaker 2004). According to data from the United Nations (UN) Department of Economic and Social Affairs Population Division (2015), the singulate mean age at marriage is now greater than 18 in the majority of countries in the region. This is relevant, as the age at which men and women form unions is influenced by social norms and expectations regarding their roles as spouse and parent – factors that are likely to change with globalization, urbanization, and rising educational attainment. Mensch, Singh, and Casterline (2006) found a marked reduction in the percent of 15-19-year-olds married throughout most LMICs over the past 30 years. These reductions were particularly striking in SSA. Even so, SSA remains the region with the highest rates of child marriage in the world (Singh and Samara 1996), for the most part driven by Western and Central Africa.[29] Western and Central Africa are also the regions with the highest percentage of women ever married by age 25, while in Eastern and Southern Africa the likelihood of

[29] Niger (Western Africa) has the highest rate of child marriage in the world, followed – within the SSA context – by Central African Republic and Chad (Central Africa). These are also the regions in which arranged marriage is more commonly practiced.

being still unmarried at 25 is higher (Mensch, Singh, and Casterline 2006).[30] Southern Africa has actually had a late marriage pattern since the early 1970s, and is now the only sub-region in SSA to exhibit non-negligible shares of never-married individuals (about 15 percent of women at age 45), partly due to labor migration (Tabutin and Schoumaker 2004).

Western Africa is also distinctive in that in most countries age at marriage has been increasing for women but not for men, likely due to changes in the practice of polygyny – an idiosyncratic feature of the region. Research suggests that in SSA the expansion of schooling had some impact on delaying women's age at marriage, yet a considerable fraction of the increase cannot be accounted for by changes in education. Conversely, rising costs of establishing a household have been found to contribute more than increasing educational attainment to men's marriage delays (Mensch, Singh, and Casterline 2006).

Differential increases in men and women's ages at first union affect inter-spouse age differences, whose variation across societies can be interpreted in terms of two interrelated factors, namely kinship structure and women's status. Casterline, Williams, and McDonald (1986) suggested that in patriarchal societies and in societies characterized by patrilineal kinship organization, the spousal age difference tends to be relatively large, and unions in which the husband is ten or more years older are relatively frequent. Conversely, in settings where the traditional social structure allows for a more equal status of spouses, or where exposure to Western family forms and modernization processes have improved the status of women – such as Southern Africa – the age difference tends to be smaller. Indeed, variation in inter-spouse age differences is also explained by marriage market – namely, age-structure – constraints. On the whole, research from SSA suggests that age differences at first marriage have narrowed, though they remain important in a subset of Western African countries such as Guinea-Bissau and Sierra Leone (Tabutin and Schoumaker 2004).

The widespread geographical heterogeneity in urbanization, globalization, and socio-cultural practices – such as child marriage, arranged marriage, polygyny, patriarchy, and patrilocality – and their differential prevalence and decline over time leads to expect heterogeneous patterns of mating by sub-region of SSA, with Western Africa and Southern Africa following the most diverse – likely, opposed – trajectories.[31]

Educational homogamy, inequality, and the social stratification system

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[30] Trends in ages at first marriage are deeply intertwined with educational expansion and urbanization patterns. For instance, Mensch, Singh, and Casterline (2006) reported that in Eastern and Southern Africa, more than four times as many women with 0 to 3 years of schooling married by age 18 as did women with 8-plus years of schooling. Similarly, 1.6 times as many women in rural areas married before age 18 as did women in urban areas.
[31] Even within sub-regions, but this is beyond the scope of this paper.

It has been widely recognized that the degree of partners' homogamy along specific socio-economic characteristics has the potential to shape different dimensions of inequality. Among these is household income inequality. Thinking about marital sorting on education – provided there is a reasonable correlation between educational attainment and later-life earnings – societies in which high-educated marry other high-educated and low-educated marry other low-educated will be more unequal than those in which high-educated marry low-educated. Increased educational assortative mating may affect inequality through changing the distribution of household configurations (or "types"), regardless of whether the increase itself is produced by shifts in shares of people with certain levels of education (so-called, structure), or changed sorting behavior (so-called, preferences). Given that household types possess different amounts of human capital – hence, different income potentials – a changed distribution of household types is expected to change inequality between types (Breen and Andersen 2012).[32]

While, as discussed in the previous sections, documenting trends and variation in educational assortative mating has taken a rather global and comparative scale (except for SSA), studies assessing implications of mating for inequality have centered primarily on high-income societies. Hu and Qian (2015) and Torche (2010) are notable exceptions from China and Latin America (Brazil, Chile, and Mexico), respectively. From this body of studies there is overwhelming agreement that educational assortative mating plays a small to negligible role in explaining trends in household income inequality. In the US context, Western, Bloome, and Percheski (2008) found that neither educational inequalities in women's incomes nor assortative mating contributed significantly to the rise in inequality. Similar results for the US are echoed in Breen and Salazar (2011), Eika, Mogstad, and Zafar (2017), and Greenwood et al. (2014). Similar conclusions were reached in the European context by Breen and Salazar (2010) for the UK and Boertien and Permanyer (2017) for a subset of 21 European countries. A minor exception to this finding is Breen and Andersen (2012), who showed that in Denmark – where inequality increased between 1987 and 2006 but educational homogamy declined – changes in assortative mating increased income inequality by about 7 percent, almost fully driven by changes in the educational distribution of men and women rather than in the propensity to choose a partner with a given level of education.

Several hypotheses have been proposed to shed light on the weak relationship between educational assortative mating and income inequality (Schwartz 2013). One postulates that increases in

[32] Note that part of this process is contingent on realizing the income potential once the couple is formed, which tends to be achieved through post-marital labor supply decisions (Breen and Andersen 2012; Gonalons-Pons and Schwartz 2017).

educational homogamy may not be large enough to produce meaningful shifts in inequality (Breen and Salazar 2011). Yet Boertien and Permanyer (2017) showed that even under extreme counterfactual scenarios, results would not change. Another hypothesis is that increases in educational homogamy among some types of couples might be offset by declines among other types of couples, such that the overall effect on inequality is negligible (Rosenfeld 2008). Alternatively, it might be the case that wives' education is not as highly correlated with earnings as one would think. This very much depends on post-sorting labor supply adjustments, but if most women exit the labor force upon union formation, the correlation between the two would be driven down. Consistent with this hypothesis is the case of Denmark, where high percentages of wives are in the labor force (higher than in the US), and changes in educational homogamy have had a larger impact on income inequality (Breen and Andersen 2012). In light of the latter hypothesis, some of the most recent literature has claimed that women's relative position within the couple and their labor supply decisions might constitute the "missing link" in explaining increases in family income inequality (Gonalons-Pons and Schwartz 2017).

Again, none of these dynamics have ever been studied in SSA. The main challenge in this context – as in many other LMICs – is the lack of good measures of household income or, even more so, the lack of measures of each partner's earnings. However, most existing surveys such as the DHS collect information of household assets that enter the computation of a wealth index which is measured at the household level. Previous research has shown that in contexts where household income or consumption is absent, wealth indices are effective indicators of long-term socio-economic position, living standards, or material well-being of households (Filmer and Pritchett 1999, 2001; McKenzie 2005; Sahn and Stifel 2000). Shimeles and Ncube (2015) have shown that this is also the case in Africa. My analysis investigates whether educational assortative mating has implications for inequality defined as inequality between households in asset possession. Scant evidence exists on whether measuring inequality through assets in SSA is a meaningful approach in the context of mating studies. Although not ideal and perhaps far from measuring an equivalent of partners' income, an approach of this kind has the potential to shed some light on the relationship between mating and inequality, thus providing some foundations for understanding the social stratification system in the African context.

2.3 Data and measures

The analysis uses pooled cross-sectional Demographic and Health Survey (DHS) data – 126 survey waves – from 39 sub-Saharan African countries (Table 2.1).[33] DHS are publicly available nationally

[33] On average, four waves per country.

representative surveys of women ages 15-49 collected by ICF International in collaboration with host country governments. Standardized questionnaires allow for comparisons across countries and survey waves. SSA countries are grouped in four regions – namely Western (14 countries), Central (8 countries), Eastern (12 countries), and Southern Africa (5 countries) – according to the classification provided by the United Nations Statistics Division (UNSD). The pooled comparative analysis offers a series of advantages over country-specific studies, including more observations, more variance on key variables, and the ability to assess pan-national regional trends (for similar analyses using DHS data see Clark and Brauner-Otto 2015; DeRose and Kravdal 2007; Reniers and Tfaily 2012; Smith-Greenaway and Trinitapoli 2014, etc.). The analysis spans a 30-year time frame, with the oldest surveys collected in 1986 in Liberia and Senegal, and the most recent survey collected in 2016 in Ethiopia. Additional details on the countries included, the number of waves, and the number of observations (couples) per wave are provided in Appendix Table A2.1.

In line with the observation that the focus on marriage cohort – rather than survey year or birth cohort – is more adequate for detecting trends in educational homogamy (Mare 1991), in this study I assess time trends over marriage cohort (MC). A similar perspective has been adopted in several prominent studies in the field (Casterline, Williams, and McDonald 1986; Smits and Park 2009; etc.). I construct ten 5-year marriage cohorts: <1970, 1970-1974, 1975-1979, 1980-1984, 1985-1989, 1990-1994, 1995-1999, 2000-2004, 2005-2009, >=2010.[34] This approach is sensible when using DHS data as surveys are collected at non-regular intervals, hence only data from selected countries are available for each survey year.

While DHS collect couple-level files in some countries, this study relies on information provided in the women's file to maximize the number of couples in the analysis.[35] I use the partnership information provided by the women to construct a couple-level dataset where wives and husbands are nested within couples.[36] Women whose marital status is missing or who provide no information on their own and their partner's educational attainment are excluded from the sample. I keep couples who are currently married or living in a cohabiting union ("living together"), and rely on the DHS definition of marital union, which includes both civil and customary marriages – as prevalent in the African context (van de Walle and Meekers 1994). In so doing, I follow previous scholarship in the claim that in settings where the definition of union is ambiguous and the process of union formation is "fluid,"

[34] Note that the first and last cohorts span more than five years for sample size reasons.
[35] The couple-level file is not available for every country, and the sample of couples would be restricted by about two-thirds.
[36] Throughout the paper I use the terms "husband" and "male partner", "wife" and "female partner", and "marriage" and "union" interchangeably.

distinguishing between formal marriages and informal unions may be impossible, hence the combination of the two constitutes the correct focus (Casterline, Williams, and McDonald 1986; Clark and Brauner-Otto 2015; Gage 1995). The sample is further restricted to couples where women are between the ages of 25 and 40. The reason is that by age 25 virtually all women have reached their highest educational level, and 95 percent of them have entered their first union, therefore reducing concerns about censoring on single marital status or education (Esteve et al. 2012). To avoid specification problems, I perform sensitivity analyses using both narrower and wider age ranges (15-49, 20-35, 30-45); results obtained are essentially the same and reported in the Appendix. As the DHS only provide data on the year of first union and include information on the education of the current partner/husband – but not any previous one – the sample is limited to couples where women have been married or have cohabited only once, i.e. about 82 percent of women (in a spirit similar to Casterline, Williams, and McDonald 1986).[37] These restrictions provide a sample of 416,038 couples with complete information on marital status, year of first union, and educational level of both partners.

The DHS includes a categorical and a continuous measure of educational attainment, namely highest level attained, and grade attained. The categorical variable is coded as follows: 0 for "no education", 1 for "primary", 2 for "secondary", and 3 for "higher." The continuous variable ranges from a minimum of 0 to a maximum of 23. While the continuous variable offers a more precise measure of schooling achievement, it ignores the importance of academic boundaries, which matter more for determining whether individuals marry "within their group." Furthermore, this latter classification captures similar stages in the educational career, even if these stages represent a different number of years across countries. Table 2.2 provides descriptive statistics on the number of couples and the highest level (panel a) and grade (panel b) attained by wives and husbands, by marriage cohort.[38] Estimates suggest that couples from the earliest marriage cohort (<1970) have on average some lower primary schooling, with husbands completing 2.7 grades, as compared to wives completing around 1.4 grades. Conversely, couples from the latest marriage cohort (>2010) possess upper primary/secondary education, with wives and husbands attaining an average of 8.3 and 9 school grades, respectively. Overall, the table shows a steep increase in educational attainment over marriage cohort, with a

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[37] DHS include a question on the total number of unions the woman has been in: "Have you been married or lived with a man only once or more than once?". All women reporting two or more unions are considered to have ever been remarried. Note that the sample is not restricted to men who have only married once. Indeed, the high prevalence of polygyny, particularly in Western Africa, suggests that many of the sampled men have married more than once (Fenske 2015; Reniers and Tfaily 2012; Smith-Greenaway and Trinitapoli 2014; Wagner and Rieger 2015).

[38] I here report both measures of schooling attainment to assess consistency and comparability, yet I rely on the categorical variable for all analyses that follow.

proportionally faster increase – yet no gender gap reversal – in wives' educational attainment, hinting at decreasing intra-household schooling inequality over time. Most importantly, a comparison between the two panels suggests a high degree of consistency between the categorical and the continuous measures. For instance, wives' averages in the 1970-1974 marriage cohort are 1.2 and 1.3 times their <1970 value for the categorical and continuous measures, respectively, while husbands' averages are 1.1 times their <1970 value for both measures. Similarly, wives' averages in the latest marriage cohort are 5.1 and 5.8 times their <1970 value for the categorical and continuous measures, respectively, while husbands' averages are 3.2 and 3.3 times their <1970 value. Appendix Table A2.2 provides some descriptive statistics on spousal differences in age by marriage cohort and shows similar patterns. While in the earliest marriage cohort the average difference is 11 years, in the latest it is reduced by about half.[39]

To measure household wealth, I rely on the International Wealth Index (IWI), the first comparable asset-based wealth index measuring the level of material well-being and standard of living in the complete developing world (Smits and Steendijk 2015). IWI is a stable and understandable yardstick for comparing the performance of societies with regard to wealth, inequality and poverty. IWI runs from 0 to 100, with 0 representing households having none of the assets and lowest quality housing, and 100 representing households having all assets and highest quality housing. Information collected on the possession of consumer durables, access to basic services and housing characteristics is entered into a factor analysis (PCA) from which the first factor is selected as the wealth index.[40]

Thanks to the inclusion of a household identifier, the IWI can be merged to the original DHS datasets. Note, however, that the IWI cannot be computed for some DHS surveys collected before (or around) 1990. It follows that the analytical sample included in the wealth analysis is reduced to 392,486 couples (~94 percent of the original sample), for a total of 112 survey waves across 38 countries – rather than 126 across 39 countries.[41] The main benefit of the IWI over the standard wealth index provided in the DHS lies in its comparability across countries and over time. As a matter of fact, the standard DHS wealth index is specific to the situation in each country at the time of the survey, making

[39] Due to a high number of missing cases in the age of the current partner/husband, the number of couples for these age-analyses is reduced to 373,831.
[40] Information on 12 assets is needed to compute the IWI of a household. These assets include seven consumer durables (possession of a TV, fridge, phone, bike, car, a cheap utensil and an expensive utensil), access to two public services (water and electricity) and three housing characteristics (number of sleeping rooms, quality of floor material and of toilet facility). For additional details on the IWI see Smits and Steendijk (2015).
[41] There is only one DHS survey for Botswana collected in 1988. The IWI for this country is not available, hence the country is not included in the wealth analysis. The survey waves that are excluded in the wealth analysis due to the unavailability of the IWI are reported in Appendix Table A1 in italic.

it a reliable measure only for households within a certain country-year combination. This is not to claim that the IWI provides a flawless measure of assets and wealth – its limitations will be discussed in the concluding remarks – yet it is more suited to studies that are comparative in nature.

2.4 Descriptive statistics

Figure 2.1 describes the types of unions prevailing in SSA in the earliest (left panel) and latest (right panel) marriage cohorts available for each country.[42] The graph reports the share of homogamous ("="), hypergamous ("H"), and hypogamous ("W") unions for the 39 countries, in a spirit similar to De Hauw, Grow, and Van Bavel (2017). The dominant pattern across cohorts is one in which the highest share of couples is homogamous followed, respectively, by hypergamous and hypogamous unions. Some exceptions are noteworthy. First, looking at the left panel we observe that the share of hypergamous couples is higher than – or very close to – the share of homogamous couples in countries such as Angola (AGO), Gabon (GAB), Mozambique (MOZ), Sao Tome and Principe (STP), and Uganda (UGA) – yet there is a trend towards declining hypergamy across cohorts observed in all five countries. Second, hypogamous unions are more prevalent than hypergamous unions in Botswana (BWA), Lesotho (LSO), Namibia (NAM), and Swaziland (SWZ), highlighting the somewhat peculiar nature of Southern African countries, and providing a first indication that assortative mating dynamics might differ by sub-region. In terms of extreme cases, Lesotho and Liberia stand out for being the countries with the highest shares of hypogamous and hypergamous unions, respectively.[43]

The two panels combined suggest significant changes in the composition of couples between the earliest and latest marriage cohort, evidencing a far narrower distribution in the right panel, driven primarily by a combination of increasing hypogamy and declining hypergamy. The coexistence of these opposing dynamics (namely, W moving the right and H moving to the left) alters the prevalence of homogamy only to a small extent. In fact, as shown in Appendix Figure A2.1 (top panel), country and sub-regional trends in the share of homogamous couples are heterogeneous – declining across Western and Central Africa and mildly increasing across Eastern and Southern regions – and point towards a decline for SSA as a whole from 0.7 to around 0.6. Hidden from these figures is, however, an assessment of the extent to which the composition of homogamous couples has changed over time,

[42] Not all ten marriage cohorts are available for each country, especially if only one survey wave per country is available. For instance, there is only one DHS for Angola, collected in 2015. As only women 25-40 are included in the sample, the oldest women were born around 1975 and entered their first union around 1990. Hence, the first marriage cohort available for Angola is the 1990-1994 one. These discrepancies are likely to create issues when analyzing trends at the country level, but less so when trends are analyzed at the sub-regional level, as mostly done in this paper.
[43] Lesotho reports the highest shares of hypogamous couples in both marriage cohorts.

i.e., whether variation along the educational distribution is responsible for observed upward or downward trends in educational homogamy.

Figure 2.2 plots the share of unions involving men and women of the same educational strata by educational level, for SSA as a whole (top panel) and by location of residence (bottom panel). The top panel points towards declining shares of couples with no education and increasing shares of couples with secondary or higher education. As the share of couples with both partners having primary education has remained virtually unchanged, this graph suggests that homogamy trends in SSA have been mostly driven by changes at the bottom and the top of the educational distribution. Specifically, the share of couples with both spouses having no education has declined from about 0.5 to 0.1, while the share of couples with both spouses having secondary or higher education has increased from 0 to 0.22 and 0.14, respectively. The steep decline in couples with no education (also shown in Figure A2.1, middle panel, by country and sub-region) thus more than offsets the weaker increase in couples with higher education (also shown in Figure A2.1, bottom panel, by country and sub-region), producing a downward overall trend in the share of homogamous couples – all levels of education combined.

Estimates by location of residence (Figure 2.2, bottom panel) show vastly different trends between urban and rural areas. While most of the decline in the share of homogamous couples with partners having no education is occurring in rural areas, increasing shares of couples with partners having secondary or higher education are driven primarily by urban areas. This is reasonable, as these areas underwent rapid industrialization earlier in time, thereby creating economic growth and job opportunities drawing people to cities, in tandem with a faster expansion of higher education and access to other public services.

Although the share of unions involving men and women of the same educational strata is a straightforward measure of educational homogamy (Mare 1991), trends in educational assortative mating based on this variable should be interpreted with caution (Schwartz and Mare 2005; Torche 2010). The reason is that variation in observed proportions in different categories of the joint distribution of partners' education is the outcome of two "forces": variation in the marginal distributions (e.g., declines in the share of women with no education over time), and variation in the association between partners' educational attainment net of marginal distributions (Torche 2010). For instance, the share of homogamous unions may simply be higher in the earliest marriage cohort because of the high concentration of husbands and wives in the "No education" category. Even given a constant association between husbands' and wives' levels of education, periods in which the marginal distributions are highly concentrated tend to produce a higher percentage of homogamous unions

(Schwartz and Mare 2005). Furthermore, the share of homogamous unions neglects information about the permeability of boundaries (Mare 1991), i.e., the ease with which individuals marry outside their own educational group. In what follows, I address some of these criticisms and explore whether the strength of the association between husbands' and wives' education has increased, or whether this trend is altered after controlling for shifts in the marginal distributions of husbands' and wives' education. In so doing, my ultimate goal is to provide a more precise categorization of educational assortative mating patterns in SSA.

2.5 Educational assortative mating

Trends

To measure educational assortative mating I follow an approach similar to Eika, Mogstad, and Zafar (2017) and Greenwood et al. (2014), based on contingency tables and marital sorting parameters. For every given marriage cohort, each cell in the contingency table gives the observed fraction of partnered households that occurs in a specific educational pairing. Positive (negative) educational assortative mating is defined as men and women with the same level of education marrying more (less) frequently than what would be expected under a marriage pattern that is random with respect to education. Marital sorting between education levels e_h and e_w is then the observed probability that a husband with education level e_h is married to a wife with education level e_w, relative to the probability under random mating with respect to education:

$$s(e_h, e_w) = \frac{Pr(E_h = e_h, E_w = e_w)}{Pr(E_h = e_h)\Pr(E_w = e_w)}$$

where E_h (E_w) denotes the education level of the husband (wife). Positive assortative mating occurs when the marital sorting parameter $s(e_{h,}e_w)$ is larger than 1 when i is equal to j. In a contingency table world, the diagonal of the contingency table describes the matches that occur when husbands and wives have the same educational level. This observed pattern of mating can be compared with the one that would obtain if husbands and wives matched randomly.[44] Taking the sum along the diagonals for each of these two types of matches, actual and random, and computing the ratio of these two sums, we obtain $s(e_{h,}e_w)$. The estimated marital sorting parameters – relative sum of diagonals – by marriage

[44] Proportions under random mating are the expected frequencies under the independence assumption (i.e. the product of the marginal distributions for husbands and wives). For explanatory purposes, contingency tables by marriage cohort for SSA as a whole are reported in Appendix Table A2.3.

cohort are plotted in Figure 2.3 by sub-region (top panel) and location of residence (bottom panel). The exact value of the sorting parameters is provided in Appendix Table A2.4.

Figure 2.3 provides evidence of positive educational assortative mating in SSA. That is, the ratios are larger than one, implying that the number of matches between husbands and wives with identical education is larger than what would occur if matching was random. Sorting parameters are higher for the latest marriage cohort relative to the earliest one, both for SSA as a whole and for each sub-region individually (top panel), suggesting that educational assortative mating has increased over subsequent cohorts. However, while for the whole SSA the marital sorting parameters increase monotonically from 1.4 to approximately 2 – meaning that in the latest marriage cohort assortative matches occur twice as often relative to a situation in which matches are formed randomly – sub-regional trends are heterogeneous. For instance, positive assortative mating in early cohorts is lower in Western Africa, yet this region experiences the steepest increase in the sorting parameter followed, in turn, by Eastern and Central Africa. Conversely, Southern Africa experiences mild increases across early cohorts, followed by a downward trend thereafter. Steep upward trends in Western Africa and relatively flat/downward trends in Southern Africa are confirmed in Appendix Figure A2.2, which tests the robustness of the findings to alternative age ranges of women. The Southern African trends that emerge from this analysis are unique within SSA, and consistent with the hypotheses outlined in the theoretical background.

The bottom panel of Figure 2.3 provides estimates of the marital sorting parameters by location of residence and shows evidence of positive educational assortative mating in both urban and rural areas. Yet, although mating in early cohorts is higher in urban areas, most of the increase in mating across cohorts is accounted for by changes in rural areas, where the sorting parameter increases monotonically from 1.3 to about 2.1. Conversely, overall trends in urban areas are fairly flat. Note that Southern Africa is the only sub-region where the sorting parameter does not follow an upward trend neither in urban nor rural areas, and where the rural-urban divide in mating patterns is less stark. As such, these findings seem consistent with the status attainment hypothesis in rural areas, and the inverted U-curve hypothesis in urban areas, where greater geographical mobility, educational expansion, cross-cultural exchange, and mass communication contribute to gradually spreading the logic of 'romantic love.'

Given that conclusions about changes in educational assortative mating are dependent on the methodology used (Blossfeld 2009; Rosenfeld 2008; Schwartz 2013; Schwartz and Mare 2005), in Appendix Figure A2.3 I present results using an alternative measure, namely Kendall's tau correlation between husband and wife's highest level attained in each 5-year marriage cohort.[45] Despite minor discrepancies, this supplementary analysis confirms my main findings, i.e., the steep increase in the tau-correlation in Western Africa, the uniqueness of Southern Africa as the only sub-region where mating has not increased, and the pivotal role of rural areas in driving SSA mating patterns.

Log-linear models

To be in line with some of the most prominent sociological literature on educational assortative mating (Mare 1991; Schwartz and Mare 2005; Smits, Ultee, and Lammers 1998; Torche 2010; etc.), I complement the above analysis with a series of log-linear models. The underlying motivation is to summarize international variation in marital sorting in the best – defined as a combination of fit and parsimony – possible way. Log-linear models are appropriate in that they provide estimates of the changing association between couples' educational characteristics while controlling for shifts in their marginal distributions (e.g., Agresti 2002). I estimate the following baseline model:

$$ln(F_{ijkl}) = \lambda + \lambda_i^H + \lambda_j^W + \lambda_k^R + \lambda_l^M + \lambda_{ij}^{HW} + \lambda_{ik}^{HR} + \lambda_{il}^{HM} + \lambda_{jk}^{WR} + \lambda_{jl}^{WM} + \lambda_{kl}^{RM} + \lambda_{ikl}^{HRM} + \lambda_{jkl}^{WRM} + \lambda_{ijk}^{HWR}$$

where H is husband's education (i=1,2,3,4), W is wife's education (j=1,2,3,4), R is sub-region (k=1,2,3,4), and M is marriage cohort (l=<1970, …,>=2010).[46] In line with findings from the previous analysis, I estimate models separately for the overall sample, urban areas, and rural areas. F_{ijkl} is the expected number of unions between husbands in education category i and wives in education category j, in sub-region k, from marriage cohort l. This baseline model captures variation in the distribution of husband's and wife's education by cohort and sub-region (λ_{ikl}^{HRM} and λ_{jkl}^{WRM}), allows the interaction between husband's and wife's education to vary by sub-region (λ_{ijk}^{HWR}), and contains all lower-order terms. In a second step I add homogamy and crossing parameters to the baseline specification to assess which model fits the data best. A homogamy model is:

$$ln(F_{ijkl}) = Baseline\ model + \gamma_{ol}^{OM}$$

[45] Kendall's tau is a measure of rank correlation, given by the difference between the number of concordant and discordant pairs of couples relative to the total number of pairs of couples. A pair of couples is said to be concordant if both the wife and husband in one couple have higher education than the wife and husband in the other couple. The pair of couples is discordant if one couple has a wife with lower education and a husband with higher education as compared to the other couple. The Kendall correlation ranges from -1 to 1 and it is closer to 1 the more similar the ranks of the spouses are in the marginal distribution of education of husbands and wives.
[46] For operationalization purposes, the two categorical variables for education – wife's education and husband's education – have been recoded as 1,2,3,4 instead of 0,1,2,3.

where O=1 if husband's education category equals wife's education category, and 0 otherwise. γ_{ol}^{OM} estimates the change in the odds of homogamy in marriage cohort l relative to the baseline year (<1970). A crossing model is:

$$ln(F_{ijkl}) = Baseline\ model + \gamma_{ijl}^{CM}$$

$$\text{where } \gamma_{ijl}^{CM} = \begin{cases} \sum_{q=j}^{i-1} \gamma_{ql}\ for\ i > j \\ \sum_{q=i}^{j-1} \gamma_{ql}\ for\ i < j \\ 0\ for\ i = j \end{cases}$$

γ_{ql} represents the change in the difficulty of crossing educational barrier q in marriage cohort l relative to the baseline year. Crossing models summarize the association between spouses' education as a series of barriers to marriage between education groups, or in terms of the relative permeability of boundaries between adjacent education groups. Hence, the crossing parameters measure the log odds of marriage for couples in adjacent education categories relative to the log odds of homogamy, net of the marginal distributions of spouses' education (Schwartz and Mare 2005; Torche 2010). Past research from the US has found these models to fit marriage data quite well (Blackwell 1998; Mare 1991).

Table 2.3 provides the model specifications and fit statistics of the log-linear models. The table is divided in three panels, for SSA as a whole (panel a), urban areas (panel b), and rural areas (panel c). I present both the deviance and the Bayesian Information Criterion (BIC) statistics for model fit, yet rely mainly on the latter due to large sample sizes that make it hardly possible to find a model that does not significantly differ from the saturated model (Raftery 1995). More negative BIC statistics indicate a better-fitting model, and differences in BIC values that are larger than 10 provide good evidence that the model with the more negative BIC fits the data better.

Model 1 is the baseline model, which assumes that the educational resemblance of spouses is time invariant – yet it is allowed to vary by sub-region.[47] This model (panel a) fits the data better than alternative general specifications that allow the association to vary by marital cohort (Model 2). Model 3 is the homogamy trend model, which parameterizes the trend as a change in the likelihood that husbands and wives share the same education level. By the BIC, adding this term does not alter the fit of the model relative to the baseline model, suggesting that the tendency for couples to marry within the same education category for SSA as a whole has not changed significantly over subsequent marriage

[47] The model in which the husband-wife association is assumed to be both time-invariant and region-invariant (i.e., HRM and WRM, excluding HWM and HWR) has a deviance of 4,365 and a positive BIC of 3,308. As the model fits the data poorly by conventional standards, it is not reported in the analysis – yet it is available upon request.

cohorts. This model, however, might conceal variation in trends across different portions of the distribution (Schwartz and Mare 2005). Model 4 hence allows the degree of homogamy to differ across the diagonal cells of the homogamy table, showing a marginal improvement over the homogamy model described by a single parameter. Model 5 significantly improves the fit of the model by including an asymmetry parameter which accounts for the possibility that men and women "marry up" or "marry down" with respect to socioeconomic characteristics.[48] Model 6 is the crossing trend model, which adds terms to capture variation in the difficulty of crossing educational boundaries across the education distribution. According to the BIC, the crossing model provides a better fit to the data than the baseline model, while it performs worse than the gender asymmetry specification. Models 7, 8, 9, and 10 are similar to 3, 4, 5, and 6, respectively, although the added parameters (O, D, A, and C) are allowed to vary by sub-region, testing the assumption that variation over marriage cohort might unfold differently across contexts. Among these, Model 7 (OMR) – allowing the homogamy parameter to vary by marriage cohort and region – significantly improves upon previous specifications and best describes the data. This finding indicates that while the tendency for couples to marry within the same education category for SSA as a whole has not changed significantly over marriage cohort (Model 3), the same is not true once sub-regional differences are accounted for.

Models 11 to 14 build on Model 7 – the preferred specification up to this point – by adding inter-cohort (Models 11 to 13) and cross-regional (Model 14) variation in diagonal, asymmetry, and crossing parameters. As Model 12 has the lowest BIC, I conclude that the best-fitting model is one that permits the homogamy parameter to vary by marriage cohort and sub-region (OMR), while allowing for an asymmetric tendency to marry up or down to vary by marriage cohort (AM). Not least, consistent with my previous findings I show that model specifications summarizing international variation in marital sorting differ between urban and rural areas, and that SSA trends are mostly driven by variation in rural areas. In line with panel a, in rural areas (panel c) Model 7 is the best fitting model, and Model 12 provides a similar level of fit. Conversely, in urban areas (panel b), the baseline model is good enough to summarize variation in the data. Overall, this analysis shows that in SSA trends in assortative mating are better described by inter-cohort and cross-regional variation in homogamy and heterogamy, rather than by changes in the degree to which couples cross educational barriers. This

[48] The variable is created as 0 if husband and wife have the same education, 1 if the husband "marries down" and 2 if the husband "marries up" – irrespective of how many educational categories there are between husband and wife (e.g., the variable is 1 both for the pairing "husband-higher education" and "wife-no education" and the pairing "husband-higher education" and "wife-secondary education").

finding is novel in itself and departs from previous scholarship in the US (Mare 1991; Schwartz and Mare 2005) and Latin America (Torche 2010).

2.6 Wealth analysis

Trends in wealth dispersion

Before assessing the implications of educational assortative mating for household wealth inequality, I begin the analysis by exploring how between-household wealth inequality has evolved over marriage cohort. Note that for this portion of the analysis the number of marriage cohorts is reduced from ten (5-year) to five (10-year) to maximize sample variability.[49] As the IWI is measured on a 0-100 scale in every country and it is comparable both between countries and over time, I measure inequality through the most straightforward measure of dispersion, i.e. the variance or standard deviation (SD). Specifically, I compute the variance of the IWI for every country-cohort combination.

Figure 2.4 provides a geographical overview of wealth dispersion (in SD) by country and marriage cohort. In a spirit similar to Figure 2.1, I provide estimates for the earliest (left panel) and latest (right panel) marriage cohorts available for each country. The map shows that wealth dispersion is on average higher in Southern Africa and has increased across cohorts throughout most of SSA. There are some exceptions to this pattern in countries such as Gabon, Nigeria, and Central African Republic, where wealth dispersion shows a downward trend. Table A2.5 in the Appendix reports estimates from an OLS regression of the IWI SD on a categorical variable for marriage cohort. Estimates show that wealth dispersion has been increasing over marriage cohort, with only marginal differences between urban and rural areas. For instance, compared to the SD in the earliest marriage cohort, the SD in IWI in the latest marriage cohort for SSA as a whole is 6 to 7 units higher (panel a). Although there is a dearth of research on patterns of wealth inequality in SSA – mostly due to the complexities inherent in measuring social and economic performance in this region (Harttgen, Klasen, and Vollmer 2013; Klasen and Blades 2013) – my findings are consistent with figures from the African Development Bank (Shimeles and Nabassaga 2018).[50] Other recent studies suggest that inequality

[49] Also, as households/couples in more recent cohorts have likely had less time to accumulate assets/wealth, by widening the horizon to ten years we are likely to obtain a more representative and balanced picture.

[50] Data on income inequality are instead more readily available and show that SSA remains one of the most unequal regions in the world. Ten of the 19 most unequal countries globally are in SSA and seven outlier African countries drive this inequality. Between 1991 and 2011, a clear bifurcation in inequality trends existed across countries in the region. Furthermore, 17 countries (predominantly agricultural economies from West Africa and a few from other regions) experienced declining inequality, whereas 12 countries, predominantly in Southern and Central Africa and economies characterized by an important oil and mining sector, recorded an inequality rise (UNDP 2017). Although asset measures largely differ from income measures, there is a good degree of

trends across countries in Africa have not leveled off, with no downward pattern emerging either with respect to the recent economic resurgence, or any other improvements in the level of human development (Bigsten 2018; Fosu 2015).

Counterfactual analysis

To assess implications of educational assortative mating for household wealth inequality, I follow a simple approach that well suits micro-level data. Specifically, I model the cohort-specific variance (VAR) of wealth

$$VAR[W] = [E(W^2) - (E(W))^2]$$

and use regression analysis to estimate counterfactual expectations reweighting the betas using either "assortative" (observed) or "random" (counterfactual) proportions from the above contingency tables. For each marriage cohort, each component of the variance in the above formula (i.e., the second moment and the squared mean) is regressed onto a series of dummies for whether the couple is homogamous with both partners having no education (reference category), both partners having primary education, both partners having secondary education, both partners having higher education, and partners having discordant levels of education (the off-diagonals).[51] After obtaining the betas, expectations are computed multiplying the betas by either the observed or counterfactual proportions. This way, for each marriage cohort I obtain an estimated variance computed under observed proportions, and an estimated variance computed under counterfactual proportions. With these quantities, I compute the share of cohort-specific inequality attributable to educational assortative mating as follows:

$$\%ineq = \frac{VAR[W]_{observed} - VAR[W]_{counterfactual}}{VAR[W]_{observed}}$$

where $VAR[W]_{counterfactual} = (VAR \middle| VAR_{mating=random})$.

How would wealth inequality change if we imposed random – instead of assortative – mating in each marriage cohort? Table 2.4 reports the variance under observed proportions and random proportions, by urban/rural location of residence (panel a) and sub-region (panel b). Estimates for SSA as a whole (panel a) show that the share of inequality attributable to mating is low, reaching at most 3.7 percent in the latest cohort. Further disaggregation unravels interesting heterogeneity, suggesting that only in

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consistency between the maps I provide and the UNDP findings on income inequality, especially for what concerns Southern African countries such as Botswana, Lesotho, Namibia, South Africa, Zimbabwe, etc.

[51] These regression estimates are not reported in the analysis (available upon request).

urban areas mating explains a share of the cohort-specific inequality, albeit low. Heterogeneity by sub-region (panel b) further reveals that the low shares accounted for by mating are driven primarily by Western Africa. Conversely, in Eastern Africa mating accounts for up to 12 percent of the cohort-specific inequality in wealth, followed in turn by Southern Africa (at most 10-11 percent) and Central Africa (at most 7 percent). High shares in Eastern Africa are aligned with the urban/rural differences identified in panel a, in that – as of 2014 – Eastern Africa has the lowest share of urban population (25 percent, against 44 percent in Western and Central Africa, and 61 percent in Southern Africa), yet it exhibits the highest urbanization rate within SSA, with an average annual increase in the urban population of 4.5 percent (UN-DESA 2015). Overall, these findings support the idea that educational assortative mating accounts for a non-negligible share of the cohort-specific inequality in wealth. Indeed, these are not sizeable coefficients, yet they still point to a link between educational assortative mating and household wealth inequality which has not been previously identified in the literature.

Can changes in mating overtime explain time trends – mostly, the increase – in wealth inequality? To answer this question, I examine what would happen to wealth inequality if couples from the latest marriage cohort matched as those in the earliest ones. Methodologically, I re-compute the variance in the latest cohort, applying the observed proportions from the earlier cohorts. However, changes in observed proportions are affected by shifts in marginal distributions. To overcome this issue, I use the Sinkhorn-Knopp algorithm (1967), an iterative procedure outlined in Mosteller (1968) and recently adopted by Greenwood et al. (2014), to construct standardized contingency tables, such that two contingency tables have the same marginal distributions associated with the rows and columns.[52] After imposing the marginal distributions of the latest marriage cohort to all preceding cohorts, I thus compute the new "observed" proportions and estimate the corresponding variances.

Table 2.5 provides results from the simulation exercise described above. The first two columns in each sub-panel rely on "unadjusted" (i.e., affected by differences in marginal distributions) observed proportions, while the last two columns rely on "adjusted" (i.e., independent of differences in marginal distributions) observed proportions obtained through the above iterative procedure. Focusing on unadjusted estimates for SSA as a whole (panel a), the first two columns suggest that wealth inequality in the latest cohort (>=2005) would be lower by about 19 percent if we imposed the observed pattern of mating from the earliest cohort (<1975), with trends very much driven by rural areas. The opposite

[52] The basic idea is to fix the marginal distributions of a contingency table and rework the internal cells such that the "new" marginal distributions are respected. Once two contingency tables have the same marginal distributions, the cells within the table can be compared. Taking a 4x4 table, this can be standardized so that each element of the two marginal distributions is ¼.

trend is observed in urban areas, where wealth inequality would actually be larger if we imposed the observed pattern of mating from the earliest cohort (by 7 percent). However, if we rely on adjusted estimates we observe that changes in mating hardly move time trends in wealth inequality, irrespective of location of residence and SSA sub-region. The bottom line of this exercise is that whatever is the share of the time trend in inequality that is accounted for by changes in mating, this is driven by shifting educational distributions rather than by assortative mating itself.

Note that – differently from income and consumption expenditure data – IWIs and asset indices in general are not adjusted for household size or other demographic characteristics of the household. The reason is that the assets used for constructing these indices consist almost exclusively of household public goods. Housing characteristics, access to basic services and durables like a TV, fridge, clock, or car tend to benefit all household members (Smits and Steendijk 2015).[53] In any case, to provide a proxy for "crowding" and evaluate whether household characteristics explain any variability in the IWI, I re-estimate IWI variances controlling for some household characteristics, namely the total number of household members (residents plus visitors), a dummy for whether the partner lives in the household or elsewhere, the total number of sons living at home, and the total number of daughters living at home. Appendix Table A2.6 replicates panel a of Table 2.4 comparing variances computed under observed and random proportions, controlling for the above-listed household characteristics. Estimated variances are almost identical to those provided in Table 2.4, in line with most of the available literature (Filmer and Scott 2012; Rutstein and Johnson 2004; Sahn and Stifel 2000). This finding suggests I am not missing any significant household-related variability in the estimation of variances.

Lastly, I conduct these analyses by country selecting the three countries where inequality has increased the most between the earliest and latest marriage cohort – namely, Guinea, Rwanda, and Uganda – and the three countries where inequality has increased the least (or has decreased) – Central African Republic, Congo, and Zimbabwe. Results show that even in these "extreme" cases, changes in mating itself explain trends in wealth inequality to a negligible extent – available upon request.

[53] As explained in Smits and Steendijk (2015), in some studies the number of sleeping rooms is divided by the number of persons in the household, to obtain an indicator of "crowding". For IWI this is not the case, as the number of rooms is meant to be an indicator of the size of the house and not of crowding. A house with three sleeping rooms is generally bigger and more expensive than a house with less sleeping rooms and this is independent of family size.

2.7 Conclusions and discussion

This study has provided a comprehensive analysis of educational assortative mating across 39 countries in sub-Saharan Africa, a region of the world that has experienced rapid socio-economic and demographic change yet has been largely neglected in the assortative mating literature. Adopting a marriage-cohort temporal perspective and computing measures that net out the confounding role of shifting educational distributions, I have shown that mating in SSA has followed rather different trajectories both by sub-region and by location of residence. Specifically, while there is evidence of positive educational assortative mating throughout SSA – i.e., men and women with the same level of education marrying more frequently than what would be expected under a marriage pattern that is random with respect to education – mating has increased over subsequent cohorts in Western, Central, and Eastern Africa, yet it has flattened out and somewhat decreased in Southern Africa. Heterogeneity is also evident in levels and relative growth, as mating was lower in Western Africa for early cohorts, yet the sub-region has witnessed the steepest increase in the marital sorting parameter. Additionally, I have shown that increases in mating have been largely driven by rural areas – where the trend for SSA as a whole is consistent with the status attainment hypothesis – while mating in urban areas has shown a mild increase followed by an incipient decline – consistent with the inverted U-curve framework and the increasing applicability of the general openness hypothesis. Overall, the documented heterogeneity – and, foremost, the diverging trends between Western and Southern Africa – is consistent with the economic (e.g., urbanization), socio-demographic (e.g., changes in families), and cultural specificities (e.g., patriarchal norms) of each sub-region.

In the second part of the analysis I have explored implications of educational assortative mating for household wealth inequality measured through the International Wealth Index. Using counterfactual simulations both within and across-cohorts, I have shown that assortative mating accounts for a non-negligible share of the cohort-specific inequality in household wealth, which ranges sub-regionally between 3 and 12 percent and is wholly driven by urban areas. Mating accounts for a higher share of inequality in Southern Africa – the most urbanized sub-region – and Eastern Africa – the sub-region that has experienced the highest rates of urbanization. Provided a link exists between mating and wealth inequality, the steepest increases in mating in rural areas would have led us to expect the share of cohort-specific inequality attributable to mating to be higher in rural areas. Empirical evidence contradicts this expectation, thus strengthening the claim that mating matters little for wealth inequality. This is further confirmed by cross-cohort simulations, which show that changes in mating over time barely move the time trends in wealth inequality. This finding echoes the solid body of evidence from high-income societies claiming that mating plays a small to negligible role in explaining trends in household income inequality (Breen and Salazar 2011; Eika, Mogstad, and Zafar 2017) and pushes to consider additional factors – missing in the present analysis – such as women's labor supply decisions (Gonalons-Pons and Schwartz 2017; Greenwood et al. 2014; etc.).

To the best of my knowledge, this is the first large-scale study focusing on trends, variation, and implications of mating in SSA. As such, it suffers from several limitations that set the stage for future research. First, the data and measures present limitations that relate to the nature and sampling frame of DHS data. As the DHS only provide data on the year of first union and include information on the education of the current partner – but not any previous one – the sample was limited to couples where women had been married only once, while no restrictions were imposed on men due to the lack of information on their marriage order. While this approach follows existing literature (e.g., Casterline, Williams, and McDonald 1986) and pretty much aligns with the claim that a focus on first marriages is what really matters for understanding mating patterns (Schwartz and Mare 2012), there is still room for improvement. However – at least as of now – there is no other dataset that would permit an analysis of mating patterns in SSA with analogous coverage.

Another possible source of concern in studies of mating is the classification of educational levels which, as stated by Blossfeld (2009), should be neither too crude nor too detailed to be informative. Ideally, it is key to define categories such that differences between them reflect well-chosen attainment levels with social significance. Only in that case increases in homogamy rates can really be interpreted as indicators of social closure, and increases in intermarriage as gains in social openness. Although the DHS collect educational variables similarly across countries and over time, making sure the above requirements are satisfied is always a challenge in broad-scale comparative studies.

Methodologically, this study – as many other studies of mating that build on cross-sectional data – takes marital matches as the starting point and attempts to explain trends and variation in mating through spouses' individual characteristics. As such, the analysis excludes all those individuals who are still single at the time of the interview. This is likely to create issues in societies with increasing single rates at the beginning of the life course. However, I believe in the SSA context this is less problematic, as getting married remains the largely predominant social norm for both men and women and virtually everyone eventually enters a union (Tabutin and Schoumaker 2004). Given the increasing proportion of never-married individuals in Southern Africa, this is likely the only sub-region in which this omission is likely to introduce some bias. Another methodological issue is tied to the scale of the analysis. As mating is ultimately determined by the availability of partners and potential matches, its functioning is more easily – and perhaps properly – understood at a finer level of analysis, such as districts or cities. As this study sought to provide an overview of patterns for the region as a whole, this ultimately boils down to the usual trade-off between breadth of analysis and level of detail.

Lastly, the wealth analysis presents some limitations that pertain to the type of measure used – which is the only available for most LMICs. The IWI has the advantage of easy reproducibility, as it builds on the same set of assets across countries. At the same time, its universality may be a drawback, as finding a small set of assets common to such a large number of surveys requires discarding a lot of the asset information gathered about any given household – at the risk of obtaining a variable that is somewhat uninformative. Despite not being dismissive of asset indices, Harttgen, Klasen and Vollmer (2013) claimed that asset indices substantially overstate the pace of poverty reduction as there is strong evidence of 'asset drift', i.e., an accumulation of assets over time, with households accumulating assets such as mobile phones and TVs without getting any less poor. Households often accumulate these assets because they are becoming relatively cheaper, preferences are shifting towards them, and households often do not dispose of them; but this does not mean that these households are any less poor as a result. Using data from South Africa, Wittenberg and Leibbrandt (2017) also showed that, despite being powerful tools for analyzing social trends, asset indices often fail an internal validity test, that is, ranking individuals with rural assets below individuals with no assets at all – thus leading to an exaggerated sense of rural deprivation and a lack of appreciation for deprivation in urban areas. Lastly, there is skepticism on whether asset indices may proxy for measures on income. In line with Sahn and Stifel (2003) and Filmer and Scott (2012), Smits and Steendijk (2015) claimed that asset indices are more indicators of longer-term, more stable, aspects of household's economic status, rather than monetary or expenditure-based welfare measures. As such, it is not clear (yet) whether it makes sense to study mating patterns in relation to analyses of inequality based on asset measures. The high degree of consistency between my results and research on mating and inequality in high-income societies provides some reliability to the findings. Yet these are certainly not indisputable, and further advances in the field will permit to assess their robustness.

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2.9 Tables

Table 2.1: Number of countries and survey waves included in the analysis, by region of sub-Saharan Africa

Regional classification of Sub-Saharan African countries			
Western	**Central**	**Eastern**	**Southern**
Benin (4)	Angola (1)	Burundi (2)	Botswana (1)
Burkina Faso (4)	Cameroon (4)	Comoros (2)	Lesotho (3)
Cote d'Ivoire (2)	Central African Republic (1)	Ethiopia (4)	Namibia (4)
Gambia (1)	Chad (3)	Kenya (6)	South Africa (1)
Ghana (6)	Congo (2)	Madagascar (4)	Swaziland (1)
Guinea (3)	Congo, DR (2)	Malawi (5)	
Liberia (3)	Gabon (2)	Mozambique (3)	
Mali (4)	Sao Tome and Principe (1)	Rwanda (5)	
Mauritania (1)		Tanzania (5)	
Niger (4)		Uganda (5)	
Nigeria (4)		Zambia (5)	
Senegal (7)		Zimbabwe (6)	
Sierra Leone (2)			
Togo (3)			
14 countries - 48 surveys	**8 countries - 16 surveys**	**12 countries - 52 surveys**	**5 countries - 10 surveys**

Notes: Regional classification from the United Nations Statistics Division (UNSD). Number of survey waves in parentheses.

Table 2.2: Summary statistics on couples' education, by marriage cohort

Marriage cohort	a. Highest level attained					b. Grade attained				
		Wife		Husband			Wife		Husband	
	N	Average	Ratio over <1970	Average	Ratio over <1970	N	Average	Ratio over <1970	Average	Ratio over <1970
<1970	4,956	0.32	.	0.54	.	4,893	1.43	.	2.71	.
		(0.011)		(0.015)			(0.055)		(0.081)	
1970-1974	12,718	0.40	1.2	0.62	1.1	12,536	1.83	1.3	3.09	1.1
		(0.008)		(0.010)			(0.042)		(0.059)	
1975-1979	25,384	0.46	1.4	0.67	1.2	24,959	2.18	1.5	3.37	1.2
		(0.007)		(0.009)			(0.036)		(0.047)	
1980-1984	40,607	0.55	1.7	0.75	1.4	40,069	2.62	1.8	3.76	1.4
		(0.006)		(0.008)			(0.032)		(0.040)	
1985-1989	55,511	0.63	2.0	0.83	1.6	54,927	3.00	2.1	4.19	1.5
		(0.006)		(0.007)			(0.030)		(0.037)	
1990-1994	74,626	0.68	2.1	0.91	1.7	73,977	3.26	2.3	4.57	1.7
		(0.006)		(0.007)			(0.030)		(0.036)	
1995-1999	82,918	0.79	2.5	1.02	1.9	82,293	3.82	2.7	5.17	1.9
		(0.006)		(0.007)			(0.032)		(0.038)	
2000-2004	69,789	0.93	2.9	1.15	2.1	69,358	4.56	3.2	5.89	2.2
		(0.007)		(0.008)			(0.038)		(0.043)	
2005-2009	37,903	1.22	3.8	1.39	2.6	37,633	6.11	4.3	7.21	2.7
		(0.010)		(0.010)			(0.055)		(0.057)	
>=2010	11,626	1.62	5.1	1.72	3.2	11,565	8.31	5.8	9.05	3.3
		(0.017)		(0.016)			(0.092)		(0.092)	
Total	416,038					412,210				

Notes: Weighted estimates using sample DHS weights. Standard errors in parentheses. "Ratio over <1970" gives the relative ratio of the value in each cohort compared to the <1970 one, i.e., the earliest.

Table 2.3: Log-linear models for the association between partners' educational attainment

Overall				Urban				Rural			
Model	***df***	**Deviance**	**BIC**	**Model**	***df***	**Deviance**	**BIC**	**Model**	***df***	**Deviance**	**BIC**
(1) HRM, WRM, HWR	153	702.55	-196.3	(1) HRM, WRM, HWR	140	322.69	-492.9	(1) HRM, WRM, HWR	132	351.12	-413.6
(2) Model 1 + HWM	109	304.12	-176.2	(2) Model 1 + HWM	98	198.01	-372.9	(2) Model 1 + HWM	91	229.50	-297.7
(3) Model 1 + OM	148	671.03	-198.5	(3) Model 1 + OM	135	310.07	-476.4	(3) Model 1 + OM	127	340.06	-395.7
(4) Model 1 + DM	133	568.06	-213.3	(4) Model 1 + DM	121	272.14	-432.8	(4) Model 1 + DM	113	283.14	-371.5
(5) Model 1 + AM	143	589.47	-250.6	(5) Model 1 + AM	130	281.16	-476.2	(5) Model 1 + AM	122	315.49	-391.3
(6) Model 1 + CM	138	581.55	-229.2	(6) Model 1 + CM	126	279.05	-455.0	(6) Model 1 + CM	118	317.93	-365.6
(7) Model 1 + OMR	133	484.49	-296.9	(7) Model 1 + OMR	121	260.25	-444.7	(7) Model 1 + OMR	112	230.72	-418.1
(8) Model 1 + DMR	76	305.49	-141.0	(8) Model 1 + DMR	67	169.02	-221.3	(8) Model 1 + DMR	61	122.83	-230.5
(9) Model 1 + AMR	113	375.44	-288.4	(9) Model 1 + AMR	102	191.73	-402.5	(9) Model 1 + AMR	94	197.46	-347.1
(10) Model 1 + CMR	96	354.91	-241.0	(10) Model 1 + CMR	86	181.18	-319.9	(10) Model 1 + CMR	79	144.77	-312.9
(11) Model 7 + DM	118	413.70	-279.5					(11) Model 7 + DM	98	193.12	-374.6
(12) Model 7 + AM	128	414.80	-337.2					(12) Model 7 + AM	107	212.17	-409.7
(13) Model 7 + CM	118	337.60	-325.6					(13) Model 7 + CM	98	188.58	-379.1
(14) Model 7 + CMR	76	224.58	-221.9					(14) Model 7 + CMR	61	106.79	-246.6

Notes: Model terms: H=husband's education; W=wife's education; R=region; M=marriage cohort; O=homogamy (reduced homogamy); D=main diagonal (expanded homogamy); A=marrying up/down (asymmetry); C=crossing parameters.

Table 2.4: Cohort-specific variance in wealth (IWI) under observed and random mating scenarios, by location of residence (top panel) and region of sub-Saharan Africa (bottom panel)

a.	**Overall**			**Urban**			**Rural**		
Marriage cohort	Variance (assortative)	Variance (random)	% ineq.	Variance (assortative)	Variance (random)	% ineq.	Variance (assortative)	Variance (random)	% ineq.
<1975	328.1	316.6	3.5%	607.3	604.9	0.4%	116.7	116.7	0.0%
1975-1984	434.8	429.1	1.3%	681.9	646.6	5.2%	154.7	156.3	-1.0%
1985-1994	512.3	514.8	-0.5%	635.9	614.2	3.4%	220.1	223.8	-1.7%
1995-2004	584.0	576.0	1.4%	560.1	533.5	4.7%	271.4	269.6	0.7%
>2005	717.4	690.7	3.7%	474.9	454.0	4.4%	356.3	348.7	2.1%

b.	**Western**			**Central**			**Eastern**			**Southern**		
Marriage cohort	Variance (assortative)	Variance (random)	% ineq.	Variance (assortative)	Variance (random)	% ineq.	Variance (assortative)	Variance (random)	% ineq.	Variance (assortative)	Variance (random)	% ineq.
<1975	323.7	331.2	-2.3%	363.5	339.0	6.8%	280.3	246.3	12.1%	914.8	819.5	10.4%
1975-1984	364.9	376.5	-3.2%	415.6	382.2	8.0%	383.9	336.5	12.3%	1084.7	999.6	7.8%
1985-1994	460.3	464.6	-0.9%	586.9	584.6	0.4%	427.5	381.5	10.8%	984.7	923.0	6.3%
1995-2004	546.9	534.4	2.3%	638.2	629.4	1.4%	503.5	444.3	11.8%	869.2	807.7	7.1%
>2005	594.3	568.0	4.4%	843.6	800.3	5.1%	653.5	593.2	9.2%	776.7	694.4	10.6%

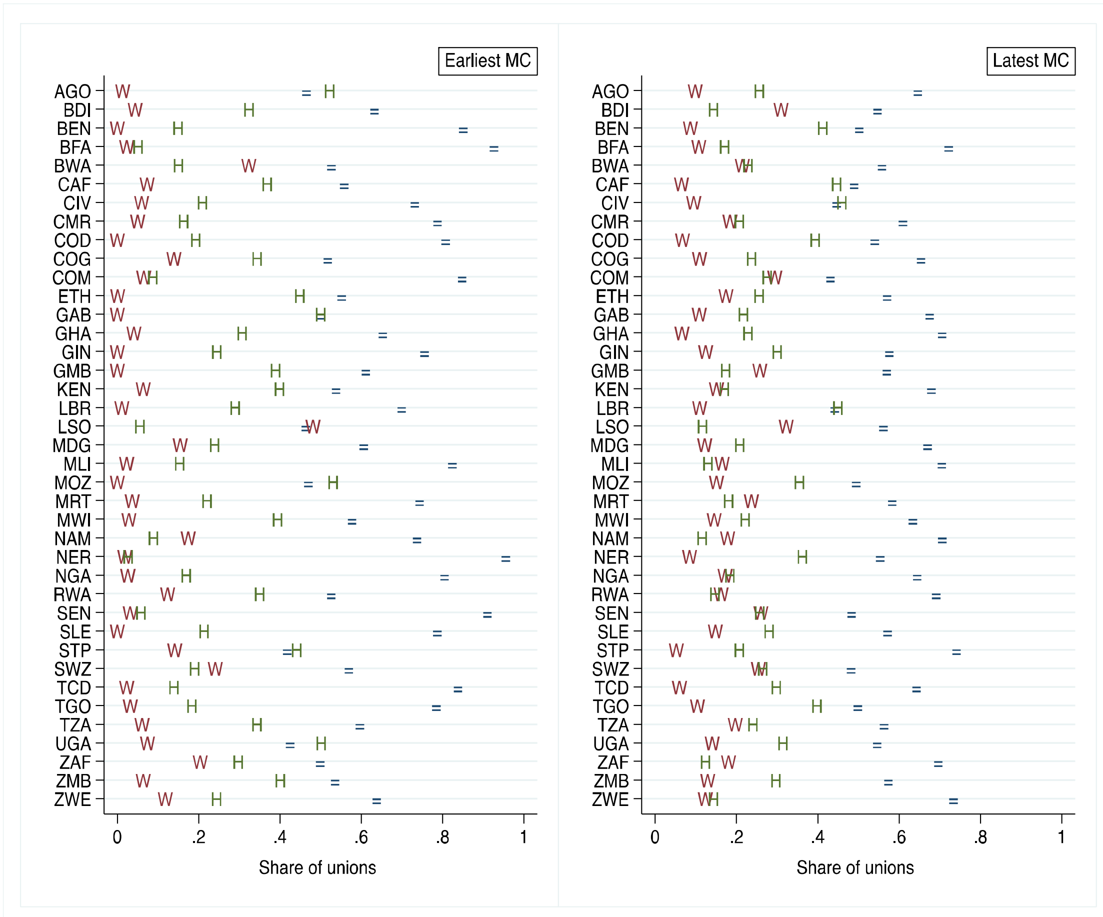
Table 2.5: Variance in wealth (IWI) for the latest marriage cohort (>=2005) under different counterfactual distributions

a.	**Overall**				**Urban**				**Rural**			
			Fixed MD (latest)				*Fixed MD (latest)*				*Fixed MD (latest)*	
Counterfactual distribution	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change
>2005	717.4	.	717.4	.	474.9	.	474.9	.	356.3	.	356.3	.
1995-2004	656.4	-8.5%	719.1	0.2%	483.4	1.8%	474.8	0.0%	330.0	-7.4%	356.3	0.0%
1985-1994	624.6	-12.9%	717.6	0.0%	490.6	3.3%	475.1	0.0%	321.0	-9.9%	355.6	-0.2%
1975-1984	598.1	-16.6%	720.8	0.5%	500.7	5.4%	475.5	0.1%	311.9	-12.4%	357.0	0.2%
<1975	578.4	-19.4%	722.5	0.7%	506.8	6.7%	476.9	0.4%	307.8	-13.6%	353.6	-0.8%

b.	**Western**				**Central**				**Eastern**				**Southern**			
			Fixed MD (latest)				*Fixed MD (latest)*				*Fixed MD (latest)*				*Fixed MD (latest)*	
Counterfactual distribution	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change	Variance latest MC	% change
>2005	594.3	.	594.3	.	843.6	.	843.6	.	653.5	.	653.5	.	776.7	.	776.7	.
1995-2004	580.6	-2.3%	596.5	0.4%	780.7	-7.5%	847.9	0.5%	562.7	-13.9%	655.3	0.3%	731.6	-5.8%	773.3	-0.4%
1985-1994	559.9	-5.8%	595.4	0.2%	743.0	-11.9%	846.4	0.3%	534.9	-18.1%	655.2	0.3%	709.5	-8.7%	776.4	0.0%
1975-1984	539.2	-9.3%	596.7	0.4%	670.1	-20.6%	851.5	0.9%	514.0	-21.3%	662.9	1.4%	688.9	-11.3%	776.8	0.0%
<1975	521.5	-12.3%	594.7	0.1%	562.8	-33.3%	817.2	-3.1%	497.4	-23.9%	664.1	1.6%	641.9	-17.4%	779.2	0.3%

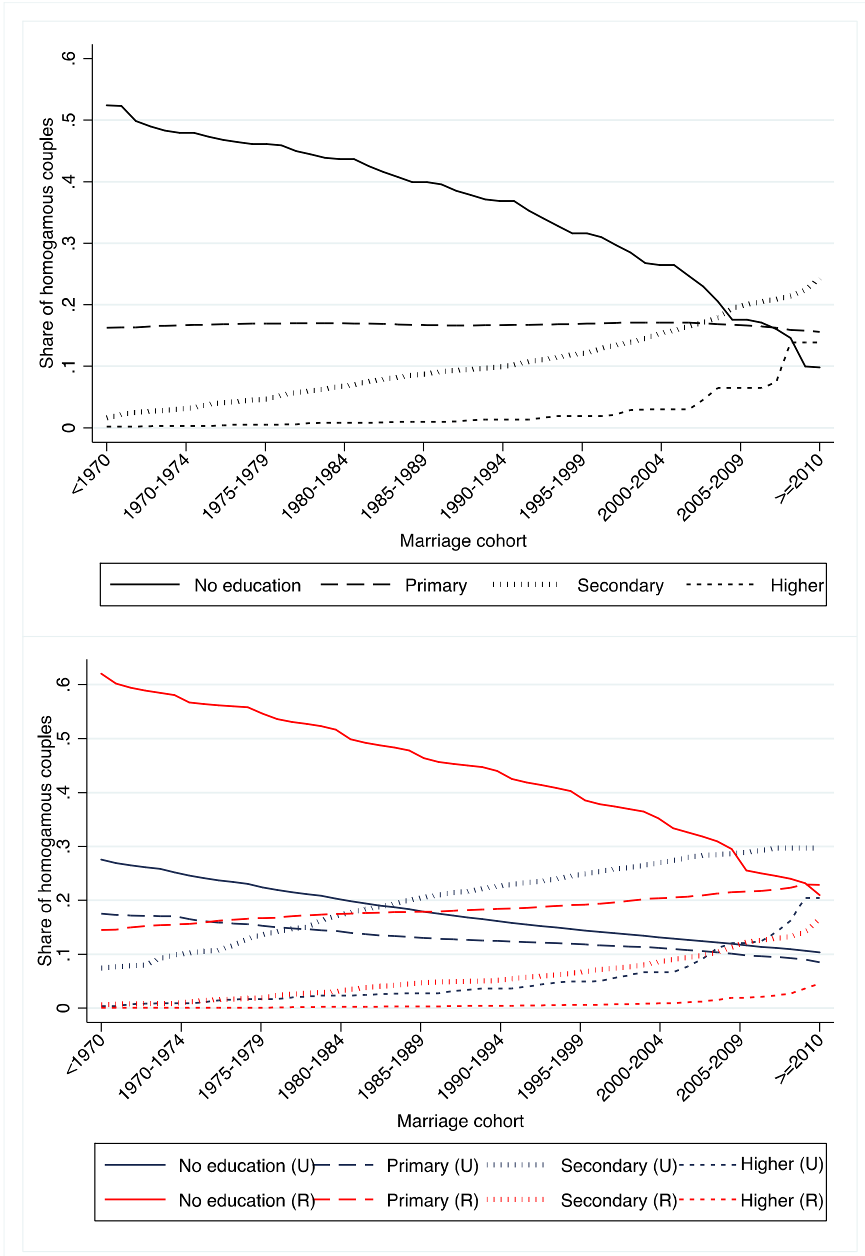
2.10 Figures

Figure 2.1: Share of homogamous (=), hypergamous (H), and hypogamous (W) unions for the earliest (left panel) and latest (right panel) marriage cohorts for each country



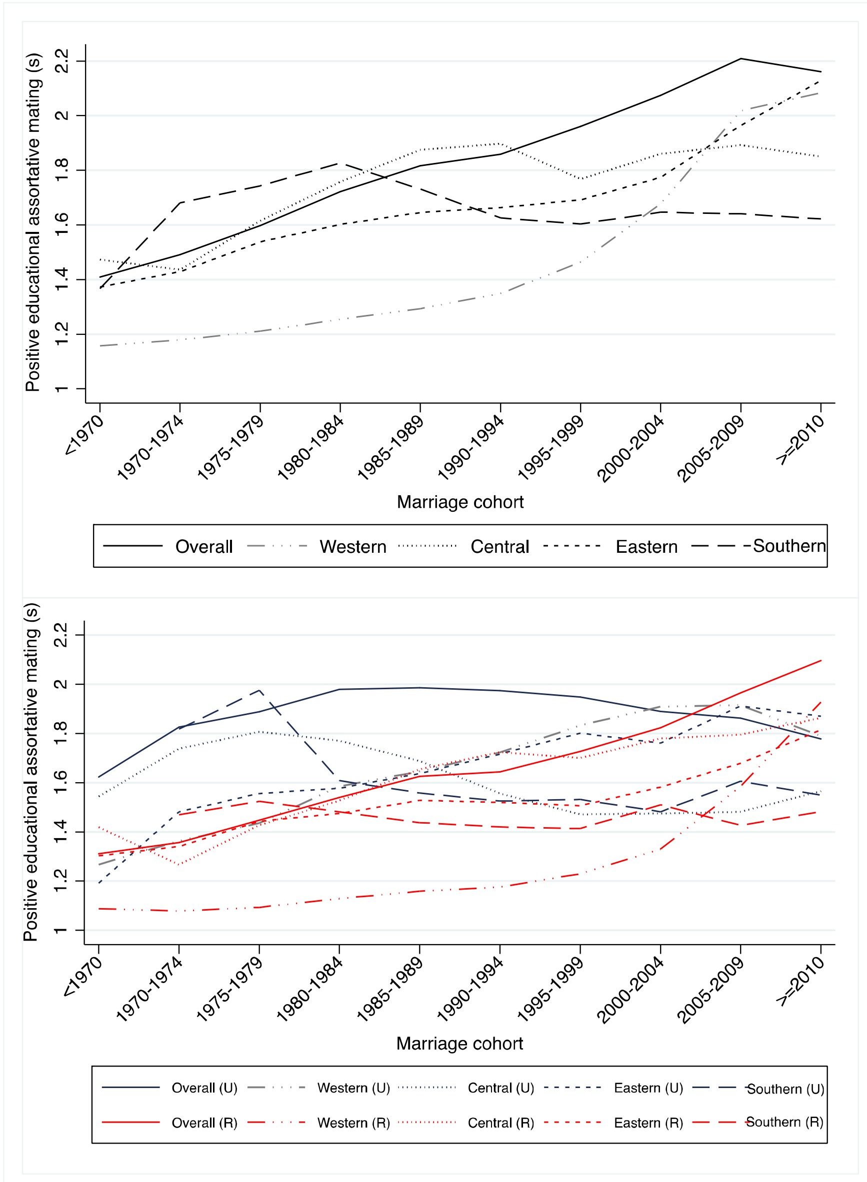
Notes: MC: marriage cohort. "W": Men<Women; "=": Men=Women; "H": Men>Women. Country codes: AGO-Angola; BDI-Burundi; BEN-Benin; BFA-Burkina Faso; BWA-Botswana; CAF-Central African Republic; CIV-Côte d'Ivoire; CMR-Cameroon; COD-Democratic Republic of the Congo; COG-Congo; COM-Comoros; ETH-Ethiopia; GAB-Gabon; GHA-Ghana; GIN-Guinea; GMB-Gambia; KEN-Kenya; LBR-Liberia; LSO-Lesotho; MDG-Madagascar; MLI-Mali; MOZ-Mozambique; MRT-Mauritania; MWI-Malawi; NAM-Namibia; NER-Niger; NGA-Nigeria; RWA-Rwanda; SEN-Senegal; SLE-Sierra Leone; STP-Sao Tome and Principe; SWZ-Swaziland; TCD-Chad; TGO-Togo; TZA-Tanzania; UGA-Uganda; ZAF-South Africa; ZMB-Zambia; ZWE-Zimbabwe.

Figure 2.2: Share of homogamous couples by educational level, for sub-Saharan Africa as a whole (top panel) and by location of residence (bottom panel)



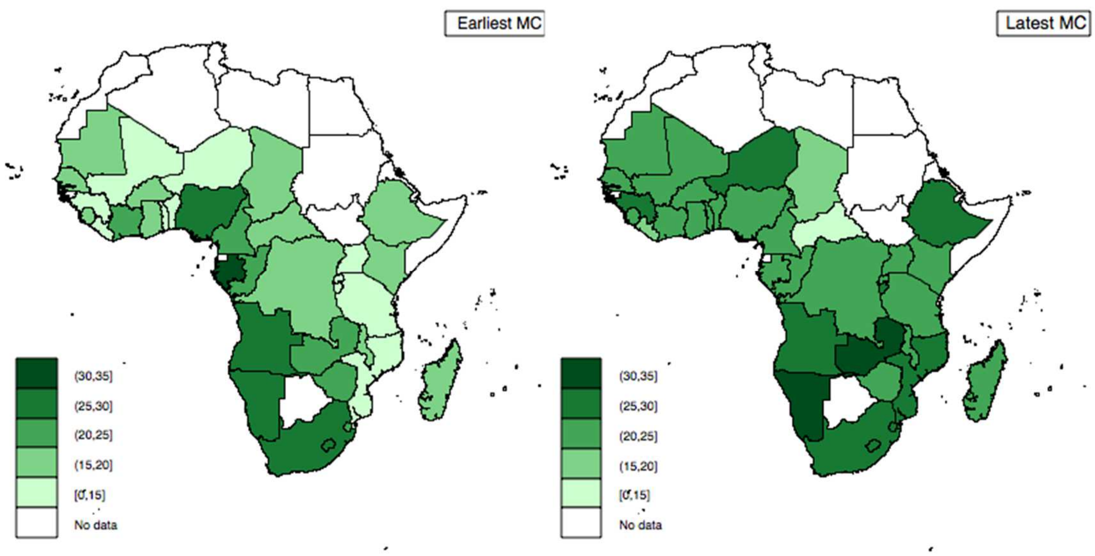
Notes: "U": urban; "R": rural.

Figure 2.3: Positive educational assortative mating (s parameter), by region of sub-Saharan Africa (top panel) and location of residence (bottom panel)



Notes: "U": urban; "R": rural.

Figure 2.4: Wealth dispersion (SD in IWI) for the earliest (left panel) and latest (right panel) marriage cohort, by sub-Saharan African country



2.11 Appendix

Table A2.1: Number of countries, survey waves, and couples per wave included in the analysis, by region of sub-Saharan Africa

Western (14 countries - 48 surveys)	Central (8 countries - 16 surveys)	Eastern (12 countries - 52 surveys)	Southern (5 countries - 10 surveys)
Benin	Angola	Burundi	Botswana
1996 (3,362)	2015 (6,764)	*1987 (2,718)*	*1988 (1,998)*
2001 (3,994)	Cameroon	2010 (5,572)	Lesotho
2006 (13,272)	*1991 (1,035)*	Comoros	2004 (1,803)
2011 (13,438)	1998 (1,403)	1996 (621)	2009 (2,059)
Burkina Faso	2004 (2,605)	2012 (1,345)	2014 (1,914)
1993 (4,346)	2011 (3,917)	Ethiopia	Namibia
1998 (4,490)	Central African Republic	2000 (3,623)	1992 (1,024)
2003 (8,682)	1994 (2,518)	2005 (4,001)	2000 (1,339)
2010 (13,082)	Chad	2011 (4,943)	2006 (1,761)
Cote d'Ivoire	1996 (2,351)	2016 (5,039)	2013 (1,690)
1994 (2,260)	2004 (1,880)	Kenya	South Africa
2011 (2,977)	2014 (6,644)	*1989 (2,568)*	1998 (2,838)
Gambia	Congo	1993 (2,453)	Swaziland
2013 (3,373)	2005 (3,210)	1998 (2,594)	2006 (1,081)
Ghana	2011 (2,635)	2003 (2,543)	
1988 (1,157)	Congo, DR	2008 (2,696)	
1993 (1,300)	2007 (5,460)	2014 (5,185)	
1998 (1,347)	2013 (11,242)	Madagascar	
2003 (1,581)	Gabon	1992 (1,377)	
2008 (1,325)	2000 (1,163)	1997 (1,608)	
2014 (2,755)	2012 (1,614)	2003 (2,113)	
Guinea	Sao Tome and Principe	2008 (4,572)	
1999 (2,551)	2008 (590)	Malawi	
2005 (2,966)		1992 (1,236)	
2012 (3,219)		2000 (3,389)	
Liberia		2004 (3,024)	
1986 (1,021)		2010 (6,470)	
2007 (1,733)		2015 (6,905)	
2013 (2,304)		Mozambique	
Mali		1997 (1,921)	
1995 (3,652)		2003 (2,923)	
2001 (4,790)		2011 (3,502)	
2006 (5,396)		Rwanda	
2012 (4,871)		1992 (2,014)	
Mauritania		2000 (2,405)	
2000 (1,763)		2005 (2,779)	
Niger		2010 (3,999)	
1992 (1,954)		2014 (4,261)	
1998 (2,209)		Tanzania	
2006 (3,188)		1991 (2,347)	
2012 (4,515)		1996 (2,288)	
Nigeria		2004 (2,966)	
1990 (3,427)		2010 (2,791)	
2003 (2,389)		2015 (3,665)	
2008 (12,081)		Uganda	
2013 (14,025)		*1988 (1,117)*	
Senegal		1995 (1,741)	
1986 (1,274)		2000 (1,829)	
1992 (1,817)		2006 (2,317)	
2005 (3,887)		2011 (2,496)	
2010 (4,846)		Zambia	
2012 (2,488)		*1992 (1,704)*	
2014 (2,553)		1996 (1,841)	
2015 (2,724)		2001 (1,862)	
Sierra Leone		2007 (1,942)	
2008 (2,561)		2013 (4,623)	
2013 (5,108)		Zimbabwe	
Togo		*1988 (1,228)*	
1988 (967)		1994 (1,714)	
1998 (2,780)		1999 (1,615)	
2013 (3,328)		2005 (2,251)	
		2010 (2,672)	
		2015 (2,964)	

Notes: Regional classification from the United Nations Statistics Division (UNSD). Number of couples per wave in parentheses. The 14 survey waves for which no IWI is available are reported in italic.

Table A2.2: Summary statistics on couples' age, by marriage cohort

Marriage cohort	N	Age			
		Wife	Husband	Diff.	Ratio over <1970
<1970	258	39.50	50.61	11.11	
		(0.069)	(0.603)	(0.594)	
1970-1974	3,285	38.49	49.26	10.77	0.97
		(0.041)	(0.193)	(0.184)	
1975-1979	11,665	36.71	46.99	10.28	0.92
		(0.039)	(0.120)	(0.111)	
1980-1984	28,545	34.96	44.78	9.82	0.88
		(0.045)	(0.083)	(0.068)	
1985-1989	49,810	33.59	42.94	9.34	0.84
		(0.041)	(0.067)	(0.051)	
1990-1994	73,873	32.53	41.32	8.79	0.79
		(0.033)	(0.052)	(0.043)	
1995-1999	84,241	30.92	39.14	8.22	0.74
		(0.027)	(0.045)	(0.040)	
2000-2004	71,309	29.25	36.80	7.55	0.68
		(0.025)	(0.046)	(0.042)	
2005-2009	38,862	28.31	34.94	6.63	0.60
		(0.027)	(0.053)	(0.048)	
>=2010	11,983	28.26	34.17	5.91	0.53
		(0.049)	(0.098)	(0.084)	
Total	373,831				

Notes: Weighted estimates using sample DHS weights. Standard errors in parentheses. "Ratio over <1970" gives the relative ratio of the value in each cohort compared to the <1970 one, i.e., the earliest.

Table A2.3: Contingency tables (actual and random mating) by marriage cohort, sub-Saharan Africa as a whole

	<1970							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.524**	**0.411**	0.051	0.148	0.001	0.016	0.000	0.001
Primary (H)	0.161	0.231	**0.157**	**0.083**	0.005	0.009	0.000	0.001
Secondary (H)	0.026	0.062	0.046	0.023	**0.016**	**0.002**	0.000	0.000
Higher (H)	0.003	0.009	0.003	0.003	0.005	0.000	**0.002**	**0.000**
Marginal	0.713		0.257		0.027		0.002	
	1970-1974							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.479**	**0.355**	0.053	0.151	0.003	0.027	0.000	0.002
Primary (H)	0.158	0.220	**0.166**	**0.094**	0.009	0.017	0.000	0.001
Secondary (H)	0.023	0.075	0.059	0.032	**0.031**	**0.006**	**0.001**	**0.000**
Higher (H)	0.003	0.012	0.005	0.005	0.008	0.001	0.003	0.000
Marginal	0.663		0.283		0.051		0.004	
	1975-1979							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.461**	**0.323**	0.049	0.148	0.004	0.039	0.000	0.004
Primary (H)	0.140	0.204	**0.170**	**0.094**	0.015	0.025	0.000	0.002
Secondary (H)	0.026	0.087	0.065	0.040	**0.047**	**0.011**	0.001	0.001
Higher (H)	0.002	0.014	0.005	0.006	0.011	0.002	**0.005**	**0.000**
Marginal	0.628		0.288		0.076		0.007	
	1980-1984							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.437**	**0.287**	0.049	0.145	0.006	0.054	0.000	0.005
Primary (H)	0.114	0.178	**0.170**	**0.090**	0.021	0.033	0.000	0.003
Secondary (H)	0.030	0.099	0.070	0.050	**0.067**	**0.019**	0.002	0.002
Higher (H)	0.003	0.020	0.006	0.010	0.016	0.004	**0.008**	**0.000**
Marginal	0.584		0.296		0.109		0.011	
	1985-1989							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.399**	**0.246**	0.049	0.138	0.006	0.064	0.000	0.006
Primary (H)	0.104	0.162	**0.169**	**0.091**	0.027	0.043	0.000	0.004
Secondary (H)	0.034	0.110	0.078	0.062	**0.088**	**0.029**	0.003	0.003
Higher (H)	0.004	0.023	0.008	0.013	0.022	0.006	**0.010**	**0.001**
Marginal	0.541		0.304		0.142		0.014	
	1990-1994							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.369**	**0.219**	0.048	0.130	0.007	0.067	0.000	0.008
Primary (H)	0.104	0.152	**0.163**	**0.090**	0.026	0.047	0.001	0.006
Secondary (H)	0.038	0.118	0.086	0.070	**0.099**	**0.036**	0.005	0.004
Higher (H)	0.004	0.027	0.009	0.016	0.026	0.008	**0.013**	**0.001**
Marginal	0.516		0.307		0.159		0.019	
	1995-1999							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.316**	**0.171**	0.050	0.121	0.007	0.072	0.000	0.010
Primary (H)	0.099	0.139	**0.172**	**0.098**	0.032	0.058	0.001	0.008
Secondary (H)	0.039	0.119	0.093	0.084	**0.122**	**0.050**	0.007	0.007
Higher (H)	0.005	0.029	0.009	0.021	0.032	0.012	**0.019**	**0.002**
Marginal	0.458		0.323		0.192		0.027	
	2000-2004							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.264**	**0.126**	0.045	0.104	0.010	0.077	0.000	0.013
Primary (H)	0.086	0.117	**0.171**	**0.096**	0.038	0.071	0.001	0.012
Secondary (H)	0.038	0.119	0.100	0.098	**0.154**	**0.073**	0.010	0.012
Higher (H)	0.005	0.032	0.009	0.027	0.039	0.020	**0.030**	**0.003**
Marginal	0.393		0.325		0.241		0.041	
	2005-2009							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.176**	**0.065**	0.043	0.073	0.013	0.074	0.001	0.021
Primary (H)	0.065	0.079	**0.170**	**0.089**	0.045	0.090	0.003	0.025
Secondary (H)	0.033	0.097	0.094	0.110	**0.200**	**0.110**	0.021	0.031
Higher (H)	0.005	0.038	0.008	0.043	0.059	0.043	**0.065**	**0.012**
Marginal	0.278		0.316		0.317		0.090	
	>2010							
	No Education (W)		Primary (W)		Secondary (W)		Higher (W)	
	Assortative	Random	Assortative	Random	Assortative	Random	Assortative	Random
No Education (H)	**0.098**	**0.024**	0.031	0.037	0.016	0.058	0.002	0.028
Primary (H)	0.037	0.036	**0.128**	**0.056**	0.050	0.087	0.006	0.043
Secondary (H)	0.023	0.064	0.083	0.100	**0.243**	**0.156**	0.048	0.077
Higher (H)	0.003	0.038	0.010	0.060	0.085	0.093	**0.139**	**0.046**
Marginal	0.161		0.252		0.394		0.194	

Table A2.4: Relative sum of diagonals of cohort-specific contingency tables (observed mating/random mating)

Marriage cohort	Relative sum of diagonals of cohort-specific contingency tables (actual mating/random mating)														
	Overall			**Western**			**Central**			**Eastern**			**Southern**		
	All	Urban	Rural	All	Urban	Rural	All	Urban	Rural	All	Urban	Rural	All	Urban	Rural
<1970	1.409	1.623	1.311	1.158	1.268	1.088	1.473	1.544	1.419	1.373	1.192	1.303	1.367	.	.
1970-1974	1.491	1.826	1.357	1.180	1.363	1.078	1.436	1.738	1.267	1.429	1.482	1.341	1.681	1.818	1.469
1975-1979	1.598	1.889	1.447	1.211	1.436	1.093	1.615	1.807	1.427	1.538	1.556	1.443	1.744	1.975	1.524
1980-1984	1.722	1.980	1.540	1.254	1.585	1.128	1.758	1.771	1.528	1.602	1.578	1.475	1.827	1.609	1.482
1985-1989	1.817	1.986	1.626	1.294	1.646	1.159	1.875	1.688	1.655	1.645	1.637	1.528	1.731	1.558	1.438
1990-1994	1.858	1.974	1.644	1.349	1.723	1.176	1.897	1.556	1.724	1.663	1.717	1.520	1.625	1.525	1.420
1995-1999	1.961	1.948	1.728	1.464	1.833	1.230	1.768	1.471	1.701	1.692	1.801	1.506	1.604	1.532	1.413
2000-2004	2.074	1.890	1.824	1.678	1.909	1.330	1.860	1.475	1.780	1.774	1.761	1.582	1.647	1.482	1.511
2005-2009	2.209	1.863	1.965	2.018	1.915	1.585	1.892	1.482	1.796	1.964	1.912	1.679	1.641	1.606	1.426
>2010	2.161	1.778	2.096	2.084	1.789	1.927	1.850	1.565	1.863	2.128	1.870	1.814	1.622	1.549	1.483

Table A2.5: Association between wealth dispersion (SD) and marriage cohort

IWI (SD)	a. All		b. Urban		c. Rural	
	(1)	(2)	(1)	(2)	(1)	(2)
Marriage cohort (Ref.<1975)						
1975-1984	2.036***	1.846***	2.441***	1.916***	1.859***	1.803***
	(0.070)	(0.068)	(0.173)	(0.126)	(0.077)	(0.079)
1985-1994	3.673***	3.412***	3.911***	3.351***	3.475***	3.390***
	(0.078)	(0.074)	(0.192)	(0.139)	(0.086)	(0.086)
1995-2004	4.984***	4.626***	4.884***	4.250***	4.807***	4.727***
	(0.080)	(0.074)	(0.198)	(0.139)	(0.089)	(0.086)
>=2005	7.055***	6.188***	6.534***	5.388***	6.958***	6.689***
	(0.086)	(0.079)	(0.203)	(0.146)	(0.095)	(0.093)
Constant	16.688***	23.289***	17.516***	23.822***	16.503***	23.087***
	(0.076)	(0.078)	(0.192)	(0.142)	(0.084)	(0.092)
Country FE	No	Yes	No	Yes	No	Yes
Obs.	392,486	392,486	126,272	126,272	266,214	266,214
R2	0.164	0.874	0.105	0.878	0.176	0.870

Table A2.6: Cohort-specific variance in wealth (IWI) under observed and random mating scenarios, estimating wealth controlling for household characteristics

	Overall			Urban			Rural		
Marriage cohort	**Variance (assortative)**	**Variance (random)**	**% ineq.**	**Variance (assortative)**	**Variance (random)**	**% ineq.**	**Variance (assortative)**	**Variance (random)**	**% ineq.**
<1975	327.8	318.9	2.7%	604.4	606.0	-0.3%	123.0	123.9	-0.7%
1975-1984	439.2	436.6	0.6%	683.1	648.5	5.1%	155.4	158.9	-2.3%
1985-1994	513.1	517.9	-0.9%	636.3	614.9	3.4%	220.3	226.2	-2.7%
1995-2004	584.1	577.9	1.1%	560.1	533.6	4.7%	271.4	272.7	-0.5%
>2005	717.4	690.7	3.7%	474.9	453.9	4.4%	356.3	351.1	1.5%

Figure A2.1: Share of homogamous couples: both partners with the same educational level (top panel), both partners with no education (middle panel), both partners with higher education (bottom panel), by country (left) and region (right) of sub-Saharan Africa

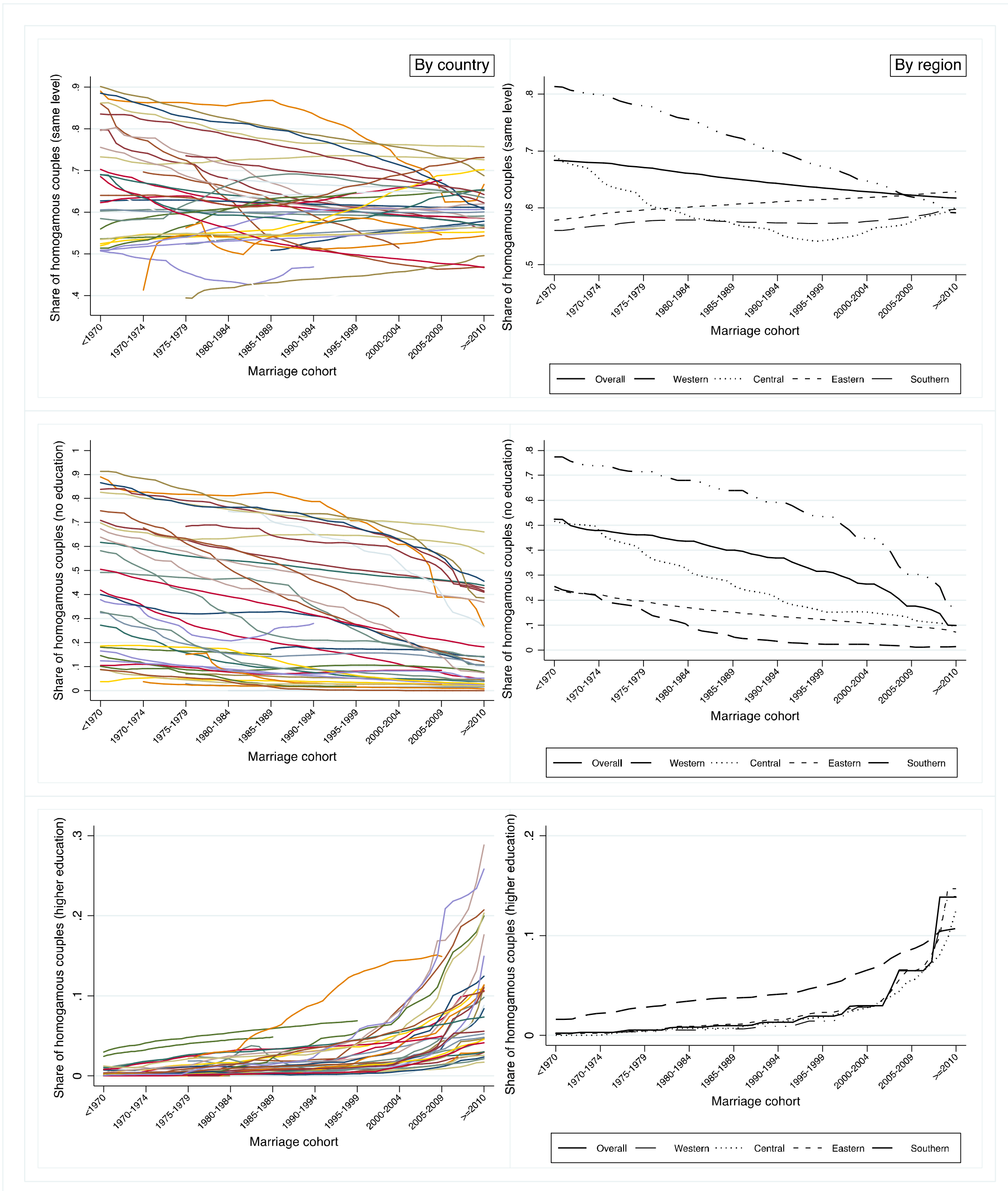


Figure A2.2: Positive educational assortative mating (s parameter), by region of sub-Saharan Africa and alternative age ranges of women: 15-49 (top), 20-35 (middle), and 30-45 (bottom)

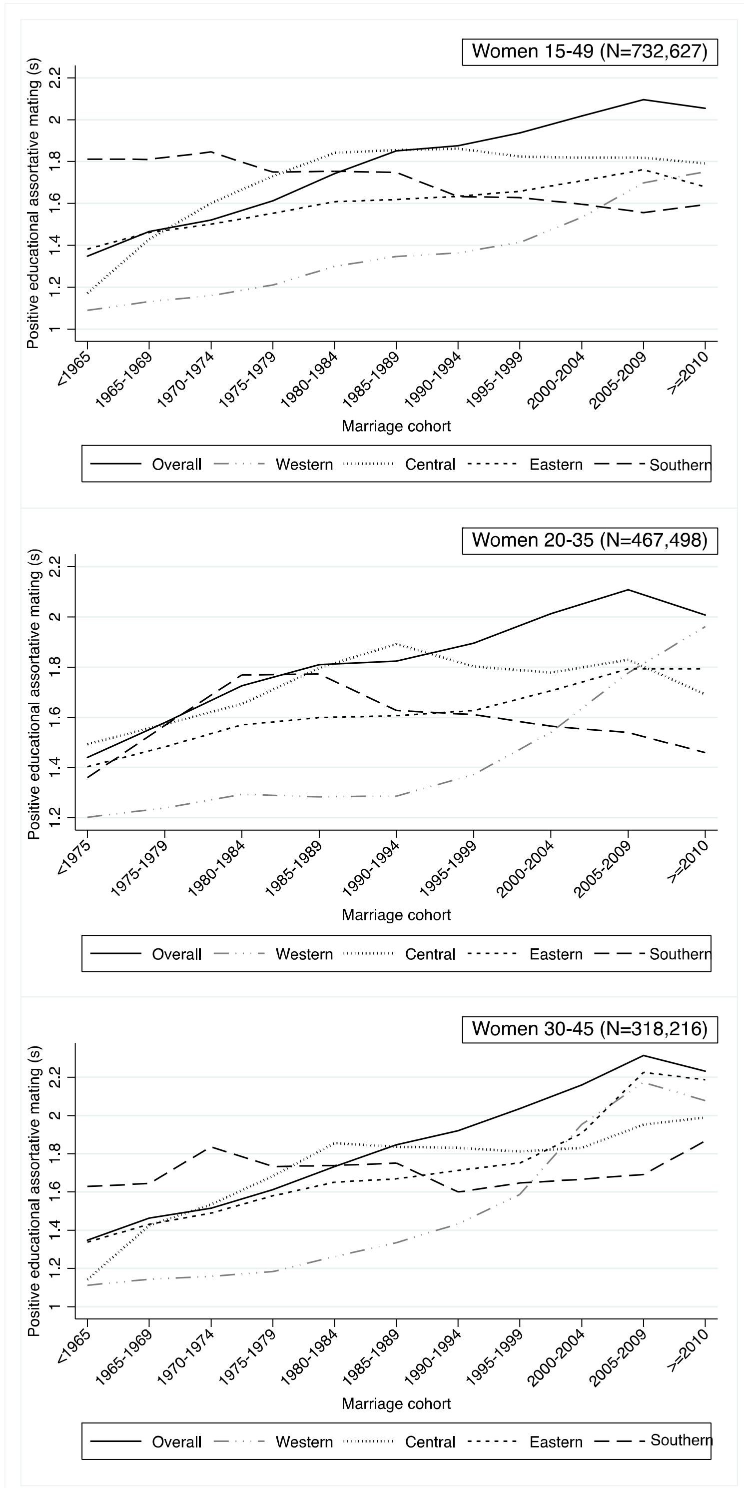
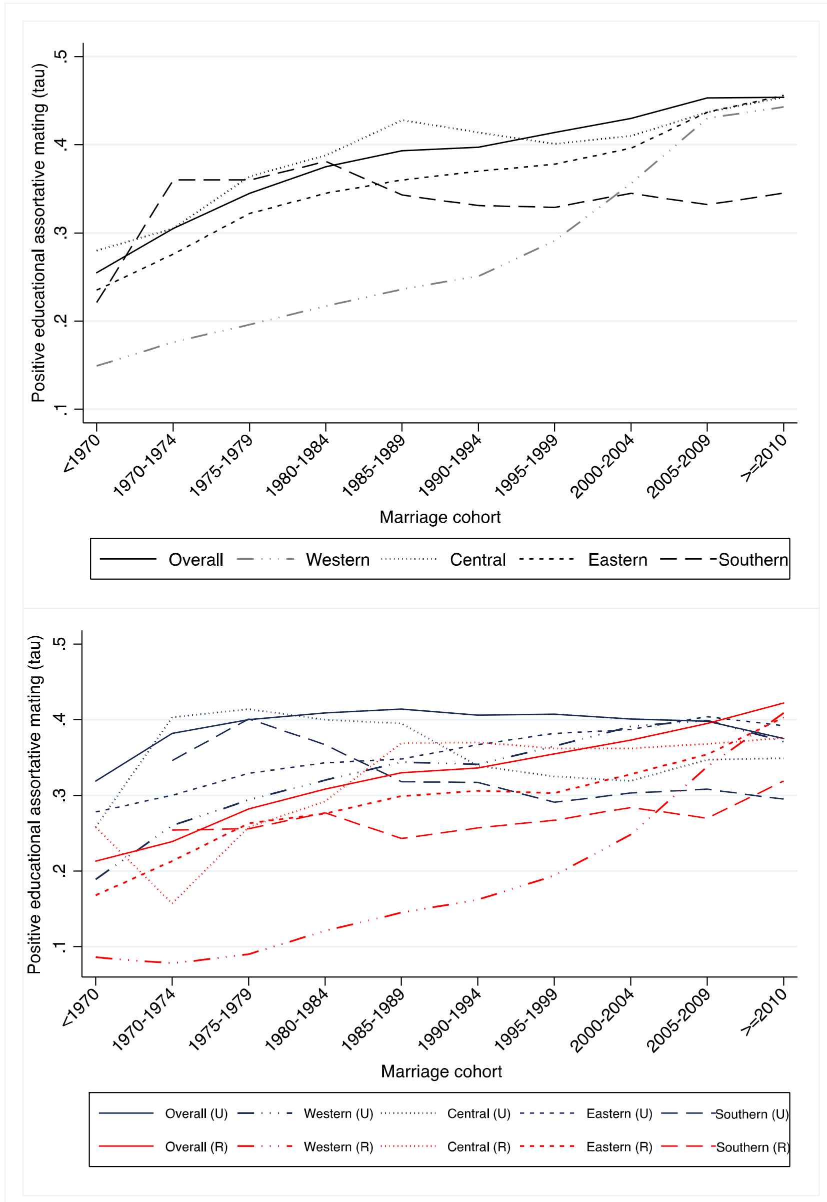


Figure A2.3: Positive educational assortative mating (τ parameter), by region of sub-Saharan Africa (top panel) and location of residence (bottom panel)



Notes: "U": urban; "R": rural.

Chapter 3. Beyond Attendance: Gendered Impacts of a Cash Transfer for Education and the Unpaid Care Burden in Rural Morocco

3.0 Abstract

This paper capitalizes on a randomized cash-transfer intervention implemented in rural Morocco between 2008 and 2010 to shed new light on the interplay between household inequality, as driven by gender and unpaid care work dynamics, and children's schooling. The study explores the effect of the cash transfer on school progression and unpaid care work, including analyses of how effects vary by gender and time spent on unpaid care work prior to intervention implementation. Results suggest that the intervention increased the likelihood of girls progressing through grades on time by approximately 6 to 10 percentage points, while it had no discernible effect for boys. Yet the benefit of the treatment on timely grade progression was reduced by a third to a half for girls engaged in unpaid care tasks, and the transfer proved ineffective in lessening the care burden. Taken together, findings suggest that as a result of the intervention girls performing unpaid care work were staying in school more but were less likely to progress on time relative to their counterparts not engaged in unpaid work. Insights from this research shed light on whether promoting gender equitable opportunities within the household might enable children to follow a more regular school path.

3.1 Introduction

Considerable academic research focuses on the socio-economic factors that predict school enrolment and attainment in both developed and developing countries. More often neglected in the scholarly debate is research delving into the factors that prevent children from progressing through grades in a timely fashion. The costs of age-grade distortions – an umbrella term that accounts for both delayed school entry and grade repetition – are very high, particularly for developing countries, where retention rates are high (Schiefelbein and Wolff 1992; Gomes-Neto and Hanushek 1994; Patrinos and Psacharopoulos 1996). Estimates for Brazil reveal that the costs of grade repetition alone represent an amount equivalent to the entire federal government contribution to first-level schooling (UNESCO 2012). Costs incurred by students in terms of lost opportunities and wasted human capital are even more significant (Manacorda 2012).

A factor that may significantly relate to the risk of not progressing through school is the amount of time children devote to unpaid care work within the household (Siddiqui and Iram 2007; El-Kogali and Krafft 2015). Household chores affect children's opportunities to learn and thrive by taking away valuable time they could spend on their education. The situation tends to be worse for girls, and somewhat exacerbated in rural settings characterized by high poverty rates, weak infrastructure, poor school quality, and large family sizes (Patrinos and Psacharopoulos 1997; Gupta 2015). For instance, data from Guatemala show that an increase in the number of siblings does not affect time devoted to domestic work for boys, while it brings about an additional four hours per week for girls (Dammert 2009b). Among other factors, large family size implies increased responsibilities for girls, more time spent on rearing children, cooking meals, washing clothes, caring for sick relatives, etc. The implications of this unequal care burden extend beyond resource-deprived households in low-income contexts. Women of all ages across all world regions suffer from the burden of unpaid care responsibilities, with particularly stark imbalances in the Middle East and North Africa (MENA) region – the geographical focus of this paper – where the female-to-male ratio of time devoted to unpaid care responsibilities approaches seven, as shown in Appendix Figure A3.1 (Ferrant, Pesando, and Nowacka 2014; World Bank 2015).

Using data from a randomized cash-transfer intervention ("Tayssir") implemented in rural Morocco between 2007 and 2010, this paper aims to shed new light on the interplay between household inequality, as driven by gender and unpaid care work dynamics, and children's schooling. Specifically, I assess the impact of the cash transfer on school progression outcomes – defined as progressing through grades in a timely fashion – allowing for heterogeneity of the treatment along socio-demographic lines such as the gender of the child and the amount of time spent by children on unpaid care work prior to intervention implementation. As unpaid care work emerges as a negative predictor of school progression, the analysis concludes with an examination of whether the cash transfer had any effect on lessening the care burden itself.

This work capitalizes on previous research from Benhassine et al. (2015),[54] who first evaluated Tayssir documenting positive and significant impacts of the program on school enrolment and attendance. My study builds on the premise that extending the focus to school progression outcomes is key for several reasons. First, enrolment and attendance do not necessarily translate into learning gains and progression through grades, which hinge upon demand-side, as well as supply-side factors such as school infrastructure, classroom structure, teachers' quality, and grade repetition policies.[55] Second, there is evidence that school progression is a key determinant of subsequent educational outcomes such as school completion (Jacob and Lefgren 2009; Glick and Sahn 2010). Therefore, in a

[54] Benhassine, N., F. Devoto, E. Duflo, P. Dupas, and V. Pouliquen. 2015. "Turning a Shove into a Nudge? A 'Labeled Cash Transfer' for Education." *American Economic Journal: Economic Policy* 7, no. 3: 86-125.

[55] While this idea has been acknowledged globally with the shift from the Millennium Development Goals (MDGs), targeting "education for all", to the Sustainable Development Goals (SDGs), stressing the value of "quality education", few scholars evaluating educational interventions embed this component in their analyses.

study that spans a two-year intervention period, school progression matters to the extent that it captures children who are in school but are exposed to the risk of not completing primary or secondary education at a subsequent point in time. This is relevant to the Moroccan context, where large percentages of youth enroll in school without completing primary education (Benhassine et al. 2015; Wagner 1994). Third, when studied in conjunction with intra-household dynamics such as the unequal allocation of unpaid care tasks, a specific focus on school progression may unravel interesting patterns. Specifically, a cash transfer of this kind – relatively small in size and not conditional on attendance[56] – may incentivize children to stay in school, although it might not be enough to help them progress through grades on time if they have competing time demands within the household. By the same token, competing time demands might not be so high as to prevent children from going to school whatsoever, while they can interfere with children's smooth progression by taking away valuable time they could devote to, for instance, out-of-school study time.

Girls in low-income contexts – such as rural settings in MENA countries – are often at higher risk of not completing primary education due to rooted traditions and moral and religious beliefs that perpetuate gender inequalities since young ages. These inequalities affect the role girls play within the household, the distribution of activities, and the amount of time spent on them, leading to unequal allocations of care tasks (World Bank 2015). In this work I bring a gender lens to the analysis of the effects of a cash transfer to investigate whether these household-level dynamics shape the effectiveness of the policy implemented. By assessing whether the impacts of the transfer on school progression differ by gender of the child and burden of unpaid care work prior to intervention implementation, I evaluate whether the disproportionate care burden affecting girls may – at least partly – explain differential educational outcomes among heterogeneous groups of children.

In general terms, evaluating whether the impacts of a cash-transfer intervention differ for defined subpopulations is crucial to identifying groups that might need ad hoc policy targeting (Dammert 2009a; Handa et al. 2010; Moffitt 2009; Vivalt 2015). In this context, a stronger effect of the treatment for children engaged in unpaid care responsibilities would point towards the idea that the intervention targeted those groups that were ex ante most exposed to the risk of following an irregular school path due to competing time demands. In other words, parents receiving the cash transfer would invest the money "wisely" within the household towards promoting a better future for the children left behind. Conversely, a weaker effect of the treatment for the aforementioned group

[56] The cash transfer was equal to approximately 5 per cent of the average household monthly consumption. In contrast, the range for traditional CCTs is between 6 per cent and 27 per cent of mean monthly household consumption – with PROGRESA being around 20 per cent (Fizbein et al. 2009).

would hint at the need to design careful policy interventions promoting more gender equitable opportunities within the household. At a deeper level, a finding of this kind would imply tackling rooted social norms that in resource-constrained contexts lead parents to make differential "investments" in sons versus daughters.

3.2 Background

The context

Morocco is a lower middle-income country with GDP per capita estimated at about 7,800 USD (PPP) in 2015. Education levels in the population are relatively low, with about 68.5 per cent of the adult population literate (Central Intelligence Agency 2015). In terms of schooling outcomes, recent UNICEF estimates (2014) suggest that 26 per cent of 5-year-olds who should be in pre-primary school are out of school, along with nearly 2 per cent of primary school-aged children, and over 16 per cent of lower secondary school-aged children.

Over the past two decades, Morocco instituted a series of successful reforms in the educational system aimed at achieving universal primary school enrolment. From 1999 to 2004, the enrolment rate increased from 79 per cent to 88 per cent at the national level, and from 58 per cent to 87 per cent in rural areas. Figure 3.1 provides a regional breakdown of Morocco, reporting the number and percentage of out-of-school children using the latest available Demographic and Health Survey (DHS, 2003-04).[57] Despite progress, the data show that the risk of being out of school at primary and lower secondary school ages was still significant as of 2004. In the 2007/2008 academic year, the year preceding the introduction of the pilot program examined here, the Ministry of Education estimated that over 90 per cent of rural children started primary school, but 40 per cent dropped out before completing the full six years of primary education (Benhassine et al. 2015), and the government became concerned again about improving enrolment and retention in school.

Micro-evidence from DHS data further reveals that the risk of being out of school was particularly high for girls, for children living in rural areas, and in regions where the poverty rate was above 30 per cent. Figure 3.2 suggests that a girl living in a poor rural community was five times more likely to be out-of-school at primary school age, and four times more likely to be out-of-school at lower secondary school age with respect to a boy living in a non-poor urban community, a gap which is likely

[57] These statistics include children of primary school age who are not in primary or secondary school, and children of lower secondary school age who are not in primary or secondary school. Also, note that at the time of the intervention Morocco was divided into 16 administrative regions, while since 2015 Morocco officially administers 12 regions.

indicative of entrenched social and cultural norms that still place rural girls at a significant disadvantage.[58]

The cash-transfer intervention

In order to address some of the concerns outlined above, in 2007 the Moroccan Higher Council of Education (CSE) launched a nationwide cash-transfer program together with the National Ministry of Education (MEN) to encourage parents to keep their children in school. With this goal in mind, the Government of Morocco partnered with a group of researchers affiliated with the Abdul Latif Jameel Poverty Action Lab (J-PAL) to evaluate "Tayssir", a two-year (2008-2010) pilot program targeted at the geographic level designed to increase student participation in primary school. Tayssir, which means "facilitation" in Arabic, made cash payments to parents with children aged 6 to 15 in targeted communities. Parents had to formally enroll each of their children into the program, and the enrolment process took place in schools. The pilot involved 318 rural primary school sectors,[59] each of which included two communities in the poorest areas (i.e., at least 30 per cent of the population living in poverty) within five of Morocco's 16 regions, namely Marrakech-Tensift-Al Haouz, Meknès-Tafilalet, l'Oriental, Souss-Massa-Draa and Tadla-Azilal (see Figure 3.1). All households with children aged 6-15 in targeted communities (i.e., municipalities) were eligible to receive the transfer, and the transfer was capped at three children per household. Ninety-seven per cent of households in the household sample had at least one child enrolled in Tayssir by the end of year two, hence program participation was nearly full.[60]

The Tayssir pilot included two versions of the program, a Labeled Cash Transfer (LCT) and a Conditional Cash Transfer (CCT). In the first version, families with children of primary school age were eligible to receive the transfer whether or not their children attended school. In practice, since enrolment in the Tayssir program happened at schools, children registering for Tayssir were enrolled in school at the same time, but the transfer was not conditional on continued enrolment. The transfer

[58] Note that it is not possible to draw estimates comparable to Figure 3.2 for school-progression outcomes such as the likelihood of being in the correct grade-for-age, as the latest DHS for Morocco does not report the grade the children were enrolled in at the time of the survey, nor their age of entrance in primary school. Besides country-level statistics supplied by UNICEF (2014), very little is known about school progression patterns at the micro-level, as no comparable dataset exists that would permit an analogous regional breakdown.

[59] Randomization was implemented at the school-sector level. A school sector includes a "main" primary school unit and several "satellite" school units (four on average). Satellite units fall under the authority of the headmaster of the main unit, and sometimes offer only lower grade classes. The pilot originally involved 320 school sectors but two (one in the control group and one in the treatment group) had to be dropped after the randomization because floods rendered them inaccessible.

[60] For more details on the experimental design, program take-up, variants of the intervention, and sampling procedure, please see Benhassine et al. (2015).

was given every two months during the 2008-2010 school years. The monthly amount per child increased as each child progressed through school, starting from 60 MAD (8 USD) for each child in grades 1 and 2 and increasing to 100 MAD (13 USD) for children in grades 5 and 6. The average transfer represented about 5 per cent of the average household's monthly consumption. In the second version of the program, the cash transfer was disbursed to parents of primary school-age children, as long as their child did not miss school for more than four times each month. The monthly transfer amount was the same as for the LCT program.

Mothers received the transfer in half of the school sectors sampled for Tayssir, while fathers received it in the other half. Hence, each school sector sampled for the study was randomly assigned to one of the following five arms: LCT to fathers (80 communities from 40 school sectors), LCT to mothers (80 communities from 40 school sectors), CCT to fathers (180 communities from 90 school sectors), CCT to mothers (178 communities from 89 school sectors), and a comparison group receiving no transfers (118 communities from 59 school sectors). Figure A3.2 in the Appendix summarizes the experimental design.

Previous evaluation of Tayssir

Benhassine et al. (2015) documented positive and significant impacts of the cash transfer on school enrolment and attendance, together with positive yet weakly significant effects on arithmetic test scores. Concerned with variation in program impact by gender of the transfer recipient (mother versus father) and type of transfer issued (LCT versus CCT), Benhassine et al. found no difference in impacts between transfers issued to fathers and transfers issued to mothers. Making cash transfers conditional did not increase the effectiveness of the program either. Likely, directing the cash transfer to mothers or adding conditionality did not substantially alter the program's impacts because Tayssir was framed as an educational support program and perceived as an endorsement of the local schools, since headmasters were responsible for enrolling families. Overall, Benhassine et al. documented that Tayssir in all forms increased parents' belief that education was a worthwhile investment.

Despite the richness of the analysis carried out in Benhassine et al., no reference is made to whether the intervention translated into smoother progression through grades for students who remained in school between baseline and endline, which I address in the present analysis. The treatment arms in the current study are reduced from five to two, i.e. any cash transfer given to parents – the treatment group (259 school sectors) – versus no cash transfer (59 school sectors) – the control group (shaded boxes in Figure A3.2 in the Appendix). The motivation behind this choice is twofold. First, the focus and policy relevance of Benhassine et al.'s study lies in the differential effectiveness of the

LCT versus the CCT, and the related implications for budgeting and government spending. Conversely, my study aims to provide a better understanding of how a cash transfer – irrespective of type – shapes parents' differential investments in boys versus girls. Second, and most importantly, Benhassine et al. reported that in the first year of the program about 50 per cent of parents did not understand the difference between the conditional versus labeled nature of the transfer, thereby evidencing large misunderstanding of program functioning and making this distinction less meaningful for my purposes.[61] As for differences in targeting mothers versus fathers, Benhassine et al. ascribed their null findings to the fact that women's independent role in financial transactions in poor households in rural Morocco is very limited. I test differential impacts of targeting mothers versus fathers and find results analogous to Benhassine et al. In the main body of the text I hence leave these distinctions aside (available upon request), thus boosting sample size.

Grade repetition in Morocco

Albeit limited, existing research suggests that in Morocco repetition rates – defined as the proportion of pupils from a cohort enrolled in a given grade in a given school-year who study in the same grade in the following year – are high, especially in primary school (Mansouri and El Amine Moumine 2017). Within primary school, recent data from the Education Policy Data Center (2014) show that students in Morocco are more likely to repeat grade 1. The repetition rate in grade 1 is 10.9 per cent for both males and females, which is 1.7 points higher than the average repetition rate across other primary grades (of 9.2 per cent).

Although there remains considerable between-school variation in terms of rules and standards behind grade repetition policies, repetition in Moroccan primary schools is guided by a national policy that combines normative rank ordering on academic performance with consideration of the capacity of higher grade level classrooms to absorb children promoted to the next grade (Wagner 1994). Through the first four years of primary school, grade promotion or repetition is determined on the basis of an annual meeting of classroom teachers. During this year-end meeting, children are ranked by classroom grade point average (GPA) – taking into account attendance and absenteeism – and then passed or failed according to a system whereby the lowest-ranked 10 per cent to 20 per cent of children in each grade level are required to repeat the entire year's academic program (Wagner 1994). Promotion rates are set yearly on the basis of projected space availability in the higher-grade levels of each school district, as well as on certain political exigencies. Only two grade repetitions are allowed by Moroccan

[61] I conducted separate analyses by type of transfer (LCT versus CCT) and found no differential impacts on the outcomes I investigate. These analyses are hence omitted in the text (available upon request).

law during the primary school years, and children exceeding this number are candidates for school expulsion or dropout. It follows that the high rates of repetition are a serious harbinger of school dropout (Agnaou 2004).

3.3 Literature review

School participation and school progression

While many studies explore the impact of cash transfer programs on school attendance and school performance (Gomes-Neto and Hanushek 1994; Cardoso and Souza 2009; Ponce and Bedi 2010; Kumara and Pfau 2011; Dubois, de Janvry, and Sadoulet 2012; Amarante, Ferrando, and Vigorito 2013; Benhassine et al. 2015; Reynolds 2015; etc.), there is surprisingly little evidence on the impact of cash-transfer interventions on measures of school progression. Some exceptions are Behrman, Parker, and Todd (2009) and Maluccio, Murphy, and Regalia (2010), who document positive and significant effects of conditional cash transfers on grade progression in Mexico and Nicaragua, respectively. Both papers aim to fill a gap in the impact evaluation literature by stressing the importance of focusing on school progression, in addition to enrolment and attendance. The latter further note that half of the estimated program effect on progression was accounted for by a reduction in the dropout and repetition rates of beneficiary children who were already in school when the program began.

As previous research has documented that school attendance and progression may be quite heterogeneous in terms of both their socio-economic determinants and their consequences (Patrinos and Psacharopoulos 1996; Pal 2004; Glick and Sahn 2010), it is puzzling that so little research related to cash-transfer interventions has focused on the latter. There are several reasons why school progression tends to be more often neglected in scholarly discourse. First, the operationalization of the measure requires more data inputs such as the age of entry in school and builds – almost by construction – on longitudinal designs or repeated cross-sections. Second, as mentioned above the determinants of school progression include more binding supply-side factors like school policies and regulations on grade repetition (Maluccio, Murphy, and Regalia 2010). Third, as academic performance is a key determinant of school progression, and most cash-transfer interventions aim at boosting attendance rather than targeting learning itself, there is a tendency to believe that impacts on school progression generally echo impacts on school performance.

Unpaid care work

Unpaid care work is defined by the United Nations (UN) as "a critical - yet largely unseen - dimension of human well-being that provides essential domestic services within households, for other households, and to community members" (UNDP 2009).[62] The literature on unpaid care work has expanded rapidly over the past decade, following two main directions. First is a more macro-level approach concerned with pooling country-level time-use data to assess gender imbalances in unpaid care work among the adult population (Razavi and Staab 2008; Budlender 2008), and the implications of these inequalities for labor market outcomes (Ferrant, Pesando, and Nowacka 2014). Ferrant, Pesando, and Nowacka document that across all regions of the world women spend, on average, three to six hours on unpaid care activities, while men spend between half an hour and two hours. North America (NA) and the Middle East and North Africa (MENA) stand out, respectively, as the most equal and the most unequal regions, with a female-to-male ratio of time spent on unpaid care work approaching two in the former case and seven in the latter (Appendix Figure A3.1). Often embedded in this line of research is a broader discussion of the importance of valuing unpaid care work to make its contribution visible and accounted for in GDP calculations (Budlender and Brathaug 2004; Folbre 2006, 2014), and the need to develop appropriate indicators to quantify the "feminization of poverty" (UNIFEM 2000).

The second strand of the literature focuses on micro-level analyses of the relationship between unpaid care work and school outcomes among youth in disadvantaged, mostly rural contexts. Authors in countries as diverse as Bolivia (Zapata, Contreras, and Kruger 2011), Egypt (Assaad, Levison, and Zibani 2010), rural Ethiopia (Admassie 2003), Guatemala and Nicaragua (Dammert 2009b), Peru (Levison and Moe 1998), rural Vietnam (Le and Homel 2015) etc. find that domestic work is associated with lower rates of school participation and attendance, and poorer academic performance. Among the main takeaways from these studies is the need to separate paid market work from unpaid care work due to their different drivers and later-life consequences. In line with this idea, using data from 186,795 families with 7 to 14-year-old children across 30 low middle-income countries, Putnick and Bornstein (2015) claim that the negative relations observed between child labor and school enrolment are much more consistent for family work and household chores as compared to paid work outside the home.

My work enriches the literature on unpaid care work in two directions. First, as there is still a paucity of micro-level research documenting time spent on unpaid care work by young children in low-income contexts,[63] this study provides a clear picture of gender imbalances in care-related activities

[62] The adjective 'unpaid' signals that the person carrying out the activity does not receive a wage, hence the work is not counted in official GDP calculations as it falls outside the production boundary in the System of National Accounts (SNA). 'Care' suggests that the activity serves people and their wellbeing, and includes both personal care and care-related activities, such as cooking, cleaning and washing clothes. 'Work' means that the activity entails expenditures of time and energy. Unpaid care work is also referred to as 'reproductive' or 'domestic' work in order to distinguish it from market-based work.

[63] Lloyd, Grant, and Ritchie (2008) and Vu (2014) are notable exceptions.

using time-use data from rural Morocco. Second, as most studies relating child domestic work and schooling outcomes focus on school participation or school performance, I here shift the focus onto how unpaid care work relates to grade progression through school.[64] In this respect, it is worth mentioning that the specific focus on unpaid care work in this context vis-à-vis paid work is tied to the far higher prevalence of the former – less than 4 per cent of out-of-school children and 2 per cent of all children in the sample report any work outside the home (Benhassine et al. 2015).

3.4 Data and methodology

Data and variables

This paper uses publicly available data.[65] Four types of data were collected as part of the randomized control trial. School participation was measured through school visits spread over the two years of the program, for all students enrolled in the study schools at the beginning of "year 0" (the academic year 2007/2008). This is labeled the "school sample", and it comprises over 47,000 students. Second, a comprehensive survey was carried out at both baseline and endline with close to 4,400 households, defined as the "household sample." Third, one child per household completed a basic arithmetic test (ASER test) during the endline household survey. Finally, the authors conducted "awareness" surveys at and around schools to measure teachers and households' understanding of the program. This study builds primarily on the baseline and endline surveys, as these are the sole sources that permit to retrieve a measure of unpaid care work from the time-use diaries.

Households were sampled as follows. For each school unit, eight households were randomly selected for a baseline survey administered in June 2008, before Tayssir was announced and before school sectors had been randomly assigned to either treatment or control. The endline survey was administered in June 2010, after exactly two years. To select the households, enumerators visited each school in spring 2008, and used the 2007/2008 school register, as well as the registers from the previous three academic years, to draw two lists: (i) the list of all households in the school's vicinity that had at least one child enrolled in school, and (ii) the list of households with no child currently enrolled in school but at least one child of school age who had enrolled at some point but dropped out within the previous three years. A total of six households were randomly selected from list 1, and two from list 2. As a consequence of this sampling strategy, the sampling frame does not include households who never

[64] Buonomo Zabaleta (2011) focuses on the relationship between child labor and school progression (as measured by grade-for-age) in Nicaragua. Yet her focus is mostly on non-housework labor.
[65] Data are available on the World Bank platform or at: https://www.aeaweb.org/articles.php?doi=10.1257/pol.20130225

enrolled any school-age children in school, though such households appear to be very rare (Benhassine et al. 2015). An average of two children per household were enrolled in the Tayssir program.

This study focuses on children aged 6-15 as the main units of analysis, and school progression as the main outcome of interest. Scholars tend to operationalize school progression through variables such as schooling-for-age (SAGE) or grade-for-age (Cascio 2005; Patrinos and Psacharopoulos 1997), which measure whether a child is in the right grade given her/his age.[66] These are powerful indicators, yet they are not suitable to the case at hand as they capture a degree of pre-treatment heterogeneity that is beyond the scope of these analyses. As I am interested in the actual grade progression over the two-year intervention period, I measure school progression as the difference between the grade the child was enrolled in at endline (2010) and the grade the child was enrolled in at baseline (2008). This way, children progressing two (or more) grades over two years were considered "on time", – regardless of whether they were in the correct grade for their age to start with – whereas children progressing zero or one grade over two years were not progressing in a timely fashion.[67] Children enrolled in school in 2008 may have dropped out between baseline and endline, hence raising challenges on the operationalization of school progression, a point I will return to in the following sections.

I measure unpaid care work at baseline and endline as the amount of time per day children spent on several unpaid care activities. The household surveys include a time-use diary recording the primary and secondary activities each child performed the day preceding the survey during fixed time intervals of 30 minutes which, summed up, account for a full 24-hour day. In line with the UN definition of unpaid care work, I consider as unpaid care work the following 10 activities: "preparing food for a meal" (FM), "preparing food for another occasion" (FO), "doing housework" (HW), "washing clothes" (WC), "other domestic activities" (OD), "shopping for the house" (SH), "going to get water" (GW), "occupied with children in the household" (CHH), "occupied with older members in the household" (EHH), and "occupied with other sick or handicapped members in the household" (SHH). Following the relevant literature (United Nations 2005), I attribute 20 minutes to the primary activity and 10 minutes to the secondary activity in order to allow for simultaneity of tasks, i.e., to account for the possibility that the 30-minute slot may not be entirely devoted to the activity listed by

[66] For instance, in a country like Morocco where the school entry age is 6, a 8-year-old in first grade is not in the correct grade for age.
[67] An analogous way of framing this concept of grade progression, perhaps more common in the literature, is in terms of grade repetition. Children progressing two grades over two years experienced no grade repetition within the two-year timeframe. Conversely, assuming there was no temporary school exit, children progressing one and zero grades over two years repeated one and two grades, respectively.

the caregiver as the primary one.[68] After adding up all of the 30-minute time slots in which a child performs one or more of the above-mentioned activities, I obtain an estimate of the amount of time per day (minutes or hours) devoted to unpaid care work.[69]

Methodology

The effect of the treatment on school progression and the subgroup analyses by baseline unpaid care work are analyzed using a series of linear probability models (LPM). As the treatment was implemented at the school-sector level, while the outcome and the main heterogeneity variable were measured at the individual level, the analysis is carried out by means of the following specification:

$$Y_{i,j} = \alpha + \beta CT_{parent_j} + X'_{i,j}\gamma + \left(UCW_{i,j} * CT_{parent_j}\right)\delta + \varepsilon_{i,j} \quad (1)$$

where $Y_{i,j}$ is the school outcome of interest for student i in school j; CT_{parent_j} (the treatment) is a dummy equal to 1 if school j falls under any of the cash-transfer typologies; $X_{i,j}$ is a vector of strata randomization dummies, child-level (including $UCW_{i,j}$), household-level, and school-level controls.[70] These include variables such as age at baseline, number of siblings, unpaid care work, mother's education, father's education, schooling status at baseline, ownership of a cellphone, access to electricity, and remoteness.[71] Due to incomplete data on grade attained, parents' education is measured through dummy variables that equal one if the parent has at least some primary education. Household composition variables are constructed using information from the household roster. As children aged 6-15 are the main units of analysis and young children are unlikely responsible for the care of far older siblings, I here restrict the focus to siblings aged 15 or less. $UCW_{i,j} * CT_{parent_j}$ is an interaction term that permits to investigate variation in program impact by means of baseline time-use data on unpaid care work (kept as a continuous variable).[72] Note that I also construct a categorical counterpart of the unpaid care work variable classified as "No UCW," "Medium UCW" (0-2 hours/day), and "High

[68] Note that the alternative approach would be to attribute the whole 30-minute slot to the activity listed as *primary*. Measures constructed using these two different approaches yield a correlation of about 0.98.

[69] Unless further specified, unpaid care work is measured in minutes per day in the descriptive statistics, and hours per day in the regression analyses.

[70] Randomization was stratified by region, school size, dropout rate, and by whether the government was planning to make improvements to school infrastructure within the two-year time frame of the evaluation (Benhassine et al. 2015).

[71] The latter household-level and school-level variables are included as they were found to be unbalanced at baseline in the original study.

[72] One significant advantage of the regression framework is the ability to estimate continuous heterogeneous treatment effects. This way, the interaction term is analogous to the idea of the marginal treatment effect (Heckman 2001; Heckman and Vytlacil 2005). The interpretation of the continuous conditional average treatment effect is the same as for any other continuous-binary interaction term, although care should be taken not to extrapolate beyond the support of the covariate.

UCW" (>2 hours/day). Some analyses using this variable are reported in the Appendix (additional analyses available upon request).

To test the effectiveness of the intervention in reducing unpaid care work within the household I adopt the following Difference-in-Difference (DD) specification:

$$UCW_{i,j,t} = \alpha + \eta CT_{parent_j} + \vartheta Post_t + \left(Post_t * CT_{parent_j}\right)\varphi + X'_{i,j,t}\zeta + \varepsilon_{i,j,t} \quad (2)$$

where $Post_t$ is a dummy for time that equals 0 in 2008 (baseline) and 1 in 2010 (endline), and φ is the coefficient of interest capturing the effect of the cash transfer on time spent on unpaid care work. Although in the context of a randomized experiment – where variables ought to be balanced at baseline – a simple LPM controlling for baseline UCW is also a viable strategy, a DD estimation strategy permits to difference out both time-invariant confounders between treatment and control units and time-trends.

For expositional clarity I stratify the sample by gender of the child. Pooled-sample analyses including a dummy for gender and a treatment/gender interaction are included in the Appendix. Strata dummies take account of stratification variables used in the randomization. Standard errors are adjusted for clustering at the school-sector level. Lastly, as the sampling procedure oversamples households with dropout children (the final household sample includes 17 per cent of households with dropout children, while those households represent only 9 per cent of the population), sampling weights are used in all statistics and analyses. This way regressions are representative of the population from which the researchers surveyed.

Sample attrition bias

From the baseline survey I was able to identify 10,889 children aged 6-15 living in the household with their parents. Among these, 951 (8.73 per cent of the baseline sample) dropped out of the study between 2008 and 2010, delivering a post-attrition sample of 9,938 children. In order to assess differences between children that dropped out and the 9,938 children who were present at both baseline and endline, Table A3.1 in the Appendix examines differential attrition by treatment status and selected socio-demographic characteristics. While attritors were not significantly different on age, number of siblings, amount of time spent on unpaid care work, and baseline schooling status, attrition was higher among control units and among boys. There is also evidence that attrition was higher in smaller households and in households where the father was slightly better educated. As control units did not receive the transfer, they were indeed the ones with the least incentives to participate in the study. Therefore, this finding raises potential concerns to keep in mind while interpreting results.[73] To check whether differential attrition yields imbalances in child and household-level characteristics, in what follows I present summary statistics by treatment group on the post-attrition sample.

3.5 Descriptive statistics

Table 3.1 provides balance checks and baseline summary statistics on the analytical sample of 9,938 children. Column 1 reports the mean of the variable in the control group, with standard deviations in brackets, while column 2 reports coefficients and standard errors (in parentheses) from an OLS regression of the left-hand side variable on a dummy for treatment, controlling for strata randomization dummies. The table shows that the groups are relatively well balanced with respect to observable characteristics, hence increasing confidence in the effectiveness of the randomization. There are significant differences in baseline schooling status, however, as the share of treated children enrolled in school at baseline was higher by roughly 4.3 percentage points.[74] Similarly, the share of treated children never enrolled in school at baseline was lower by about 2.4 percentage points. In all analyses below, I condition on baseline schooling status in order to ensure these differences do not drive my results. In terms of sample composition, Table 3.1 suggests that out of the 9,938 children 46.6 per cent were girls, and children at baseline were on average 10.4 years old, had 1.7 siblings aged 15 or less, and spent around 47 minutes per day on unpaid care work activities. Also, 16 per cent of children had fathers with at least some primary education, while this same figure for mothers was around 4.9 per cent only. As for children's schooling status, their average entrance age in primary school was 6.5, and 76 per cent of children were enrolled in school as of June 2008. Children enrolled in school were on average in their third grade of primary school.

Table 3.2 provides descriptive statistics on the amount of time per day boys and girls spent on each of the 10 unpaid care activities selected from the time-use diary. The table shows pronounced differences between boys and girls, suggesting that in the poorest regions of rural Morocco girls disproportionately suffer from the burden of unpaid care work within the household. Overall, mean gender differences in the total amount of time spent on unpaid care work (UCW) are stark, with boys devoting approximately 18 minutes per day as compared to girls who spent one additional hour. This is particularly true for activities such as preparing food for a meal (FM), doing housework (HW),

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[73] In the presence of post-treatment data on the 951 children who attrited, attrition concerns could be reduced by means of imputation techniques. Absent these data, any imputation would be based on either baseline characteristics or endline characteristics of students who did not attrite. This is not a valid strategy as it relies on the assumption that attritors would respond to the intervention the same way as non-attritors.

[74] This finding aligns with what found by Benhassine et al. (2015) in the main study.

washing clothes (WC), and other related domestic tasks. The only activity to which boys devoted marginally more time was shopping for the house, a finding which is in line with restrictions to the freedom of movement that women face in several MENA countries. Similarities also emerge when dealing with sick or handicapped members within the household, though for these activities averages are very low in absolute values. With reference to Table 3.1, note that the unpaid care work variable is balanced at baseline, hence making it possible to use it as a moderator of the treatment. Appendix Table A3.2 further shows that all components of the UCW variable do not statistically differ by treatment status.

It is important to stress, however, that many children – approximately 77 per cent of boys and 48 per cent of girls – did not engage in any form of unpaid care work, hence the averages are driven down by these "artificial" zeroes. In Table A3.3 in the Appendix I run t-tests for gender differences in means for the sub-sample of children aged 6-15 who reported spending a non-zero amount of time on each activity. These tests show that averages are statistically higher for girls across the majority of the activities, and working girls spend twice as much time on unpaid care work per day relative to working boys (149 against 78 minutes, respectively). Figure 3.3 provides a clearer picture of boys' and girls' participation in each unpaid care activity and conveys the idea that not only girls spent on average more time on UCW as compared to boys, but also the percentage of girls engaged in each task was far higher, except for shopping for the house (SH). For instance, the female-to-male participation ratio reaches up to 9 and 7.5 for housework and washing clothes, respectively.

Figure 3.4 shows the total minutes per day spent on UCW by number of siblings (top panel) and by age of the child (bottom panel). Unpaid care work increases with number of siblings for both boys and girls, yet the gradient is clearer for girls. While girls with no siblings devoted about an hour to unpaid care tasks, girls with three or more siblings spent an additional 23 minutes.[75] Patterns of unpaid care work by age are strikingly different by gender. While the age profile of UCW for girls is steep and upward sloping – reaching almost three hours per day at age 15 – boys spend an average of 15 to 20 minutes irrespective of age. These descriptive relationships provide prima facie evidence that gender and unpaid care work are closely intertwined dimensions, and hint at the existence of rooted gender inequalities within the household that place rural girls at a disadvantage.

[75] Figure A3.3 in the Appendix is analogous to Figure 3.4 (top panel) in the paper, yet the number of siblings is replaced by the overall number of children (ages 0-15) present in the household. Gradients of UCW in this latter scenario are less clearly identifiable.

3.6 Results

School outcomes are operationalized along two axes of analysis. First is the extensive margin of school participation, which I define as being in school versus being out of school. To capture the extensive margin of school participation, I construct a dummy for dropout since 2008 which equals 1 if the child was enrolled in school at baseline but dropped out between 2008 and 2010 (i.e., the child was not in school anymore as of the endline survey, june 2010), and 0 otherwise. Second is the intensive margin of school participation, which I define as progressing through grades on time versus stalling, conditional on remaining in school. To capture the intensive margin of school participation, I construct a timely grade progression dummy, which equals 1 if the child progressed two (or more) grades over the two-year intervention period, and 0 if the child progressed 0 or 1 grades over two years.

Extensive margin: dropout

Table 3.3 (panel a) provides LPM estimates of the effect of the cash transfer on the dropout dummy variable. First, as captured by the constant term, the share of girls that dropped out between baseline and endline was substantially higher than that of boys (0.17 against 0.11, respectively). In line with findings from the original study, the cash transfer had large and significant impacts on school dropout, decreasing it by approximately 4.3 and 7.3 percentage points for boys and girls, respectively (model 1).[76] Adding controls does not alter the magnitude and significance of the estimates to a big extent (model 2). Note that in relative terms results are substantially similar by gender, as a drop of 4.7 percentage points off of a base rate of about 11.4 per cent in the male control group entailed a reduction in boys' dropout rate of roughly 41 per cent, while a drop of 7.3 percentage points off of a base rate of about 17.1 per cent in the female control group entailed a reduction in girls' dropout rate of about 43 per cent.[77]

Neither sibship size nor unpaid care work emerge as strong predictors of girls' likelihood of dropping out, while unpaid care work more strongly predicts boys' likelihood of dropping out. Specifically, one additional hour of unpaid care work is associated with an increase in boys' probability of dropping out by 1.8 percentage points. Given the stark gender differences in unpaid care work documented above, this finding is somewhat surprising, yet consistent with the idea that UCW might

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[76] This is the only result that replicates findings from the original study. As the post-attrition sample and the operationalization of most variables (foremost, unpaid care work) partly differ from Benhassine et al. (2015), the effect size is not identical, yet qualitatively the same. Importantly, Benhassine et al. did not present separate analyses by gender, nor did they account for treatment-baseline UCW interactions.

[77] Also, pooled estimates with a dummy for gender and a gender/treatment interaction show that the effect of the treatment on dropout was not statistically different for boys and girls (Panel a, Table A3.4 in the Appendix).

be so ingrained in girls' daily lives that it does not operate as a determining factor for whether they are able to go to school or not. Conversely, boys' likelihood of staying in school versus dropping out might be more sensitive to smaller increases in UCW. Lastly, estimates in panel a reveal no differential treatment effect by baseline unpaid care work, suggesting that the effect of the intervention on dropout operated similarly for all children.

Intensive margin: timely grade progression

Table 3.3 (panel b) provides LPM estimates of the effect of the cash transfer on the timely grade progression dummy variable. Note that in this specification children enrolled in school in 2008 but dropping out before 2010 (i.e., 707 children) are excluded – the problems that arise adopting this approach are discussed in the next section. While the cash transfer significantly reduced dropout for both boys and girls, these estimates provide initial evidence that the effects on timely grade progression operated along gender lines. While the intervention increased the likelihood of girls progressing through grades by approximately 6 percentage points – equivalent to a 10 per cent increase relative to the mean in the control group – it had no discernible effect for boys.

Differently from panel a, both sibship size and unpaid care work emerge as negative predictors of girls' timely grade progression. Specifically, one additional sibling and one additional hour of unpaid care work are associated with a decrease in girls' likelihood of timely grade progression by 2.2 and 3.4 percentage points, respectively. Moreover, the significant coefficient on the interaction between unpaid care work and the treatment points at subgroup differences among girls. The negative coefficient suggests that the benefit of the treatment on timely grade progression was halved for girls engaged in each additional hour of unpaid care work prior to intervention implementation. Figure 3.5 plots the predicted proportion of girls progressing through grades on time based on estimates from panel b, yet replacing the continuous UCW variable with its categorical counterpart (No, medium, high) – detailed regression estimates for boys and girls using UCW as a categorical variable are reported in Appendix Table A3.5.[78] This figure shows that the cash transfer was effective in increasing girls' timely grade progression in the "No UCW" group only, and the benefit of the treatment was cut proportionally more, the more was the time girls spent on UCW at baseline. For girls spending more than two hours per day on UCW, the benefit of the treatment was cut by 100 per cent (i.e., it was reduced to zero).

In light of the non-significant coefficient on the interaction term between UCW and the treatment dummy in the dropout specification (panel a), estimates from Table 3.3 as a whole suggest

[78] The distributions of children that fall in each UCW category is, approximately, 77 (No), 18 (Medium), and 5 (High) per cent for boys, and 48 (No), 23 (Medium), and 29 (High) per cent for girls.

that as a result of the treatment, girls engaged in UCW were staying in school more – rather than dropping out – but were not progressing through grades at a pace comparable to girls who were not engaged in analogous tasks (or at least, not as much).

Sample selectivity and robustness checks

Despite suggestive, these findings raise some methodological concerns. First, as grade progression is only observed for kids who remained in school throughout the observation period, OLS estimates on the restricted sample (i.e., the sample excluding the 707 dropouts) may be biased and inconsistent. The reason is that kids who remained in school (6,981) were likely a non-random sample of all children enrolled in school at baseline (7,688). Second, as the cash transfer is found to affect dropout, excluding dropouts "breaks" the RCT, as it entails differentially excluding treatment versus control units from the sample. To deal with these issues, I follow three different strategies.

First, I bound the estimates assigning extreme values of grade progression to kids who dropped out. Specifically, I assume that, had they remained in school, kids who dropped out would have either progressed on time (dummy=1), or at a sub-optimal pace (dummy=0). Although both scenarios are unlikely – the second one perhaps more realistic – this exercise is useful to the extent that it provides a lower-bound and an upper-bound to the effect of the treatment. Table 3.4 shows that keeping dropouts as progressing on time (panel a) weakens the effect of the treatment for both boys and girls, yet some evidence of a differential treatment effect by baseline UCW persists for girls. Conversely, keeping dropouts as not progressing on time (panel b) strengthens the effect of the treatment for both boys and girls (reasonable, as it conflates the estimates with the findings on dropout documented in Table 3.3, panel a), although a differential treatment effect by baseline UCW status is observed for girls only. This latter coefficient suggests that the beneficial effect of the treatment for girls engaged in each additional hour of UCW was cut by about a third.

Second, I conduct more formal tests of attrition – defined in this case as leaving school between 2008 and 2010 (rather than leaving the study sample altogether) – to assess whether attrition is random. I perform a Becketti, Gould, Lillard, and Welch (BGLW) test, whose goal is to determine whether those who subsequently leave the sample differ in their initial "behavior" relationships between predetermined (observed) variables and baseline outcomes (Alderman et al. 2001; Becketti et al. 1988).[79] As the test strongly rejects the null hypothesis that attrition is random ($p>F=0.000$;

[79] The test needs to be repeated for each baseline outcome of interest and works as follows: an outcome at baseline (y(0)) is regressed on a set of predetermined variables (x(0)), an attrition dummy (A) for whether the child subsequently leaves the sample, and the attrition dummy interacted with the explanatory variables (A*x(0)). An F-test of the joint significance of coefficients on A and (A*x(0)) suggests whether there is evidence to reject

Appendix Table A3.6), I reweight the sample using Inverse Probability Weighting (IPW) and re-estimate the restricted model on timely grade progression (Table 3.5, panel a). The underlying idea with IPW is to give more weight to children whose observed characteristics are more similar to those of children who attrited.[80] Estimates using IPW somewhat weaken the effect of the treatment for boys, while they strengthen it for girls. According to the IPW specification for girls, the cash transfer increased timely grade progression by about 10 percentage points.

Third, as the above methodology deals with sample selectivity based on observable characteristics only, I implement a 2-step Heckman correction approach (Table 3.5, panel b) which relies on identifying a variable – or a set of variables – that affect the selection equation but are, at least theoretically, orthogonal to the second-stage error term. Following the literature (Carneiro, Heckman, and Vytlacil 2011; Schwartz, Stiefel, and Wiswall 2013), I hypothesize that the time and distance (in km) from the household to the school potentially affect whether children go and attend school, but not whether they progress through grades on time (or, at least, not through channels other than going to school).[81] These two variables have strong predictive power in the selection equation, while the Inverse Mills Ratio (IMR) added to the second-stage does not turn out as statistically significant, thus suggesting that there is little evidence of selection on unobservables. This set of estimates reveals that the cash transfer increased timely grade progression for girls by 7.7 percentage points, and that the benefit of the treatment was cut in half for girls engaged in UCW. Also, these estimates provide neater evidence that the effect of the cash transfer was statistically different between boys and girls.

Unpaid care work

Table 3.6 shows results from a DD specification on unpaid care work.[82] Results point to two main findings. First, time spent on unpaid care work increased overtime for girls by approximately half an hour, while it decreased for boys by less than 3 minutes, even after accounting for age and the same set of controls included in the above set of models. In other words, there is evidence that as children grow older, gender inequalities in unpaid care work widen. Second, time spent on unpaid care work

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the null hypothesis that attrition is random. For more details and applications of the test to longitudinal data from low- and middle-income countries, see Alderman et al. (2001).

[80] For boys, variables that most strongly predict the likelihood of dropping out of school between 2008 and 2010 – hence entering the estimation of weights – are age (+), treatment status (-), UCW (+), having electricity in the household (-), and having access to a water network connection (-). For girls, these factors boil down to age (+) and treatment status (-) only – see Appendix Table A3.7.

[81] As always, the validity of the exclusion restriction can be criticized along several directions. Note that in this study this strategy is only adopted as a further robustness check, not as a primary finding.

[82] Appendix Table A3.8 reports analogous results from a LPM on endline UCW controlling for baseline UCW (in a spirit similar to the other set of estimates presented in this study).

did not change as a result of the intervention, as witnessed by the non-significant coefficient on the time-treatment (Post x CT) interaction term.

3.7 Additional analyses

Mechanisms and explanations

Summarizing all this evidence, how can we make sense of these findings? First, given that test scores are a – perhaps the most – proximate determinant of grade progression, we would expect the cash transfer to have had a stronger impact on girls' test scores. In line with Benhassine et al. (2015), I find that the cash transfer had in fact weak effects on test scores measured as the proportion of children who could correctly recognize digits and numbers and perform simple operations such as subtraction and division (Appendix Table A3.9). If anything, evidence is more in favor with the transfer being more effective in improving test scores for boys. In this respect, three observations are worth making. First, the endline ASER test was only administered to one child per household aged 6-12 at baseline, hence the sample size for these analyses is greatly reduced. Second, the ASER test only measures numeric and arithmetic capabilities. In light of existing literature showing that also in low-income contexts boys are typically better at maths while girls perform better in other domains such as literacy (Dickerson, McIntosh, and Valente 2015), a test covering a broader set of skills would be desirable to evaluate alternative channels. Third, measurement error remains a potential issue plaguing the validity of test score measures. Hence, the combination of small sample size, limited test coverage, and possible measurement error might account for the limited role of test scores in explaining differential grade progression outcomes.

Other explanations could lie behind the significant effect of the transfer on girls' timely grade progression.[83] For instance, girls might have been more likely than boys to be pulled out of school by their families in the first place, hence the cash transfer – increasing in amount as a child progressed through school – could have had a proportionately larger impact on the former group. A second hypothesis may posit that grade progression improved for girls due to a more sizeable reduction in girls' absenteeism, another important predictor of grade repetition. This finding is somewhat consistent with Benhassine et al. (2015) – who reported a reduction in absenteeism following the intervention – and with my additional analyses by gender of the child showing that as a result of the treatment absenteeism decreased for both boys and girls (more for girls), yet the benefit for girls engaged in UCW

[83] As the treatment did not reduce time spent on unpaid care work, reduction in unpaid care work is not a plausible mechanism.

was reduced (analyses available upon request). In a similar vein, girls might have spent more time doing homework and participating in extracurricular activities organized by the school following the implementation of the program. Yet this explanation is less likely to hold true as more dedication to homework plausibly translates into improved test scores, a finding which is at odds with Benhassine et al. Alternatively, the implementation of the cash transfer program might have changed headmaster's behavior, yet data on these variables are not available. Most likely, the cash transfer affected girls' progression differentially as parents interpreted the introduction of a program sponsored by the Ministry of Education as a positive signal about the value of girls' education. Consistent with Benhassine et al., parental beliefs regarding the returns to education – measured as the subjective probability that a child who completes a certain level of education gets a job once adult – dramatically increased following the intervention, especially for girls. For girls expected to complete junior high school or high school, the cash transfer program led to large positive changes in the perceived returns to education (Appendix Table A3.10).[84] This hypothesis also aligns with findings from previous research showing that parents respond to interventions that increase the perceived returns to education by boosting their children's motivation and effort in school (Jensen 2010; Jensen 2012; Nguyen 2008).

Limitations

One important limitation of this study deals with the possibility that the amount of unpaid care work children engage in reflects an explicit choice of parents who respond to their children's cognitive endowments by investing more in the child who is more successful in school. Stated otherwise, parents might value unpaid care work as an option that empowers less academically-oriented offspring. Data limitations – such as the lack of data on test scores at baseline – prevent us from exploring whether, at the very least, test scores at baseline differ between kids engaged and not engaged in UCW. Some evidence consistent with this idea emerges from analyses of endline scores limited to the control group. These suggest that the proportion of kids who answered correctly was lower the higher the engagement in UCW, with the pattern limited to girls (Appendix Table A3.11).

Adopting an instrumental variable approach would be a suitable approach to deal with the endogenous nature of UCW. This route is however challenging, as the instrument would need to proxy for the demand for household work and be exogenous to household decisions on schooling. A number of papers have proposed instrumental variables to identify the causal impact of child labor on

[84] In households where the average amount of time spent on unpaid care work at baseline – average among all children aged 6-15 at baseline present in the household – was more than two hours per day, the positive effect of the treatment was reduced to almost zero.

educational outcomes in the context of developing countries, with most of the focus being on market work. An exception is Assaad, Levison, and Zibani (2010), who proxied for the demand for domestic work in rural Egypt using households' access to basic public services such as piped water, piped sewage disposal, and garbage collection. In a similar spirit I attempted to use information on household access to services (well, piped water, etc.) and household composition (ratio of girls over boys aged 16 to 18 in the household) as instruments for unpaid care work, though the first stage did not prove strong enough.[85] As the primary aim of this paper is to uncover the differential causal impact of the intervention on different types of students defined by baseline characteristics – rather than to estimate the causal effect of unpaid care work on progression – I leave this issue aside. Yet it is important to keep in mind that this concern might underlie a different mechanism. Further research should take up this important concern.

3.8 Conclusions and discussion

In this study I have used secondary data from a cluster-randomized control trial to evaluate the impact of a cash-transfer program that was launched in 2007 with the aim of increasing the rural primary school completion rate in rural Morocco. My first finding suggests that a cash transfer explicitly tied to an educational purpose – yet not expressly supporting learning – affected school progression on top of enrolment and attendance by stemming dropout (extensive margin) and increasing the likelihood of timely grade progression (intensive margin). While the former effect operated similarly for boys and girls, the latter was null for boys. In other words, the cash transfer was effective in reducing boys' dropout, though it did not alter their grade progression path.

My second finding suggests that while the effect of the intervention on dropout operated equally for children who performed unpaid care work versus those who did not, the beneficial effect of the treatment on timely grade progression was cut by a half to a third for girls overburdened by household chores. Therefore, there is reason to believe that – as a result of the treatment – girls engaged in unpaid care tasks were staying in school more but were less likely to progress on time. This claim lends itself to two implications. First, it points to the need to design policy interventions with the potential to help children attend and progress through school in a timely manner, as a monetary

[85] In Appendix Table A3.12 I nonetheless reported an analysis of the determinants of UCW, including the above-mentioned (potential) instruments. Interestingly, a higher ratio of girls over boys aged 16-18 in the household is associated with a sharp reduction in hours per day spent on UCW for girls but not for boys, suggesting that older girls "relieve" younger girls from some of these responsibilities. Also, having access to a well or an individual network connection, and a pipe to sanitation network for used water in the household are negatively associated with UCW for all kids, yet coefficients are on the whole bigger in magnitude for girls.

"nudge" might not be sufficient to simultaneously achieve both goals. Second, it stresses the importance of tackling rooted gender inequalities within the household by directing policies to vulnerable children that may need ad hoc targeting.

Finally, despite unpaid care work emerges as a barrier to school progression, estimates show that a cash transfer of this kind was not effective in reducing unpaid care work, neither for boys nor for girls. Although the cash transfer system was not designed to address any form of child work in the first place, it is still surprising that the policy boosted school enrolment, attendance, and progression without affecting household chores. Once again, this finding suggests that policymakers concerned with improving children's academic outcomes should take into more careful consideration the mechanisms whereby these outcomes are produced and reinforced – often originating in the family and attributable to within-household inequalities – and craft policies accordingly.

My findings complement and enrich knowledge on the effectiveness of Tayssir by shedding light on an academic outcome that rarely receives adequate consideration both in the education literature and in the policy debate. Moreover, while the main policy implication in Benhassine et al. relates to the idea of promoting cheaper cash transfers tied to education ("nudge") over more expensive programs conditional on attendance ("shove"), I here suggest that policies aimed at improving school outcomes should target both the intensive and the extensive margin of school participation, and take into more careful consideration the within-household pathways that promote or hinder some of these outcomes. A concern that I share with Benhassine et al. is the extent to which the cash transfer impacts would persist in the long run, as the documented effects relate to an intervention that was implemented over a two-year period and evaluated shortly thereafter. To the extent that the positive impacts of the cash transfer on girls' progression are due to an increased estimate of the returns to education, long-run impacts would be hampered, for instance, if the program led parents to temporarily overestimate those returns.

Overall, this study has demonstrated the need to focus on school progression outcomes, and to advance knowledge on the interplay between household inequality, as driven by gender and unpaid care work dynamics, and children's schooling. The study has also stressed the added value of capitalizing on the many hundreds impact evaluations conducted in low and middle-income countries over the past decade to shed new light on dynamics that may have been neglected in previous analyses. Addressing questions concerning variation in and mechanisms of program impact is crucial to both promoting transparency and guiding the replicability and scalability of successful programs, and greatly increases the returns yielded from valuable, yet hugely expensive large-scale interventions. Future research should push forward the field in its use and sharing of large-scale intervention data toward greater insights, and thus better outcomes, in areas such as women's economic empowerment and child development in low and middle-income countries.

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3.10 Tables

Table 3.1: Summary statistics and balance checks for the post-attrition sample

Variable	Mean control	Diff. treatment-control	Obs.
	(1)	(2)	(3)
Female	0.466	0.027	9,938
	[0.499]	(0.018)	
Age	10.44	0.047	9,938
	[2.493]	(0.064)	
Number of siblings (0-15)	1.701	-0.009	9,938
	[0.965]	(0.044)	
Number of children in the HH (ages 0-5)	0.806	-0.009	9,938
	[0.876]	(0.040)	
Number of children in the HH (ages 0-15)	3.575	-0.035	9,938
	[1.341]	(0.066)	
Household size	7.273	-0.052	9,938
	[2.151]	(0.099)	
Unpaid care work (mins/day)	47.50	-2.321	9,938
	[87.08]	(3.326)	
Mother has primary education (or more)	0.049	-0.000	9,938
	[0.216]	(0.010)	
Father has primary education (or more)	0.164	0.009	9,938
	[0.370]	(0.023)	
Never enrolled in school (june 2008)	0.086	-0.024*	9,938
	[0.280]	(0.009)	
Currently enrolled in school (june 2008)	0.761	0.043**	9,938
	[0.426]	(0.016)	
Share of children in the HH enrolled in school (june 2008)	0.606	0.031*	9,938
	[0.269]	(0.015)	
Current grade (june 2008)	3.190	0.023	7,688
	[1.649]	(0.050)	
Age of entry in primary school (for enrolled in june 2008)	6.449	0.013	7,688
	[1.019]	(0.045)	

Source: Baseline survey – post-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. Column (1) reports the mean of the variable listed on the left in the control group, with standard deviations in brackets; column (2) reports the coefficient from an OLS regression of the left-hand side variable on a dummy for treatment, accounting for strata randomization dummies.
$^{+}$ *p*<0.1; * *p*<0.05; ** *p*<0.01; *** p<0.001.

Table 3.2: Average time per day (mins/day) spent on unpaid care work, by activity. Children aged 6-15, by gender

Activity		**Overall**	**Boys**	**Girls**
Preparing food for a meal (FM)	Mean	6.344	2.956	9.948
	[SD]	[21.64]	[13.99]	[27.08]
Preparing food for another occasion (FO)	Mean	0.461	0.139	0.804
	[SD]	[4.857]	[2.004]	[6.648]
Doing housework (HW)	Mean	14.67	2.049	28.08
	[SD]	[40.36]	[15.06]	[52.63]
Washing clothes (WC)	Mean	4.645	0.672	8.869
	[SD]	[19.39]	[6.213]	[26.45]
Other domestic activities (OD)	Mean	6.581	2.153	11.29
	[SD]	[26.85]	[14.43]	[34.98]
Shopping for the house (SH)	Mean	0.934	1.093	0.765
	[SD]	[6.343]	[6.981]	[5.582]
Get water (GW)	Mean	9.909	8.026	11.91
	[SD]	[27.27]	[25.01]	[29.35]
Occupied with children in the HH (CHH)	Mean	2.270	0.577	4.071
	[SD]	[16.98]	[6.982]	[23.16]
Occupied with elderly in the HH (EHH)	Mean	0.195	0.093	0.303
	[SD]	[4.460]	[2.473]	[5.875]
Occupied with other sick/handicapped in the HH (SHH)	Mean	0.063	0.031	0.097
	[SD]	[2.007]	[1.119]	[2.643]
Unpaid care work (UCW)	Mean	46.07	17.79	76.14
	[SD]	[83.30]	[43.35]	[102.8]
Obs.		9,938	5,101	4,837

Source: Baseline survey – post-attrition sample.
Notes: Sampling weights used. Standard deviations in brackets.

Table 3.3: LPM on dropout by the end of year 2 among those enrolled in school at baseline (panel a) and timely grade progression through school (panel b), by gender

a. Dropout	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	-0.047**	-0.046**	-0.041**	-0.073***	-0.080***	-0.077***
	(0.016)	(0.015)	(0.014)	(0.016)	(0.016)	(0.017)
Number of siblings		0.002	0.002		0.008	0.008
		(0.004)	(0.004)		(0.006)	(0.006)
Unpaid care work (hrs/day)		0.018*	0.037		-0.001	0.002
		(0.008)	(0.023)		(0.005)	(0.013)
CT x Unpaid care work			-0.023			-0.004
			(0.024)			(0.014)
Constant	0.114***	-0.227***	-0.233***	0.171***	-0.450***	-0.453***
	(0.015)	(0.034)	(0.034)	(0.015)	(0.045)	(0.046)
Controls	No	Yes	Yes	No	Yes	Yes
Obs.	4,256	4,256	4,256	3,432	3,432	3,432

b. Timely grade progression	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.021	0.020	0.026	0.061+	0.054+	0.088*
	(0.025)	(0.024)	(0.027)	(0.032)	(0.032)	(0.038)
Number of siblings		-0.002	-0.002		-0.022*	-0.022*
		(0.009)	(0.009)		(0.011)	(0.011)
Unpaid care work (hrs/day)		0.019	0.043		-0.034***	0.003
		(0.012)	(0.028)		(0.009)	(0.020)
CT x Unpaid care work			-0.029			-0.044*
			(0.031)			(0.022)
Constant	0.617***	0.456***	0.450***	0.620***	0.376***	0.347***
	(0.022)	(0.066)	(0.066)	(0.029)	(0.071)	(0.074)
Controls	No	Yes	Yes	No	Yes	Yes
Obs.	3,936	3,936	3,936	3,045	3,045	3,045

Source: Baseline and endline surveys – post-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. Estimates from panel a from specification (1) pooled by gender show that there is no differential treatment effect by gender of the child (p=0.132); estimates from panel b from specification (1) pooled by gender show that there is no differential treatment effect by gender of the child (p=0.168).
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table 3.4: Bounding of estimates for timely grade progression. Dropouts assumed to progress on time (panel a) and to not progress at all (panel b)

	Timely grade progression					
a. Dropout coded as 1	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.004	0.003	0.011	0.029	0.022	0.054+
	(0.023)	(0.022)	(0.024)	(0.027)	(0.027)	(0.033)
Number of siblings		-0.002	-0.002		-0.017+	-0.017+
		(0.009)	(0.009)		(0.010)	(0.010)
Unpaid care work (hrs/day)		0.023*	0.051*		-0.029***	0.003
		(0.011)	(0.024)		(0.007)	(0.016)
CT x Unpaid care work			-0.033			-0.039*
			(0.027)			(0.018)
Constant	0.659***	0.390***	0.381***	0.684***	0.268***	0.241***
	(0.020)	(0.062)	(0.062)	(0.025)	(0.063)	(0.065)
b. Dropout coded as 0	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.051*	0.049*	0.052*	0.103***	0.104***	0.135***
	(0.025)	(0.025)	(0.026)	(0.029)	(0.029)	(0.034)
Number of siblings		-0.004	-0.004		-0.025*	-0.025*
		(0.009)	(0.009)		(0.010)	(0.010)
Unpaid care work (hrs/day)		0.006	0.016		-0.029***	0.002
		(0.012)	(0.030)		(0.008)	(0.019)
CT x Unpaid care work			-0.012			-0.037+
			(0.032)			(0.021)
Constant	0.545***	0.617***	0.614***	0.511***	0.714***	0.688***
	(0.023)	(0.065)	(0.065)	(0.026)	(0.072)	(0.074)
Controls	No	Yes	Yes	No	Yes	Yes
Obs.	4,256	4,256	4,256	3,432	3,432	3,432

Source: Baseline and endline surveys – post-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. Estimates from panel a from specification (1) pooled by gender show that there is no differential treatment effect by gender of the child (p=0.365); estimates from panel b from specification (1) pooled by gender show that there is a differential treatment effect by gender of the child (p=0.081).
$^{+}$ *p*<0.1; * *p*<0.05; ** *p*<0.01; *** p<0.001.

Table 3.5: Inverse Probability Weighting (panel a) and Heckman two-step correction (panel b)

	Timely grade progression					
a. IPW	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.020	0.015	0.004	0.101+	0.097+	0.139*
	(0.053)	(0.049)	(0.052)	(0.061)	(0.053)	(0.066)
Number of siblings		0.006	0.006		-0.058*	-0.058*
		(0.020)	(0.020)		(0.026)	(0.026)
Unpaid care work (hrs/day)		0.024	-0.042		-0.065**	0.010
		(0.026)	(0.050)		(0.020)	(0.039)
CT x Unpaid care work			0.070			-0.062+
			(0.057)			(0.033)
Constant	0.579***	0.283*	0.293*	0.506***	0.093	0.051
	(0.050)	(0.124)	(0.125)	(0.066)	(0.157)	(0.159)
Controls	No	Yes	Yes	No	Yes	Yes
Obs.	3,936	3,936	3,936	3,045	3,045	3,045
b. Heckman	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.021	0.024	0.028	0.077**	0.056*	0.087**
	(0.024)	(0.022)	(0.023)	(0.027)	(0.025)	(0.029)
Number of siblings		-0.003	-0.002		-0.017+	-0.018+
		(0.009)	(0.009)		(0.009)	(0.009)
Unpaid care work (hrs/day)		0.012	0.028		-0.037***	-0.000
		(0.013)	(0.032)		(0.008)	(0.018)
CT x Unpaid care work			-0.019			-0.043*
			(0.034)			(0.019)
Constant	0.652***	0.401***	0.400***	0.678***	0.454***	0.439***
	(0.081)	(0.101)	(0.101)	(0.078)	(0.098)	(0.098)
Lambda (IMR)	0.022	0.117	0.113	0.120	0.044	0.029
	(0.108)	(0.083)	(0.083)	(0.082)	(0.070)	(0.070)
Controls	No	Yes	Yes	No	Yes	Yes
Censored obs.	320	320	320	387	387	387
Uncensored obs.	3,936	3,936	3,936	3,045	3,045	3,045

Source: Baseline and endline surveys – post-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. Estimates from panel a from specification (1) pooled by gender show that there is a differential treatment effect by gender of the child (p=0.089); estimates from panel b from specification (1) pooled by gender show that there is a differential treatment effect by gender of the child (p=0.078).
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

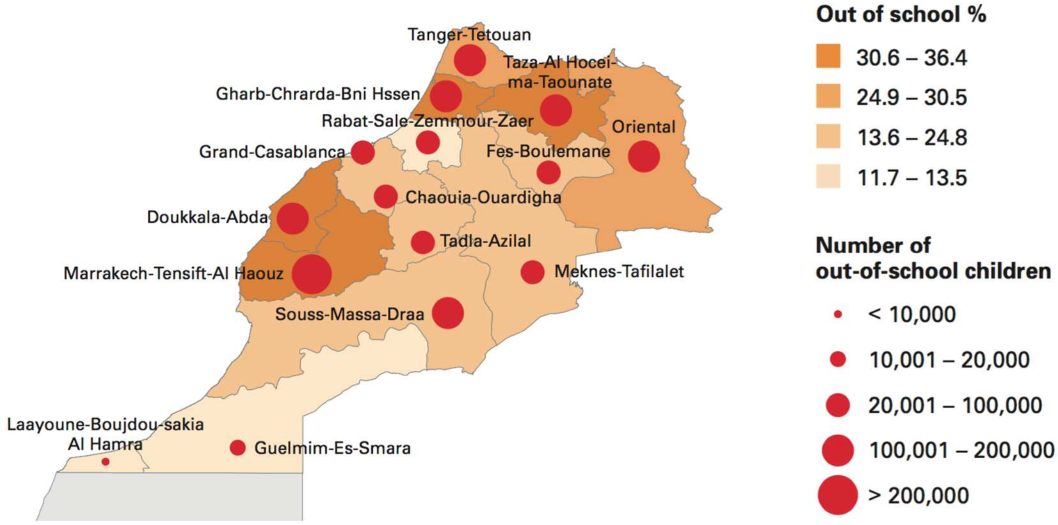
Table 3.6: Difference-in-Difference (DD) estimation predicting time (hours) per day spent on unpaid care work, by gender

	Boys		Girls	
	(1)	(2)	(1)	(2)
Time (Post)	-0.040	-0.040	0.532***	0.532***
	(0.041)	(0.041)	(0.118)	(0.118)
Treatment (CT)	-0.009	0.002	-0.114	-0.082
	(0.041)	(0.040)	(0.097)	(0.083)
Post x CT	0.037	0.037	0.033	0.033
	(0.046)	(0.046)	(0.128)	(0.128)
Number of siblings		0.012		-0.042*
		(0.009)		(0.021)
Age (at baseline)		0.030***		0.232***
		(0.003)		(0.008)
Constant	0.304***	0.143*	1.359***	-0.343**
	(0.037)	(0.062)	(0.090)	(0.132)
Controls	No	Yes	No	Yes
Obs.	10,202	10,202	9,674	9,674

Source: Baseline and endline surveys – post-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. Number of observations is doubled as there are two observations per child (2008 and 2010).
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

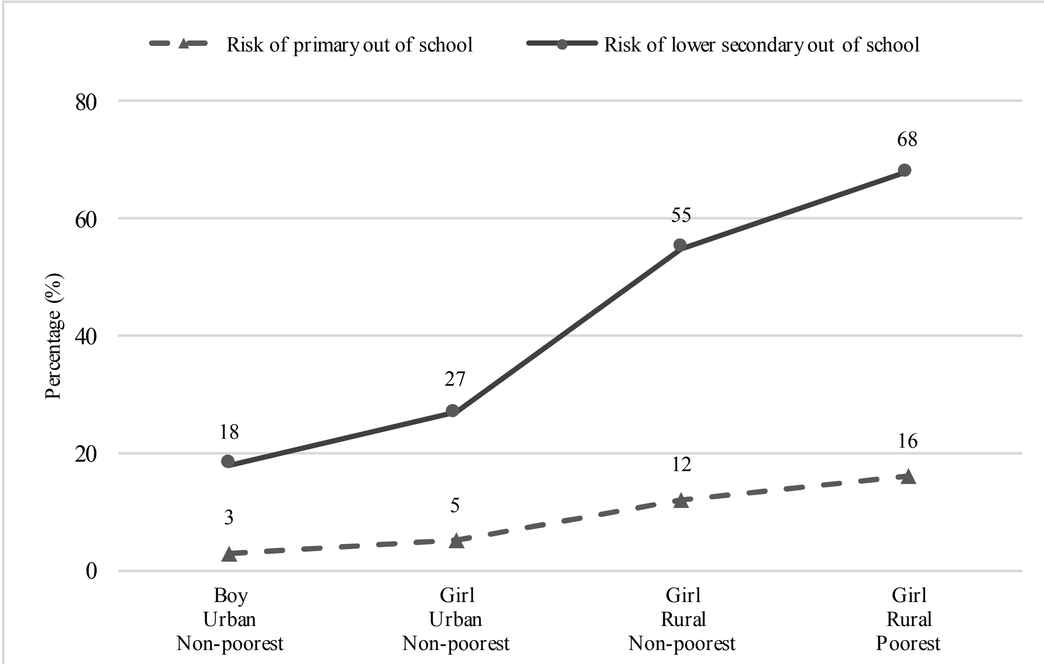
3.11 Figures

Figure 3.1: Number and percent of out-of-school children by region of Morocco (DHS 2003-04)



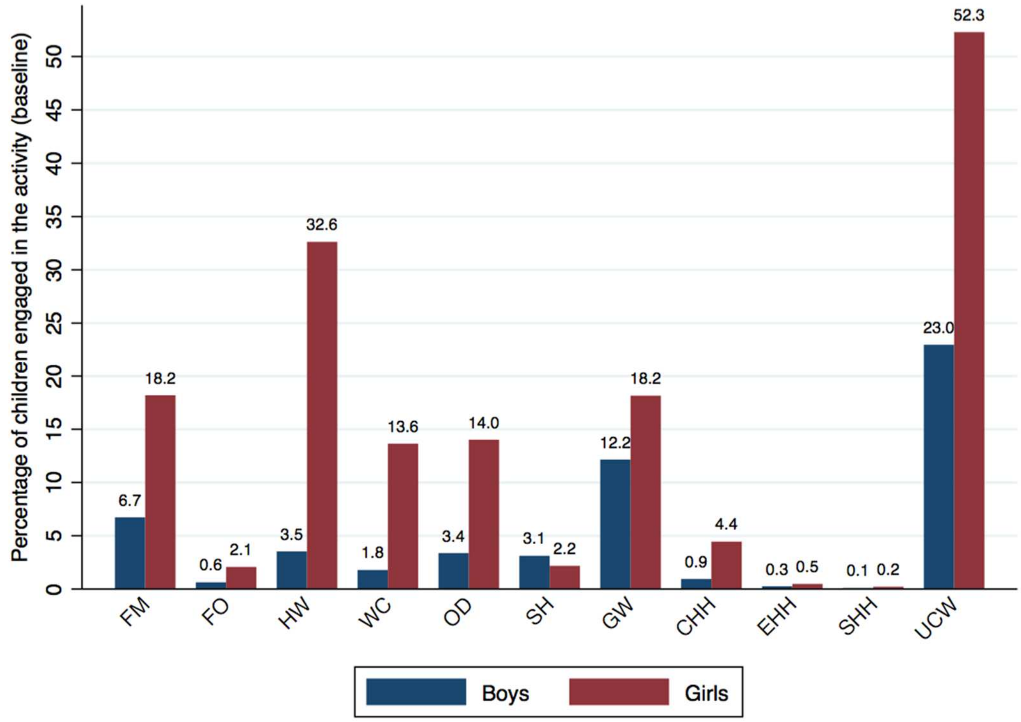
Source: Morocco's most recent Demographic and Health Survey (DHS), i.e., DHS 2003-04.
Notes: These statistics include children of primary school age who are not in primary or secondary school, and children of lower secondary school age who are not in primary or secondary school.

Figure 3.2: Cumulative out-of-school risk according to selected characteristics (DHS 2003-04)



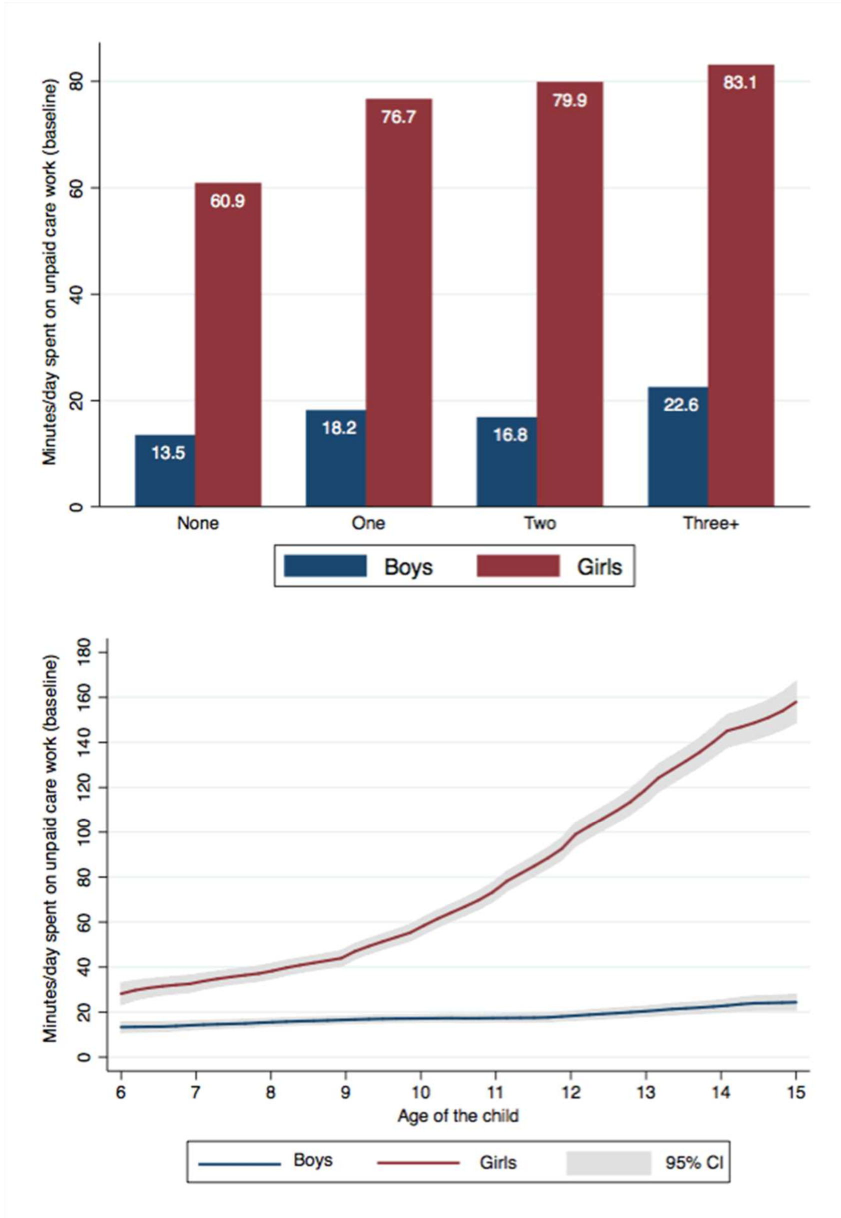
Source: Morocco's most recent Demographic and Health Survey (DHS), i.e., DHS 2003-04.
Notes: Estimates computed using multivariate logistic regression.

Figure 3.3: Percentage of children performing each unpaid care work activity, by gender



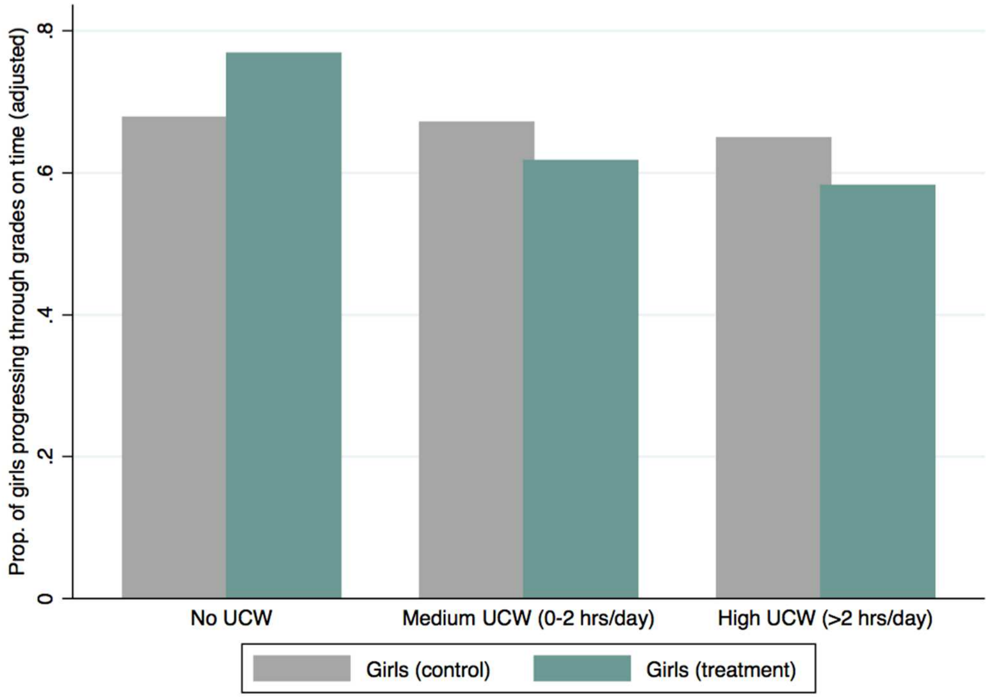
Source: Baseline survey – post-attrition sample.

Figure 3.4: Minutes per day spent on unpaid care work by gender and number of siblings (top panel), and by gender and age of the child (bottom panel)



Source: Baseline survey – post-attrition sample.

Figure 3.5: Predicted girls' grade progression, by treatment status and categories of unpaid care work



Source: Baseline and endline survey – post-attrition sample.
Notes: Corresponding detailed regression estimates reported in Table A.5 in the Appendix. Adjusted proportions computed assuming controls at their mean values. Tests for statistical significance of the coefficients between treatment and control groups: T-No UCW vs C-No UCW: p>F=0.037; T-Medium UCW vs C-Medium UCW: p>F=0.596; T-High UCW vs C-High UCW: p>F=0.623. Tests for statistical significance of the coefficients between treatment groups: T-No UCW vs T-Medium UCW: p>F=0.093; T-No UCW vs T-High UCW: p>F=0.069; T-Medium UCW vs T-High UCW: p>F=0.753.

3.12 Appendix

Table A3.1: Differential attrition by treatment status and selected characteristics

Variable	No attrition	Diff. attrition - no attrition	Obs.
	(1)	(2)	(3)
Treatment (CT)	0.817	-0.063*	10,889
	[0.386]	(0.029)	
Female	0.485	-0.039*	10,889
	[0.500]	(0.019)	
Age	10.45	-0.118	10,889
	[2.486]	(0.076)	
Number of siblings (0-15)	1.678	-0.078	10,889
	[0.974]	(0.056)	
Number of children in the HH (ages 0-5)	0.808	0.080	10,889
	[0.850]	(0.050)	
Number of children in the HH (ages 0-15)	3.537	-0.013	10,889
	[1.337]	(0.073)	
Household size	7.179	-0.458***	10,889
	[2.036]	(0.089)	
Unpaid care work (mins/day)	45.95	-2.023	10,889
	[83.22]	[3.441]	
Mother has primary education (or more)	0.048	0.020	10,889
	[0.215]	(0.013)	
Father has primary education (or more)	0.167	0.037+	10,889
	[0.373]	(0.022)	
Never enrolled in school (june 2008)	0.068	0.015	10,889
	[0.253]	(0.011)	
Currently enrolled in school (june 2008)	0.791	-0.016	10,889
	[0.407]	(0.018)	
Current grade (june 2008)	3.196	-0.064	8,385
	[1.632]	(0.063)	

Source: Baseline survey – pre-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. Column (1) reports the mean of the variable listed on the left in the non-attrition group, with standard deviations in brackets; column (2) reports the coefficient from an OLS regression of the left-hand side variable on a dummy for attrition, accounting for strata randomization dummies.
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table A3.2: Summary statistics and balance checks for the unpaid care work activities

Activity	**Mean control**	**Diff. treatment-control**	**Obs.**
	(1)	(2)	(3)
Preparing food for a meal (FM)	7.327	-1.124	9,938
	[24.15]	(0.865)	
Preparing food for another occasion (FO)	0.487	-0.001	9,938
	[4.433]	(0.142)	
Doing housework (HW)	14.32	0.293	9,938
	[42.15]	(1.433)	
Washing clothes (WC)	5.144	-0.497	9,938
	[21.48]	(0.622)	
Other domestic activities (OD)	5.730	0.835	9,938
	[23.81]	(0.784)	
Shopping for the house (SH)	0.732	0.172	9,938
	[5.041]	(0.171)	
Get water (GW)	10.58	-0.741	9,938
	[27.43]	(1.350)	
Occupied with children in the HH (CHH)	2.958	-1.238+	9,938
	[21.81]	(0.694)	
Occupied with elderly in the HH (EHH)	0.095	0.069	9,938
	[2.777]	(0.084)	
Occupied with other sick/handicapped in the HH (SHH)	0.122	-0.088	9,938
	[3.476]	(0.091)	

Source: Baseline survey – post-attrition sample.
Notes: Time measured in minutes per day. Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. Column (1) reports the mean of the variable listed on the left in the control group, with standard deviations in brackets; column (2) reports the coefficient from an OLS regression of the left-hand side variable on a dummy for treatment, accounting for strata randomization dummies.
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table A3.3: Gender differences in means (t-test) in time spent on unpaid care work, by activity. Children aged 6-15 who report spending a *non-zero* amount of time on the activity

Activity	Boys		Girls		Diff.
	Obs.	Mean	Obs.	Mean	(s.e.)
Preparing food for a meal (FM)	343	42.87	880	57.17	-14.30***
					(2.467)
Preparing food for another occasion (FO)	32	22.81	100	40.50	-17.69***
					(4.775)
Doing housework (HW)	180	60.90	1,578	88.41	-27.51***
					(4.641)
Washing clothes (WC)	91	39.23	660	66.19	-26.96***
					(4.056)
Other domestic activities (OD)	172	66.94	678	83.27	-16.33***
					(4.578)
Shopping for the house (SH)	159	35.62	105	32.22	3.396
					(2.382)
Get water (GW)	620	66.41	878	66.52	-0.107
					(1.931)
Occupied with children in the HH (CHH)	48	62.29	215	89.50	-27.21**
					(9.585)
Occupied with elderly in the HH (EHH)	13	36.41	23	65.80	-29.39
					(17.81)
Occupied with other sick/handicapped in the HH (SHH)	5	34.67	10	47.00	-12.33
					(17.57)
Unpaid care work (UCW)	1,172	78.46	2,529	149.1	-70.62***
					(3.177)

Source: Baseline survey – post-attrition sample.
Notes: Time measured in minutes per day. Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used.
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table A3.4: Model specifications on the pooled sample with a dummy for gender and a treatment-gender interaction term. Dropout since 2008 (panel a) and timely grade progression (panel b)

	a. Dropout				**b. Timely grade progression**			
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Treatment (CT)	- 0.055***	- 0.058***	-0.044**	-0.042**	0.039+	0.034	0.015	0.024
	(0.012)	(0.012)	(0.015)	(0.014)	(0.022)	(0.022)	(0.025)	(0.026)
Number of siblings		0.006	0.006	0.006		-0.011	-0.011	-0.011
		(0.004)	(0.004)	(0.004)		(0.007)	(0.007)	(0.007)
Unpaid care work (hrs/day)		0.008+	0.008+	0.015		-0.016*	-0.016*	0.016
		(0.004)	(0.004)	(0.011)		(0.007)	(0.007)	(0.017)
CT x Unpaid care work				-0.009				-0.038*
				(0.012)				(0.018)
Female		0.046***	0.074***	0.070***		0.051***	0.013	-0.004
		(0.008)	(0.019)	(0.020)		(0.013)	(0.031)	(0.032)
CT x Female			-0.033	-0.029			0.046	0.065+
			(0.020)	(0.022)			(0.034)	(0.035)
Constant	0.136***	- 0.343***	- 0.355***	- 0.357***	0.619***	0.415***	0.432***	0.423***
	(0.011)	(0.030)	(0.031)	(0.031)	(0.020)	(0.053)	(0.054)	(0.055)
Controls	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Obs.	7,688	7,688	7,688	7,688	6,981	6,981	6,981	6,981

Source: Baseline and endline surveys – post-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used.
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table A3.5: LPM on dropout by the end of year 2 among those enrolled in school at baseline (panel a) and timely grade progression through school (panel b), by gender. Unpaid care work measured in categories

a. Dropout	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	-0.047**	-0.046**	-0.038**	-0.073***	-0.082***	-0.079***
	(0.016)	(0.015)	(0.015)	(0.016)	(0.016)	(0.021)
Number of siblings		0.002	0.002		0.008	0.008
		(0.004)	(0.004)		(0.006)	(0.006)
Unpaid care work (Ref.: None)						
Medium (0-2 hrs/day)		0.019	0.046		-0.011	-0.008
		(0.012)	(0.035)		(0.012)	(0.038)
High (>2 hrs/day)		0.026	0.072		0.006	0.019
		(0.022)	(0.079)		(0.017)	(0.045)
CT x Unpaid care work (Ref.: CT x None)						
CT x Medium			-0.033			-0.003
			(0.037)			(0.041)
CT x High			-0.056			-0.015
			(0.083)			(0.048)
Constant	0.114***	-0.227***	-0.234***	0.171***	-0.444***	-0.447***
	(0.015)	(0.034)	(0.034)	(0.015)	(0.045)	(0.047)
Controls	No	Yes	Yes	No	Yes	Yes
Obs.	4,256	4,256	4,256	3,432	3,432	3,432
b. Timely grade progression	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.021	0.019	0.021	0.061+	0.056+	0.090*
	(0.025)	(0.024)	(0.028)	(0.032)	(0.032)	(0.043)
Number of siblings		-0.002	-0.002		-0.022*	-0.022*
		(0.009)	(0.009)		(0.011)	(0.011)
Unpaid care work (Ref.: None)						
Medium (0-2 hrs/day)		0.012	0.010		-0.067**	-0.007
		(0.023)	(0.054)		(0.023)	(0.059)
High (>2 hrs/day)		0.061	0.103		-0.110***	-0.029
		(0.038)	(0.106)		(0.027)	(0.067)
CT x Unpaid care work (Ref.: CT x None)						
CT x Medium			0.002			-0.071
			(0.058)			(0.064)
CT x High			-0.050			-0.096+
			(0.112)			(0.058)
Constant	0.617***	0.456***	0.454***	0.620***	0.381***	0.351***
	(0.022)	(0.066)	(0.067)	(0.029)	(0.072)	(0.075)
Controls	No	Yes	Yes	No	Yes	Yes
Obs.	3,936	3,936	3,936	3,045	3,045	3,045

Source: Baseline and endline surveys – post-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used.
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table A3.6: BLGW test for non-random attrition

BLGW test	Enrolled in school at baseline (1)	Current grade at baseline (2)
Dropout since 2008 (A)	-0.230***	1.037*
	(0.042)	(0.426)
Female	-0.042***	-0.011
	(0.010)	(0.026)
Female x A	0.037***	0.002
	(0.011)	(0.111)
Age	-0.038***	0.593***
	(0.003)	(0.008)
Age x A	0.034***	-0.121***
	(0.003)	(0.035)
Number of siblings (0-15)	-0.039***	0.012
	(0.005)	(0.014)
Number of siblings (0-15) x A	0.037***	0.167**
	(0.005)	(0.061)
Unpaid care work (hrs/day)	-0.088***	-0.018
	(0.004)	(0.013)
Unpaid care work (hrs/day) x A	0.087***	0.012
	(0.005)	(0.055)
Mother has primary education (or more)	0.013	0.187***
	(0.017)	(0.054)
Mother has primary education (or more) x A	-0.031	0.098
	(0.020)	(0.221)
Father has primary education (or more)	0.037***	0.118***
	(0.011)	(0.032)
Father has primary education (or more) x A	-0.046***	0.289*
	(0.013)	(0.128)
Household size	0.001	-0.012+
	(0.003)	(0.007)
Household size x A	0.001	0.027
	(0.003)	(0.028)
HH owns a cellphone	0.008	0.108***
	(0.010)	(0.031)
HH owns a cellphone x A	-0.003	-0.024
	(0.011)	(0.103)
HH has electrical network connection	0.038***	0.087*
	(0.011)	(0.034)
HH has electrical network connection x A	-0.037**	0.288**
	(0.012)	(0.104)
HH has one well (or more)	0.037*	-0.020
	(0.015)	(0.040)
HH has one well (or more) x A	-0.030+	0.123
	(0.018)	(0.143)
HH gets water from individual network connection	0.003	0.050
	(0.014)	(0.040)
HH gets water from individual network connection x A	-0.022	0.045
	(0.014)	(0.144)
HH gets water from natural source	-0.009	-0.048
	(0.013)	(0.032)
HH gets water from natural source x A	0.009	0.068
	(0.014)	(0.105)
Pipe to sanitation network for used water in the HH	-0.009	0.095
	(0.034)	(0.097)
Pipe to sanitation network for used water in the HH x A	0.006	-0.718*
	(0.039)	(0.364)
Constant	1.273***	-2.887***
	(0.032)	(0.093)
Obs.	9,938	7,688
F-test (Joint significance of A and A(x(0)))	110.9	4.48
P>F	0.000	0.000

Source: Baseline and endline surveys – post-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used.
$^{+}$ *p*<0.1; * *p*<0.05; ** *p*<0.01; *** p<0.001.

Table A3.7: Statistically significant baseline predictors of attrition (defined as dropping out of school between 2008 and 2010)

Strongest baseline predictors of dropout	
Boys	**Girls**
Age (+)	Age (+)
Treatment (-)	Treatment (-)
UCW (+)	
HH has electrical network connection (-)	
HH gets water from individual network connection (-)	

Table A3.8: LPM on endline unpaid care work, controlling for baseline

	Boys			Girls		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.025	0.026	0.032	-0.051	-0.023	-0.047
	(0.031)	(0.031)	(0.031)	(0.085)	(0.084)	(0.085)
Unpaid care work (hrs/day, baseline)		0.151***	0.136***		0.199***	0.088***
		(0.032)	(0.031)		(0.018)	(0.020)
Number of siblings			-0.001			-0.040
			(0.012)			(0.033)
Constant	0.267***	0.221***	-0.054	1.867***	1.591***	-0.056
	(0.026)	(0.027)	(0.071)	(0.076)	(0.078)	(0.187)
Controls	No	No	Yes	No	No	Yes
Obs.	5,101	5,101	5,101	4,837	4,837	4,837

Source: Baseline and endline surveys – post-attrition sample.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used.
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table A3.9: LPM on ASER endline test scores

a. Digits	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.010	0.009	0.013	0.011	0.008	-0.001
	(0.008)	(0.008)	(0.009)	(0.018)	(0.017)	(0.018)
Number of siblings		-0.003	-0.003		-0.003	-0.002
		(0.004)	(0.004)		(0.005)	(0.005)
Unpaid care work (hrs/day)		0.006+	0.019*		-0.005	-0.013
		(0.004)	(0.008)		(0.005)	(0.013)
CT x Unpaid care work			-0.015+			0.010
			(0.008)			(0.015)
Constant	0.974***	0.871***	0.868***	0.964***	0.868***	0.877***
	(0.007)	(0.030)	(0.030)	(0.016)	(0.040)	(0.039)
b. Numbers	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.023	0.018	0.015	0.043+	0.036	0.036
	(0.016)	(0.016)	(0.016)	(0.025)	(0.023)	(0.025)
Number of siblings		-0.011	-0.011		-0.018*	-0.018*
		(0.009)	(0.009)		(0.008)	(0.008)
Unpaid care work (hrs/day)		0.002	-0.005		0.004	0.004
		(0.008)	(0.020)		(0.007)	(0.015)
CT x Unpaid care work			0.009			-0.000
			(0.021)			(0.017)
Constant	0.923***	0.644***	0.646***	0.891***	0.572***	0.571***
	(0.014)	(0.052)	(0.053)	(0.022)	(0.059)	(0.059)
c. Subtraction	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.057	0.045	0.045	0.032	0.014	-0.029
	(0.045)	(0.044)	(0.046)	(0.050)	(0.048)	(0.055)
Number of siblings		-0.011	-0.011		-0.006	-0.005
		(0.014)	(0.014)		(0.014)	(0.014)
Unpaid care work (hrs/day)		-0.029	-0.029		-0.002	-0.041+
		(0.024)	(0.051)		(0.010)	(0.021)
CT x Unpaid care work			0.001			0.048*
			(0.059)			(0.024)
Constant	0.459***	-0.205*	-0.205*	0.474***	-0.223*	-0.185+
	(0.040)	(0.097)	(0.097)	(0.046)	(0.097)	(0.102)
d. Division	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Treatment (CT)	0.065+	0.060	0.069+	-0.019	-0.031	-0.074
	(0.038)	(0.037)	(0.038)	(0.042)	(0.041)	(0.050)
Number of siblings		-0.001	-0.001		-0.006	-0.005
		(0.013)	(0.013)		(0.013)	(0.013)
Unpaid care work (hrs/day)		0.003	0.030		-0.002	-0.040+
		(0.024)	(0.045)		(0.011)	(0.022)
CT x Unpaid care work			-0.032			0.048+
			(0.053)			(0.025)
Constant	0.327***	-0.551***	-0.557***	0.392***	-0.315***	-0.277**
	(0.034)	(0.089)	(0.089)	(0.038)	(0.089)	(0.094)
Controls	No	Yes	Yes	No	Yes	Yes
Obs.	1,694	1,694	1,694	1,562	1,562	1,562

Source: Baseline and endline surveys – post-attrition sample. ASER test administered to one child aged 6-12 at baseline per household during endline household survey.
Notes: Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used.
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table A3.10: LPM on parental beliefs on children's returns to education

Probability of being employed once adult for:	Boys		Girls	
	Mean in control group	Effect of the treatment	Mean in control group	Effect of the treatment
Who does not complete primary school	0.257	0.004	0.015	-0.003
	[0.181]	(0.008)	[0.069]	(0.003)
Who completes primary school	0.272	0.006	0.013	0.001
	[0.174]	(0.007)	[0.066]	(0.004)
Who completes junior high school	0.323	0.014	0.025	0.012*
	[0.186]	(0.010)	[0.100]	(0.006)
Who completes high school	0.475	0.010	0.129	0.022+
	[0.248]	(0.014)	[0.243]	(0.012)

Source: Endline survey – post-attrition sample.
Notes: Unit of observation is the household. Standard deviations in brackets. Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. No controls added. Note that respondents were not asked for a probability between 0 and 1. They were asked to choose between five categories (no chance, few chances, 50 percent chance, lots of chances, and certain chance). Probabilities of 0, 0.25, 0.5, 0.75, and 1 are imputed to these categories, respectively.
$^{+}$ p<0.1; * p<0.05; ** p<0.01; *** p<0.001.

Table A3.11: Endline ASER test scores by levels of unpaid care work at baseline (in hours/day), control group only

ASER scores (endline)	Knows digits		Knows numbers		Knows subtraction		Knows division	
	Boys	Girls	Boys	Girls	Boys	Girls	Boys	Girls
Unpaid care work (Ref.: None)								
Medium (0-2 hrs/day)	0.008	-0.027	-0.046	-0.054	0.034	-0.125+	-0.012	-0.221**
	(0.016)	(0.021)	(0.037)	(0.057)	(0.091)	(0.070)	(0.081)	(0.070)
High (>2 hrs/day)	0.037	-0.079+	0.069+	-0.036	-0.196	-0.185*	-0.009	-0.204*
	(0.022)	(0.044)	(0.035)	(0.067)	(0.163)	(0.089)	(0.169)	(0.097)
Constant	0.965***	0.987***	0.924***	0.921***	0.459***	0.546***	0.330***	0.466***
	(0.007)	(0.012)	(0.014)	(0.026)	(0.034)	(0.045)	(0.034)	(0.048)
Controls	No	No	No	No	No	No	No	No
Obs.	316	274	316	274	316	274	316	274

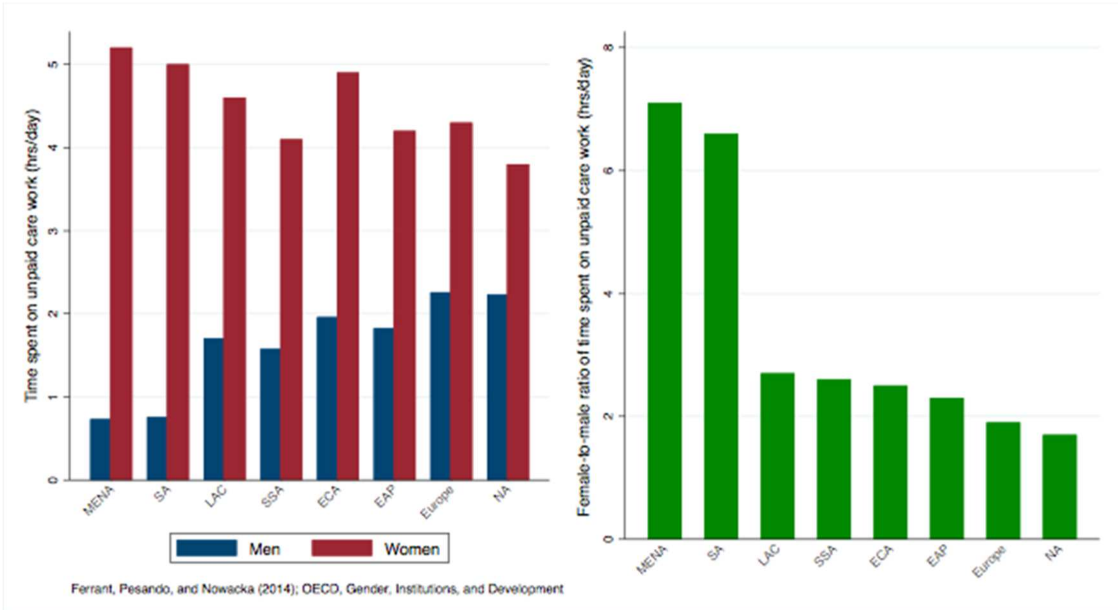
Source: Baseline and endline surveys – post-attrition sample. ASER test administered to one child aged 6-12 at baseline per household during endline household survey.
Notes: Unpaid care work measured in hours per day. Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. Strata randomization dummies included in the estimation but omitted from the table. Test scores are all dichotomous variables. Estimation limited to control units only.
$^{+}$ $p<0.1$; * $p<0.05$; ** $p<0.01$; *** $p<0.001$.

Table A3.12: Child and household-level predictors of unpaid care work at baseline (in hours/day)

Unpaid care work (baseline)	**Boys**			**Girls**		
	(1)	(2)	(3)	(1)	(2)	(3)
Age	0.021***	0.021***	0.021***	0.287***	0.287***	0.286***
	(0.005)	(0.005)	(0.005)	(0.010)	(0.010)	(0.010)
Number of siblings (0-15)		0.039**	0.037**		0.037	0.035
		(0.014)	(0.014)		(0.028)	(0.028)
Household size		-0.007	-0.007		-0.021	-0.022
		(0.005)	(0.005)		(0.014)	(0.014)
Ratio of number of girls over boys (ages 16-18)		-0.017	-0.013		-0.113**	-0.112**
		(0.015)	(0.015)		(0.035)	(0.035)
Father has primary education (or more)		-0.017	-0.003		-0.061	-0.036
		(0.032)	(0.032)		(0.070)	(0.070)
Mother has primary education (or more)		0.003	0.028		-0.188+	-0.150
		(0.044)	(0.044)		(0.103)	(0.104)
HH has one well (or more)			0.006			-0.127+
			(0.035)			(0.075)
HH gets water from individual network connection			-0.075*			-0.042
			(0.034)			(0.084)
HH gets water from natural source			0.123**			0.112+
			(0.037)			(0.064)
Pipe to sanitation network for used water in theHH			-0.094			-0.337*
			(0.058)			(0.151)
HH has electrical network connection			-0.032			-0.133*
			(0.028)			(0.060)
Constant	0.074	0.085	0.070	-1.728***	-1.500***	-1.417***
	(0.047)	(0.055)	(0.061)	(0.098)	(0.134)	(0.137)
Obs.	5,101	5,101	5,101	4,837	4,837	4,837

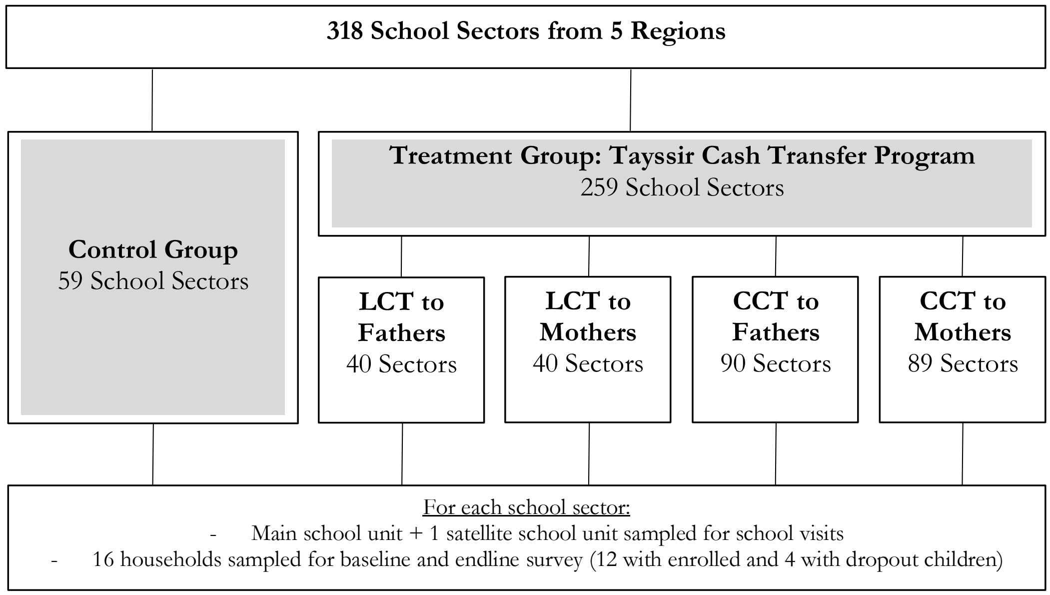
Source: Baseline survey – post-attrition sample.
Notes: Unpaid care work measured in hours per day. Standard errors (in parentheses) clustered at the school-sector level. Sampling weights used. Strata randomization dummies included in the estimation but omitted from the table. Age, number of siblings, household size, and the ratio of girls over boys (ages 16-18) in the household are continuous variables. The remaining predictors are dichotomous variables.
$^{+}$ *p*<0.1; * *p*<0.05; ** *p*<0.01; *** p<0.001.

Figure A3.1: Hours per day spent on unpaid care work by men and women (left) and female-to-male ratio of time spent on unpaid care work (right) across different world regions



Source: Ferrant, Pesando, and Nowacka (2014). OECD, Gender, Institutions, and Development Database.
Notes: The left graph reports the number of hours per day spent on unpaid care work by men (blue) and women (red); the right graph reports the female-to-male ratio of time spent on unpaid care work. Regions of the world: Middle East and North Africa (MENA), South Asia (SA), Latin America and the Caribbean (LAC), sub-Saharan Africa (SSA), Eastern Europe and Central Asia (ECA), East Asia and Pacific (EAP), Europe, and North America (NA).

Figure A3.2: Experimental design.



Notes: Shaded boxes refer to the treatment arms considered in the present analysis, while white boxes refer to treatment arms considered in Benhassine *et al.* (2015).