Essays in Applied Microeconomics

by

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ABSTRACT

ESSAYS IN APPLIED MICROECONOMICS

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This thesis consists of three Essays in Applied Microeconomics. Chapter 1 focuses on childhood chronic poverty estimates that look beyond a count approach. Chapter 2 examines the link between early childhood chronic poverty experiences and adult outcomes. Chapter 3 examines the relationship between women legislators in Africa and foreign aid allocations.

In addressing the question of which child suffers greater chronic poverty, Chapter 1 looks beyond a count-based approach by paying attention to poverty measurement approaches that account for the timing, spacing and severity of poverty spells. I compare chronic poverty experiences between groups of children based on race, age of mother at birth, region, type of household, parental educational attainment and experiences of parental marital dissolution. Not surprisingly, non-whites suffer more chronic poverty than whites. This study shows that this difference is significantly increased when the timing and spacing of poverty spells are accounted for.

Chapter 2 investigates the association between chronic poverty experiences from birth to age 10 and later life outcomes at age 25 and 30 using chronic poverty measures that account for the timing, spacing and severity of poverty spells. After controlling for correlates of childhood poverty, the results reveal that assessing the link between chronic child

poverty and adverse outcomes in adulthood based solely on time spent poor, ignoring critical aspects of chronic poverty, gives misleading estimates of the extent of damage suffered by adults who experienced chronic poverty as young children.

Chapter 3 examines the recent rise in the share of women legislators globally. We document a strong and statistically robust relationship: an increase in the share of women legislators from 15 to 20 percent is associated with an increase of about 4.3 percent in aid conditional on current levels of aid. We also show that the most effective policy instrument to implement higher women representation in national legislators in Africa is through reserved seats for women. We estimate that reserving seats for women in parliaments in a recipient country is associated with about a 53 percent increase in aid receipts.

To my wonderful mother, Benedicta.

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TABLE of CONTENTS

Abstract ii

Dedication iv

Acknowledgements v

List of Figures viii

List of Tables ix

1 Childhood Chronic Poverty Estimations: Looking Beyond a Count Index 1

1.1 Introduction 1
1.2 Literature Review 5
1.3 Measures of child chronic poverty 8
1.3.1 Notation 8
1.3.2 The Count Index 10
1.3.3 The Foster (2009) measure 11
1.3.4 The Bossert et al. (2012) measure 12
1.3.5 The Hoy and Zheng (2011) measure 13
1.4 Results and Discussion 16
1.4.1 Data and Sample Statistics 16
1.4.2 Results: Child Chronic Poverty levels by background characteristics 20
1.4.2.i Chronic poverty levels using the Count Index 20
1.4.2.ii Chronic poverty levels using the F, BCD and HZ Indices 22
1.4.3 Results: Descriptive Regression Analysis 31
1.4.4 Equally-Distributed-Equivalent (EDE) poverty gaps 33
1.5 Conclusion 38
1.6 Appendix 42

2 Chronic Child Poverty and Later Life Outcomes **50**
2.1 Introduction 50
2.2 Literature Review 54
2.3 Measures of Child Chronic Poverty 58
2.3.1 Notation 58
2.3.2 The Count Index 59
2.3.3 The Foster (2009) measure of chronic poverty 60
2.3.4 The Bossert, Chakravarty and d'Ambrosio (2012) measure of chronic poverty 61
2.3.5 Hoy and Zheng (2011) measure of chronic poverty 62
2.4 Data and Methodology 65
2.4.1 Data and Summary Statistics 65
2.4.2 Childhood Chronic Poverty Levels 67
2.4.3 Methodology 70
2.5 Results and Discussion 75
2.5.1 Childhood Chronic Poverty Experience and Completed Education . 76
2.5.2 Childhood Chronic Poverty Experience and Adult Health Status . . 78
2.5.3 Childhood Chronic Poverty Experience and the Likelihood of Teen Births 78
2.5.4 Childhood Chronic Poverty Experience and Labor Market Success . 79
2.5.5 Childhood Chronic Poverty Experience and Adult Poverty Status . 81
2.5.6 Childhood Chronic Poverty Experience and Forming Own Households 82
2.5.7 The Choice between Chronic Poverty Measures 82
2.5.8 Robustness Checks 83
2.6 Conclusion 86
2.7 Appendix 96

3 Women legislators in Africa and foreign aid **109**
3.1 Introduction 109
3.2 Background 116
3.3 Donor Aid Priorities 127
3.4 Gender Selectivity of Aid Allocations 129
3.5 Robustness Checks 138
3.6 Conclusion 140
3.7 Way forward/Future Research 141

References **149**

List of Figures

1.1 HZ weight functions for different values of the sensitivity parameter δ . . . 45

2.1 Average Completed Years of Schooling by Birth Order 74

3.1 Share of women legislators globally (1960-2018) 111
3.2 Share of women legislators vs. GDP per capita (1960–2017) 112
3.3 Share of women legislators vs. men/women-leader perceptions (1994–2016) 119
3.4 Share of women legislators vs. men/women-leader perceptions for Non-African countries (1994–2016) 120
3.5 Share of women legislators vs. men/women-leader perceptions for African countries (1994–2016) 121
3.6 Share of women legislators and quota systems in Africa (1960-2018) 124
3.7 Share of women in parliaments for African countries with reserved seats (1970-2016) 145

List of Tables

1.1 Childhood Chronic poverty estimated by the Count Index (percent) 27
1.2 Childhood Chronic poverty estimated by the F measure 28
1.3 Childhood Chronic poverty estimated by the BCD measure 29
1.4 Childhood Chronic poverty estimated by the HZ measure 30
1.5 Association between childhood chronic poverty and background characteristics 34
1.6 Equally Distributed Equivalent (EDE) poverty gaps 38
1.7 Equivalence scales by different family sizes 46
1.8 Childhood Chronic poverty estimated by the F measure (Test of significance between groups) 47
1.9 Childhood Chronic poverty estimated by the BCD measure (Test of significance between groups) 48
1.10 Childhood Chronic poverty estimated by the HZ measure (Test of significance between groups) 49

2.1 Descriptive Statistics: Adult Outcomes by Childhood Poverty Status 67
2.2 Childhood Chronic Poverty, by Birth Cohort Year 71
2.3 Association between Childhood Chronic Poverty and Adult Outcomes . . . 89
2.4 Association between Childhood Chronic Poverty and Completed Schooling 90
2.5 Association between Childhood Chronic Poverty and Good Health 91
2.6 Association between Childhood Chronic Poverty and Teen Birth 92
2.7 Association between Childhood Chronic Poverty and Formation of Own Household 93
2.8 Association between Childhood Chronic Poverty and Employment Status . 94
2.9 Association between Childhood Chronic Poverty and Adult Poverty Status . 95
2.10 Equivalence scales by different family sizes 99
2.11 Correlation between Pairs of Poverty Indices 99
2.12 Likelihood of Exiting the PSID Survey 100
2.13 Association between Childhood Chronic Poverty and Adult Outcomes . . . 101
2.14 Association between Childhood Chronic Poverty and Adult Outcomes . . . 102

2.15 Association (unstandardized coefficient) between Childhood Chronic Poverty and Completed Schooling . . . 103
2.16 Association between Childhood Chronic Poverty and Good Health . . . 104
2.17 Association between Childhood Chronic Poverty and Teen Birth . . . 105
2.18 Association between Childhood Chronic Poverty and Formation of Own Household . . . 106
2.19 Association between Childhood Chronic Poverty and Employment Status . . . 107
2.20 Association between Childhood Chronic Poverty and Adult Poverty Status 108

3.1 Share of women legislators and men/women-leader perception (1994–2016) . . . 122
3.2 Share of women legislators and quota systems in Africa (1960–2018) . . . 126
3.3 Aid selectivity in Africa: share of women legislators (1990-2017) . . . 136
3.4 Aid selectivity in Africa: quota-systems (1990-2017) . . . 137
3.5 Aid selectivity in Africa: share of women legislators excluding Rwanda (1990-2017) . . . 139
3.6 Aid selectivity in Africa: share of women legislators—only Sub-Saharan Africa countries (1990-2017) . . . 140
3.7 Variables and their definition . . . 146
3.8 Countries who have implemented Gender Quotas/Reserved Seats in parliaments (Lower Chamber) . . . 147
3.9 How reserved seat quotas are implemented (African countries) . . . 148

Chapter 1

Childhood Chronic Poverty Estimations: Looking Beyond a Count Index

1.1 Introduction

Children are more vulnerable to the effects of household poverty than adults because of its long-run impacts on several outcomes spanning from adolescence to adulthood. Childhood poverty experiences affect cognitive and non-cognitive skill formation including, human capital accumulation, neural development, mental and emotional well being of the child, nutrition of the child, and the ability of parents' to nurture their children (McLoyd, 1990; Evans and Schamberg, 2009; Milligan and Stabile, 2011). Several empirical works have documented the long-run consequences of growing up poor on future adult success. The findings suggest that children with early poverty experiences end up with worse outcomes such as poor cognitive development, low employment opportunities, lower productivity, lower educational attainment, lower income levels, poor adult health, higher propensities of teen births, etc (Gregg et al., 1999; Suryadarma et al., 2009; Evans and Schamberg, 2009; Wagmiller and Adelman, 2009; Isaacs and Magnuson, 2011; Duncan et al., 2012; Ratcliffe and McKernan, 2012; Schoon et al., 2012; Dickerson and Popli, 2016; Ratcliffe and Kalish, 2017).

There is a great deal of research estimating the level of chronic poverty suffered by children. However, the usual approach is mostly based on the time spent poor in conjunction with a duration cut-off fraction, i.e., the count index. This approach does not account for the size of the gaps (i.e., depth of poverty in a given spell) and the temporal pattern of poverty spells experienced by an individual. This can lead to misleading comparisons of chronic poverty levels. For instance, two individuals may spend the same fraction of periods in poverty, but one individual will have poverty experiences occurring very early in life while the other individual experiences poverty later in life. A chronic poverty measure should differentiate between the poverty experiences of such individuals, assigning a greater index to the individual with earlier poverty spell experiences. Empirical evidence suggests that economic hardships experienced in the early years of life has a greater impact on future success than that experienced in later years (Duncan et al., 1998; Duncan et al., 2012; Schoon et al., 2012; Ratcliffe and Kalish, 2017). Moreover, Heckman and Kautz's (2013) review of the evidence on the effectiveness of early intervention programs in promoting later life success also shows that intervention programs before age three led to improved skill formation and improved IQ. Again, for two individuals experiencing the same periods in poverty, the individual with breaks in poverty spells should be treated differently (assigned less poverty) than the individual who experiences the same number of poverty spells uninterrupted by periods of non-poverty.

Approximately 1 in 3 people living in poverty in the United States is a child below the age of 18.[1] Child poverty in the United States affects not only the skills development, coping skills and well-being of children but also costs the economy billions of dollars a year in lost earnings and expenses related to crime and poor quality of health (Holzer et al., 2007). This study describes the chronic poverty experiences of children born in the United States from the late 1960s to the late 1980s, following them from birth through to age 10 using the Panel Study of Income Dynamics (PSID) longitudinal dataset. The alternative measures of chronic poverty examined in this paper are Foster's (2009) measure (abbreviated as F), Bossert et al.'s (2012) measure (abbreviated as BCD) and Hoy and Zheng's (2011) measure (abbreviated as HZ). Each of the measures differs in the way it accounts for the closeness and timing of poverty spells. The F measure emphasizes the number of periods spent in poverty, the BCD measure emphasizes the closeness of poverty experiences by assigning greater weights to poverty spells that occur in a string of two or more poverty spells and the HZ measure emphasizes the timing and closeness of poverty spells by assigning greater weights to earlier poverty spells and poverty spells that occur closer in time to each other. For comparison purposes, I also estimate chronic poverty

[1] Source-https://www.census.gov/library/publications/2018/demo/p60-263.html

levels using the traditional count index. This study is the first to document the poverty experiences of children in a developed nation using these chronic poverty measures.[2] By accounting for poverty gaps together with the timing and spacing of poverty spells, I show that the count index leads to misleading and very conservative estimates of chronic poverty levels for children growing up in the United States.

This paper begins with a description of the chronic poverty measures used in this study, followed by a descriptive analysis of the level of chronic poverty suffered by different groups based on race, age of mother at birth, region of residence, type of household a child is born into, parental educational attainment and experiences of parental marital dissolution. This is relevant because numerous studies have demonstrated the importance of early childhood environments in shaping the abilities of children and the implications for later life outcomes (Cunha et al., 2006; Cunha and Heckman, 2007; Huggett et al., 2011; Shonkoff et al., 2012). The results from this exercise demonstrate how chronic poverty levels for children born in the United States differ among different sub-populations.

The descriptive analysis is followed by a regression analysis of the relationship between childhood chronic poverty and family background characteristics. The results from this descriptive analysis cannot be interpreted as demonstrating causality of poverty. However, it does show patterns in chronic poverty differences across groups. Thus, the regression analysis provides a useful framework to understand these patterns concisely and can help to provide policy makers with important inputs in the design and implementation of anti-

[2] Hoy et al. (2012) analyzed the commonalities and differences which exist between the F, BCD and HZ indices in the measurement of lifetime poverty.

poverty programs.

The remainder of this paper is organized as follows. Section 2 reviews the literature on chronic poverty measurement. Section 3 explains the poverty measures employed in this paper to estimate child chronic poverty levels. Section 4 describes the data and findings from the descriptive and regression analyses. The study ends with a discussion of the implications of this paper's findings for future research.

1.2 Literature Review

How poverty should be defined and measured has received considerable attention among economists, political scientists, sociologists, anthropologists, neuro-scientists and other social scientists. The measure of living standards can be quantitative (income, consumption, and wealth) or qualitative (access to health care, housing, sanitation facilities, information, education, and clothing). To measure a single spell of poverty, i.e., snapshot poverty, the popular FGT (Foster et al., 1984) class of poverty gap measure is frequently adopted. Snapshot poverty focuses on the present standard of living of an individual ignoring the influence of past or future poverty experiences on the current level of hardship suffered (Calvo and Dercon, 2009). In recent years, research has stressed the need to extend snapshot poverty measurement to address both the multidimensionality and lifetime (dynamic) aspects of poverty, with the latter often referred to as chronic poverty.

Chronic poverty measures have been developed based on an aggregation of snapshot poverty levels over a sequence of periods into a single index of poverty. These chronic poverty measures can be broadly classified into two categories; the *permanent-income* approach and the *spells* approach. An early attempt to measure poverty over time is the *permanent-income* approach, which computes an average of all incomes over the lifetime (called the permanent income). It identifies a person as chronically poor if the permanent income is below a corresponding poverty line (Rodgers and Rodgers, 1993; Hill and Jenkins, 1999; Jalan and Ravallion, 2000; Valletta, 2006). The permanent-income approach implicitly assumes that income from non-poor periods will compensate the periods of low income by accounting for the potential saving and borrowing behaviour of individuals over their lifetime. With this approach, chronic poverty is the level of poverty an individual experiences as if his/her income (or consumption) in every period equals their permanent income (or consumption). The second approach, which is the *spells* approach, measures a person's level of chronic poverty by focusing on the distribution of poverty spells over an individual's lifetime (Calvo and Dercon, 2009; Hoy and Zheng, 2011; Bossert et al., 2012; Gradin et al., 2012; Dutta et al., 2013) or time spent in poverty (the count index) or both (Foster, 2009; Alkire et al., 2017).

The recent literature noted above extends the more common, earlier approach to measuring chronic poverty. Earlier work focused on the count index i.e., the number of times spent in poverty within a given period (Duncan et al., 1984; Gaiha, 1989; Gaiha and Deolalikar, 1993; Whelan et al., 2003; Suryadarma et al., 2009; Ratcliffe and McKernan, 2010; Ratcliffe, 2015; Ratcliffe and Kalish, 2017).[3] Identifying chronically poor persons

[3]With the exception of Gaiha (1989), Gaiha and Deolalikar (1993) and Suryadarma et al. (2009), the studies are based on the US PSID dataset.

using this approach can lead to misleading comparisons of the actual impact of individuals' chronic poverty experiences since the size of the poverty gaps and the timing and spacing of poverty spells are not considered. For instance, consider one individual who spends half of his/her lifetime poor and another individual who spends a third of his/her lifetime poor. Suppose the second person experiences larger gaps of poverty.[4] Without accounting for the size of their poverty gaps, the first individual is deemed more economically deprived by the count index. However, accounting for the poverty gaps in addition to the time spent poor can provide evidence that the second person suffers more chronic poverty than the first person. Again, consider two individuals who spend number of periods in poverty. Suppose the first individual has all poverty spells occurring consecutively while the other individual's poverty spells are separated by periods of non-poverty. The poverty experience of the latter is more transitory in nature and hence has less extreme implications on well-being. This study focuses only on chronic poverty measures that differ in terms of how early, close and recurring poverty spells are treated in intertemporal poverty measurement. A more detailed description of these measures is presented in section three below.

In related work on chronic poverty measurement, other scholarly works examine the transitions of poverty in a given population (Bane and Ellwood, 1986; Stevens, 1999; Baulch and Hoddinott, 2000; Jenkins, 2000; Jenkins et al., 2003; Finnie and Sweetman, 2003; Layte and Whelan, 2003; Valletta, 2006; Ratcliffe and McKernan, 2010; Kim, 2019). Relying on poverty and non-poverty spell experiences, they model the mean duration of

[4]Gaiha (1989) show that chronically poor persons (i.e., persons who spend all periods poor) are not necessarily the poorest, i.e., those with wider poverty gaps.

completed poverty spells, identify individuals or groups who are likely to fall in and out of poverty, then compute poverty entry, exit and re-entry rates and their determinants in a given population. The evidence from such analyses assists policy makers in the design of programs targeted at preventing recurring poverty entries and promoting poverty exits instead of the traditional approach of providing monetary benefits to people who are currently poor. A similar analysis is carried out in this paper where I examine which groups of children are more likely to experience chronic poverty.

1.3 Measures of child chronic poverty

This section describes three recently developed measures of chronic poverty used in this paper; the Foster (2009) measure, the Hoy and Zheng (2011) measure, and the Bossert et al. (2012) measure, as well as the count index. The measurement of chronic poverty is taken in two steps. The first step involves identifying who is poor among a given population based on a choice of a poverty criterion (e,g., consumption, income, nutrition levels, etc). In the second step, the poverty experiences of the poor individual are summed into an overall index of chronic poverty.

1.3.1 Notation

First, consider some useful notations. Consider an individual i who lives for T periods. In this study, I estimate the level of chronic poverty suffered by a child in the first ten years of life (i.e., T = 10) and not all T years of life. In each period t = 1, 2,...,T, individual i= 1,2,...,N has a level of income y_i^t. Each period's level of income of the individual is compared with a pre-determined poverty threshold $0 < z^t < \infty$. Individual i is considered poor in period t if his/her income level y_i^t is strictly less than z^t. In any given period t for an individual i, the poverty gap is defined as $G_i^t = z^t - y_i^t$ and the relative poverty gap is given as $g_i^t = \frac{z^t - y_i^t}{z^t}$. From Table 1.7, a family of three needs a minimum of \$8,573 to meet their basic needs in the year 1985. Suppose the annual family income is \$8,000, then the poverty gap G, for the family is \$573 and the relative poverty gap g, is 0.07 for the year 1985.

The Foster et al. (FGT) measure of snapshot poverty in period t for individual i is given as

$$p_i^t = p(y_i^t; z^t) = \begin{cases} \left(1 - \dfrac{y_i^t}{z^t}\right)^\alpha & \text{if } y_i^t < z^t \\ \\ 0 & \text{if } y_i^t \geq z^t \end{cases} \tag{1.1}$$

where the choice of α is typically restricted to values $\{0, 1, 2\}$. α equal to 0 gives the incidence of poverty used in the count index; that is, p_i^t is equal to 1 if y_i^t is below the poverty threshold z^t and 0 if the income y_i^t is at least as large as the poverty threshold z^t. α = 1 gives the size of the normalized poverty gap while α = 2 provides a measure of the intensity of poverty. From the illustrative example above, suppose the family income falls from \$8,000 to \$7,427 such that the poverty gap doubles from \$573 to \$1,146. For α = 2, the size of the poverty gap for a family income of \$8,000 is \$328,329 and the size of the poverty gap for a family income of \$7,427 is \$1,313,316 which is four times the poverty gap of the former. Hence, greater importance is given to poorer individuals with wider poverty gaps for α = 2.

1.3.2 The Count Index

This approach does not address the extent to which a person's income is below the poverty line. It is derived by assigning α equal to 0 in equation 1 such that for each period t, a poverty spell is assigned the value "1" and a non-poverty spell is assigned the value "0". For each individual i, counting the number of $1s$ (i.e., number of periods poor) divided by the total number of periods of observation T gives the fraction of time individual i spent poor. Identifying the chronically poor individual with this approach is based on the fraction of time an individual's income was below the poverty threshold over some specified number of periods. The periods spent in poverty need not be consecutive and there is no consensus regarding the appropriate cut-off fraction (i.e., what fraction of time a person must be in poverty in order to be considered to suffer from chronic poverty). The greater the duration cut-off fraction, the fewer the people considered to be suffering chronic poverty and vice versa.

1.3.3 The Foster (2009) measure

The F measure of chronic poverty uses a dual cut-off spells approach in measuring levels of poverty. First, a person is considered to experience poverty in a given period or spell only if his income falls below the poverty line ($y_i^t < z^t$). Secondly, a person must spend at least some minimum fraction of time in poverty (τ), say for example 30 percent (i.e., τ=0.3), in order to be considered to have experienced chronic poverty. If the second cut-off is met, then each period that a person's income is below the poverty line contributes to the chronic poverty experience according to the size of the relative poverty gap, $\frac{z^t - y_i^t}{z^t}$, in that period.

Formally, for an individual i who spent s out of T periods poor, the F measure of chronic poverty over time is given as

$$P_i^F(y_i^t; z^t) = \begin{cases} \frac{1}{T}\sum_{t=1}^{T} p_i^t & \text{if } s/T \geq \tau \\ \\ 0 & \text{if } s/T < \tau \end{cases} \tag{1.2}$$

where p_i^t is the FGT measure of snapshot poverty described in equation 1. The F measure addresses chronic poverty through the required minimum fraction of time that a person must experience poverty (τ) in order to be classified as having experienced chronic poverty. For persons who meet or exceed this threshold, all poverty spells are assigned equal weights regardless of the temporal pattern; that is, if a person experiences four spells of poverty out of ten periods, the F measure is not sensitive to whether these four periods occurred consecutively, relatively bunched together, or evenly spread out over the ten periods. The other chronic poverty measures described below are sensitive in different ways to the pattern of poverty spells experienced by an individual.

1.3.4 The Bossert et al. (2012) measure

The BCD measure of chronic poverty evaluates the persistence in poverty with a focus on the contiguity of poverty spells. First, a person is considered to experience poverty in a given period or spell only if his income falls below the poverty line ($y_i^t < z^t$). Then for all persons who experience at least one poverty spell, each period that a person's income is below the poverty line contributes to the chronic poverty experience according to the size of the relative poverty gap, $\frac{z^t - y_i^t}{z^t}$, in that period. However, the contribution of a poverty spell that occurs in a string of two or more consecutive poverty spells is greater than the contribution of a poverty spell that occurs in isolation. Additionally, the more poverty spells experienced consecutively, the greater the contribution of each of the poverty spells occurring in that string of contiguous poverty spells to the chronic poverty experience of an individual.

Formally, the BCD measure of chronic poverty over time for an individual i is given as

$$P_i^{BCD}(y_i^t; z^t) = \frac{1}{T}\sum_{t=1}^{T} \gamma^{k-1} p_i^t \tag{1.3}$$

where p_i^t is the FGT measure of snapshot poverty described in equation 1. k is the (maximal) number of consecutive periods including the t^{th} period with a positive poverty gap and γ is a measure of the sensitivity to chronic poverty with $\gamma \geq 1$. If γ equals 1, then the BCD is similar to the F where all poverty spells are assigned equal weights regardless of the sequence of poverty/non-poverty spells in which it occurs. For poverty spells that occur in a string of k consecutive spells, γ^{k-1} is the weight assigned to each of the poverty spell. The BCD measure addresses chronic poverty with a focus on the duration of consecutive poverty spells; that is, for any given γ > 1, the contribution of each poverty spell that occurs in a string of four consecutive spells is greater than the contribution of each poverty spell that occurs in a string of two consecutive spells. The BCD measure is not sensitive to the "closeness" of poverty spells if the poverty spells are not consecutive. The HZ measure which is described below addresses poverty spells which are non-consecutive but occur closely in time to each other.

1.3.5 The Hoy and Zheng (2011) measure

The HZ measure addresses both chronic poverty and early poverty concerns with a set of weights applied to poverty spells. A person is considered to experience poverty in a given period or spell only if his income falls below the poverty line (y_i^t < z^t). Then for all persons who experience at least one poverty spell, each period that a person's income is below the poverty line contributes to the chronic poverty experience according to the size of the relative poverty gap, $\frac{z^t - y_i^t}{z^t}$, in that period. The contribution of a poverty spell that occurs earlier in life is greater than the contribution of a poverty spell that occurs later. Also, poverty spells that occur closer in time to each other have a greater contribution to the chronic poverty experience of an individual compared with poverty spells that are interrupted with many periods of non-poverty.

Formally, the HZ measure of chronic poverty for an individual i is given as[5]

$$P_i^{HZ}(y_i^t; z^t) = \sum_{t=1}^{T} \beta(t,T) p_i^t \tag{1.4}$$

where $\beta(t,T)$ are weights assigned to per-period poverty and non-poverty spells. To satisfy the early poverty axiom of the HZ measure, the weight function must be decreasing in t. Likewise, the weight function must be concave in t to satisfy the chronic poverty axiom. Consider the weight function below:

$$\beta(t,T) = \left(1 - \frac{t}{T+1}\right)^{\delta}, \quad 0 \leq \delta \leq 1 \tag{1.5}$$

The weights are normalized to sum to one; that is, each period $t's$ weight is obtained by dividing $\beta(t,T)$ by the sum of the weights over time $\sum_{t=1}^{T} \beta(t,T)$. The parameter δ is a measure of the sensitivity to early poverty and chronic poverty. The smaller (bigger) the value of δ, the lesser (greater) the contribution of an early poverty spell and the greater (lesser) the contribution of poverty spells that occur closely to each other. Figure 1.1 in the

[5] Hoy and Zheng's (2011) measure has an additional component which involves the level of poverty suffered by an individual over his/her lifetime, where the average income over an entire period is compared with a corresponding poverty line. One can choose the weight on this term to be zero which is what this study adopts.

appendix displays the shapes of this weight function for δ = 0.2, 0.5 and 1.0. There is a trade-off between early and chronic poverty concerns that the HZ measure accommodates. Earlier poverty experience in a sense takes priority over closeness of poverty spells; that is, if a person over four periods has a normalized poverty gap profile (3/5, 0, 3/5, 0) and another person has a normalized poverty gap profile (0, 3/5, 3/5, 0), the HZ chronic poverty level of the former will be greater than the HZ chronic poverty level of the latter because of the earlier poverty experience. This treatment is different from the F and BCD measures. The second person will have the greater BCD chronic poverty level while both persons will have equal F chronic poverty levels. The HZ measure is similar to the BCD measure in a sense that a person with a normalized poverty gap profile (0, 3/5, 3/5, 0) will be assigned a greater chronic poverty level compared with a person with a normalized poverty gap profile (3/5, 0, 0, 3/5). Note that compared to the former profile, the latter profile has the poverty spells spread out one period earlier and one period later. Although the latter profile has an earlier poverty spell, the spreading out of poverty spells takes precedence and so the latter contains less chronic poverty according to the axioms of the HZ.

Often researchers are interested in obtaining a population measure of poverty, i.e., an aggregate measure of poverty for a group of individuals (e.g., all households with non-white heads or all households from a given region). This is useful to examine the extent of chronic poverty suffered by different groups within a population and to design policies targeted at them. For a population of size N, the individual chronic poverty levels (P_i^F, P_i^{BCD}, P_i^{HZ}) are summed across all N individuals to arrive at an overall measure of aggregate chronic poverty. The aggregate chronic poverty measure for a population of size N is given as:

$$P(y;z) = \frac{1}{N}\sum_{i=1}^{N} P_i \tag{1.6}$$

where P_i is the chronic poverty level for individual i. Unless otherwise stated, α in equation 1 is equal to 1 in this paper.

1.4 Results and Discussion

1.4.1 Data and Sample Statistics

The analyses are based on 30 years of data from the Panel Study of Income Dynamics (PSID) of the United States. PSID is an ongoing study which began in 1968 with a nationally representative sample of 5000 families consisting of 15,000 individuals. This survey has followed siblings and parents for over fifty years spanning from 1968 collecting information such as educational attainment, labor supply, health, income levels, household wealth, among others, on all household members annually.

The poor are identified in this study using the official definition of poverty in the United States. A child is described as poor in any period if his/her equivalent household income is below the appropriate poverty threshold. Total family income in the PSID is defined as the sum of taxable income from all sources of the husband, wife and all other earners in the household plus transfer incomes. Poverty thresholds in the United States are determined by the Census Bureau based on money income before taxes (excluding capital gains or non-cash benefits). They vary by family size and composition but not geographically and they are updated annually for inflation using the Consumer Price Index for All Urban Consumers (CPI-U). The equivalent household income is derived using equivalence scales implied by the poverty thresholds published by the US Census Bureau.[6] For instance, a family size of one compared to a family size of four implies an equivalence scale of 2 (refer to Table 1.7 in the appendix).

It is often argued that the standard of living should be measured quantitatively using consumption because it reflects the amount of goods and services purchased by an individual or household. However, data on consumption is not readily available, and in instances where the data is present, it is unlikely to be accurate if the information is collected using surveys. This is because, collection of the data is based on recall and it is generally easier for people to remember earned incomes in a previous year or month than the amount of money they spend and the forms of expenses made. Therefore, in the absence of reliable data on consumption, researchers rely on income as a proxy. Although high incomes imply more consumption of goods and services, low incomes may not necessarily imply low consumption levels because of prior savings and/or access to tax credits and child benefits which boost consumption levels. In the PSID data, total income is comprised of earned income, investment income and transfer income making it an appropriate alternative for consumption.

[6] It is worth noting that the poverty thresholds vary by the number of children in the household. However, it does not vary by the gender of the children in the household. Therefore, it makes the assumption that resources are shared equally amongst boys and girls.

I follow children from birth to age ten. Only children with complete information on their first 10 years of life are included. If a household head withdraws from the survey, then information on every member of the household (including children) becomes unavailable. However, children who move from their original households to other households due to a variety of reasons are still followed and their information is available. Overall, the sample includes 5,557 children over 21 birth cohort periods from 1968 to 1988.[7] For instance, a child aged 1 in 1968 will be followed from 1968 to 1977, a child aged 1 in 1969 will be followed from 1969 to 1978,..., a child aged 1 in 1988 will be followed from 1988 to 1997.[8] Each child has a unique person ID that is used to identify them in each year that they are present in the survey. Hence, I am able to merge the information of a child over the ten year period using that unique person identifier from the yearly PSID release data. In addition, every family that a child resides in has a unique family identifier. Therefore, I am able to connect multiple children who were born in different years to a single family.

Differences in chronic poverty experiences for children belonging to varying sociodemographic groups are investigated. The characteristics (obtained at time of birth of the child) include race of the household head, marital status and gender of the household head, age of mother, age of the household head, level of education of the household head and region of residence. Throughout the ten-year period, the family structure and household

[7]The number of birth cohorts studied is dictated by the availability of yearly data from the PSID. After the year 1997, PSID interviews are conducted every two years. PSID investigators in a technical paper, Andreski et al. (2008) caution researchers about the use of income data from two years ago collected in their biennial interviews. Beginning in 2003, PSID stopped asking respondents about incomes from two years ago due to great recalling error. Hence, data after 1997 are not included in this study.

[8]New-born children up to the age of 1 are assigned age 1 in the PSID data set.

heads for some children changed due to parental marital separation or divorce.[9] Thus, the difference in chronic poverty levels of children who encountered parental marital dissolution (classified as switchers in this study) and children growing up in stable households (classified as non-switchers in this study) is explored.[10]

On classification of race, 3,359 children (representing 60 percent) are born into households with white heads while 2,198 children (representing 40 percent) are born into households with non-white heads (i.e., Blacks, Hispanics, Asians, etc) in the sample. 4,090 children (representing 74 percent) have household heads who completed at least 12th grade while 1,467 children (representing 26 percent) have heads who completed less than 12th grade. Children are grouped according to the age of household head and age of mother as of the date of birth. 3,463 (representing 62.3 percent), 1,912 (representing 34.4 percent) and 182 (representing 3.3 percent) children have household heads aged in the category 26+ years, 20-25 years and 16-19 years respectively while 2,664 (representing 48 percent), 2,443 (representing 44 percent) and 450 (representing 8 percent) children are born to mothers in the age categories 26+ years, 20-25 years and 16-19 years respectively. On the basis of region of residence classification, 837 children (representing 15.1 percent) reside in the North-East Region, 1,423 children (representing 25.6 percent) reside in the North-Central Region, 2,451 children (representing 44.1 percent) reside in the South Region and 846

[9]Other causes of household switches found in the data include death of parent(s) and/or move-ins with a relative. I focus on parental divorces because it constitutes about 95 percent of the disruptions in the initial household environment of children.

[10]A switcher is a child who at time of birth resided in a two-parent household, but due to a parental divorce or separation had to live with a single parent at any point during the first ten years of life. On the other hand, a non-switcher is a child who remained in the same type of household that they were born into (whether a one-parent or two-parent household) over the first ten years of their life.

children (representing 15.2 percent) reside in the West Region. Further, 935 children (representing 17 percent) in the sample experienced disruptions in the household environment due to parental divorce and separation. Out of the 5,557 children, 4,703 are born into two parent households while 854 children are born into lone parent households, bringing the percentage of children living with both parents to 85 percent in the sample.

1.4.2 Results: Child Chronic Poverty levels by background characteristics

1.4.2.i Chronic poverty levels using the Count Index

In this subsection, the chronic poverty levels as estimated by the different poverty measures are provided for the different sub-populations. For the count index, three definitions of chronic poverty are analyzed; poor at least three out of ten periods, poor at least five out of ten periods and poor at least eight out of ten periods. The chronic poverty rates are presented in Table 1.1. For all cohorts, 24.6 percent of the children spent at least three out of ten periods poor, 17.3 percent spent at least five out of ten periods poor whilst 9.5 percent spent at least eight out of ten periods poor. Across all columns, children living in lone parent households suffer more chronic poverty compared with children living with both parents. Whilst 40.2 percent of children born into lone parent households spent at least 8 periods poor, only 3.9 percent of children born into two parent households spent the same number of periods poor. Moreover, children with non-white household heads suffer greater chronic poverty compared with children with white household heads regardless of the sex of the household head. For female headed households, 66.2 percent of non-whites spent at least five periods poor whilst 39 percent of whites spent at least five periods poor. For male headed households, 20.5 percent of non-whites spent at least five periods poor whilst 4.7 percent of whites spent at least five periods poor.

Children with more educated heads suffer less chronic poverty compared with their counterparts whose parents have lower educational levels. 5.1 percent of children whose household heads have at least a high school degree spent at least eight periods poor whilst 21.8 percent of children whose household heads have less than a high school degree spent the same number of periods poor. For regional comparisons, children residing in the South Region suffer the greatest chronic poverty followed by children residing in the North Central Region and then children residing in the West Region with children residing in the North East Region having the least chronic poverty experience. For instance, 22.2 percent of children residing in the South Region, 17.5 percent of children residing in the North Central Region, 10.3 percent of children residing in the West Region and 9.8 percent of children residing in the North East Region spent at least five periods poor. The percentage of switchers who suffer chronic poverty is higher than the percentage of non-switchers who suffer chronic poverty. Whilst 23.4 percent of children who experienced parental marital dissolution spent at least five periods poor, 16.1 percent of children with no such experience spent at least five periods poor. However, the rank between the two groups is reversed for the cut-off fraction 8/10, which indicates that switchers do not spend a lot of time in poverty as they start off from a two parent household and transition into a lone parent household due to parental divorce. Children born to teen parents are found to suffer greater chronic poverty compared with children born to parents aged 20 years and above for all duration cut-off fractions. Whilst 29.6 percent of children born to teen mothers spent five or more periods poor, 19.7 percent of children born to mothers aged 20-25 years and 13.1 percent of children born to mothers aged 26+ years spent the same number of periods poor.

1.4.2.ii Chronic poverty levels using the F, BCD and HZ Indices

The F measure emphasizes the number of periods spent in poverty without taking into account different temporal patterns of poverty experiences. However, unlike the count index, it takes into account the size of the poverty gap for each year of poverty experienced. The chronic poverty levels according to the F measure for the different sub-populations of children are reported in Table 1.2. The first column provides the level of chronic poverty for the duration cut-off of τ = 0.00, where all children who spend at least one period poor over the ten-year period are assigned a positive F chronic poverty index. In the second column, τ equals 0.25 and all children who spend at least three out of ten periods poor are assigned a positive F chronic poverty index. Children who spend five or more periods poor are assigned a positive F chronic poverty index in the third column for τ = 0.50 whilst children who spend eight or more periods poor are assigned a positive F chronic poverty index in the fourth column for τ = 0.75. In the fifth column, children who spend all ten periods in poverty are assigned a positive F chronic poverty index for τ = 1.00.

The comparison of the degree of F chronic poverty levels between groups for the most part is similar to the findings from the count index with a few exceptions. Across all duration cut-off fractions, children born in the South Region have the highest F chronic poverty indices, followed by children born in the North-Central Region and children born in the North-East Region. Children born in the West Region have the least F chronic poverty index even though they suffered more chronic poverty than children living in the North-East Region according to the count index. This is an indication that, compared with children in the West Region, children living in the North-East Region have relatively deep poverty gaps in the few years they spent poor. The chronic poverty levels for children born to teen parents is greater than the chronic poverty levels for children born to parents aged 20 years and above, however, the rank reverses for τ equal 1.00. Thus, children born to teen parents who spend all ten periods poor have relatively smaller poverty gaps compared with children born to older parents who spend all ten periods poor. The results that switchers suffer less chronic poverty than non-switchers in some cases with the count index is reinforced with the F index. Switchers have the lowest chronic poverty level for τ = 0.75 (i.e., spend at least eight periods poor) and τ = 1.00 (i.e., spend all ten periods poor).

Chronic poverty levels are estimated by the BCD measure for different γ values; 1.0, 1.2, 1.5, 1.7 and 2.0. These are presented in columns 1, 2, 3, 4 and 5 of Table 1.3 respectively. The BCD measure emphasizes the closeness of poverty spells by assigning greater weights to poverty spells that occur in a string of two or more consecutive periods. However, unlike the HZ measure, the BCD measure applies no additional weight to poverty spells that are closer together unless they are contiguous. For a given string of two or more consecutive poverty spells, the larger is the parameter γ, the greater the contribution of each of those poverty spells to the BCD chronic poverty level of an individual. All children who spend at least one period poor are assigned a positive BCD chronic poverty index for all values of γ. In the first column, γ equals 1.0 and all poverty spells are assigned equal weights regardless of when and how they occur.

Like the F index, the comparison of the degree of BCD chronic poverty levels between groups is similar to the findings from the count index with a few exceptions. When all poverty spells are assigned equal weights in column 1, children who experienced parental marital separation have higher chronic poverty levels than children with no such experience. However, as continuous poverty spells are assigned greater weights, the rank between the two groups is reversed, indicating that the poverty spells of switchers may be less contiguous compared to non-switchers. Children born to teen parents suffer the most BCD chronic poverty for γ = 1.0 and γ = 1.2 only. Therefore, the impact of consecutive poverty spell experiences for teen parents implies less chronic poverty for the BCD index when the impact parameter gamma is higher. Note that the difference in the BCD chronic poverty levels between some groups (e.g., lone parent vs two parent households, white vs non-white households, less educated household heads vs more educated household heads) is heightened as consecutive poverty spells receive greater emphasis (i.e., as γ increases).

Chronic poverty levels are estimated by the HZ measure for five different δ values; 0.0, 0.2, 0.5, 0.8 and 1.0. They are presented in columns 1, 2, 3, 4 and 5 of Table 1.4 respectively. The HZ measure emphasizes the closeness of poverty spells by assigning greater weights to poverty spells that occur closer in time to each other without the requirement that poverty spells be consecutive as required by the BCD measure. It also emphasizes another aspect of the timing of poverty spells by assigning greater weights to poverty spells that occur earlier in life. In column 1, δ equals 0.0 and the HZ measure has zero sensitivity to early or close poverty spells. As the value of δ is increased, the HZ measure becomes more sensitive to early poverty spells and less sensitive to close poverty spells. Again, rankings between groups based on HZ poverty levels are mostly consistent with the findings from the count index.

The HZ chronic poverty levels for children born to parents aged 20 years and above is not very sensitive to the value of the parameter δ while that for children born to teen parents increases as sensitivity to early poverty spells increases and sensitivity to close poverty spells decreases. The result that children of teen parents experience poverty much earlier in life is not surprising because teen parents are more likely to be in school and less likely to be gainfully employed at the time of birth. When the weight on late poverty spells is increased and closeness of poverty spells becomes more important, the HZ chronic poverty index for switchers (i.e., children who experienced parental divorces) increases while that of non-switchers decreases. This provides evidence in support of the economic hardships that children face following parental divorce. This result is consistent with Valletta (2006) who finds that more than one third of poverty entries in the United States is due to divorces. As sensitivity to early poverty spells increases, the HZ chronic poverty index for children with female household heads increases while that of children with male household heads decreases regardless of race. This indicates that compared with the poverty experiences of children with male household heads, the poverty experiences of children with female household heads occur much earlier. Across all the chronic poverty measures, it is observed that non-whites suffer more chronic poverty than whites. The results reveal that the difference between the two groups is especially prevalent for households with male heads than for households with female heads even though the chronic poverty levels for households with female heads are much higher.

Additional tests reveal that the difference in the estimated chronic poverty levels between different groups of children are statistically significant (at 1 percent and 5 percent) with a few exceptions. These are reported in the appendix (Tables 1.8, 1.9 1.10). For instance, the difference in the estimated chronic poverty levels between children growing up in the West Region and children growing up in the North-Central Region are not statistically significant for all chronic poverty measures. In addition, the difference in the estimated BCD chronic poverty levels between children born to teen parents and children born to parents aged 26 years and above are not statistically significant.

Table 1.1: Childhood Chronic poverty estimated by the Count Index (percent)

	Duration cut-off fraction		
	3/10	5/10	8/10
By Household Type			
Two-parent (4,703)	15.88	9.67	3.89
Lone-parent (854)	72.72	59.48	40.16
By Age of Household Head			
26+ (3,463)	18.51	12.99	6.87
20-25 (1,912)	32.74	23.27	13.08
16-19 (182)	55.49	37.36	20.88
By Age of Mother			
26+ (2,664)	18.28	13.14	7.51
20-25 (2,443)	27.88	19.65	10.56
16-19 (450)	44.44	29.56	15.11
By Household Head's Education			
Less than 12th grade (1,467)	48.88	36.54	21.75
At least 12th grade (4,090)	15.92	10.44	5.06
By Race of Household Head			
White (3,359)	10.90	6.28	2.26
Non-white (2,198)	45.59	34.21	20.47
By Race and Sex of Household Head			
White female (154)	54.55	38.96	16.88
Non-white female (660)	78.94	66.21	46.36
White male (3,205)	8.80	4.71	1.56
Non-white male (1,538)	31.27	20.48	9.36
By Region of Residence			
North-East (837)	14.58	9.80	4.90
North-Central (1,423)	24.31	17.50	10.40
South (2,451)	31.25	22.24	12.16
West (846)	15.84	10.28	4.61
By Household Switches			
Non-switchers (4,622)	22.13	16.10	9.48
Switchers (935)	36.90	23.42	9.41
All Cohorts (5,557)	24.62	17.33	9.47

Notes: Sample sizes are in brackets.

Source: Author PSID 1968-1997.

Table 1.2: Childhood Chronic poverty estimated by the F measure

	τ				
	0.00	0.25	0.50	0.75	1.00
By Household Type					
Two parent (4,703)	0.040	0.034	0.027	0.015	0.006
Lone parent (854)	0.262	0.256	0.239	0.193	0.124
By Age of Head					
26+ (3,463)	0.054	0.049	0.043	0.030	0.017
20-25 (1,912)	0.101	0.095	0.084	0.061	0.036
16-19 (182)	0.157	0.146	0.120	0.081	0.022
By Age of Mother					
26+ (2,664)	0.055	0.051	0.045	0.033	0.020
20-25 (2,443)	0.085	0.079	0.069	0.040	0.029
16-19 (450)	0.120	0.112	0.093	0.060	0.013
By Head's Education					
Less than 12th grade (1,467)	0.149	0.142	0.128	0.095	0.053
At least 12th grade (4,090)	0.047	0.041	0.035	0.023	0.013
By Race and Sex of Head					
White female (154)	0.158	0.148	0.128	0.073	0.041
Non-white female (660)	0.295	0.289	0.273	0.226	0.147
White male (3205)	0.023	0.017	0.013	0.006	0.002
Non-white male (1,538)	0.077	0.071	0.058	0.035	0.015
By Region of Residence					
North-East (837)	0.037	0.033	0.028	0.018	0.008
North-Central (1,423)	0.076	0.071	0.062	0.047	0.029
South (2,451)	0.098	0.092	0.082	0.058	0.033
West (846)	0.034	0.027	0.023	0.013	0.004
By Household Switches					
Non-switchers (4,622)	0.070	0.064	0.058	0.043	0.027
Switchers (935)	0.094	0.084	0.069	0.036	0.010
All Cohorts (5,557)	0.074	0.068	0.060	0.042	0.024

Notes: Sample sizes are in brackets.

Source: Author, PSID 1968-1997.

Table 1.3: Childhood Chronic poverty estimated by the BCD measure

	γ				
	1.0	1.2	1.5	1.7	2.0
By Household Type					
Two parent (4,703)	0.040	0.096	0.416	1.079	4.025
Lone parent (854)	0.262	0.960	5.939	17.361	71.703
By Age of Head					
26+ (3,463)	0.054	0.166	0.913	2.582	10.387
20-25 (1,912)	0.101	0.326	1.858	5.309	21.540
16-19 (182)	0.157	0.389	1.715	4.448	16.538
By Age of Mother					
26+ (2,664)	0.055	0.179	1.028	2.951	12.032
20-25 (2,443)	0.085	0.269	1.517	4.319	17.495
16-19 (450)	0.120	0.302	1.302	3.307	11.937
By Head's Education					
Less than 12th grade (1,467)	0.149	0.490	2.805	8.000	32.359
At least 12th grade (4,090)	0.047	0.135	0.713	1.996	7.994
By Race and Sex of Head					
White female (154)	0.158	0.416	2.075	5.774	23.22
Non-white female (660)	0.295	1.114	6.993	20.50	84.822
White male (3,205)	0.023	0.047	0.170	0.412	1.443
Non-white male (1,538)	0.077	0.209	1.007	2.706	10.392
By Region of Residence					
North-East (837)	0.037	0.099	0.486	1.325	5.162
North-Central (1,423)	0.076	0.248	1.423	4.088	16.678
South (2,451)	0.098	0.314	1.763	5.008	20.214
West (846)	0.034	0.078	0.325	0.829	3.033
By Household Switches					
Non-switchers (4,622)	0.070	0.229	1.330	3.828	15.634
Switchers (935)	0.094	0.227	0.943	2.363	8.452
All Cohorts (5,557)	0.074	0.229	1.265	3.581	14.426

Notes: Sample sizes are in brackets.

Source: Author, PSID 1968-1997.

Table 1.4: Childhood Chronic poverty estimated by the HZ measure

	δ				
	0.0	0.2	0.5	0.8	1.0
By Household Type					
Two parent (4,703)	0.040	0.039	0.039	0.038	0.038
Lone parent (854)	0.262	0.265	0.269	0.273	0.275
By Age of Head					
26+ (3,463)	0.054	0.054	0.054	0.054	0.054
20-25 (1,912)	0.101	0.102	0.102	0.103	0.103
16-19 (182)	0.157	0.159	0.162	0.165	0.168
By Age of Mother					
26+ (2,664)	0.055	0.055	0.055	0.055	0.055
20-25 (2,443)	0.085	0.086	0.086	0.086	0.086
16-19 (450)	0.120	0.121	0.122	0.123	0.124
By Head's Education					
Less than 12th grade (1,467)	0.149	0.150	0.151	0.152	0.153
At least 12th grade (4,090)	0.047	0.047	0.047	0.046	0.046
By Race and Sex of Head					
White female (154)	0.158	0.161	0.164	0.168	0.171
Non-white female (660)	0.295	0.298	0.303	0.307	0.310
White male (3,205)	0.023	0.023	0.022	0.022	0.022
Non-white male (1,538)	0.077	0.076	0.075	0.074	0.073
By Region of Residence					
North-East (837)	0.037	0.037	0.037	0.037	0.037
North-Central (1,423)	0.076	0.076	0.076	0.076	0.076
South (2,451)	0.098	0.099	0.099	0.100	0.100
West (846)	0.034	0.034	0.034	0.034	0.034
By Household Switches					
Non-switchers (4,622)	0.070	0.070	0.071	0.072	0.073
Switchers (935)	0.094	0.092	0.089	0.085	0.084
All Cohorts (5,557)	0.074	0.074	0.074	0.074	0.075

Notes: Sample sizes are in brackets.

Source: Author, PSID 1968-1997.

1.4.3 Results: Descriptive Regression Analysis

In this section, I investigate whether there are conditional correlations between background characteristics such as household type, age of mother, household head's education, race, region of residence, and parental divorces and early childhood chronic poverty. For instance, 45.6 percent of non-whites experienced at least one poverty spell over the ten-year period whilst 10.9 percent of whites experienced at least one poverty spell over the ten-year period. Moreover, the fraction of non-whites in the entire sample is 40 percent compared to 57 percent of non-whites in the South Region. It is therefore not obvious from the descriptive analysis if the high chronic poverty estimates for children in the South Region is driven entirely by the higher proportion of non-white individuals in that region. Therefore I model in a single regression, the level of chronic poverty (estimated by each chronic poverty measure) as a function of background characteristics. This exercise shows which family characteristics are most highly associated with a particular chronic poverty measure. This is useful in identifying who suffers chronic poverty but not for determining causes of chronic poverty.

In all regression models, standard errors are clustered at the household level to correct for multiple children born into the same household. The test results are presented in Table 1.5. In column 1, the dependent variable is a dummy variable equal to 1 if a child spends five or more periods in poverty and 0 otherwise. In column 2, the dependent variable is the chronic poverty level of a child estimated by the F measure for the duration cut-off fraction of 0.50. The dependent variable in column 3 is the chronic poverty level of a child estimated by the BCD measure for the parameter γ equal 1.5. In the last column, the dependent variable is the chronic poverty level of a child estimated by the HZ measure for the parameter δ equal 0.5.

The results reveal that the type of household a child is born into, race of household head, educational attainment of household head and changes in family structure are key predictors of childhood chronic poverty. A child born into a lone parent household is more likely to experience chronic poverty compared with a child born into a two parent household. From column 3, the predicted BCD chronic poverty index for children in lone parent households is 4.8 points higher than the predicted BCD chronic poverty index for children in two parent households. Children whose parents completed 12 grades of school or more are less likely to experience chronic poverty whilst children with non-white household heads are more likely to experience chronic poverty. The predicted HZ chronic poverty index for children with non-white heads is 0.05 points higher than children with white heads whilst the predicted F chronic poverty index for children with more educated heads is 0.05 points lower than the predicted F chronic poverty level for children with less educated heads.

Compared with children residing in the North East Region, children residing in the West Region are less likely to experience chronic poverty whilst children residing in the South Region and North Central Region are more likely to experience chronic poverty. Compared with children residing in the North East Region, the predicted BCD chronic poverty index for children residing in the South Region is 0.5 points higher whilst the predicted BCD chronic poverty index for children residing in the West Region is 0.3 points lower. Children who experienced a disruption in the family environment due to parental divorces are found to be more likely to experience chronic poverty. Specifically, the predicted F chronic poverty index for switchers is 0.04 points higher than the predicted F chronic poverty index for non-switchers (i.e., children growing up in stable households). Children with teen mothers are more likely to experience chronic poverty for the count index, however, their predicted BCD chronic poverty index is 0.63 points lower compared to children born to mothers aged 26 years and above. This further reiterates the findings from the descriptive analysis that the poverty experience of children born to teen mothers are less contiguous.

The results reveal the independent associations between being non-white and growing up in the South Region to childhood chronic poverty experiences. The coefficient of non-white is statistically significant at 1 percent for all chronic poverty measures while the coefficient of the South Region is significant at 1 percent for all measures except for the count index (significant at 10 percent). Therefore, growing up in the South Region and being non-white are both individually associated with chronic poverty, with race having the greater size associations.

1.4.4 Equally-Distributed-Equivalent (EDE) poverty gaps

The HZ poverty index is an ordinal measure of chronic poverty as are the other measures presented in this paper. From Table 1.4, the HZ Index for the two parent households is

Table 1.5: Association between childhood chronic poverty and background characteristics

	Chronic poverty measure			
	Count Index	F Index	BCD Index	HZ Index
Household Type (Two parent)				
Child is born into a lone parent household	2.475***	0.186***	4.830***	0.201***
	(0.149)	(0.012)	(0.466)	(0.011)
Age of mother (26+ years)				
Mother is aged 20-25 years	0.156*	0.002	-0.042	0.006*
	(0.115)	(0.005)	(0.154)	(0.004)
Mother is aged 16-19 years	0.458***	0.006	-0.632***	0.021***
	(0.180)	(0.009)	(0.229)	(0.008)
Head's Education (Less than 12th grade)				
Head completed at least 12th grade	-1.204***	-0.049***	-1.076***	-0.055***
	(0.119)	(0.007)	(0.227)	(0.006)
Race of Head (White)				
Head is Non-white	1.327***	0.042***	1.013***	0.045***
	(0.149)	(0.006)	(0.174)	(0.006)
Region (North-East)				
Child resides in the North-Central Region	0.494***	0.022***	0.670***	0.024***
	(0.249)	(0.007)	(0.214)	(0.007)
Child resides in the South Region	0.261*	0.015***	0.419***	0.019****
	(0.224)	(0.006)	(0.197)	(0.006)
Child resides in the West Region	-0.279	-0.013***	-0.285**	-0.013***
	(0.287)	(0.006)	(0.158)	(0.006)
Change in family structure (Non-switcher)				
Child is a switcher	1.278***	0.039***	0.432***	0.046***
	(0.147)	(0.006)	(0.133)	(0.005)
Constant	2.820***	0.032***	0.598***	0.042***
	(0.236)	(0.007)	(0.229)	(0.007)
Regression Statistics				
N	5557	5557	5557	5557
R-squared	0.328	0.326	0.218	0.392
P value	0.000	0.000	0.000	0.000

Notes: The dependent variables are the different chronic poverty measures. Cut-off fraction is 5/10 for the count index. τ = 0.50 for F index; γ = 1.5 for BCD index and δ = 0.5 for HZ index. Robust standard errors in parentheses are clustered at the household level. Column 1 presents the log odds-ratio and standard errors from a logistic regression. Columns 2 to 4 presents OLS marginal effects and standard errors. Omitted categories are in parentheses. The R-squared from the logistic regression (Pseudo R-squared) and OLS regressions (adjusted R-squared) are not directly comparable. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID 1968 - 1997

0.269 while that of lone parent households is 0.039 (for δ=0.5). Even though the index of the latter group is about seven times the index of the former, it is not clear to what extent one can say that children born in lone parent households experience seven times the chronic poverty experienced by children born in two parent households. Therefore, it is useful to create a measure of poverty that allows for comparisons between groups in a cardinal sense. To do this, I follow the approach of Hoy et al. (2012) to define a money metric cost of poverty, i.e., the equally distributed equivalent (EDE) poverty gap.[11] The EDE poverty gap for any chronic poverty measure is the size of poverty gap which, if distributed equally to all individuals and in all periods of life, would produce the same level of poverty for that measure as is generated by the actual lifetime profiles of poverty experienced across a population (Duclos et al., 2010).

The equally distributed equivalent poverty gap for a group of size N (who live for T periods) is the hypothetical size of a poverty gap that, if suffered by each person in each period, generates the same amount of poverty for a relevant measure (in this case, F, BCD or HZ) as generated from the actual poverty profiles of the N individuals. The equally distributed equivalent poverty gap measure reflects a common cardinal interpretation of the amount of poverty generated by each index. Suppose the equivalent poverty gap for households with non-white heads is \$600 per person per period and the equivalent poverty gap for households with white heads is \$200 per person per period, then one can conclude that children who reside in households with non-white heads suffer three times as much

[11] Hoy et al. (2012) follow the approach developed by Duclos et al. (2010) which in turn relies on the seminal measurement of the cost of social welfare and inequality developed in Kolm (1969) and Atkinson (1970).

poverty as children who reside in households with white heads. A detailed description of the EDE approach is provided in the appendix.

The EDE poverty gap for different groups generated from the F, BCD and HZ measures are presented in Table 1.6. The EDE values are computed using 1985 poverty lines, hence all poverty lines and incomes are normalized accordingly. Results are provided for only race and household type based on the findings from the descriptive regression analysis that these two characteristics have the greatest size associations with early childhood chronic poverty. From column 2, if every child born into a two parent household were to experience a poverty gap of \$151.75 in each period, it will yield the same F chronic poverty level of 0.027 (τ=0.5) computed from their actual household income profiles over the ten year period. Similarly, the EDE poverty gap generated from the HZ index for children born into lone parent households is \$1,508.17; that is, if every child born into a single parent household were to experience a poverty gap of \$1,508.17 in each period, it would yield the same HZ chronic poverty index of 0.269 (δ=0.5) computed from their actual household income profiles over the ten year period. In column 1, I present the raw average poverty gap over the poverty experiences of all individuals over all periods. Although, the average poverty gap takes into account the depth of poverty in any period, it is not sensitive to the distribution of poverty spells and so does not display the chronic poverty characteristics highlighted in this paper. Including the raw average poverty gap helps to make a sharp comparison of poverty gap differentials across groups for measures that account for chronic poverty versus those that do not.

The EDE poverty gap for the F index is lower than the raw average poverty gap for both household types. However, this should not be interpreted as an indication of less chronic poverty suffered. This is because the EDE poverty gap for the F index (with τ=0.5) is the poverty gap averaged over time for all persons who spent at least five periods in poverty, hence a subset of the overall average poverty gap for all sample children. Likewise, the absolute size of the EDE poverty gap generated by the BCD is typically very small since incurring poverty consecutively in each of the ten periods as is done in the hypothetical computation of the EDE is emphasized by this index. What is of great concern is the ratio of this EDE value between two poverty profiles (i.e., of two household types). The EDE poverty gap measure is a cardinal measure of poverty, thus, the ratio of the poverty gaps are comparable across measures. For example, although the size of the EDE poverty gap is lower when using the BCD measure than for the average poverty gap for both two parent and lone parent households, the ratio of the EDE poverty gap of lone parent to two parent is substantially higher with the BCD index (ratio equals 13.70 compared to 6.09 for the average poverty gap measure). From the last column of Panel A, it is observed that children born into lone parent households suffer seven times as much poverty as children born into two parent households.

When accounting for chronic poverty, the non-white to white poverty differential is significantly increased as seen across columns 2 to 4. From panel B, the average raw poverty gap for non-whites is \$1,687.73 while that of whites is \$334.96 yielding a ratio of 5.03. However, the ratio of the poverty gap is increased to 11.79 for the BCD index and 6.89 for the F index. These results provide evidence in support of the need to account for the distribution of poverty spells when estimating levels of chronic poverty to avoid misleading comparisons across groups and individuals.

Table 1.6: Equally Distributed Equivalent (EDE) poverty gaps

		EDE poverty gap		
	Average Poverty Gap	F Index	BCD Index	HZ Index
Panel A: By Household Type				
Two parent	\$499.29	\$151.75	\$66.66	\$214.00
Lone parent	\$3,042.31	\$1,347.42	\$913.42	\$1,508.17
Ratio	6.09	8.88	13.70	7.05
Panel B: By Race of Head				
White	\$334.96	\$97.93	\$36.10	\$155.50
Non-white	\$1,687.73	\$674.89	\$425.643	\$780.61
Ratio	5.03	6.89	11.79	5.02

Notes: EDE poverty gaps are measured in dollars. They are calculated using 1985 poverty lines; poverty lines and incomes are normalized accordingly. τ = 0.50 for the F index; γ = 1.5 for the BCD index and δ = 0.5 for the HZ index.
Source: Author, PSID 1968-1997

1.5 Conclusion

Using data from the PSID, this paper compares the child chronic poverty rates for the first ten years of children's lives in the United States using three recently developed chronic poverty measures that account for the gaps and distribution of poverty, i.e., the Foster (2009) measure, the Bossert et al. (2012) measure and the Hoy and Zheng (2011) to those obtained by the widely used count poverty index. The Foster (2009) measure emphasizes the number of periods spent in poverty without taking into account different temporal patterns of poverty experiences. However, unlike the count index, it takes into account the size of the poverty gap for each year of poverty experienced. The Bossert et al. (2012) measure emphasizes the closeness of poverty by assigning greater weights to poverty spells that occur in a string of two or more poverty spells and the Hoy and Zheng (2011) measure emphasizes the timing and spacing of poverty spells by assigning greater weights to earlier poverty spells and poverty spells that occur closer in time to each other. Applying these chronic poverty measures to a sample of children born in the United States (between 1968 and 1988) who are followed from birth to age 10, this study highlights the role that assumptions made regarding the treatment of the timing, spacing and severity of poverty play in addressing the question of who suffers more chronic poverty. Chronic poverty levels are compared between groups of children based on race, age of mother at birth, region of residence, household head's educational attainment, type of household a child is born into (i.e., lone parent or two parent) and whether a child's parents got divorced or separated.

For the most part, rankings between groups based on chronic poverty levels are consistent across the various poverty measures studied. Children born into lone-parent households suffer more chronic poverty compared with children born into two-parent households; children residing in the South Region suffer more chronic poverty compared with children living in other Regions; and children with less educated heads (i.e., less than a high school degree) suffer more chronic poverty compared with children with educated heads. Non-whites suffer more chronic poverty than whites and the difference is significantly increased when the timing and spacing of poverty spells are accounted for. Put differently, excluding information about the gaps and distribution of poverty leads to very conservative estimates of differences across groups based on chronic poverty levels.

Rather surprisingly, children born to teen parents have higher chronic poverty levels than those with older parents at birth only when poverty gaps are unaccounted for or all poverty spells are treated equally regardless of the temporal pattern of poverty/non-poverty spells. Children residing in the West Region have lower chronic poverty levels than those residing in the North-East Region, except for when poverty gaps are unaccounted for in the calculation of chronic poverty levels (i.e., in the case of the count index). Finally, this study determines whether there are independent associations between family background characteristics and early childhood chronic poverty. Such an exercise is important for policy considerations. Although this exercise is not a demonstration of causality of poverty, it does highlight possible mechanisms through which chronic poverty occurs for children. The findings imply that the mechanisms vary, with type of household a child is born into and race of household head having the greatest size associations. Further, the results reveal that being non-white and growing up in the South Region have independent associations with childhood chronic poverty.

These findings provide input for policy makers in the design of anti-poverty programs targeted at specific sub-populations in the country. Consider the following example. The United States has one of the highest childhood poverty rates among several Western industrialized countries (Gornick and Jäntti, 2010). In Canada and Denmark, new mothers can expect up to 15 weeks and 52 weeks of paid maternity leave respectively. Sweden has one of the most generous parental leave programs where both the new mother can split up to

120 weeks of paid leave with her partner at 80 percent of the mother's salary. In contrast, there is no law in place requiring businesses and corporations to offer paid maternity leave to their employees in the United States. Therefore, new mothers in the US would benefit from some kind of parental monetary allowance while away from work to lower the risk of slipping into poverty. In addition, support programs like the Earned Income Tax Credits (EITC) should give generous credits to families, especially people of color and single mothers, with very young children.

The suggestions provided in this paper do not imply that the introduction of policies to boost incomes during the early childhood years is the only strategy to combat childhood chronic poverty. Training programs for parents with less than a high school degree as well programs to boost parental employment are also likely to be important. The findings from this research point future empirical work in the direction of childhood chronic poverty estimations that is over and beyond a count approach.

This study is limited by the use of the uni-dimensional income poverty measure for its child poverty analysis. Although the findings are based on a rich longitudinal data set, further work should be conducted to analyze how identifying the poor using a non-monetary multidimensional deprivation approach may affect the results. Such work has demonstrated that the direction of comparisons between groups can depend on both the application and the measures used. For example, Alkire and Foster (2011) find that the percentage of poor African Americans decreased while the percentage of poor Hispanics increased when the definition of poverty is switched from an income based measure to a multidimensional non-income poverty measure. Main and Bradshaw (2012), also identifying deprived children (ages 8-14) in the UK using subjective poverty measures, find deprived children in some income non-poor families and non-deprived children in some income poor families. Chzhen et al. (2016) and Qi and Wu (2019) find that multidimensional deprivation rates were higher than income poverty rates for most children in China and the United Kingdom respectively, indicating that the income poverty measure may underestimate the severity of child poverty.

1.6 Appendix

Equally-Distributed-Equivalent (EDE) poverty gap measure

A detailed description of the EDE approach described in Section 1.4.4 of the main paper is explained below.

Let g represent the vector of poverty gaps associated with an income profile. For two income profiles a = (2, 5, 5, 2), b = (5, 2, 2, 5) and a constant poverty line of 5, the poverty gap profile is given as; g_a = (3, 0, 0, 3) and g_b = (0, 3, 3, 0). Let $\bar{\mathbf{g}}$ be the vector formed by inserting the average poverty gap in each period of life hence, $\bar{\mathbf{g}}_a$= $\bar{\mathbf{g}}_b$= (1.5, 1.5, 1.5, 1.5). The average value of the poverty gaps over the four periods is 1.5. The EDE poverty gap is therefore the single valued poverty gap (in this case 1.5) which if incurred in each period of life would yield the same level of poverty as that of the original profiles a and b. Let this value be represented by $\widehat{g}$.

Following Hoy et al. (2011), the EDE poverty gap is obtained by two equations: For the F index,

$$P_F = \widehat{p} \tag{1.7}$$

and

$$\widehat{g} = z(P_F)^{1/\alpha} \tag{1.8}$$

For the BCD index,

$$P_{BCD} = \gamma^{T-1}\widehat{p} \tag{1.9}$$

and

$$\widehat{g} = z\left(\frac{P_{BCD}}{\gamma^{T-1}}\right)^{1/\alpha} \tag{1.10}$$

For the HZ index,

$$P_{HZ} = \widehat{p} \tag{1.11}$$

and

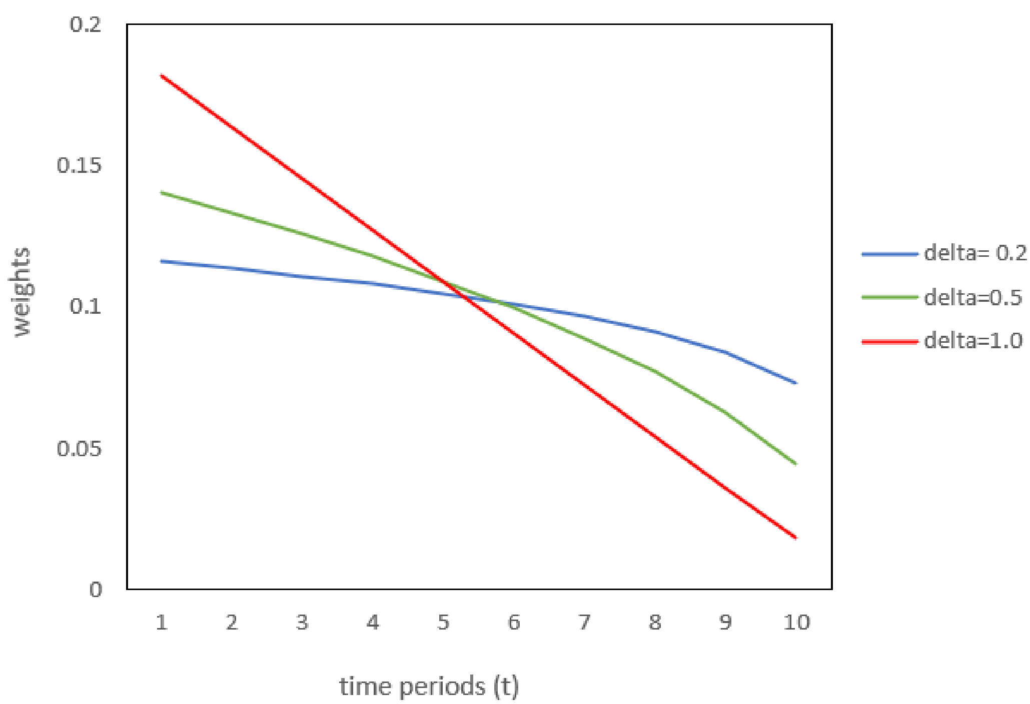
$$\widehat{g} = z(P_{HZ})^{1/\alpha} \tag{1.12}$$

where $\widehat{p}$ is the per-period normalized poverty gap generated by EDE and $\widehat{g}$ is the EDE poverty gap. Note that, for the BCD measure, experiencing poverty in each period in the hypothetical profile results in the highest sensitivity to chronic poverty where the maximal number of consecutive periods poor is now T. Again, unless otherwise stated α = 1.

Consider an individual who lives for four periods with income profile (2, 5, 5, 2) and a constant poverty line of z = \$5. Using equations 2 to 4 from the paper, the $F(\tau = 0)$, BCD (γ=2) and HZ (δ = 0.5) indices are given as 0.300, 0.300 and 0.293 respectively. Applying equations 7 and 8 for the F index would generate an EDE value of \$1.500, that is, experiencing a poverty gap of \$1.500 in each period for four periods would generate the same poverty index as the original income profile that yields the F index of 0.30. Applying equations 9 and 10 for the BCD index would generate an EDE value of \$0.188. Thus, if the individual experiences a poverty gap of \$0.188 in each of the four periods, it will generate the same poverty index as the original income profile that yields the BCD index of 0.60. Finally with the HZ index, applying equations 11 and 12 would generate the EDE value of \$1.465. Experiencing a poverty gap of \$1.465 in each period for four periods would generate the same poverty index as the original income profile that yields the HZ index of 0.293.

Similarly, consider another individual with an income profile, (5, 2, 2, 5), over four periods. The $F(\tau = 0)$, BCD (γ=2) and HZ (δ = 0.5) indices are given as 0.30, 0.60 and 0.307 respectively using equations 2 to 4 from the main paper. Applying equations 7 to 12 would generate the EDE poverty gap for this individual as \$1.500, \$0.375 and \$1.535 for the F, BCD and HZ respectively. For each of the two individuals, the average poverty gap is \$1.50. Therefore, when the temporal pattern of poverty spells is not considered, the two individuals appear to have equal poverty costs for the same time spent in poverty. From the BCD index, the EDE poverty gap for the latter (\$0.188) is twice the EDE poverty gap for the former ($0.375). Therefore, we can conclude that the second individual suffers two times as much poverty as the first individual. This exercise reveals the effects that the actual pattern of poverty experiences have on the cost of poverty computed by each poverty measure. It also illustrates how the poverty differential between groups becomes large when the distribution of poverty is taken into account in the estimation of chronic poverty.

Figure 1.1: HZ weight functions for different values of the sensitivity parameter δ



The blue line graphs the HZ weight function for δ equals 0.2. The green line graphs the HZ weight function for δ equals 0.5 and the red line graphs the HZ weight function for δ equals 1.0
Source: Author's own

Table 1.7: Equivalence scales by different family sizes

Family size	Poverty Line	Implied equivalence scale
1	5,593	1.00
2	7,231	1.29
3	8,573	1.53
4	10,989	1.96
5	13,007	2.33
6	14,696	2.63
7	16,656	2.98
8	18,512	3.31
9 or more	22,083	3.95

Source: Author's calculation of Equivalence scales derived from 1985 poverty thresholds published by the United States Census Bureau.

Table 1.8: Childhood Chronic poverty estimated by the F measure (Test of significance between groups)

	τ				
	0.00	0.25	0.50	0.75	1.00
By Household Type					
Two parent (4,703)	0.040	0.034	0.027	0.015	0.006
Lone parent (854)	0.262 (0.00)	0.256 (0.00)	0.239 (0.00)	0.193 (0.00)	0.124 (0.00)
By Age of Head					
26+ (3,463)	0.054	0.049	0.043	0.030	0.017
20-25 (1,912)	0.101 (0.00)	0.095 (0.00)	0.084 (0.00)	0.061 (0.00)	0.036 (0.00)
16-19 (182)	0.157 (0.00)	0.146 (0.00)	0.120 (0.00)	0.081 (0.00)	0.022 (0.54)
By Age of Mother					
26+ (2,664)	0.055	0.051	0.045	0.033	0.020
20-25 (2,443)	0.085 (0.00)	0.079 (0.00)	0.069 (0.00)	0.049 (0.00)	0.029 (0.05)
16-19 (450)	0.120 (0.00)	0.112 (0.00)	0.093 (0.00)	0.060 (0.00)	0.013 (0.12)
By Head's Education					
Less than 12th grade (1,467)	0.149	0.142	0.128	0.095	0.053
At least 12th grade (4,090)	0.047 (0.00)	0.041 (0.00)	0.035 (0.00)	0.023 (0.00)	0.013 (0.00)
By Race and Sex of Head					
White female (154)	0.158	0.148	0.128	0.073	0.041
Non-white female (660)	0.295 (0.00)	0.289 (0.00)	0.273 (0.00)	0.226 (0.00)	0.147 (0.00)
White male (3205)	0.023	0.017	0.013	0.006	0.002
Non-white male (1,538)	0.077 (0.00)	0.071 (0.00)	0.058 (0.00)	0.035 (0.00)	0.015 (0.00)
By Region of Residence					
North-East (837)	0.037	0.033	0.028	0.018	0.008
North-Central (1,423)	0.076 (0.00)	0.071 (0.00)	0.062 (0.00)	0.047 (0.00)	0.029 (0.00)
South (2,451)	0.098 (0.00)	0.092 (0.00)	0.082 (0.00)	0.058 (0.00)	0.033 (0.00)
West (846)	0.034 (0.71)	0.027 (0.46)	0.023 (0.47)	0.013 (0.41)	0.004 (0.21)
By Household Switches					
Non-switchers (4,622)	0.070	0.064	0.058	0.043	0.027
Switchers (935)	0.094 (0.00)	0.084 (0.00)	0.069 (0.09)	0.036 (0.24)	0.010 (0.00)
All Cohorts (5,557)	0.074	0.068	0.060	0.042	0.024

Notes: P values are in brackets and represents the test of whether the difference in chronic poverty estimates between groups is statistically significant. For each category, the reference group is listed first.

Source: Author, PSID 1968-1997.

Table 1.9: Childhood Chronic poverty estimated by the BCD measure (Test of significance between groups)

	γ				
	1.0	1.2	1.5	1.7	2.0
By Household Type					
Two parent (4,703)	0.040	0.096	0.416	1.079	4.025
Lone parent (854)	0.262 (0.00)	0.960 (0.00)	5.939 (0.00)	17.361 (0.00)	71.703 (0.00)
By Age of Head					
26+ (3,463)	0.054	0.166	0.913	2.582	10.387
20-25 (1,912)	0.101 (0.00)	0.326 (0.00)	1.858 (0.00)	5.309 (0.00)	21.540 (0.00)
16-19 (182)	0.157 (0.00)	0.389 (0.00)	1.715 (0.02)	4.448 (0.08)	16.538 (0.17)
By Age of Mother					
26+ (2,664)	0.055	0.179	1.028	2.951	12.032
20-25 (2,443)	0.085 (0.00)	0.269 (0.00)	1.517 (0.01)	4.319 (0.01)	17.495 (0.02)
16-19 (450)	0.120 (0.00)	0.302 (0.00)	1.302 (0.23)	3.307 (0.59)	11.937 (0.97)
By Head's Education					
Less than 12th grade (1,467)	0.149	0.490	2.805	8.000	32.359
At least 12th grade (4,090)	0.047 (0.00)	0.135 (0.00)	0.713 (0.00)	1.996 (0.00)	7.994 (0.00)
By Race and Sex of Head					
White female (154)	0.158	0.416	2.075	5.774	23.220
Non-white female (660)	0.295 (0.00)	1.114 (0.00)	6.993 (0.00)	20.503 (0.00)	84.822 (0.00)
White male (3,205)	0.023	0.047	0.170	0.412	1.443
Non-white male (1,538)	0.077 (0.00)	0.209 (0.00)	1.007 (0.00)	2.706 (0.00)	10.392 (0.00)
By Region of Residence					
North-East (837)	0.037	0.099	0.486	1.325	5.162
North-Central (1,423)	0.076 (0.00)	0.248(0.00)	1.423 (0.00)	4.088 (0.00)	16.678 (0.00)
South (2,451)	0.098 (0.00)	0.314 (0.00)	1.763 (0.00)	5.008 (0.00)	20.214 (0.00)
West (846)	0.034 (0.71)	0.078 (0.42)	0.325 (0.32)	0.829 (0.29)	3.033 (0.27)
By Household Switches					
Non-switchers (4,622)	0.070	0.229	1.330	3.828	15.634
Switchers (935)	0.094 (0.00)	0.227 (0.95)	0.943 (0.02)	2.363 (0.00)	8.452 (0.00)
All Cohorts (5,557)	0.074	0.229	1.265	3.581	14.426

Notes: P values are in brackets and represents the test of whether the difference in estimates between groups is statistically significant. For each category, the reference group is listed first.

Source: Author, PSID 1968-1997.

Table 1.10: Childhood Chronic poverty estimated by the HZ measure (Test of significance between groups)

	δ				
	0.0	0.2	0.5	0.8	1.0
By Household Type					
Two parent (4,703)	0.040	0.039	0.039	0.038	0.038
Lone parent (854)	0.262 (0.00)	0.265 (0.00)	0.269 (0.00)	0.273 (0.00)	0.275 (0.00)
By Age of Head					
26+ (3,463)	0.054	0.054	0.054	0.054	0.054
20-25 (1,912)	0.101 (0.00)	0.102 (0.00)	0.102 (0.00)	0.103 (0.00)	0.103 (0.00)
16-19 (182)	0.157 (0.00)	0.159 (0.00)	0.162 (0.00)	0.165 (0.00)	0.168 (0.00)
By Age of Mother					
26+ (2,664)	0.055	0.055	0.055	0.055	0.055
20-25 (2,443)	0.085 (0.00)	0.086 (0.00)	0.086 (0.00)	0.086 (0.00)	0.086 (0.00)
16-19 (450)	0.120 (0.00)	0.121 (0.00)	0.122 (0.00)	0.123 (0.00)	0.124 (0.00)
By Head's Education					
Less than 12th grade (1,467)	0.149	0.150	0.151	0.152	0.153
At least 12th grade (4,090)	0.047 (0.00)	0.047 (0.00)	0.047 (0.00)	0.046 (0.00)	0.046 (0.00)
By Race and Sex of Head					
White female (154)	0.158	0.161	0.164	0.168	0.171
Non-white female (660)	0.295 (0.00)	0.298 (0.00)	0.303 (0.00)	0.307 (0.00)	0.310 (0.00)
White male (3,205)	0.023	0.023	0.022	0.022	0.022
Non-white male (1,538)	0.077 (0.00)	0.076 (0.00)	0.075 (0.00)	0.074 (0.00)	0.073 (0.00)
By Region of Residence					
North-East (837)	0.037	0.037	0.037	0.037	0.037
North-Central (1,423)	0.076 (0.00)	0.076 (0.00)	0.076 (0.00)	0.076 (0.00)	0.076 (0.00)
South (2,451)	0.098 (0.00)	0.099 (0.00)	0.099 (0.00)	0.100 (0.00)	0.100 (0.00)
West (846)	0.034 (0.71)	0.034 (0.70)	0.034 (0.69)	0.034 (0.68)	0.034 (0.68)
By Household Switches					
Non-switchers (4,622)	0.070	0.070	0.071	0.072	0.073
Switchers (935)	0.094 (0.00)	0.092 (0.00)	0.089(0.01)	0.085 (0.03)	0.084 (0.07)
All Cohorts (5,557)	0.074	0.074	0.074	0.074	0.075

Notes: P values are in brackets and represents the test of whether the difference in estimates between groups is statistically significant. For each category, the reference group is listed first.

Source: Author, PSID 1968-1997.

Chapter 2

Chronic Child Poverty and Later Life Outcomes

2.1 Introduction

Early childhood experience has been found to be important for later life outcomes. In particular, childhood poverty experiences have been strongly linked to adult outcomes. This paper focuses on the early chronic poverty experiences of children and investigates empirically its association with later life outcomes. I describe the early chronic poverty experiences of children born in the United States from the late 1960s to the early 1970s, following them from birth through to age 10 using the Panel Study of Income Dynamics (PSID) longitudinal dataset. The alternative measures of chronic poverty studied are the Foster (2009) measure (abbreviated as F), the Bossert et al. (2012) measure (abbreviated as

BCD) and the Hoy and Zheng (2011) measure (abbreviated as HZ). Each of the measures differs in the way it accounts for the closeness and timing of poverty spells. The F and BCD measures address chronic poverty concerns while the HZ measure addresses both chronic and early poverty concerns in their measurement of intertemporal poverty. For comparison purposes, I also estimate childhood chronic poverty levels using the traditional count poverty index; i.e., the number of times spent in poverty within a given period. The adult outcomes examined in this study comprise; completed years of schooling, adult health status, labor market success, teenage childbearing, adult poverty status and the formation of own households measured as late as ages 25 and 30. This study is the first to document the link between childhood poverty experiences and later life outcomes using these recently developed chronic poverty measures.

The results from this study indicate that time spent living in poverty as a child still matters in explaining poor outcomes in adulthood even after controlling for a host of demographic and family characteristics often correlated with childhood poverty. I find large and robust associations between chronic child poverty and completed schooling, adult health, employment, teen birth and adult poverty status. The evidence suggests that the choice between one of the chronic poverty measures when analyzing the long-run consequences of childhood chronic poverty depends on the adult outcome of interest, at least using the United States PSID data set between the late 1960s and early 2000s.

The study of the associations between childhood chronic poverty and later life success have been limited because of the non-existence of surveys that follow children from birth into adulthood collecting information on childhood family incomes and later life outcomes. However, with the recent availability of several longitudinal surveys around the world, research has shown that individuals with early poverty experiences end up with worse outcomes such as low employment opportunities, lower productivity, lower educational attainment, poor cognitive development, lower income levels, poor adult health, higher propensities of teen and non-marital births, criminal arrests and adult poverty experiences (Gregg et al. 1999; Suryadarma et al., 2009; Evans and Schamberg, 2009; Wagmiller and Adelman, 2009; Isaacs and Magnuson, 2011; Ratcliffe and McKernan, 2012; Schoon et al., 2012; Dickerson and Popli, 2016; Ratcliffe and Kalish, 2017). The measure of childhood poverty in these studies is either the average family income or the count index, that is, the number of times a child was observed to be poor over a specified period of time.

There are limitations to the commonly used measures of childhood chronic poverty noted above. The average family income does not allow for the identification of who is poor in each period nor the number of times an individual's income fell below the poverty line within a specified period of time. The count index indicates the number of times an individual spent poor but does not account for the depths and distribution of poverty. Both can lead to misleading estimates and comparisons of chronic poverty levels across individuals and groups. For instance, Asiamah (2021) using the PSID dataset compares chronic poverty levels across different socio-economic groups, including race, and show that the difference in estimates of chronic poverty levels between non-whites and whites is significantly heightened when the depth of poverty and temporal patterns of poverty are accounted for in the measurement of chronic poverty.[1]

Consider two individuals who spend equal periods in poverty. One individual may have poverty experiences occurring very early in life while the other individual experiences poverty later in life. Economic hardships experienced in the early years of life have been found to have a greater impact on future success than those experienced in later years (Duncan et al., 1998; Duncan et al., 2012; Schoon et al., 2012; Ratcliffe and Kalish, 2017). Moreover, Heckman and Kautz's (2013) review of the evidence on the effectiveness of early intervention programs in promoting later life success shows that intervention programs before age three led to improved skills formation and improved IQ. Thus, a measure of chronic poverty should differentiate between the poverty experiences of the two individuals, assigning a greater index to the individual with early poverty spell experiences.

The rest of this paper is organized as follows. Section 2 reviews the literature on chronic poverty measurement and its association with later life success. Section 3 explains the poverty measures employed in this paper to estimate child chronic poverty levels. Section 4 describes the data and methodology employed for this study. Section 5 presents the findings from the regression analyses. The study ends with a discussion of the implications of this paper's findings for future research.

[1]The relative childhood poverty between different groups for the periods of study in this paper (1968-1982) is qualitatively similar to the relative childhood poverty between different groups for the periods of study (1968-1997) in Asiamah (2021).

2.2 Literature Review

Chronic poverty experiences in early childhood have been strongly linked to adult outcomes. For example, it has been argued that increases in income for low income families leads to improved mental health for both mother and child in Canada (Milligan and Stabile, 2011) and improved school achievements for children in the United States (Dahl and Lochner, 2005). Economic hardship has been found to increase the psychological stress and depression among parents which may lead to a more coercive and strict style of parenting (McLoyd, 1990). This style of parenting is linked to poor verbal development amongst children (Parker et al., 1999). After controlling for family characteristics, Duncan et al. (1994) find quantitatively large income effects on intelligence test scores (positive) and behavior problem scores (negative) for children at age five. Children growing up in low income families are also less likely to have access to proper health care (in the absence of public insurance) because their parents often lack health insurance (Duncan and Brooks-Gunn, 1997). The resulting poor childhood health leads to increased risk of poor health as an adult (Freedman et al., 1999; Barker et al., 2002) which can influence labor market outcomes and adult socio-economic status negatively.

How poverty should be defined and measured has received considerable attention among economists, political scientists, sociologists, anthropologists, neuroscientists and other social scientists. The measure of living standards can be quantitative (income, consumption, wealth, etc) or qualitative (access to health care, access to housing, access to information, access to education, etc). To measure a single spell of poverty, i.e., snapshot poverty, the popular FGT (Foster et al., 1984) class of poverty gap measure is frequently adopted. Snapshot poverty focuses on the present standard of living of an individual ignoring the influence of past or future poverty experiences on the current level of hardship suffered (Calvo and Dercon, 2009). In recent years, research has stressed the need to extend snapshot poverty measurement to address both the multidimensionality and lifetime (dynamic) aspects of poverty, with the latter often referred to as chronic poverty.

Chronic poverty measures have been developed based on an aggregation of snapshot poverty levels over a sequence of periods into a single index of poverty. These chronic poverty measures can be broadly classified into two categories; the *permanent-income* approach and the *spells* approach. An early attempt to measure poverty over time is the *permanent-income* approach, which computes an average of all incomes over the lifetime (called the permanent income). It identifies a person as chronically poor if the permanent income is below a corresponding poverty line (Rodgers and Rodgers, 1993; Hill and Jenkins, 1999; Jalan and Ravallion, 2000; Valletta, 2006). The permanent-income approach implicitly assumes that income from non-poor periods will compensate the periods of low income by accounting for the potential saving and borrowing behaviour of individuals over their lifetime. With this approach, chronic poverty is the level of poverty an individual experiences as if his/her income in every period equals their permanent income. The second approach, which is the *spells* approach, measures a person's level of chronic poverty by focusing on the distribution of poverty spells over an individual's lifetime (Calvo and

Dercon, 2009; Hoy and Zheng, 2011; Bossert et al., 2012; Gradin et al., 2012; Dutta et al., 2013) or time spent in poverty (the count index) or both (Foster, 2009; Alkire et al., 2017).

The bulk of existing empirical research on the associations between childhood chronic poverty and adult outcomes focus on the count index as a poverty measure (Gregg et al., 1999; Suryadarma et al., 2009; Wagmiller and Adelman, 2009; Evans and Schamberg 2009; Ratcliffe and McKernan, 2012; Ratcliffe and Kalish, 2017). This can lead to very conservative estimates of these associations because some critical aspects of poverty are not taken into account. For instance, consider one individual who spends half of his/her lifetime poor and another individual who spends a third of his/her lifetime poor. Suppose the second person experiences larger poverty gaps.[2] Without accounting for the size of their poverty gaps, the first individual is deemed more economically deprived by the count index. However, accounting for poverty gaps in addition to the time spent poor can provide evidence that the second person suffers more chronic poverty than the first person. This is evident in the findings from Asiamah (2021) that, children born to teen mothers only suffer more chronic poverty than children born to older mothers when poverty gaps are ignored (i.e., when chronic poverty is described using only the number of times spent poor within a specified period).

Again, consider two individuals who spend equal time in poverty. Suppose the first individual has all poverty spells occurring consecutively while the other individual's poverty spells are separated by periods of non-poverty. The poverty experience of the latter is more

[2]Gaiha (1989) show that individuals who suffered chronic poverty (i.e., spend all periods under study poor) are not necessarily those with wider poverty gaps.

transitory in nature. Hence, its negative impact on later life outcomes is likely to be less severe. Thus, it is pivotal to differentiate between these distinct chronic poverty experiences. For comparison purposes, this study focuses only on chronic poverty measures that differ in terms of how early, close and recurring poverty spells are treated in the measurement of chronic poverty. A more detailed description of these measures is presented in section three below.

Socioeconomic and demographic characteristics often correlated with childhood poverty are included as covariates in the regression models. This is relevant because numerous studies have demonstrated the importance of early childhood environments in shaping the abilities of children and accounting for a substantial variation in their later life outcomes (Currie and Hyson, 1999; Smith and Haddad, 2000; Barker et al., 2002; Currie and Moretti, 2003; Behrman and Rosenzweig, 2004; Cunha et al., 2006; Cunha and Heckman, 2007; Black et al., 2007; Currie, 2009; Huggett et al., 2011; Shonkoff et al., 2012).

Prior literature has found strong associations between childhood poverty (measured using the count index) and adult outcomes, yet no study to date has empirically examined these associations using the chronic poverty measures adopted in this study. With the recent development of dynamic poverty measures, this present study aims to fill this gap in the empirical literature by using the United States nationally representative longitudinal data to examine the strength of the association between these "new" chronic poverty measures and adult success.

2.3 Measures of Child Chronic Poverty

This section describes three recently developed measures of chronic poverty used to measure childhood chronic poverty levels in this paper; the Foster (2009) measure, the Bossert et al. (2012) measure and the Hoy and Zheng (2011) measure. The traditional count poverty index is also described in this section. For the poverty measures considered, the measurement of chronic poverty is done in two steps. First, the poor is identified in a given population based on a choice of a poverty criterion (e.g., income, consumption, access to health care, malnutrition levels, etc). In the second step, the poverty experiences of the poor individual over a given number of years are summed into an overall index of chronic poverty.

2.3.1 Notation

First, consider some useful notations. Consider an individual i who lives for T periods. The level of chronic poverty suffered by a child is estimated for the first ten years of life (i.e., T = 10) and not all T years of life. In each period t = 1,2,...,T, individual i= 1,2,...,N has a level of income x_i^t. Each period's level of income of the individual is then compared with a pre-determined poverty threshold $0 < z^t < \infty$. Individual i is identified as poor in period t if his/her income level x_i^t is strictly less than the poverty line z^t. In any given period t for an individual i, the poverty gap is defined as $G_i^t = z^t - x_i^t$ and the relative poverty gap is given as $g_i^t = \frac{z^t - x_i^t}{z^t}$.

The FGT measure of snapshot poverty in period t for individual i is given as

$$p_i^t = p(x_i^t; z^t) = \begin{cases} \left(1 - \dfrac{x_i^t}{z^t}\right)^{\alpha} & \text{if } x_i^t < z^t \\ \\ 0 & \text{if } x_i^t \geq z^t \end{cases} \tag{2.1}$$

where the choice of α is typically restricted to values $\{0,\ 1,\ 2\}$. α equal to 0 gives the incidence of poverty such that p_i^t is equal to 1 if x_i^t is below the poverty threshold z^t and 0 if the income x_i^t is at least as large as the poverty threshold z^t. α = 1 gives the size of the normalized poverty gap whilst α = 2 provides a measure of the intensity of poverty. Unless otherwise stated in this paper, α is equal to 1.

2.3.2 The Count Index

This approach does not address the extent to which a person's income is below the poverty line. It involves setting α equal to 0 in equation 1, such that a poverty spell is assigned the value "1" and a non-poverty spell is assigned the value "0". Identifying the chronically poor individual with this approach is based on the fraction of time an individual's income was below the poverty threshold over time. The fraction of time individual i spent poor is derived by dividing the number of $1s$ (i.e., the number of poverty spells) by the total number of periods of observation T. The time spent poor need not be consecutive and there is no consensus regarding the appropriate cutoff fraction. For instance, one may require an individual to spend at least one-third of the time poor or spend at least half of the time poor to be identified as chronically poor. The greater the duration cutoff fraction, the fewer the people considered to be suffering chronic poverty and vice versa.

2.3.3 The Foster (2009) measure of chronic poverty

The F measure of chronic poverty uses a dual cutoff spells approach in measuring levels of poverty. The first cutoff is in the income space, i.e., setting a poverty line z^t and identifying an individual as poor in a given period if $x_i^t < z^t$. The second cutoff is in the duration line ($0 \leq \tau \leq 1$) which specifies the minimum fraction of time that must be spent in poverty to be identified as chronically poor. For an individual i who spends q out of T periods poor, the F measure of chronic poverty over time is given as

$$P_i^F(x_i^t; z^t) = \begin{cases} \dfrac{1}{T}\sum_{t=1}^{T} p_i^t & \text{if } q/T \geq \tau \\ \\ 0 & \text{if } q/T < \tau \end{cases} \tag{2.2}$$

where p_i^t is the FGT measure of snapshot poverty described in equation 1. The F measure is sensitive to the cutoff duration line τ in that the smaller the value of τ, the more persons are considered as chronically poor and vice-versa. The F measure addresses chronic poverty concerns through the duration cutoff line, i.e., the sufficient time spent in poverty. For persons who pass the threshold τ, all poverty spells are assigned equal weights regardless of whether they occur closely in time to each other, occur consecutively or occur in isolation. The two chronic poverty measures that are described below are sensitive in different ways to the temporal pattern of poverty spells experienced by an individual.

2.3.4 The Bossert, Chakravarty and d'Ambrosio (2012) measure of chronic poverty

The BCD measure evaluates the persistence in poverty with a focus on the duration of poverty spells in the sense that, ceteris paribus, consecutive poverty spells are assigned greater weights than isolated poverty spells. Moreover, the more poverty spells experienced consecutively, the greater the weight assigned to each of the poverty spells occurring in that string of contiguous poverty spells. The BCD measure for individual i is given as

$$P_i^{BCD}(x_i^t; z^t) = \frac{1}{T}\sum_{t=1}^{T}\gamma^{k-1}p_i^t \tag{2.3}$$

where k is the (maximal) number of consecutive periods including the t^{th} period with positive poverty gaps and p_i^t is the FGT measure of snapshot poverty. γ is a measure of the sensitivity to chronic poverty with $\gamma \geq 1$. γ^{k-1} is the set of weights assigned to consecutive poverty spells. If γ equals 1, then each poverty spell is weighted equally regardless of the sequence of poverty/non-poverty spells in which it occurs for a given k. Conversely, γ > 1 assigns more weight to consecutive poverty spells than isolated poverty spells. For example, for γ equal 2, the weight for each poverty spell for the normalized poverty gap profile (0, 3/5, 3/5, 0) is 2 while the weight for each of the poverty spells for the normalized poverty gap profile (0, 3/5, 0, 3/5) is 1. For any given value of γ, the bigger the k, the greater the weight assigned to each poverty spell that occurs in a string of two or more consecutive periods. For example, for γ equal 2, the weight of each poverty spell occurring in a string of 3 consecutive spells for normalized poverty gap profile (0, 3/5, 3/5, 3/5, 0) is 4 while the weight of each of the poverty spells occurring in a string of 2 consecutive spells for normalized poverty gap profile (0, 3/5, 3/5, 0, 3/5) is 2. Note that the BCD measure is not sensitive to the "closeness" of poverty spells if they are not contiguous; e.g., normalized poverty gap profiles (3/5, 0, 3/5, 0, 0) and (3/5, 0, 0, 0, 3/5) will generate the same BCD index value. Unlike the F measure, the BCD measure does not rely on a duration cutoff fraction to describe the chronically poor. All individuals with at least one poverty spell are assigned a positive BCD index value.

2.3.5 Hoy and Zheng (2011) measure of chronic poverty

The HZ poverty measure addresses both chronic poverty and early poverty concerns with a set of weights applied to poverty spells. Firstly, the measure assigns greater weights to earlier poverty spells such that, ceteris paribus, an individual who experiences poverty earlier in life has a higher HZ poverty index (i.e., the early poverty axiom). Consider two individuals who live for three periods; early, middle-age and old age with normalized poverty gap profiles (3/5, 0, 0) and (0, 0, 3/5). The F and BCD measures will assign equal poverty to both individuals, even though the poverty experience of the first individual occurred earlier. The HZ measure, on the other hand, will assign a higher poverty to the individual with the first profile.

Secondly, the HZ measure assigns greater weight to poverty spells that occur closer to each other all else constant (i.e., the chronic poverty axiom). For two individuals with normalized poverty gap profiles A = (3/5, 0, 0, 3/5) and B = (0, 3/5, 3/5, 0), the HZ poverty index for B will be higher than A for any given set of weights that is strictly concave in t. This treatment is similar to the BCD measure. However, when profile B is compared with another profile C = (3/5, 0, 3/5, 0), C will now be poorer than B according to the HZ because of the earlier poverty experience. This illustration depicts the trade-off between early and chronic poverty concerns that the HZ chronic measure accommodates, i.e., earlier poverty in a sense takes priority over closeness of poverty spells. Moreover, a symmetric spread out of poverty spells in a person's lifetime reduces lifetime poverty. Consider the profile (0, 3/5, 0, 3/5, 0). If the poverty spells are spread out uniformly to become (3/5, 0, 0, 0, 3/5), the HZ measure will assign a higher chronic poverty to the first profile because, although the second profile has an experience of poverty in the first period, the two poverty spells are separated by three periods of non-poverty. For an individual i, the HZ measure is defined as

$$P_i^{HZ}(x_i^t; z^t) = \sum_{t=1}^{T} \beta(t, T) p_i^t \tag{2.4}$$

where $\beta(t,\,T)$ are weights assigned to per-period poverty and non-poverty spells. These weights are normalized to sum to one. To satisfy the early poverty axiom of the HZ measure, the weight function must be decreasing in t. Likewise, the weight function must be concave in t to satisfy the chronic poverty axiom. Consider the weight function below:

$$\beta(t,T) = \left(1 - \frac{t}{T+1}\right)^{\delta}, \quad 0 \leq \delta \leq 1 \tag{2.5}$$

The smaller (bigger) the value of δ, the less (more) sensitive the HZ measure is to early poverty concerns and the more (less) sensitive the HZ measure is to chronic poverty concerns. Each period $t's$ weight is obtained by dividing $\beta(t,T)$ by the sum of the weights over time $\sum_{t=1}^{T}\beta(t,T)$. The weight function $\beta(t,T)$ is linear and flat for δ equal 0 with equal weights $1/T$ assigned to all poverty spells. For δ equal to 1, $\beta(t,T)$ is linear and decreasing in t with the weight on the first period being the highest (only the early poverty axiom is satisfied). For any δ strictly between 0 and 1, $\beta(t,T)$ is strictly concave and the HZ measure addresses both early and chronic poverty concerns in their measure of lifetime poverty.

2.4 Data and Methodology

2.4.1 Data and Summary Statistics

The analyses are based on over 35 years of data from the Panel Study of Income Dynamics (PSID) of the United States. PSID is an ongoing study which began in 1968 with a nationally representative sample of 5000 families consisting of 15,000 individuals. This survey has followed children and parents for over fifty years spanning from 1968 collecting information such as educational attainment, labor supply, health, income levels, household wealth, among others, on all parties annually.

The poor are identified in this study using the official definition of poverty in the United States. An individual is described as poor in any period if his/her equivalent household income is below the appropriate poverty threshold. Poverty thresholds in the United States are determined by the Census Bureau based on money income before taxes (excluding capital gains or non-cash benefits). They vary by family size and composition and are updated annually for inflation using the Consumer Price Index for All Urban Consumers (CPI-U). The equivalent household income is derived using equivalence scales implied by the poverty thresholds published by the US Census Bureau. For instance, a family size of one compared to a family size of four implies an equivalence scale of about 2 (refer to Table 2.9 in the Appendix).

Childhood chronic poverty is estimated from birth to age 10 whilst adult outcomes are evaluated at ages 25 and 30. Overall, the final sample consists of 1,047 individuals and 728 individuals followed from birth to age 25 and age 30 respectively over six birth cohort periods from 1968 to 1973. Any loss of observations is due to attrition and missing information on responses to the adult outcomes of interest. Summary statistics for the adult outcomes are provided in Table 2.1. On average, ever poor individuals (i.e., individuals who experienced at least one poverty spell in the first ten years of life) spend an average of 11.7 years in school compared with an average of 12.9 years spent in school by individuals who never experienced poverty in the first ten years of life. 20 percent of ever poor individuals are found to be poor as an adult compared with 3 percent of individuals who had zero poverty spell experience in their early childhood. A similar pattern is observed for employment status where 63percent of ever poor individuals are found to be currently employed compared with 72 percent of individuals who never experienced early childhood poverty. Additionally, 93 percent of ever poor individuals reported a good adult health status compared with 97 percent for individuals with no early poverty spell experiences. 24 percent of ever poor individuals had a child before they turned age 20 compared with 8 percent for individuals who had zero poverty spell experiences in the first ten years of life. The average birth order for ever poor individuals is 3.2 whilst the average birth order for "never poor" children is 2.4. A description of the data set and variables are available in the Appendix.

Table 2.1: Descriptive Statistics: Adult Outcomes by Childhood Poverty Status

	Never Poor (Mean or %)	Ever Poor (Mean or %)	All (Mean or %)
Completed schooling	12.9 years	11.7 years	12.5 years
Currently working (%)	72.3	62.8	68.9
Good Health (%)	96.9	92.8	95.4
Form own Households (%)	69.3	70.7	69.8
Teen Birth (%)	8.3	23.7	13.8
Currently poor (%)	3.2	20.1	8.7
Birth Order	2.4	3.2	2.7
Mother's age at birth	26.2 years	25.5 years	25.9 years
Age of first-time moms	22.8 years	20.8 years	22.3 years

Notes: Ever Poor individuals spent at least one of the first ten years of life poor. Never poor individuals have zero poverty spell experiences over the first ten years of life. See the Appendix for an explanation of the variables.
Source: Author, PSID Data 1968-2003.

2.4.2 Childhood Chronic Poverty Levels

In this subsection, chronic poverty levels as estimated by the different poverty measures are reported for the six different birth cohorts.[3] For the count index, four definitions of chronic poverty are analyzed; poor at least one out of ten periods; poor at least three out of ten periods, poor at least five out of ten periods and poor at least eight out of ten periods. These chronic poverty rates are presented in Panel A of Table 2.2. For all birth cohorts, 35.9 percent of individuals spent at least one period poor, 19.5 percent of individuals spent at least three periods poor, 12.8 percent of individuals spent at least five periods poor, and 5.5 percent of individuals spent at least eight out of ten periods poor. For all duration cutoff fractions, persons born in 1968 have the highest chronic poverty rate. For the duration cut off fraction of 5/10, 17.3 percent of individuals born in 1968, 13.3 percent of individuals

[3]Correlation between pairs of poverty indices are available in the Appendix (Table 2.10).

born in 1970, 11.9 percent of individuals born in 1972 and 1973, 10.7 percent of individuals born in 1971 and 7.8 percent of individuals born in 1969, suffered chronic poverty in their early childhood years.

Chronic poverty levels for birth cohorts estimated by the F measure are reported in panel B of Table 2.2. The F measure emphasizes the number of periods spent in poverty without considering different temporal patterns of poverty experiences. However, unlike the count index, it takes into account the size of the poverty gap for each year of poverty experienced. In column 1, τ = 0.00 and all individuals who spent at least one period poor over the ten-year period are assigned a positive F chronic poverty index. In the second column, τ equal 0.25 and all individuals who spent at least three out of ten periods poor are assigned a positive F chronic poverty index. Individuals who spent five or more periods poor are assigned a positive F poverty index in the third column (τ = 0.50) whereas individuals who spent eight or more periods poor are assigned a positive F poverty index in the last column (τ = 0.75). The comparison of the degree of F chronic poverty levels between birth cohorts for the most part is similar to the findings from the count index with a few exceptions. For instance, for τ = 0.00, individuals born in 1969 have the least F chronic poverty index even though they suffered more chronic poverty than individuals born in 1973 according to the count index (for the duration cutoff fraction of 1/10). This indicates that compared with persons born in 1969, individuals born in 1973 have relatively deep poverty gaps in the years they spent poor.

Chronic poverty levels are estimated by the BCD measure for different γ values; 1.2, 1.5, 1.7 and 2.0. These are presented in panel C of Table 2.2. The BCD measure emphasizes the closeness of poverty spells by assigning greater weights to poverty spells that occur in a string of two or more consecutive periods. However, unlike the HZ measure, the BCD measure applies no additional weight to poverty spells that are closer together unless they are contiguous. For a given string of two or more consecutive poverty spells, the larger is the parameter γ, the greater the contribution of each of those poverty spells to the BCD chronic poverty level of an individual. Like the F index, the comparison of the degree of BCD chronic poverty levels between birth cohorts is similar to the findings from the count index with a few exceptions. In panel B when all poverty spells are weighted equally (i.e., τ=0.00), the 1973 birth cohort had a greater chronic poverty level compared with the 1969 birth cohort. However, when consecutive poverty spells are assigned greater weights the rank between the two groups reverses, suggesting that individuals born in 1973 have fewer consecutive poverty spell experiences on average compared with individuals born in 1969. Therefore, different assumptions made regarding the contribution of each poverty spell to the overall level of chronic poverty of an individual leads to different ranking of groups by chronic poverty levels.

Chronic poverty levels are estimated by the HZ measure for different δ values; 0.2, 0.5, 0.8 and 1.0. These are presented in panel D of Table 2.2. The HZ measure emphasizes the closeness of poverty spells by assigning greater weights to poverty spells that occur closer in time to each other without the requirement that poverty spells be consecutive as required by the BCD measure. It also emphasizes another aspect of the timing of poverty spells by assigning greater weights to poverty spells that occur earlier in life. As the value of δ is increased, the HZ measure becomes more sensitive to early poverty spells and less sensitive to close poverty spells. Similar to the results from the BCD measure, children born in 1968 have the greatest HZ index, followed by children born in 1970 with children born in 1973 having the least HZ index. As sensitivity to early poverty increases, the HZ chronic poverty index for the 1968, 1970 and 1971 cohorts increases while the HZ index for the 1973 cohort decreases. Therefore, poverty experiences for the 1973 cohort occur much later while individuals born in 1968, 1970 and 1971 experience poverty much earlier. The HZ chronic poverty index for the 1969 and 1972 birth cohorts is not very sensitive to the value of the parameter δ. For all sample children, the HZ chronic poverty index increases as sensitivity to early poverty is increased.

2.4.3 Methodology

Adult outcomes are modelled as a function of childhood chronic poverty, socioeconomic characteristics and household controls.

The baseline model is expressed as:

$$y_i = a_o + b\mathbf{Index}_i + c\mathbf{X}_i + d\mathbf{HH}_i + v_i \tag{2.6}$$

where y_i represents the different adult outcome variables of interest for individual i; that is, completed years of schooling, employment status, general health status, formation

Table 2.2: Childhood Chronic Poverty, by Birth Cohort Year

Panel A: Count Index	N	1/10	3/10	5/10	8/10
1968 Cohort	279	42.3	24.1	17.3	9.4
1969 Cohort	126	31.7	15.0	7.8	4.8
1970 Cohort	174	35.6	19.0	13.3	6.3
1971 Cohort	159	35.8	18.8	10.7	2.5
1972 Cohort	175	33.1	18.2	11.9	5.1
1973 Cohort	134	30.6	17.9	11.9	1.5
All Cohorts	1,047	35.9	19.5	12.8	5.5
Panel B: F Index		0.00	0.25	0.50	0.75
1968 Cohort	279	0.055	0.048	0.041	0.028
1969 Cohort	126	0.032	0.027	0.020	0.014
1970 Cohort	174	0.050	0.044	0.038	0.024
1971 Cohort	159	0.040	0.033	0.025	0.011
1972 Cohort	175	0.046	0.041	0.035	0.020
1973 Cohort	134	0.035	0.030	0.023	0.003
All Cohorts	1,047	0.045	0.039	0.032	0.018
Panel C: BCD Index		1.2	1.5	1.7	2.0
1968 Cohort	279	0.148	0.728	1.992	7.822
1969 Cohort	126	0.075	0.331	0.873	3.285
1970 Cohort	174	0.126	0.575	1.521	5.775
1971 Cohort	159	0.097	0.374	0.980	3.822
1972 Cohort	175	0.116	0.515	1.353	5.115
1973 Cohort	134	0.056	0.113	0.176	0.328
All Cohorts	1,047	0.111	0.487	1.286	4.917
Panel D: HZ Index		0.2	0.5	0.8	1.0
1968 Cohort	279	0.056	0.058	0.059	0.060
1969 Cohort	126	0.033	0.033	0.033	0.033
1970 Cohort	174	0.051	0.051	0.052	0.052
1971 Cohort	159	0.040	0.041	0.042	0.042
1972 Cohort	175	0.046	0.046	0.046	0.046
1973 Cohort	134	0.034	0.032	0.031	0.031
All Cohorts	1,047	0.045	0.046	0.046	0.047

Notes: The Count Index represent the percentage of children who suffered chronic poverty. For the count index, chronically poor children are identified using four cutoff fractions; 1/10 (poor at least once), 3/10 (poor at least three times), 5/10 (poor at least five times) and 8/10 (poor at least eight times). Four different values of τ; 0.00, 0.25, 0.50 and 0.75 are considered for the F Index. Four different values of γ; 1.2, 1.5, 1.7 and 2.0 are considered for the BCD Index. Four different values of δ; 0.2, 0.5, 0.8 and 1.0 are considered for the HZ Index.
Source: Author, PSID Data 1968-2003.

of own household, adult poverty status and the likelihood of teenage birth.

The different chronic poverty indices for individual i; that is, the F, BCD, HZ and the count index comprise the variable $\mathbf{Index}_i$. As another measure of early childhood economic deprivation, $\mathbf{Index}_i$ in equation 6 is replaced with the average poverty gap over the first ten years of life for individual i (APG_i). Whilst the APG only takes into account the depth of poverty in any period and the count index only accounts for the duration of poverty spells, the F, BCD and HZ indices account for the depth of poverty as well as the duration, timing and closeness of poverty spells but in distinct ways. Exploring the different measures in this study is useful to make a sharp comparison of the predictive performance of different childhood chronic poverty measures.

The vector $\mathbf{X}_i$ includes certain time-invariant characteristics of the child including gender (male or female), race (white or non-white) and birth order (continuous). Theoretically, Becker and Lewis (1973) and Becker and Tomes (1976) show the trade-off that exists between the number of children (i.e., quantity of children) and the amount of resources in the form of goods and time that parents devote to each child (i.e., quality of children). Empirically, studies have found both negative (Black et al., 2005; Booth and Kee, 2009; Hatton and Martin, 2009) and positive (Ejrnaes and Portner, 2004) birth order effects on outcomes including educational attainment, adult health status, teen birth, labor supply earnings, etc. A positive birth order effect is likely to occur in situations where older siblings are forced to leave school to find some work to support the younger siblings and the family at large. Including this variable in the regression models will address the long-term consequences of resource dilution within a household. Figure 2.1 depicts a negative relationship between average completed years of schooling and birth order for individuals in the sample. The negative relationship still persists when the sample is restricted to individuals with at least one poverty spell experience or to individuals with no early childhood poverty experience. However, the pattern is not strictly monotonic.

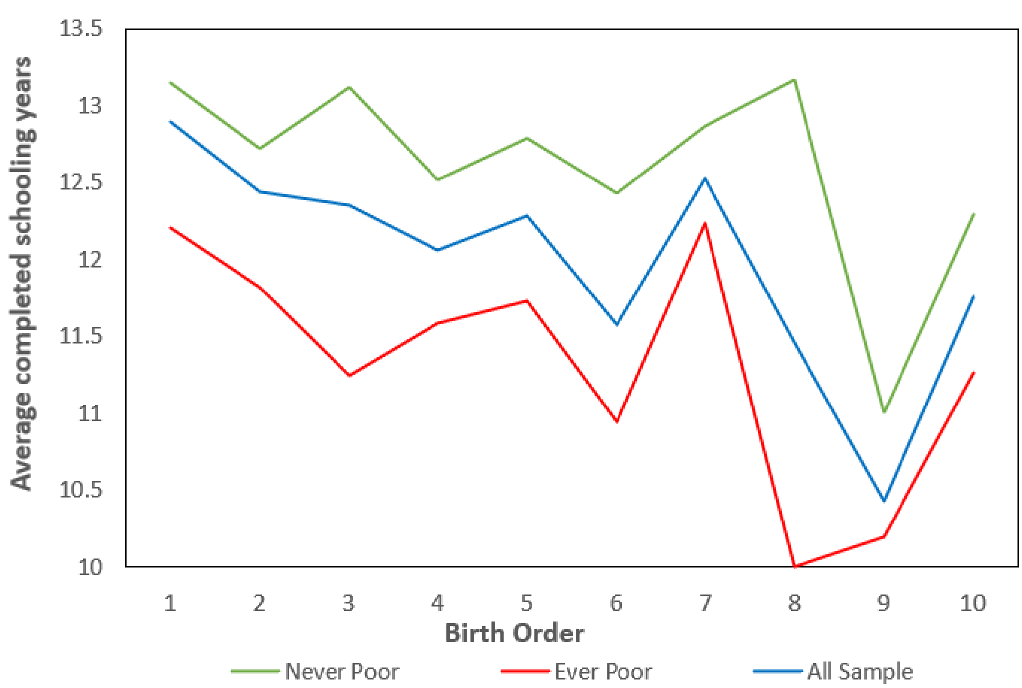
The variables included in $\mathbf{HH}_i$ are a vector of household and background characteristics for individual i including: completed education of the household head which is usually for the father in a two parent household and for the mother in a single parent household; age of mother at birth (continuous); region of birth (North-East, North-Central, South or West); year of birth; whether the household head is an immigrant; and whether the individual experienced disruptions in the household environment due to parental marital dissolution growing up.[4] Transitioning to live with a single parent for the most part is accompanied with a fall in household income. The resulting economic hardships can force children to leave school early which can lead to lower occupational attainment and wages, and adult poverty experiences. Parental marital separations also affect the amount of time that parents spend with their children (Ram and Hou, 2003). In addition, marital break-up is often found to be preceded by parental conflict, parental alcohol/drug abuse, physical and mental abuse of spouses and children (White, 1990) which are all indicators of inadequate parenting. These deficiencies in parenting lead to poor outcomes for the child (Hetherington and Stanley-Hagan, 1999; Amato, 2000; Ram and Hou, 2003). Thus, including this variable in

[4]Other causes of household disruptions found in the data include death of parent(s) and/or move-ins with a relative. I focus on parental divorces because it constitutes about 95 percent of the disruptions in the initial household environment of children in the data set.

the regression models addresses two key channels through which divorces in early childhood years affect children's future success; directly through deficiencies in parenting and unstable environments and more indirectly through the effect of poverty.

The main coefficient of interest is b. If the hypothesis that early childhood chronic poverty is linked to poor childhood development is correct, b should be negative for good outcomes and positive for bad outcomes. v_i is the classical stochastic error term. In all regression models, standard errors are clustered at the household level to correct for multiple children born into the same household.

Figure 2.1: Average Completed Years of Schooling by Birth Order



The blue line graphs the average completed years of schooling by birth order for the full sample. The red line graphs the average completed years of schooling by birth order for individuals who experienced at least one poverty spell in the first 10 years and the green line graphs the average completed years of schooling by birth order for individuals who had zero poverty spell experiences in the first 10 years of life. Data is from the PSID.

2.5 Results and Discussion

This section presents the results from the regression estimations. For the F index, the choice parameter is τ = 0.0 and all individuals who spent atleast one out of the first ten years of life in poverty will be assigned a positive poverty index as it is the case for the BCD and HZ indices. For the BCD index, the choice parameter is γ = 1.2. A value of 1.2 ensures that consecutive poverty spell experiences are emphasized without exaggerating them. For the HZ index, the choice parameter is δ = 0.2. A value of 0.2 ensures that poverty spells that occur closely in time to each other are assigned greater weight compared to poverty spells that are separated by many periods of poverty. With this, I am able to compare the performance of chronic poverty measures that differ in terms of how the closeness of poverty experiences are treated. Although both the BCD and HZ indices account for the closeness of poverty, the BCD index is not sensitive to the closeness of poverty spells if they are not consecutive.[5] For the count index, the cutoff fraction is 5/10 (i.e., an individual suffers chronic poverty if they spent at least five out of first ten periods of their life poor). All the parameter values are exogenous. The association between early childhood chronic poverty and adult outcomes when no household and background characteristics are controlled for in the regressions is presented in Table 2.3. Once additional controls are introduced, the associations are attenuated. This implies that poor outcomes in adulthood operate through several correlates of childhood poverty such as, age of mother

[5]Regression results (not shown in this paper) reveal that the results are robust to different values of the choice parameters.

at birth, birth order, region of birth, whether household head is an immigrant, whether the individual experienced parental divorces during childhood, etc.

In Table 2.3, childhood chronic poverty levels estimated by the F measure are presented in panel A, childhood chronic poverty levels estimated by the BCD measure are presented in panel B and childhood chronic poverty levels estimated by the HZ measure are presented in panel C. In panel D, the chronic poverty measure is the count index. In panel E, the chronic poverty measure is the average poverty gap over the 10 years. I report the standardized coefficients for each covariate in order to compare the relative strength of their association with the adult outcomes.[6] The standardized coefficient represent the standard deviation changes in a given outcome associated with a one standard deviation increase in a given covariate. The higher the absolute value of the standardized coefficient, the stronger the association. Hence, a standardized coefficient of -0.8 has a stronger association than a standardized coefficient of 0.5.

2.5.1 Childhood Chronic Poverty Experience and Completed Education

The association of levels of childhood chronic poverty with completed schooling by age 25 is presented in column 1 of Table 2.3. The results indicate that on average, individuals who suffered chronic poverty very early in life are less likely to complete more years of school

[6] Standardized coefficients are the coefficients that you get if the variables in the regression model are all converted to z-scores before running the analysis. The unstandardized coefficients are available in the Appendix (Tables 2.15 to Tables 2.20).

by age 25. The result is statistically significant for all poverty measures at 1 percent. The size of the associations with years of completed schooling is greatest with the F and HZ indices. The F and HZ indices have a similar coefficient estimate of -0.20 whereas the BCD index has the least size association with the number of years spent in school by an individual at age 25 with a coefficient estimate of -0.13.

From Table 2.4, the results show that females are found to be more likely to spend more years in school compared with males. Children with older mothers also attain more years of schooling on average. A negative and significant birth order effect on educational attainment is observed across all columns. Interacting birth order with age of mother reveals that when later born children have older mothers at birth, they complete more years of school. Older mothers have more financial resources at time of birth, are partners with fathers of higher quality (Aizer et al., 2018) and have lower levels of psychosocial stress (Kingston et al., 2012), a condition found to reduce the incidence of deficient parenting behaviors (Reid and Meadows-Oliver, 2007). This result highlights the role that age of mother plays in mediating the effects of birth order on the educational attainment of children. Individuals who experienced parental marital dissolution during the early childhood years complete fewer years of school on average. Across all columns, experiencing parental divorces in childhood is negatively associated with years spent in school by a child; coefficient estimates range between -0.05 and -0.07. Having a household head with at least a high school degree is positively associated with years of completed schooling. Second generation citizens (i.e., children of immigrant(s)) spend more years in school on average compared to third and higher generation citizens (i.e., children born in the United States with both parents born in United States). Compared with children born in the North East Region, children born in the South Region complete fewer years of school.

2.5.2 Childhood Chronic Poverty Experience and Adult Health Status

The association of levels of childhood chronic poverty with adult health status at age 25 is presented in column 2 of Table 2.3. The BCD index has the least size association with adult health status (coefficient estimate is -0.10) whereas the count index has the greatest size association with adult health status (coefficient estimate is -0.15). From Table 2.5 when additional controls are included in the regression, we find that all other covariates except race and the educational attainment of the child's household head are statistically insignificant in the health regressions. Having an educated household head whilst growing up is positively associated with the likelihood that an individual reports a good health; the conditional size correlation ranges between 0.18 and 0.21 across all columns.

2.5.3 Childhood Chronic Poverty Experience and the Likelihood of Teen Births

The association between childhood chronic poverty levels and the propensity that an individual has an early birth is presented in column 3 of Table 2.3. Individuals with early chronic poverty experiences are more likely to have their first child before the age 20. When all other covariates are included in the regression in Table 2.6, we lose the statistical significance on some of the coefficients of the poverty indices. This suggests that the likelihood that an individual has a teen birth may be associated with other unfavorable background conditions and/or poverty experiences in later childhood years other than childhood chronic poverty experiences during the first ten years of life.[7] Rather unexpectedly, the coefficient of the BCD index is negative and significant ($p< 0.10$).

Gender and race of an individual, the educational attainment of the household head, region of birth and age of mother have significant associations with the likelihood that an individual has a teen birth. Females are more likely to have an early first birth by age 19 compared with males and non-whites are more likely to have a teen birth compared with whites. Children with educated household heads have a lower propensity to have a teen birth compared with children with less educated heads (i.e., less than a high school degree). Individuals who had younger mothers at birth are found to be more likely to have an early first birth. Compared with individuals born in the North East Region, the association between the propensity of teen births and being born in other Regions is roughly 0.15 points higher.

2.5.4 Childhood Chronic Poverty Experience and Labor Market Success

Column 6 of Table 2.3 presents the results from the employment regressions. The association of childhood chronic poverty with the likelihood of being employed at age 30 is

[7]For instance, Duncan et al. (2012) using the PSID data, found that non-marital births were strongly correlated with average family income from age 11 to 15 than family income averaged from birth to age 10.

reported. The F has the greatest size association with the employment status of an individual with a coefficient estimate of -0.133 which is slightly higher than the coefficient estimate of the HZ index of -0.132. The count index has the least size association of -0.105 followed by the average poverty gap with a coefficient estimate of -0.119. From Table 2.8, with additional controls, the coefficients of the count index and the average poverty gap are both statistically insignificant. The individual's own years of completed schooling are controlled for in the employment regressions and the results show a positive and significant association with employment status. This is consistent with recent evidence on the importance of own education on labor market success (Oreopoulos, 2007; Fischer et al., 2016; Brunello et al., 2016; Hofmarcher et al., 2019) measured in terms of earnings or employment status.

According to human capital theory (Becker, 1964; Grossman, 1972), health plays an important role in an individual's labor supply decision. I control for the current health status of individuals in the employment regressions. Consistent with the literature (Scott et al., 1977; Currie and Madrian, 1999 review of the US literature; Cai and Kalb, 2006; Cai, 2010), the results indicate a positive and significant association between an individual's health status and the likelihood of being employed. Compared with males, females are found to be less likely to be working. A negative and significant ($p<0.05$) birth order effect on adult employment status is found across all columns. Individuals whose parents separated or divorced very early in their childhood life are less likely to be working at age 30. This result is consistent with the evidence on the many mechanisms through which divorces affect children. The disruption does not only lead to economic hardships in the short run; it also affects the educational attainment and the labor market success of the child in the long run.

2.5.5 Childhood Chronic Poverty Experience and Adult Poverty Status

The relationship between childhood chronic poverty experiences and the likelihood that an individual experiences an adult poverty spell at age 30 is presented in column 6 of Table 2.3. The results indicate a 'vicious cycle' of poverty. After accounting for years of completed schooling, the current health and employment status of an individual in Table 2.9, individuals with chronic poverty experiences in the first ten years of their childhood life are found to be more likely to experience a poverty spell as a young adult. From Table 2.9, the coefficients of the poverty indices are positive and statistically significant (except for the BCD index). The conditional size correlation is 0.121 for the F index, 0.044 for the BCD index, 0.119 for the HZ index, 0.163 for the count index and 0.110 for the average poverty gap.

Being economically active is negatively associated with an adult poverty spell experience. Moreover, having a good health status lowers the risk of adult poverty experiences with a coefficient estimate of -0.10 across all columns. The results also indicate that, an additional year spent in school by an individual significantly reduces the likelihood of being poor with a coefficient estimate between 0.22 and 0.26 across all columns. Oreopoulos

(2007) and Hofmarcher et al. (2019) found similar results. Females are more likely to experience a poverty spell at age 30 compared with males. Similarly, compared with whites, non-whites have a greater chance of experiencing an adult poverty spell at age 30.

2.5.6 Childhood Chronic Poverty Experience and Forming Own Households

The results presented in column 4 of Table 2.3 indicate no significant associations between the likelihood that an individual forms their own household and early childhood chronic poverty for all the chronic poverty measures studied. Table 2.7 include additional controls. From Table 2.7, we find that females are more likely to form their own households by age 25 compared with males; and non-whites are less likely to form their own households compared with whites. Individuals born to older mothers at birth are also less likely to form their own households and live independently by age 25. Compared with individuals who grew up in the North East Region, individuals who grew up in other Regions are more likely to form their own households by age 25.

2.5.7 The Choice between Chronic Poverty Measures

From the above sections, there is no single chronic poverty index that has the greatest size association across the various adult outcome variables studied; that is, no poverty index "performs best" across all outcome variables. I compare the different regression models under study using the coefficient of determination, R-squared, together with the statistical significance of the poverty indices. The R-squared measures the strength of the relationship between the different models (i.e., models involving the different poverty indices) and the dependent variable. It explains the percentage of the variance in the adult outcomes that the poverty indices explain. The R-squared values are reported in Table 2.3 for all outcomes. The HZ index has the highest R-squared in the health regression and the coefficient is significant at 1 percent. The F index has the highest R-squared in the employment regression and the coefficient is significant at 1 percent. Both the F and HZ indices have the highest R-squared in the schooling and the likelihood of forming own household regressions. The count index has the highest R-squared in the propensity of teen birth and adult poverty experience regressions. Therefore, no chronic poverty measure consistently explains the variation observed across all the adult outcomes studied. Two of the sophisticated poverty measures, the F and HZ indices, that account for the more nuanced aspects of poverty such as poverty gaps and the spacing of poverty, work best in explaining most of the adult outcomes examined in this study. It is worth noting that the weights assigned to poverty and non-poverty spells in the F and HZ indices sum up to 1, thus, the range of their values are similar.

2.5.8 Robustness Checks

The robustness of the results is explored in many ways, first by examining different specifications of the outcome variables. Alternatively specifying the child's education as a binary variable equal to 1 if a child completed at least 12 grades of school and 0 otherwise and labor market success as a binary variable equal to 1 if an individual is currently in the labor force were tried. The regression results from these specifications are qualitatively similar to those reported in this paper. In addition, the results from the adult health regressions are robust to good health specified as excellent or very good. The results are available in Table 2.13 of the Appendix.

Research in the areas of economics and epidemiology offer support for the fetal origins hypothesis which indicates that, prenatal conditions have significant effects on the development of an individual ranging from infancy through to adulthood (see Almond and Currie, 2011 for a review of the literature). Although the HZ measure assigns greater weight to earlier poverty spells (for bigger values of the parameter δ), it is still likely the case that the harmful effect of experiencing poverty *in utero*, a critical period of life, is being subdued by the averaging of poverty spells over the years. Therefore, I examine the association between adult outcomes and being poor in the prenatal year. The results, provided in the Appendix Table 2.13, indicate that individuals whose families experienced poverty in the prenatal year are less likely to report a good adult health status (p<0.05) and more likely to experience poverty (p<0.10) in early adulthood. The results indicate no significant associations between prenatal poverty experience and completed schooling years, employment or the propensity of teenage births. Therefore, although poverty in the prenatal period is linked to worse adult outcomes, poverty experiences in the years that follow are also harmful to a child's development.

The sensitivity of the results to different childhood time periods is tested by evaluating separately, the association of chronic poverty experiences from birth to age 5 with adult outcomes and the association of chronic poverty experiences from age 6 to 10 with adult outcomes. The results, provided in the Appendix (panel B of Table 2.13), indicate that chronic poverty experiences up to age 5 matter more for the schooling and health outcomes. This is consistent with the findings of Duncan et al. (2012) who find that family income from ages 0 to 5 matter more in explaining completed schooling than family income from ages 6 to 10. The results also reveal that chronic poverty experiences beyond age 5 has a greater size association with poor labor market outcomes and adult poverty experiences. Thus, chronic poverty regardless of its timing during the first ten years of life is linked to adverse outcomes in adulthood. The coefficients of the poverty indices measured from age 6 to 10 and from birth to age 5 are both statistically insignificant in the teen birth and formation of own household regressions.

I test the robustness of the results to attrition. For the first ten years of life, 1,549 children were present in the PSID survey. However, by age 25, only 1,047 of them remained as respondents yielding an attrition rate of 32 percent.[8] Does being poor predict attrition? Are the poor more likely to exit the PSID survey? If this assertion holds, then it suggests that the estimates of the association between early chronic poverty and later life outcomes found in this study may well be a lower bound. The result provided in Table 2.12 of the Appendix, indicates that being poor in the 10th year increases the log odds of exiting the PSID survey at age 25 by 0.52 points (p<0.01). In spite of missing some poor children,

[8]It is worth mentioning that there were some children who were missing even though their families responded; indicating that they lost their lives. This group make up less than 1 percent of the sample.

I find significantly large associations between early childhood chronic poverty and adult outcomes. This makes the findings in this study robust. It also indicates that the extent of harm suffered by children born in the United States between the late 60s and early 70s due to chronic poverty may be greater than imagined.

2.6 Conclusion

This paper compares the performance of recently developed measures of chronic poverty with measures that do not take into account the timing, spacing and closeness of poverty spells. The association of chronic poverty experienced between birth and age 10 with adverse outcomes in adulthood are examined. The chronic poverty measures examined comprise the Foster (2009) measure, the Bossert, Chakravarty and D'Ambrosio (2012) measure, and the Hoy and Zheng (2011) measure. The most common two poverty measures employed in the empirical poverty literature, the count index and the average poverty gap are also employed. The Foster (2009) measure emphasizes the number of periods spent in poverty without taking into account different temporal patterns of poverty experiences. However, unlike the count index, the Foster (2009) accounts for the gaps in poverty. The Bossert et al. (2012) measure emphasizes the closeness of poverty by assigning greater weights to poverty spells that occur in a string of two or more contiguous poverty spells. The Hoy and Zheng (2011) emphasizes the timing and spacing of poverty spells by assigning greater weights to earlier poverty spells and poverty spells that occur closer in time to each other.

Using PSID data from the late 1960s to the early 2000s, this paper is the first to document the association between chronic poverty experiences of children and their later life outcomes using these three recently developed chronic poverty measures. I find that time spent living in poverty as a child still matters in explaining poor adult outcomes even after controlling for a host of demographic and family background characteristics. The results suggest large and robust associations between early childhood chronic poverty and completed schooling, adult health status, employment status, propensity of teen births and adult poverty status measured as late as ages 25 and 30.

However, there is no single measure that explains the variation observed in all of the adult outcomes according to the coefficient of determination. The results reveal that the count index does well in explaining the variation in the propensity of teen births and adult poverty experiences; whilst the Hoy and Zheng (2011) and Foster (2009) indices do well in explaining the variation in the schooling, health, labor market and the likelihood of living independently outcomes. Since each poverty measure captures different properties of childhood poverty, it may be the case that different outcomes simply depend on different characteristics of childhood poverty. More research using different data sets should be done before any firm conclusions on this matter can be reached.

Beyond establishing relationships between childhood poverty measures and adult outcomes, I find other results of interest. In particular, the finding that there exists a negative effect of parental divorces beyond that of reducing family income and pushing children into poverty on some adult outcomes is of major interest. The results suggest that children whose parents separated or divorced in their early childhood years complete fewer years of school and are more likely to be poor as young adults. This result is consistent with the evidence on the many mechanisms through which divorces affect children. Compared with children from other Regions, children born in the South Region fare worse as young adults. This suggests that economic difficulties lie ahead of them. Racial disparities are also observed across all outcomes studied. Children of color are the least successful in escaping poverty. The results also reflect an inter-generational transmission of poverty. Children with early chronic poverty experiences are found to be more likely to experience an adult poverty spell.

The findings from this research point future empirical work in the direction of childhood chronic poverty estimations that is over and beyond a count approach when examining its long run impacts on later life achievements. For example, the National Longitudinal Survey of Youth may be combined with these recently developed poverty measures to examine outcomes not included in the Panel Study of Income Dynamics such as cognitive test scores, non-cognitive skill assessments, high school grades, occupation choices, health behaviors, substance use, etc.

Table 2.3: Association between Childhood Chronic Poverty and Adult Outcomes

	Schooling (years)	Good Health	Teen Birth	Form Own Households	Working Now	Currently Poor
Panel A						
F Index	-0.197***	-0.143**	0.176***	-0.035	-0.133***	0.302***
R-squared	0.038	0.013	0.022	0.001	0.012	0.083
Panel B						
BCD Index	-0.129***	-0.098**	0.067**	-0.020	-0.127**	0.190**
R-squared	0.016	0.007	0.003	0.000	0.012	0.035
Panel C						
HZ Index	-0.197***	-0.147***	0.174***	-0.034	-0.132***	0.296***
R-squared	0.038	0.014	0.022	0.001	0.012	0.080
Panel D						
Count Index	-0.171***	-0.152**	0.243***	-0.002	-0.105**	0.371***
R-squared	0.029	0.013	0.044	0.000	0.007	0.130
Panel E						
APG	-0.190***	-0.113**	0.182***	-0.020	-0.119**	0.293***
R-squared	0.035	0.007	0.024	0.000	0.009	0.081
No. of Obs	1,041	1,044	1,047	1,047	720	728

Notes: τ = 0.00 for F; γ = 1.2 for BCD, δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. Standardized coefficients are reported in each column; they represent the standard deviation changes in a given outcome associated with a one standard deviation increase in a given covariate. The constant term is included in the regressions but not reported in this table. The education outcome is estimated using a Linear Regression model. All other outcome variables are estimated using a Logistic Regression model. Employment and Adult poverty status are measured at age 30. Standard errors are clustered at the household level. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.4: Association between Childhood Chronic Poverty and Completed Schooling

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	-0.104***				
BCD Index		-0.052**			
HZ Index			-0.104***		
Count Index				-0.077**	
APG					-0.095***
Child is a 'switcher'	-0.060*	-0.068*	-0.062*	-0.063*	-0.059*
Child is female	0.085***	0.085***	0.084***	0.087***	0.085***
Child is Non-white	-0.004	-0.021	-0.004	-0.012	-0.009
Head's Education	0.120***	0.129***	0.119***	0.125***	0.123***
Second generation citizen	0.051*	0.053*	0.051*	0.052*	0.052*
Age of Mother	0.392*	0.182*	0.390*	0.395*	0.395*
Age of $Mother^2$	-0.349	-0.360	-0.348	-0.344	-0.349
Birth Order	-0.402**	-0.403**	-0.403**	-0.379**	-0.386**
Birth Order$\times$Age of Mother	0.320*	0.309	0.320*	0.290	0.306
Region (Ref: North East)					
North Central	-0.045	-0.047	-0.046	-0.045	-0.045
South	-0.088*	-0.095**	-0.088*	-0.094**	-0.089*
West	0.003	0.005	0.003	-0.002	0.004
Regression Statistics					
Obs	1,041	1,041	1,041	1,041	1,041
R-squared	0.087	0.081	0.087	0.083	0.086
P-value	2.09e-17	1.58e-17	3.53e-17	1.64e-16	4.20e-18

Notes: The dependent variable is the completed years of schooling estimated using a Linear regression model. τ = 0.00 for F; γ = 1.2 for BCD, δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. Standardized coefficients are reported in each column; they represent the standard deviation changes in years of completed schooling associated with a one standard deviation increase in a given covariate. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent

Source: Author, PSID Data 1968-2003.

Table 2.5: Association between Childhood Chronic Poverty and Good Health

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	-0.158**				
BCD Index		-0.096**			
HZ Index			-0.162***		
Count Index				-0.174**	
APG					-0.125**
Child is a 'switcher'	0.047	0.037	0.045	0.052	0.048
Child is female	0.017	0.019	0.017	0.022	0.018
Child is Non-white	0.205**	0.176*	0.209**	0.205**	0.181**
Head's Education	0.201**	0.215**	0.199**	0.199**	0.210**
Age of Mother	0.369	0.386	0.366	0.354	0.369
Age of $Mother^2$	-0.501	-0.501	-0.500	-0.476	-0.487
Birth Order	-0.144	-0.134	-0.148	-0.084	-0.098
Birth Order×Age of Mother	0.189	0.159	0.194	0.130	0.142
Region (Ref: North East)					
North Central	0.062	0.008	0.014	0.019	0.013
South	0.223	0.043	0.058	0.047	0.050
West	0.242	0.047	0.042	0.040	0.046
Regression Statistics					
Obs	1,044	1,044	1,044	1,044	1,044
R-squared	0.056	0.050	0.057	0.057	0.051
P-value	0.054	0.093	0.049	0.030	0.102

Notes: Dependent variable is a binary variable equal to 1 if an individual is in good health. This outcome is estimated using a logistic regression model. τ = 0.00 for F; γ = 1.2 for BCD, δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. Standardized coefficients are reported in each column; they represent the standard deviation changes in adult health status associated with a one standard deviation increase in a given covariate. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent

Source: Author, PSID Data 1968-2003.

Table 2.6: Association between Childhood Chronic Poverty and Teen Birth

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	0.017				
BCD Index		-0.065*			
HZ Index			0.014		
Count Index				0.086*	
APG					0.029
Child is a 'switcher'	0.048	0.055	0.049	0.040	0.046
Child is female	0.312***	0.311***	0.312***	0.310***	0.312***
Child is non-white	0.212***	0.231***	0.214***	0.192***	0.210***
Head's Education	-0.142***	-0.149***	-0.142***	-0.135***	-0.141***
Age of Mother	-0.992***	-1.019***	-0.993***	-0.962***	-0.987***
Age of $Mother^2$	0.886**	0.900**	0.887**	0.868**	0.884**
Birth Order	0.341	0.322	0.340	0.324	0.337
Birth Order×Age of Mother	-0.225	-0.186	-0.223	-0.231	-0.227
Region (Ref: North East)					
North Central	0.156**	0.157**	0.156**	0.151*	0.155**
South	0.149*	0.158*	0.150*	0.144	0.147*
West	0.139**	0.135**	0.139**	0.148**	0.140**
Regression Statistics					
Obs	1,047	1,047	1,047	1,047	1,047
R-squared	0.187	0.190	0.187	0.193	0.188
P-value	0.000	0.000	0.000	0.000	0.000

Notes: Dependent variable is a binary variable equal to 1 if an individual had a baby before age 20. This outcome is estimated using a logistic regression. τ = 0.00 for F; γ = 1.2 for BCD, δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. Standardized coefficients are reported in each column; they represent the standard deviation changes in the propensity of teen births associated with a one standard deviation increase in a given covariate. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.7: Association between Childhood Chronic Poverty and Formation of Own Household

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	-0.017				
BCD Index		0.011			
HZ Index			-0.015		
Count Index				0.018	
APG					-0.011
Child is a 'switcher'	0.049	0.047	0.048	0.045	0.048
Child is female	0.132***	0.131***	0.132***	0.131***	0.132***
Child is non-white	-0.213***	-0.220***	-0.214***	-0.221***	-0.215***
Head's Education	-0.053	-0.049	-0.053	-0.048	-0.052
Age of Mother	-0.885***	-0.882***	-0.885***	-0.879***	-0.883***
Age of $Mother^2$	0.748**	0.749**	0.748**	0.754**	0.748**
Birth Order	0.379	0.385	0.379	0.379	0.381
Birth Order×Age of Mother	-0.307	-0.319	-0.308	-0.316	-0.311
Region (Ref: North East)					
North Central	0.173***	0.172***	0.173***	0.172***	0.173***
South	0.182***	0.179***	0.182***	0.178***	0.181***
West	0.136***	0.138***	0.137***	0.138***	0.137***
Regression Statistics					
Obs	1,047	1,047	1,047	1,047	1,047
R-squared	0.068	0.068	0.068	0.068	0.068
P-value	0.000	0.000	0.000	0.000	0.000

Notes: Dependent variable is a binary variable equal to 1 if an individual formed their own household. This outcome is estimated using a logistic regression model. τ = 0.00 for F; γ = 1.2 for BCD, δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. Standardized coefficients are reported in each column; they represent the standard deviation changes in the formation of own households associated with a one standard deviation increase in a given covariate. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.8: Association between Childhood Chronic Poverty and Employment Status

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	-0.097*				
BCD Index		-0.107*			
HZ Index			-0.097*		
Count Index				-0.017	
APG					-0.082
Child is a 'switcher'	-0.083*	-0.087*	-0.084*	-0.093*	-0.084*
Child is female	-0.256***	-0.259***	-0.256***	-0.254***	-0.257***
Child is non-white	0.053	0.053	0.053	0.023	0.045
Child's Education	0.134**	0.139***	0.134**	0.135**	0.135**
Child is in good health	0.109***	0.109***	0.109***	0.104***	0.108***
Age of Mother	0.119	0.142	0.044	0.146	0.129
Age of $Mother^2$	-0.317	-0.336	-0.320	-0.324	-0.322
Birth Order	-0.722**	-0.746**	-0.723**	-0.692**	-0.708**
Birth Order×Age of Mother	0.786*	0.809**	0.788*	0.728*	0.771*
Region (Ref: North East)					
North Central	-0.016	-0.016	-0.016	-0.018	-0.015
South	-0.068	-0.067	-0.067	-0.078	-0.071
West	-0.101	-0.067	-0.101	-0.097	-0.099
Regression Statistics					
Obs	719	719	719	719	719
R-squared	0.111	0.114	0.111	0.106	0.110
P-value	0.000	0.000	0.000	0.000	0.000

Notes: Dependent variable is a binary variable equal to 1 if an individual is currently employed. This outcome is estimated using a logistic regression model. τ = 0.00 for F; γ = 1.2 for BCD, δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. Standardized coefficients are reported in each column; they represent the standard deviation changes in employment status associated with a one standard deviation increase in a given covariate. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.9: Association between Childhood Chronic Poverty and Adult Poverty Status

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	0.121**				
BCD Index		0.044			
HZ Index			0.119**		
Count Index				0.163**	
APG					0.110*
Child is a 'switcher'	0.075	0.088*	0.078	0.067	0.076
Child is female	0.188**	0.184**	0.187**	0.169**	0.189**
Child is non-white	0.167**	0.202***	0.167**	0.149*	0.176**
Child's Education	-0.252***	-0.257***	-0.252***	-0.236***	-0.254***
Child is working	-0.141***	-0.147***	-0.142***	-0.159***	-0.145***
Child is in good health	-0.102**	-0.095**	-0.102**	-0.104*	-0.101**
Age of Mother	-0.840	-0.866	-0.847	-0.837	-0.837
Age of $Mother^2$	0.801	0.793	0.807	0.795	0.795
Birth Order	0.180	0.141	0.180	0.120	0.155
Birth Order×Age of Mother	-0.021	0.052	-0.020	0.020	-0.004
Region (Ref: North East)					
North Central	0.088	0.090	0.089	0.077	0.086
South	0.065	0.082	0.065	0.069	0.067
West	0.069	0.057	0.068	0.081	0.066
Regression Statistics					
Obs	728	728	728	728	728
R-squared	0.255	0.246	0.254	0.264	0.252
P-value	0.000	0.000	0.000	0.000	0.000

Notes: Dependent variable is a binary variable equal to 1 if an individual is currently poor. This outcome is estimated using a logistic regression model. τ = 0.00 for F; γ = 1.2 for BCD, δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. Standardized coefficients are reported in each column; they represent the standard deviation changes in adult poverty status associated with a one standard deviation increase in a given covariate. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent

Source: Author, PSID Data 1968-2003.

2.7 Appendix

Description of Data

The Panel Study of Income Dynamics (PSID) of the United States is an ongoing study which began in 1968 with a nationally representative sample of 5000 families consisting of 15,000 individuals. This survey has now followed siblings and parents for over thirty years. Questions about educational attainment, labor supply, health, income levels, household wealth, among others are posed in each round. Interviews are conducted either by telephone, in person or by mail. The average length of an interview is 20 minutes. The interview period is roughly between February and October of each year. A majority of interviews are conducted in Spring.

Total family income in the PSID is defined as the sum of taxable income from all sources of the husband, wife and all other earners in the household plus transfer incomes. Transfer income includes aid to dependent children (ADC), aid to families with dependent children (AFDC), aid to dependent children with unemployed fathers (ADCU), social security, retirement compensation, unemployment compensation, workmen's compensation, alimony, child support, gifts from relatives, etc. Any form of non-cash benefits such as food stamps, housing subsidies, etc are excluded from the definition of total income.

I follow children from birth to age 30. Only children with complete information are included in this study. If a household head withdraws from the survey, then information on every member of the household (including children) becomes unavailable. However, children who move from their original birth households to other households due to several reasons are still followed and their information are available.

Description of variables

Years of completed schooling - Equal to the most recent report of highest schooling years completed by an individual at age 25.

Good health - Equal to 1 if health status is reported as excellent, very good or good, and 0 if health status is reported as poor or fair at age 25.

Working now - Equal to 1 if an individual is currently working and 0 if the status of employment is reported as unemployed/looking for work/student/housewife/living in an institution/disabled/other at age 30.

Teen birth - Equal to 1 if an individual had a child before age 20, and 0 otherwise.

Form own households - Equal to 1 if an individual is the head of the household or wife (including cohabitants), and 0 otherwise by age 25.

Currently poor - Equal to 1 if an individual's equivalent household income is below the appropriate poverty line (i.e., poor), and 0 otherwise (i.e., non-poor) at age 30.

Household Head's Education - Equal to 1 if the household head of an individual completed at least a high school degree, and 0 otherwise.

Switcher - Equal to 1 if an individual experienced disruptions in the household environment due to parental marital divorce/separation during the first ten years of life and 0 otherwise.

Age of mother - It is a continuous measure of the age of mother at time of birth of an individual.

Second generation citizens - Equal to 1 if at least one parent of an individual is an immigrant (i.e., born outside the United States), and 0 otherwise.

Completed High School - Equal to 1 if an individual completed at least 12 grades of school, and 0 otherwise.

Table 2.10: Equivalence scales by different family sizes

Family size	Poverty Line	Implied equivalence scale
1	5,593	1.00
2	7,231	1.29
3	8,573	1.53
4	10,989	1.96
5	13,007	2.33
6	14,696	2.63
7	16,656	2.98
8	18,512	3.31
9 or more	22,083	3.95

Source: Author's calculation of Equivalence scales derived from 1985 poverty thresholds published by the United States Census Bureau.

Table 2.11: Correlation between Pairs of Poverty Indices

	$F_{0.00}$	$F_{0.25}$	$F_{0.50}$	$F_{0.75}$	$BCD_{1.2}$	$BCD_{1.5}$	$BCD_{2.0}$	$HZ_{0.2}$	$HZ_{0.5}$	$HZ_{1.0}$	HCI_3	HCI_5	APG
$F_{0.00}$	1.00												
$F_{0.25}$	0.98	1.00											
$F_{0.50}$	0.94	0.95	1.00										
$F_{0.75}$	0.78	0.79	0.83	1.00									
$BCD_{1.2}$	0.88	0.89	0.90	0.88	1.00								
$BCD_{1.5}$	0.74	0.75	0.77	0.86	0.95	1.00							
$BCD_{2.0}$	0.65	0.66	0.69	0.80	0.90	0.99	1.00						
$HZ_{0.2}$	0.99	0.98	0.93	0.78	0.89	0.74	0.65	1.00					
$HZ_{0.5}$	0.98	0.96	0.92	0.77	0.88	0.74	0.65	0.99	1.00				
$HZ_{1.0}$	0.96	0.94	0.90	0.75	0.86	0.72	0.64	0.98	0.99	1.00			
HCI_3	0.78	0.81	0.68	0.45	0.57	0.40	0.32	0.78	0.77	0.75	1.00		
HCI_5	0.80	0.82	0.87	0.58	0.66	0.50	0.41	0.80	0.80	0.78	0.78	1.00	
APG	0.96	0.95	0.90	0.73	0.83	0.68	0.59	0.95	0.92	0.88	0.77	0.79	1.00

Notes: The table presents the correlation (pearson) coefficients between pairs of poverty indices. $F_{0.00}$, $F_{0.25}$, $F_{0.50}$, and $F_{0.75}$ represent the F Index for duration cutoff fractions, τ equal 0.00, 0.25, 0.50, and 0.75 respectively. $BCD_{1.2}$, $BCD_{1.5}$, and $BCD_{2.0}$ represent the BCD Index for different sensitivity parameters, γ equal 1.20, 1.50, and 2.00. $HZ_{0.2}$, $HZ_{0.5}$, and $HZ_{1.0}$ represent the HZ Index for different sensitivity parameters, δ equal 0.2, 0.5 and 1.0. HCI_3 and HCI_5 represent the head count index for the duration cutoff fractions 3/10 and 5/10 respectively.
Source: Author, PSID Data 1968-2003.

Table 2.12: Likelihood of Exiting the PSID Survey

Poor in the last wave (10th year)	0.517***
	(0.151)
Regression Statistics	
R-squared	0.015
P value	0.000
N	1,549

Notes: The dependent variable is a binary variable equal to 1 if an individual was present by the 10th year but left the PSID survey by age 25. This outcome is estimated using a logistic regression model. Birth cohort dummy variables are included but are not reported. Household level cluster-robust standard errors are reported in parentheses. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.13: Association between Childhood Chronic Poverty and Adult Outcomes

	Schooling (years)	Good Health	Teen Birth	Form Own Households	Working Now	Currently Poor
			Panel A			
Poor in prenatal year	-0.039	-0.196**	0.039	0.020	-0.006	0.113*
R-squared	0.096	0.059	0.188	0.068	0.106	0.250
			Panel B (Age 0-5 years)			
F Index	-0.100***	-0.180***	-0.015	-0.014	-0.071	0.105**
R-squared	0.103	0.061	0.187	0.068	0.109	0.237
BCD Index	-0.087**	-0.139**	-0.049	0.001	-0.054	0.066
R-squared	0.101	0.056	0.189	0.068	0.108	0.232
HZ Index	-0.099***	-0.182***	-0.012	-0.015	-0.070	0.103**
R-squared	0.102	0.061	0.187	0.068	0.109	0.236
			Panel C (Age 6-10 years)			
F Index	-0.077***	-0.101*	0.041	-0.016	-0.090*	0.144**
R-squared	0.099	0.049	0.188	0.0684	0.111	0.248
BCD Index	-0.062**	-0.069	0.010	-0.009	-0.100**	0.117*
R-squared	0.098	0.046	0.187	0.068	0.113	0.243
HZ Index	-0.076***	-0.102*	0.039	-0.012	-0.093*	0.144**
R-squared	0.099	0.049	0.188	0.068	0.111	0.248
			Panel D (Age 1-10 years)			
F Index	-0.104***	-0.158**	0.017	-0.017	-0.097*	0.121**
R-squared	0.087	0.054	0.187	0.068	0.111	0.255
BCD Index	-0.052**	-0.096**	-0.065*	0.011	-0.107*	0.044
R-squared	0.081	0.050	0.190	0.068	0.114	0.246
HZ Index	-0.104***	-0.162***	0.014	-0.015	-0.097*	0.119**
R-squared	0.087	0.057	0.187	0.068	0.111	0.254
N	1,041	1,044	1,047	1,047	720	728
Full controls	yes	yes	yes	yes	yes	yes

Notes: Standardized coefficients are reported in each column; they represent the standard deviation changes in a given outcome associated with a one standard deviation increase in a given covariate. All other covariates and the constant term are included in the regressions but not reported in this table. The education outcome is estimated using a Linear Regression model. All other outcome variables are estimated using a Logistic Regression model. Employment and Adult poverty status are measured at age 30. Standard errors are clustered at the household level. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.14: Association between Childhood Chronic Poverty and Adult Outcomes

	Completed High School	Good Health	Currently in the labor force
Panel A			
F Index	-0.141***	-0.060	-0.150**
R-squared	0.113	0.036	0.124
Panel B			
BCD Index	-0.062*	-0.037	-0.163**
R-squared	0.103	0.035	0.129
Panel C			
HZ Index	-0.142***	-0.062	-0.147**
R-squared	0.113	0.036	0.123
Panel D			
Count Index	-0.196***	-0.113**	-0.018
R-squared	0.121	0.040	0.111
Panel E			
APG	-0.142***	-0.062	-0.144**
R-squared	0.113	0.036	0.122
N	1,041	1,044	724
Full controls	yes	yes	yes

Notes: Standardized coefficients are reported in each column; they represent the standard deviation changes in a given outcome associated with a one standard deviation increase in a given covariate. τ = 0.00 for F; γ = 1.2 for BCD, δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. All other covariates and the constant term are included in the regressions but not reported in this table. All outcome variables are estimated using a Logistic Regression model. Currently in the labor force is measured at age 30. Standard errors are clustered at the household level. Significance levels: $*$: 10 percent $**$: 5 percent $* * *$: 1 percent

Source: Author, PSID Data 1968-2003.

Table 2.15: Association (unstandardized coefficient) between Childhood Chronic Poverty and Completed Schooling

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	-2.890***				
	(0.883)				
BCD Index		-0.384**			
		(0.162)			
HZ Index			-2.886***		
			(0.898)		
Count Index				-0.843***	
				(0.257)	
APG					-4.36***
					(0.390)
Child is a 'switcher'	-0.421*	-0.475*	-0.432*	-0.368	-0.412*
	(0.242)	(0.243)	(0.242)	(0.247)	(0.243)
Child is female	0.461***	0.461***	0.460***	0.466***	0.461***
	(0.162)	(0.163)	(0.162)	(0.161)	(0.162)
Child is non-white	-0.024	-0.119	-0.023	0.001	-0.050
	(0.216)	(0.214)	(0.217)	(0.209)	(0.215)
Head's Education	0.684***	0.737***	0.682***	0.631***	0.703***
	(0.204)	(0.205)	(0.204)	(0.205)	(0.205)
Second generation citizen	1.058*	1.115*	1.057*	1.032*	1.076*
	(0.614)	(0.617)	(0.615)	(0.610)	(0.615)
Age of Mother	0.173*	0.182*	0.173*	0.160	0.175*
	(0.099)	(0.101)	(0.098)	(0.100)	(0.099)
Age of $Mother^2$	-0.003	-0.003	-0.003	-0.002	-0.003
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Birth Order	-0.670**	-0.672**	-0.671**	-0.579**	-0.643**
	(0.262)	(0.264)	(0.262)	(0.258)	(0.262)
Birth Order×Age of Mother	0.014*	0.014	0.014*	0.012	0.014
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
Region (Ref: North East)					
North Central	-0.273	-0.284	-0.275	-0.282	-0.270
	(0.281)	(0.281)	(0.281)	(0.279)	(0.281)
South	-0.486*	-0.523**	-0.484*	-0.509**	-0.491*
	(0.257)	(0.260)	(0.257)	(0.255)	(0.258)
West	0.0234	0.0408	0.0254	-0.0101	0.0325
	(0.265)	(0.268)	(0.266)	(0.265)	(0.265)
Regression Statistics					
N	1,041	1,041	1,041	1,041	1,041
R-squared	0.087	0.081	0.087	0.090	0.086
P-value	2.09e-17	1.58e-17	3.53e-17	1.64e-16	4.20e-18

Notes: The dependent variable is the completed years of schooling estimated using a Linear regression model. Household level cluster-robust standard errors are reported in parentheses. τ = 0.00 for F; γ = 1.2 for BCD and δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. The parameter estimate and standard errors of APG have been multiplied by 10^4. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.16: Association between Childhood Chronic Poverty and Good Health

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	-3.146**				
	(1.224)				
BCD Index		-0.508**			
		(0.258)			
HZ Index			-3.229***		
			(1.233)		
Count Index				-1.097***	
				(0.406)	
APG					-4.069**
					(2.039)
Child is a 'switcher'	0.236	0.183	0.223	0.351	0.241
	(0.456)	(0.454)	(0.455)	(0.449)	(0.460)
Child is female	0.068	0.074	0.065	0.057	0.068
	(0.319)	(0.319)	(0.319)	(0.319)	(0.319)
Child is Non-white	0.843**	0.718*	0.858**	0.885**	0.741**
	(0.383)	(0.374)	(0.387)	(0.378)	(0.372)
Head's Education	0.819**	0.871**	0.813**	0.724*	0.854**
	(0.371)	(0.374)	(0.371)	(0.382)	(0.370)
Age of Mother	0.117	0.121	0.116	0.0905	0.116
	(0.142)	(0.142)	(0.141)	(0.140)	(0.142)
Age of $Mother^2$	-0.003	-0.003	-0.003	-0.002	-0.003
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Birth Order	-0.172	-0.158	-0.177	-0.005	-0.117
	(0.439)	(0.435)	(0.440)	(0.441)	(0.437)
Birth Order×Age of Mother	0.006	0.005	0.006	0.002	0.004
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
Region (Ref: North East)					
North Central	0.062	0.034	0.060	0.055	0.054
	(0.490)	(0.489)	(0.489)	(0.483)	(0.493)
South	0.223	0.167	0.228	0.197	0.195
	(0.546)	(0.535)	(0.547)	(0.538)	(0.543)
West	0.242	0.265	0.241	0.216	0.259
	(0.622)	(0.626)	(0.621)	(0.616)	(0.624)
Regression Statistics					
N	1,044	1,044	1,044	1,044	1,044
Pseudo R-squared	0.056	0.050	0.057	0.063	0.051
P-value	0.054	0.093	0.049	0.030	0.102

Notes: Dependent variable is a binary variable equal to 1 if an individual is in good health. This outcome is estimated using a logistic regression model. Household level cluster-robust standard errors are reported in parentheses. τ = 0.00 for F; γ = 1.2 for BCD and δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. The parameter estimate and standard errors of APG have been multiplied by 10^4. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.17: Association between Childhood Chronic Poverty and Teen Birth

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	0.393				
	(0.863)				
BCD Index		-0.391*			
		(0.211)			
HZ Index			0.306		
			(0.854)		
Count Index				0.318	
				(0.261)	
APG					1.060
					(1.440)
Child is a 'switcher'	0.273	0.309	0.277	0.232	0.262
	(0.233)	(0.234)	(0.233)	(0.239)	(0.234)
Child is female	1.375***	1.373***	1.375***	1.375***	1.376***
	(0.223)	(0.223)	(0.223)	(0.222)	(0.223)
Child is non-white	0.988***	1.077***	0.993***	0.931***	0.977***
	(0.240)	(0.236)	(0.240)	(0.245)	(0.239)
Head's Education	-0.655***	-0.690***	-0.657***	-0.614**	-0.653***
	(0.241)	(0.240)	(0.241)	(0.241)	(0.240)
Age of Mother	-0.355***	-0.366***	-0.355***	-0.348***	-0.353***
	(0.126)	(0.122)	(0.126)	(0.127)	(0.126)
Age of $Mother^2$	0.005**	0.006**	0.005**	0.005**	0.005**
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Birth Order	0.460	0.436	0.459	0.435	0.455
	(0.323)	(0.322)	(0.323)	(0.321)	(0.321)
Birth Order×Age of Mother	-0.008	-0.007	-0.008	-0.008	-0.008
	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)
Region (Ref: North East)					
North Central	0.758**	0.764**	0.758**	0.755**	0.754**
	(0.381)	(0.380)	(0.381)	(0.382)	(0.381)
South	0.667*	0.709*	0.669*	0.656*	0.658*
	(0.387)	(0.382)	(0.387)	(0.387)	(0.388)
West	0.900**	0.870**	0.898**	0.930**	0.901**
	(0.434)	(0.435)	(0.435)	(0.435)	(0.434)
Regression Statistics					
N	1,047	1,047	1,047	1,047	1,047
Pseudo R-squared	0.187	0.190	0.187	0.189	0.188
P-value	0.000	0.000	0.000	0.000	0.000

Notes: Dependent variable is a binary variable equal to 1 if an individual had a baby before age 20. This outcome is estimated using a logistic regression. Household level cluster-robust standard errors are reported in parentheses. τ = 0.00 for F; γ = 1.2 for BCD and δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. The parameter estimate and standard errors of APG have been multiplied by 10^4. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent

Source: Author, PSID Data 1968-2003.

Table 2.18: Association between Childhood Chronic Poverty and Formation of Own Household

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	-0.331				
	(0.815)				
BCD Index		0.059			
		(0.194)			
HZ Index			-0.290		
			(0.808)		
Count Index				-0.095	
				(0.210)	
APG					-0.340
					(1.380)
Child is a 'switcher'	0.242	0.231	0.240	0.249	0.240
	(0.199)	(0.200)	(0.199)	(0.204)	(0.199)
Child is female	0.508***	0.507***	0.508***	0.509***	0.508***
	(0.142)	(0.142)	(0.142)	(0.142)	(0.142)
Child is non-white	-0.868***	-0.896***	-0.871***	-0.867***	-0.876***
	(0.186)	(0.182)	(0.186)	(0.186)	(0.186)
Head's Education	-0.213	-0.198	-0.212	-0.218	-0.208
	(0.171)	(0.169)	(0.171)	(0.171)	(0.170)
Age of Mother	-0.277***	-0.276***	-0.277***	-0.279***	-0.277***
	(0.099)	(0.099)	(0.099)	(0.099)	(0.099)
Age of $Mother^2$	0.004**	0.004**	0.004**	0.004**	0.004**
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Birth Order	0.448	0.455	0.448	0.457	0.451
	(0.292)	(0.291)	(0.292)	(0.291)	(0.291)
Birth Order×Age of Mother	-0.010	-0.010	-0.010	-0.010	-0.010
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Region (Ref: North East)					
North Central	0.738***	0.735***	0.737***	0.737***	0.737***
	(0.226)	(0.227)	(0.226)	(0.226)	(0.226)
South	0.713***	0.700***	0.713***	0.710***	0.710***
	(0.216)	(0.216)	(0.216)	(0.215)	(0.216)
West	0.772***	0.777***	0.772***	0.768***	0.774***
	(0.259)	(0.259)	(0.259)	(0.260)	(0.259)
Regression Statistics					
N	1,047	1,047	1,047	1,047	1,047
Pseudo R-squared	0.068	0.068	0.068	0.068	0.068
P-value	0.000	0.000	0.000	0.000	0.000

Notes: Dependent variable is a binary variable equal to 1 if an individual formed their own household. This outcome is estimated using a logistic regression model. Household level cluster-robust standard errors are reported in parentheses. τ = 0.00 for F; γ = 1.2 for BCD and δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. The parameter estimate and standard errors of APG have been multiplied by 10^4. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.19: Association between Childhood Chronic Poverty and Employment Status

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	-2.104*				
	(1.229)				
BCD Index		-0.643*			
		(0.338)			
HZ Index			-2.118*		
			(1.239)		
Count Index				-0.044	
				(0.312)	
APG					-2.815
					(1.957)
Child is a 'switcher'	-0.439*	-0.465*	-0.447*	-0.495*	-0.447*
	(0.252)	(0.252)	(0.251)	(0.254)	(0.252)
Child is female	-1.039***	-1.053***	-1.039***	-1.029***	-1.041***
	(0.214)	(0.215)	(0.214)	(0.216)	(0.214)
Child is non-white	0.237	0.239	0.237	0.0886	0.201
	(0.266)	(0.264)	(0.266)	(0.260)	(0.261)
Child's Education	0.126**	0.131***	0.126**	0.127**	0.127**
	(0.051)	(0.051)	(0.051)	(0.050)	(0.051)
Child is in good health	1.026***	1.030***	1.028***	0.971**	1.015***
	(0.373)	(0.374)	(0.373)	(0.377)	(0.373)
Age of Mother	0.043	0.051	0.044	0.052	0.046
	(0.168)	(0.167)	(0.167)	(0.170)	(0.168)
Age of $Mother^2$	-0.002	-0.002	-0.002	-0.002	-0.002
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Birth Order	-0.963**	-0.996**	-0.964**	-0.914**	-0.942**
	(0.431)	(0.433)	(0.431)	(0.429)	(0.431)
Birth Order×Age of Mother	0.029*	0.030**	0.029*	0.026*	0.028*
	(0.015)	(0.015)	(0.015)	(0.015)	(0.015)
Region (Ref: North East)					
North Central	-0.067	-0.068	-0.069	-0.078	-0.067
	(0.341)	(0.341)	(0.340)	(0.340)	(0.340)
South	-0.283	-0.281	-0.281	-0.327	-0.295
	(0.323)	(0.323)	(0.324)	(0.322)	(0.322)
West	-0.592	-0.589	-0.591	-0.560	-0.580
	(0.360)	(0.360)	(0.360)	(0.358)	(0.358)
Regression Statistics					
N	719	719	719	719	719
Pseudo R^2	0.111	0.114	0.111	0.106	0.110
P-value	0.000	0.000	0.000	0.000	0.000

Notes: Dependent variable is a binary variable equal to 1 if an individual is currently employed. This outcome is estimated using a logistic regression model. Household level cluster-robust standard errors are reported in parentheses. τ = 0.00 for F; γ = 1.2 for BCD and δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. The parameter estimate and standard errors of APG have been multiplied by 10^4. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Table 2.20: Association between Childhood Chronic Poverty and Adult Poverty Status

	Model 1	Model 2	Model 3	Model 4	Model 5
F Index	2.651**				
	(1.289)				
BCD Index		0.215			
		(0.247)			
HZ Index			2.617**		
			(1.288)		
Count Index				1.398***	
				(0.462)	
APG					3.730*
					(1.960)
Child is a 'switcher'	0.514	0.590*	0.530	0.338	0.517
	(0.368)	(0.356)	(0.366)	(0.384)	(0.368)
Child is female	0.822**	0.799**	0.818**	0.777**	0.819**
	(0.369)	(0.361)	(0.368)	(0.361)	(0.367)
Child is non-white	0.850**	1.019***	0.854**	0.584	0.898**
	(0.418)	(0.396)	(0.418)	(0.460)	(0.412)
Child's Education	-0.269***	-0.271***	-0.268***	-0.240***	-0.268***
	(0.081)	(0.080)	(0.081)	(0.082)	(0.081)
Child is working	-0.836***	-0.859***	-0.837***	-0.971***	-0.855***
	(0.291)	(0.292)	(0.291)	(0.300)	(0.291)
Child is in good health	-0.662**	-0.672**	-0.662**	-0.764**	-0.662**
	(0.337)	(0.333)	(0.337)	(0.349)	(0.335)
Age of Mother	-0.322	-0.328	-0.324	-0.260	-0.318
	(0.244)	(0.244)	(0.243)	(0.243)	(0.246)
Age of $Mother^2$	0.005	0.005	0.005	0.005	0.005
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Birth Order	0.346	0.286	0.348	0.268	0.309
	(0.650)	(0.648)	(0.649)	(0.650)	(0.655)
Birth Order×Age of Mother	-0.003	0.0004	-0.003	-0.002	-0.002
	(0.022)	(0.022)	(0.022)	(0.022)	(0.022)
Region (Ref: North East)					
North Central	0.423	0.442	0.424	0.383	0.417
	(0.542)	(0.531)	(0.542)	(0.586)	(0.541)
South	0.300	0.397	0.301	0.263	0.319
	(0.552)	(0.530)	(0.553)	(0.598)	(0.545)
West	0.463	0.407	0.457	0.606	0.450
	(0.593)	(0.583)	(0.594)	(0.643)	(0.589)
Regression Statistics					
N	728	728	728	728	728
Pseudo R-squared	0.255	0.246	0.254	0.277	0.252
P-value	0.000	0.000	0.000	0.000	0.000

Notes: Dependent variable is a binary variable equal to 1 if an individual is currently poor. This outcome is estimated using a logistic regression model. Household level cluster-robust standard errors are reported in parentheses. τ = 0.00 for F; γ = 1.2 for BCD and δ=0.2 for HZ and cutoff fraction is 5/10 for the count index. North East is the baseline Region. The parameter estimate and standard errors of APG have been multiplied by 10^4. All regressions include a constant term and birth cohort dummy variables which are not reported. See the Appendix for an explanation of all variables. Significance levels: $*$: 10 percent $**$: 5 percent $***$: 1 percent
Source: Author, PSID Data 1968-2003.

Chapter 3

Women legislators in Africa and foreign aid

3.1 Introduction

The Chamber of Deputies in Rwanda which is currently made up of over 60 percent of women, has been the national parliament with the highest share of women legislators worldwide for over ten years now. Many national parliaments in Africa have a share of women legislators that is substantially above the world average such as Senegal (43%), South Africa (41%), Mozambique (39%), and many other countries.[1] What makes this fact puzzling is that it cannot be attributed to an electorate that values gender equality and having women in political leadership positions. The population share in Africa that agrees with

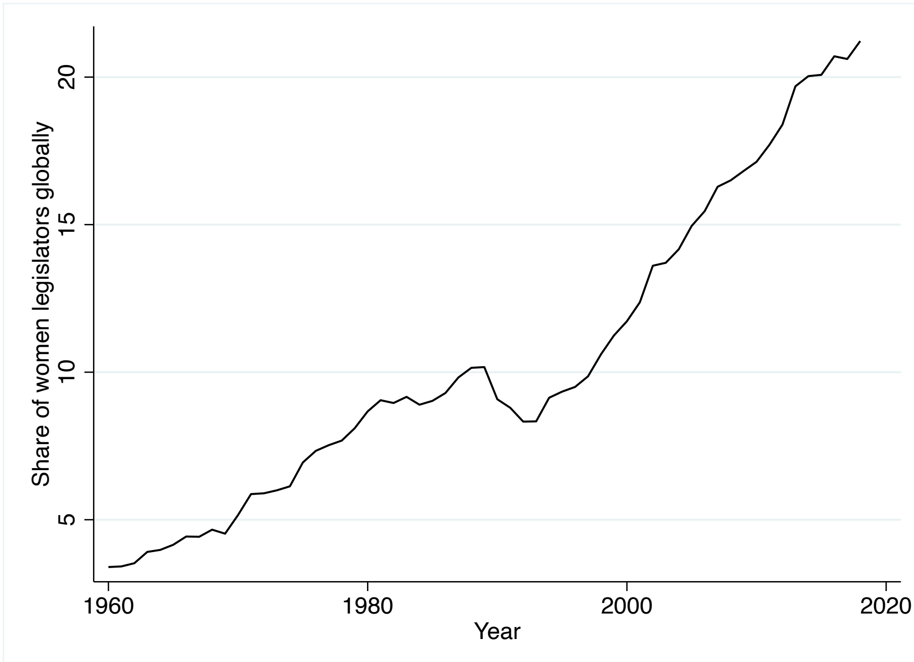
[1] Between 2012 and 2016, the last five years covered in this study, 22 national parliaments in Africa have an average share of women legislators that is above the world average of 20 percent during this time period.

the statement that "men make better political leaders than women do" is 62 percent, and this is substantially higher than the average of 43 percent for non-African countries.[2] In stark contrast, "inclusive economic growth ... and promoting social inclusion" and commitments to "ensure gender equality and women's and girls' empowerment" have become important goals of the international donor-recipient community over the last few decades. The words in the quotation marks are all written in the *first* paragraph of the Addis Ababa Financing for Development Action Agenda (United Nations, 2015), which was endorsed by the UN General Assembly in July 2015. Conscious efforts have been made by the United Nations to bridge the global gender inequality gap in the last few decades with notable movements like the 1995 Beijing Conference on Women, the Convention on the Elimination of all Forms of Discrimination against Women (CEDAW) and the various rounds of Financing for Development Action Agendas, where gender equality and women empowerment have successively moved up in the priority ranking over the years. If an electoral story cannot explain the high level of women participation in many African parliaments, then the question arises whether the priority shift in the international donor-recipient community towards gender issues is part of the explanation. The goal of the current paper is to investigate whether there is a link between the gender composition in parliaments and foreign aid allocations in Africa.

Women are underrepresented in most national parliaments globally. Figure 3.1 displays

[2]The men/women-leader question was asked in several rounds of the World Value Survey (Inglehart, Haerpfer, Moreno, Welzel, Kizilova, Diez-Medrano, Lagos, Norris, Ponarin and Puranen, 2014).

Figure 3.1: Share of women legislators globally (1960-2018)

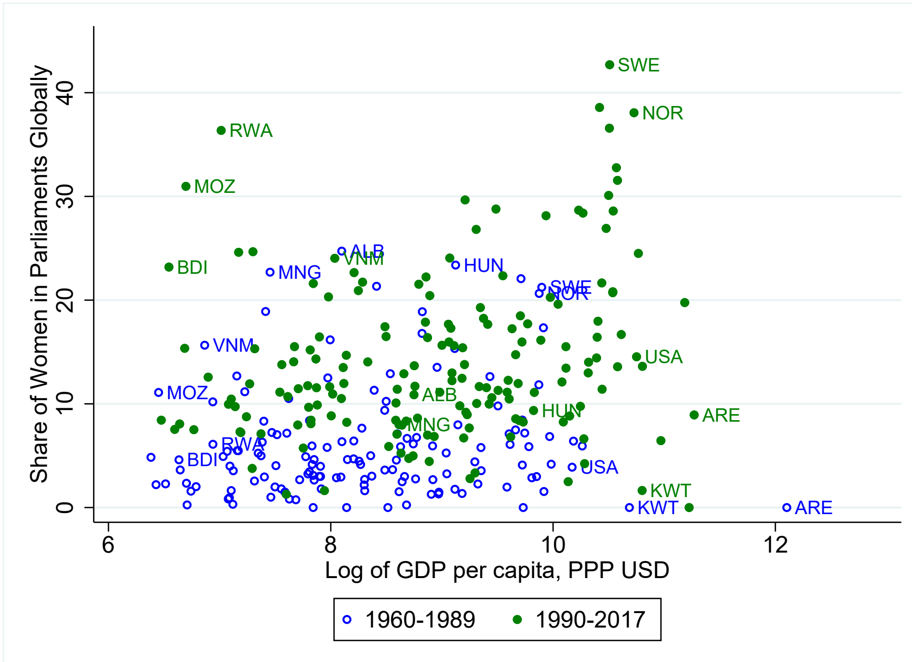


Note: This figure shows the development of the average share of women legislators in national parliaments across 200 countries between 1960 and 2018 based on a new data set created by the authors.

the average share of women legislators for the period 1960 to 2018.[3] The figure depicts a positive trend in the share of women legislators. The average share of women is currently set at about 23 percent from below 10 percent in the 1990s and from as low as 5 percent in the 1970s. The sharp decline in the share of women legislators observed after 1990 can be attributed to the collapse of the Soviet Union as quotas kept women participation in the Soviet Union and its satellite states in Eastern Europe high (Saxonberg, 2000; Inglehart et al., 2003). In order to get a better understanding of this dynamics, Figure 3.2 differen-

[3]We extracted the share of women legislators from html and pdf files made available to us by the Inter-Parliamentary Union (IPU) and combined that data with the Varieties of Democracy (Vdem) database (Coppedge et al., 2020; Pemstein et al., 2020). This produces to our knowledge the most comprehensive data set on the gender composition in national legislators covering 200 countries between 1960 and 2018.

Figure 3.2: Share of women legislators vs. GDP per capita (1960–2017)



Note: This figure shows a scatter plot between the average share of women legislators and income per capita (taken from PWT 9.1) for two periods, the 1960–1989 period (hollow dots) and the 1990–2017 period (full dots).

tiates between the Soviet era (1960–1989) and the post-Soviet era (1990–2017). It plots the share of women legislators against the log of GDP per capita for these two time periods. We observe that the range in the share of women legislators is largely independent of income. For the period 1960-1989, we observe that Western European countries and Eastern European countries associated with the Soviet Union (e.g., Norway, Sweden, Hungary, Albania, Mongolia) have the greatest share of women legislators among all countries. In Africa, the share of women legislators is at most 10 percent for countries like Mozambique, Burundi and Rwanda. In the post-Soviet era, we observe three significant changes in the distribution of the share of women legislators worldwide. First, a sharp decline in the share of women legislators in Eastern European countries (e.g., Albania, Hungary); second, a continual increase in the share of women legislators in Western European countries (e.g., Sweden and Norway); and third, a spike in the share of women legislators for countries in Africa and a few other developing countries. For instance, the share of women legislators in Burundi's parliament rose from 5 to 22 percent whilst the share of women legislators in Albania's parliament fell from 25 to 10 percent in the post Soviet era. Sweden's share of women legislators increased from 21 during 1960-1989 to 43 percent for the period 1990-2017. Norway also experienced a rise in the share of women legislators from about 21 in the Soviet era to about 38 percent for the period 1990-2017. For this period 1990-2017, low income countries like Rwanda and Mozambique rank high globally. From an average of 2 percent in 1970 and an average of 8 percent in 1990, women currently fill 21 percent of parliamentary seats in Africa. In 2016, women made up about 64 percent of legislators in Rwanda, the highest in the world, and Senegal had the second highest record in Africa with 43 percent of its parliamentary seats filled with women.

In this paper we focus on the role of international donors and development agencies as a possible explanation for this dynamic in Africa. What makes this explanation plausible is that gender equality and women empowerment has successively risen in the priority ranking of aid donors over the last few decades. Given that many African countries receive considerable amounts of foreign aid, the question arises whether there is a link between the share of women members of parliament and foreign aid. We find a strong and robust relationship between the lagged share of women and current aid conditional on lagged aid using recipient-period panel data. We estimate that an increase in the share of women legislators by 10 percent for a recipient country is associated with an increase of 1.3 percent in aid. The result is robust to controls that are typically used in aid allocation regressions. Further, the results indicate that the increased women representation in Africa's parliaments is achieved through gender quotas in the form of reserved seats. We estimate that introducing reserved seats for women in parliament leads to 53 percent in additional aid. Our results further reveal that although democratic countries receive more aid, donors do not tailor their gender-selective aid towards more democratic African countries.

A paper that relates closely to the paper here is the cross-sectional study by Bush (2011), which examines the factors that influence the adoption of gender quotas in national parliaments. The author finds a positive correlation between international development assistance and the likelihood that a country adopts gender quotas suggesting that a larger aid-dependence is a key factor affecting the implementation of gender quotas. Our study uses this insight as a starting point and uses panel data to exploit the time variation in aid flows to analyze aid-selectivity from a recipient country perspective where we show that conditional on existing aid amounts, recipient countries get remunerated with more aid if they increase the share of women legislators or if they adopt reserved seats for women in parliament.[4] This suggests that aid-selectivity in terms of gender equality may incentivize recipient countries to adopt policies that increase the share of women legislators. We believe that this effect is properly identified as we control for current aid-levels and country fixed effects, and use an appropriate lag structure in our regressions. Hicks et al. (2016)

[4]Without controlling for existing aid flows, our coefficient on the share of women legislators almost doubles, which would be consistent with the notion that countries with larger aid amounts are more likely to implement policies that foster gender equality.

study the relationship between the gender composition of national parliaments in donor countries and aid flows from donors and show that a higher share of women legislators in a donor country leads to an increase in aid efforts both in total and as a percentage of GDP. Our paper, in contrast, focuses on the gender composition of national legislatures in aid-recipient countries. With this, the current paper relates to the broader literature on aid-selectivity investigating the policy- and poverty selectivity of aid allocations (Dollar and Levine, 2006; Annen and Knack, 2018; Eichenauer and Knack, 2016; Knack et al., 2011; Annen and Moers, 2017; Annen and Knack, 2020; World Bank, 2005; Burnside and Dollar, 2000; Alesina and Weder, 2002). This literature examines the extent to which aid flows are targeted to recipient countries with sound economic and political institutions. For example, Annen and Knack (2020) show that the policy-selectivity of aid has increased substantially starting in the early 1990s. They estimate that since the year 2000 more than half of the global aid budget is allocated by policy-selective donors. Note that most of this literature examines aid-selectivity from a donor-perspective by focusing on donor specific allocation decisions with the purpose to assess aid and donor quality for example (i.e. Knack et al., 2011; Birdsall and Kharas, 2010; Easterly and Williamson, 2011; Roodman, 2012). In the study here, we focus on the recipient-perspective as it sheds light on the incentive structure for recipients to implement policy reforms produced by policy-selective aid as formally analyzed in Annen and Knack (2020). Given the large differences in donor motivations, it is important to know whether overall the adoption of certain policies "pays off" in terms of additional aid amounts for aid-recipient countries. We provide evidence in support of policy-selective aid for policies that improve gender equality and women empowerment in aid-recipient countries. Finally, the paper also relates to a literature that describes the gender composition in national legislatures across countries such as Wängnerud (2009), Sawer (2000), Hughes and Paxton (2019), and Saxonberg (2000). This literature documents the evolution of women's participation in all aspects of the political process such as voting, grassroot women's mobilization, organizing and joining political parties, and getting elected to parliament and the policies (e.g., the type of electoral system, political party ideology, political party nomination process, the strength of women's movement) that affect it.

The remainder of the paper is structured as follows. Section 2 provides some background regarding the electoral attitudes towards gender equality in Africa, and the role of quota systems in shifting the gender composition in favor of women legislators. Section 3 documents how gender equality and women empowerment has increased in the priority ranking of the international donor-recipient community in the last two decades. Section 4 presents the quantitative analysis presenting the aid-selectivity regressions. The final section concludes the paper.

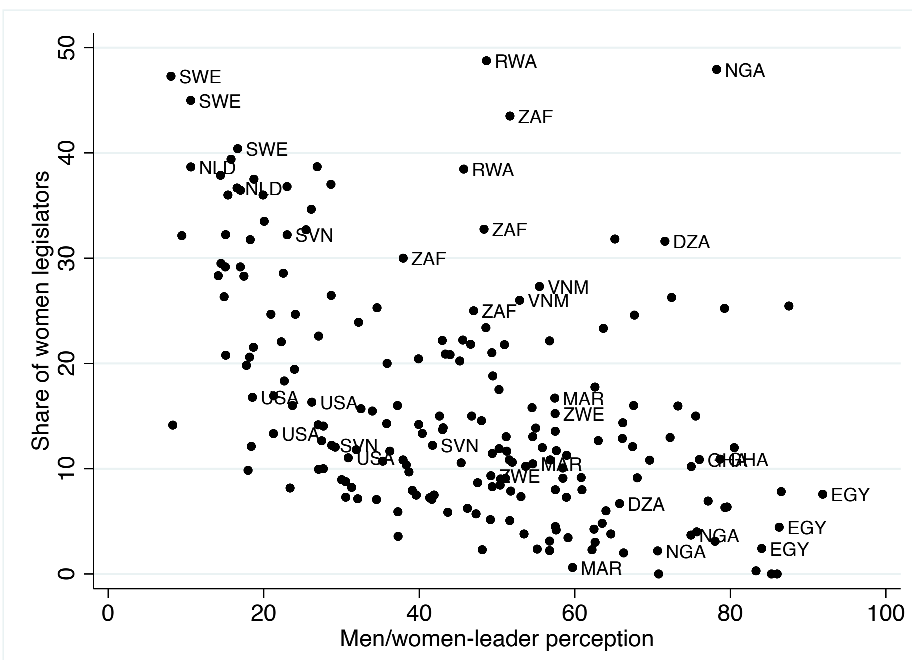
3.2 Background

Citizens in low income countries usually have their priorities set on access to food (i.e., three square meals), potable water, good roads, stable supply of electricity, medical supplies and health facilities, and the alleviation of poverty and are less likely to have discussions surrounding issues like gender equality, global warming, etc. Moreover, African societies are mostly known to be patriarchal and religious, whereby discriminatory practices exist within the family unit and negative stereotyping of women exist in the public space (Tøraasen, 2017). The status of women is low in most African countries. Therefore it is puzzling that low-middle income countries (especially those in Africa) are recording greater shares of women legislators starting from the 1990s. Research has shown that the beliefs and attitudes of electorates influence the number of women who run for and who get elected into political positions (Paxton and Kunovich, 2003; Inglehart et al., 2003; Arceneaux, 2001). For instance, the religious (Muslim) leaders in Touba, the second largest city in Senegal, rebelled against the adoption of women quotas in Senegal and presented an all-male list of 100 candidates for the local elections held in 2014 (Tøraasen, 2017). At the outset of the 2003 war in the Democratic Republic of Congo, the Congolese government and male decision makers opposed the involvement of women in the peace negotiations with the excuse that women are not fighters and so cannot make any significant contributions in drafting the peace agreement (Mpoumou, 2004).

To elucidate voter attitudes towards gender inequality we use the World Value Survey (Inglehart et al., 2014). The World Value Survey began in 1981 and has 6 completed waves: wave 1: 1981-1984, wave 2: 1990-1994, wave 3: 1995-1998, wave 4: 1999-2004, wave 5: 2005-2009, wave 6: 2010-2014. The purpose of this survey (WVS) is to measure people's beliefs and values across countries, how they change over time, and what impact they have on their social and political life. For example, the survey asks respondents for their views regarding men vs. women as political leaders. Beginning in 1995, respondents were asked whether they "(a) Agree strongly, (b) Agree, (c) Disagree, (d) Strongly disagree, (e) Don't know" with the statement: "Men make better political leaders than women do". We believe that the responses to this question yield a useful proxy for the feminist attitude of the electorate across countries and time. We coded responses that either strongly agree or agree as one and all other responses as zero and then calculated the average per country for each survey year. For African countries, that measure equals 62 percent on average, which is substantially above the average of 43 percent for non-African countries. The relationship between the share of women in parliament and the share of respondents who agree that men make better political leaders than women is presented in Figure 3.3.

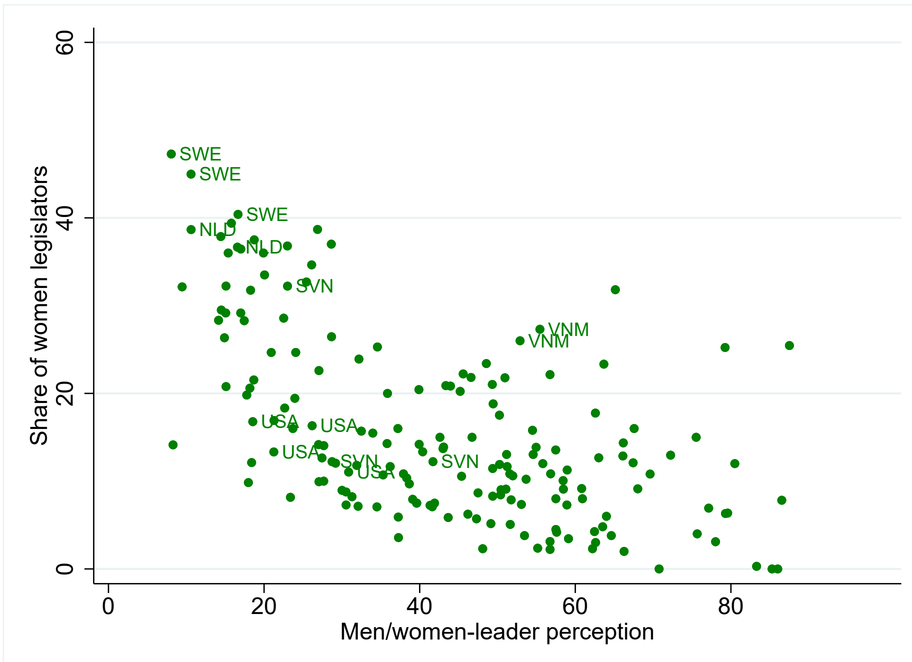
This figure shows overall a fairly strong negative correlation (correlation coefficient of -0.493) between the share of women legislators and the share of respondents who believe men make better political leaders than women which is expected. However, upon closer inspection one can see that there are qualitative differences between African- and non-African countries. For Non-African countries, the results show a negative relationship between women legislators and men/women-leader perceptions (see Figure 3.4). For countries in Africa, in contrast, we observe a positive association between these two variables (see Figure 3.5). For example in Rwanda, when 45 percent of survey respondents believe that men make better political leaders, the share of women legislators is set at 38 percent and when 49 percent of survey respondents believe that men make better political leaders, the share of women legislators is set at 48 percent. We observe a similar pattern for South Africa and Zimbabwe. Table 3.1 investigates this relationship further using OLS and country fixed effect (FE) regressions. Column (1) confirms the overall negative correlation that can be seen in Figure 3.3. This regression also includes a dummy variable for Africa, which is positive and statistically significant at 1 percent. Conditional on men/women-leader perceptions, African countries have a share of women legislators that is 9.2 percentage points higher than non-African countries, which highlights again our main point for this paper that an electoral story cannot (fully) explain the high women participation in African legislators. Columns (2), (3), and (4) show regression results from regressions that use sub-

Figure 3.3: Share of women legislators vs. men/women-leader perceptions (1994–2016)



Note: This figure shows a scatter plot between the share of women legislators and "Men/women-leader perception" which is the share of respondents in a country who strongly agree or agree with the statement "Men make better political leaders than women do" taken from the World Value Surveys (WVS).

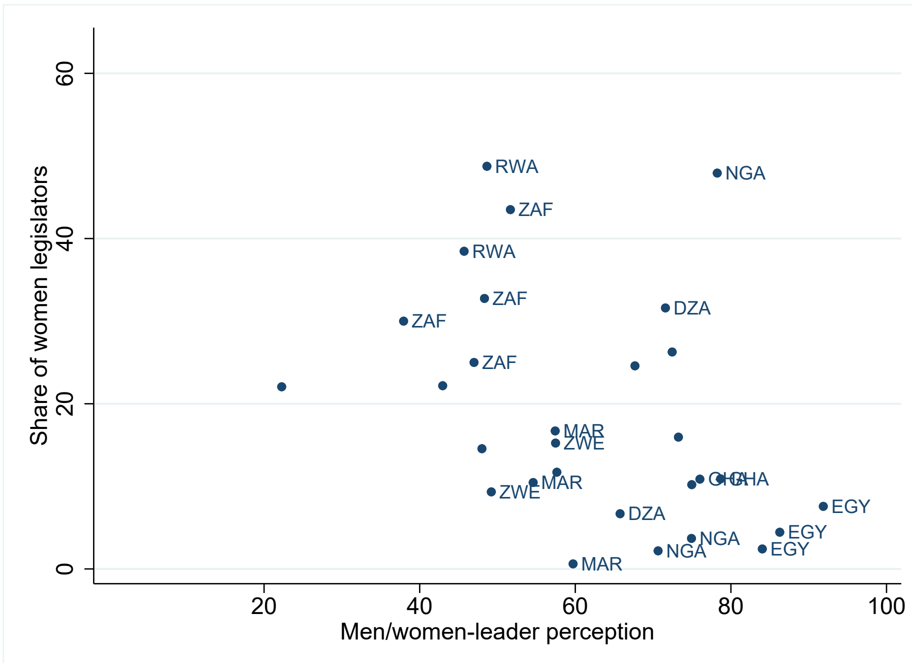
Figure 3.4: Share of women legislators vs. men/women-leader perceptions for Non-African countries (1994–2016)



Note: This figure shows for Non-African countries, a scatter plot between the share of women legislators and "Men/women-leader perception" which is the share of respondents in a country who strongly agree or agree with the statement "Men make better political leaders than women do" taken from the World Value Surveys (WVS).

samples and include country fixed effects. Columns (2) and (3) divide the full sample into upper and lower income countries respectively with the median GDP per capita being the divider, and Column (4) uses a sample that only includes African countries. These regressions show that the coefficient for the men/women-leader perception conditional on country fixed effects changes the sign depending on whether we run this regression among upper or lower income countries. In particular, for lower income countries and African countries this correlation is positive, which is contrary to what one expects if beliefs and attitudes of the electorate should explain the gender composition in national legislators. For African countries, the coefficient is large and statistically significant at the 10 percent level. Note,

Figure 3.5: Share of women legislators vs. men/women-leader perceptions for African countries (1994–2016)



Note: This figure shows for African countries a scatter plot between the share of women legislators and "Men/women-leader perception" which is the share of respondents in a country who strongly agree or agree with the statement "Men make better political leaders than women do" taken from the World Value Surveys (WVS).

however, that the results in Column (4) should be taken with a grain of salt as this regression has only very few observations. We conclude that the number of women who run and win elections in Africa can hardly be linked to its citizens becoming more supportive of women being in politics.

Another fact that supports this claim is that the adoption of gender quotas plays an important role in explaining the share of women legislators in Africa. Given that the status of women is low in these countries, women won't be elected into parliament in competitive elections. Instead, we observe what has been termed the "fast track" approach to the

Table 3.1: Share of women legislators and men/women-leader perception (1994–2016)

	(1)	(2)	(3)	(4)
Men/women-leader perception	-0.3198***	-0.3492***	0.1949	1.4841*
	(0.0372)	(0.1084)	(0.2544)	(0.7293)
Africa	9.2111***			
	(2.5010)			
Constant	29.4304***	31.9081***	1.3095	-73.0014
	(1.8267)	(4.4380)	(14.6054)	(45.1348)
Country FE	No	Yes	Yes	Yes
N	197	147	58	29
R-squared	0.32	0.12	0.02	0.25
F statistic	38.68	10.37	0.59	4.14

Dependent variable is the share of women legislators in the lower chamber of national parliaments (Source: IPU and V-Dem). "Men/women-leader perception" is the share of respondents in a country who strongly agree or agree with the statement "Men make better political leaders than women do" (Source: WVS). Column (1) uses all countries with Men/women-leader perception data, Columns (2) and (3) include countries with GDP per capita above and below the median GDP per capita respectively. Column (4) includes African countries only. Recipient level cluster-robust standard errors are reported in parenthesis. Significance levels : $*$: 10 percent $**$: 5 percent $***$: 1 percent

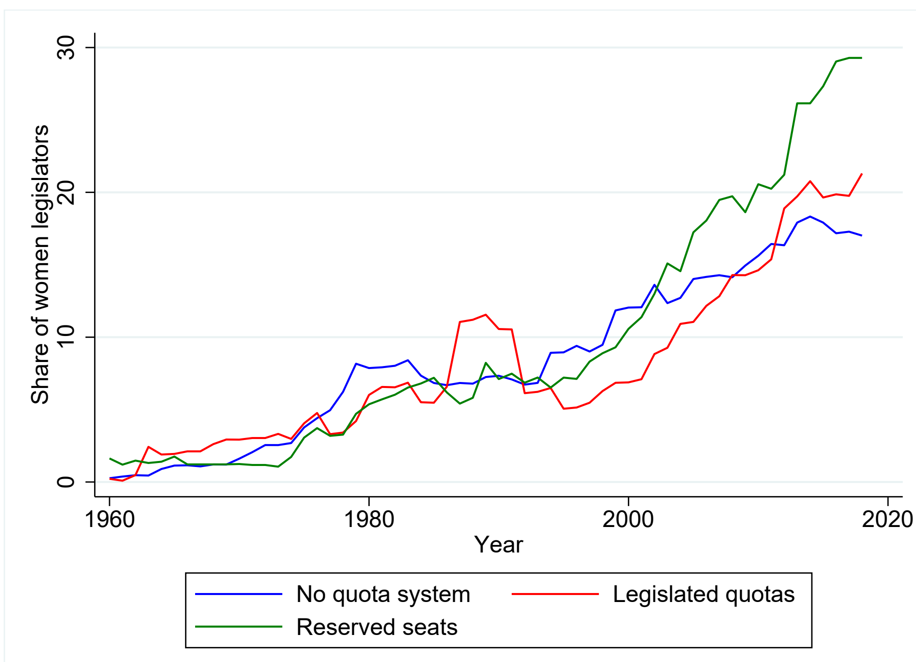
growing women's representation in politics, that is the adoption of legislated candidate quotas and seats reserved for women in parliaments. Gender quotas come in three forms: Voluntary party quotas, legislated candidate quotas, and reserved seats. Under a voluntary party quota system, political parties self-regulate the gender composition of the candidate party lists that they submit to the Electoral Commission. For instance, the African National Congress (ANC) party in South Africa has adopted a 50 percent gender quota for the party candidate lists that they submit to the electoral commission for their local and national elections. Under a legislated candidate quota system, political parties are required by law to regulate the gender composition of candidate lists submitted for elections. In some countries, state funding is provided to political parties who fulfill this quota requirement on their party lists. Countries may or may not impose legal sanctions on political parties who fail to meet this requirement. For example, in Senegal, the electoral commission (CENA) has the authority to reject the lists and exclude parties from competing in elections in case of non-compliance of the quota requirement (i.e. strong quota enforcement). Guinea imposes no legal sanctions on political parties that fail to meet the candidate quota requirement on their party lists (i.e. weak quota enforcement). Finally under the reserved seat system, a specified number of seats in the legislature is reserved for women. The women seats are either filled through elections involving women-only ballots (e.g. , Uganda) or filled through appointments made by leaders of the political parties (e.g. , Tanzania).[5] While the first two quota systems regulate the gender composition among the candidates, the reserved seat system targets the women electees directly. Thus, reserved seat quotas are more certain than candidate list quotas in achieving increased women participation in parliaments, particularly in societies where cultural barriers exist towards women to exercise their political rights. For instance, Burundi reserves 30 percent of its parliaments seats for women and requires that 1 in 4 candidates must be a woman. During the 2015 elections in Burundi, women made up 22 percent of the candidate lists but only 15 percent of the elected representatives were women. The remaining 15 percent had to be filled through co-optation; that is, women who obtained at least 5 percent of the votes cast but did not win the elections were selected into parliament (Brand, 2018).

We focus on legislated candidate quotas and reserved seats in this study because they are implemented at the national level. Data on legislated candidate quota and reserved seat is

[5]Table 3.9 in the appendix show how reserved seats are implemented in African countries.

taken from the Varieties of Democracy (V-dem) database (Coppedge et al., 2020; Pemstein et al., 2020). The list of countries who have implemented gender quotas are presented in Table 3.8 of the Appendix. Overall, 61 countries have implemented gender quotas in the form of legislated candidate party lists whilst 20 countries have implemented gender quotas in the form of special seats reserved for women in parliament globally. 16 of the 61 countries (representing 26 percent) with legislated candidate quotas are in Africa whereas 12 of the 20 countries (representing 60 percent) with reserved seats quotas are in Africa. The majority of the adoption of reserved seats in Africa occurred after the 1990s, even

Figure 3.6: Share of women legislators and quota systems in Africa (1960-2018)



The blue line graphs the average share of women legislators for countries in Africa without a quota system. The red line graphs the average share of countries with legislated candidate quotas and the green line is for countries with reserved seats. Data on quotas is from the Varieties of democracy (Vdem) database.

though some countries had reserved seats for women in the 1970s (e.g., Sudan in 1974 and

Tanzania in 1975). For instance, Kenya adopted reserved seats in 1997, Morocco in 2002, Rwanda in 2003, Somalia in 2004, Burundi in 2005, Zimbabwe in 2013 and Mauritania in 2013. Legislated candidate quotas were also enacted into law in Africa beginning in the 2000s. For example, candidate quotas were adopted in Djibouti in 2003, in Niger in 2004, in Angola in 2008, in Senegal in 2012 and in Guinea in 2013.

Figure 3.6 compares the evolution of the gender composition of African legislators for countries without a quota system, countries with a legislated quota system (with weak or strong enforcement), and countries with reserved seats for women. The figure reveals that between 1960 and 1990 the gender composition develops fairly similar in all countries irrespective of the quota system a country would eventually adopt. Also, countries with a legislated quota system follow fairly closely the dynamics of countries without a quota-system except for the last few years. For countries with reserved seats, in contrast, we observe a strong upward divergence that starts in the year 2000. As indicated earlier, this is the time where many countries started to implement reserved seats. Note that the average share of women legislators in African countries without a quota system equals 17 percent in 2016, which is 4 percentage points below the 2016 global average. For countries with reserved seats for women that average is 30 percent. Table 3.2 shows that reserved seats for women and legislated quotas with strong punishments substantially increase the shares of women legislators in Africa. This table reports regression results that include country- and year fixed effects. Conditional on those, legislated quotas with strong punishments and reserved seats each increase the share of women by about 10 percentage points. There is no such effect for quota systems that are weakly enforced. Figure 3.4 and Table 3.2 both highlight the importance of quota systems in increasing the share of women legislators in Africa. The question, of course, remains why countries in which the status of women is low will adopt such quota systems.

Table 3.2: Share of women legislators and quota systems in Africa (1960–2018)

	(1)	(2)	(3)
Legislated quota weak	-0.9691 (2.5063)		0.3859 (2.3964)
Legislated quota strong	8.4819** (3.3319)		10.0254*** (3.1305)
Reserved seats		8.5073** (3.2978)	8.9551*** (3.3107)
Country FE	Yes	Yes	Yes
N	2597	2597	2597
R-squared	0.49	0.51	0.52

Note: Dependent variable is the share of women legislators in the lower chamber of national parliaments. "Quota weak" and "quota strong" refer to countries with a legislated quota system with weak and strong enforcement respectively. "Reserved seats" refers to countries with reserved seats for women in their national legislature. The quota-system data is taken from Vdem. All regressions include country - and year fixed effects that are not reported. Recipient level cluster-robust standard errors are reported in parenthesis. Significance levels : $*$: 10 percent $**$: 5 percent $***$: 1 percent

In this paper we focus on the role of international donors and development agencies as a possible explanation for this dynamic in Africa. What make this explanation plausible is that gender equality and women empowerment has successively risen in the priority ranking of aid donors over the last few decades as we show in the next section. Given that many African countries receive considerable amounts of foreign aid, the natural question arises whether there is a link between the share of women legislators and foreign aid.

3.3 Donor Aid Priorities

In 1980 the United Nations Convention on the Elimination of all Forms of Discrimination Against Women (CEDAW) treaty was adopted by the UN General Assembly. A majority of countries have ratified this document (i.e., made it official) and agreed to be bound by its provision. Every four years, world leaders are expected to submit a report to the CEDAW committee highlighting measures that they have put in place to fulfill the mandate of the treaty. CEDAW-committee members debate on the reports and make recommendations to the world leaders on how they can continue to eradicate and eliminate discrimination against of all forms women in their respective countries. Between 1980 and 2015, all African countries except Somalia and Sudan have ratified the treaty.[6]

In 1995, the fourth world conference on women was held in Beijing, dubbed "Beijing Declaration and Platform for Action." This conference brought together world leaders, international organizations, and the media to advance the goals of equality, development, and peace for all women. The twelve strategic objectives and actions identified at the conference relate to poverty, education, health, the environment, and politics among others all targeted at improving the situation of women in these areas. One of the strategic objectives and actions deals with "Women in Power and Decision-making." All governments were encouraged to "take measures to ensure women's equal access to and full participation in power structures and decision-making". This includes the following:

[6] The median ratification year is 1989, the earliest year of ratification was done by Cape Verde in 1980 and the latest year of ratification was done by South Sudan in 2015.

- Establish gender balance in governmental bodies and committees, the judiciary, and all governmental and public administration positions.
- Protect and promote the equal rights of women to engage in political activities and to freely associate.
- Recognize that shared work and parenting between women and men promote women's increased position in public life.
- Monitor and evaluate progress on the representation of women through regular collection and analysis of data.

Political parties were also advised to:

- Examine party structures and procedures to eliminate discrimination against women's participation.
- Develop initiatives to encourage women's participation and incorporate gender issues in their political agenda.

The OECD (2018) report on aid allocations reveal that there has been an increase in gender-targeted aid allocations. Several donors have made gender equality a core priority in their aid spending. For Sweden, Iceland, Ireland, Canada, Belgium, Australia, Netherlands, and New Zealand, at least 50 percent of their aid allocations were targeted at programs and initiatives that had gender equality as the objective. Canada introduced a feminist international development assistance policy document known as the Feminist International Assistance Policy (FIAP) in 2017. This aid action policy emphasizes the economic and political participation of women and girls over other avenues of gender equality in the countries that Canada awards development assistance to (Morton et al., 2020). For the period 2015-2016, seven of the ten countries that received the largest amounts of funding for programs targeting gender equality as the primary objective were in Africa (i.e., Tanzania, Ethiopia, DR Congo, Kenya, Mali, Mozambique, Uganda), and one third of these aid amounts was awarded to the government and civil society sector.

In the next section, we evaluate the selectivity of aid allocations to the share of women legislators in Africa. That is, conditional on being previous aid recipients, do recipient countries in Africa who have high women representation in parliaments receive additional aid amounts? If the results answers this question affirmatively, then it provides some evidence on the adoption of reserved seats and legislated candidate quotas in Africa where cultural barriers exist for women in the public space.

3.4 Gender Selectivity of Aid Allocations

This study focuses on the share of women legislators in the lower chamber of national parliaments. We created the data set by extracting the gender composition information from many HTML and PDF files each capturing one election that were made available to us by the Inter Parliamentary Union (IPU). IPU tracks elections for all its members irrespective of the quality of the members' elections. We complemented our data with the "Varieties of Democracy" database (Vdem) (Coppedge et al., 2020; Pemstein et al., 2020) producing – to our knowledge – the most comprehensive data set on the gender composition in national parliaments covering 200 countries between 1960 and 2018.[7] As the number of legislators in parliament is only available for election years, we fill in data for the non-election years for a maximum of 5 years after an election is held. It is important to note that most aid-recipient countries included in our sample are not full fledged democracies. We mea-

[7] Combining the two data sources produces a data set that covers 22 more countries than the Vdem data set on its own.

sure democracy by an assessment of the competitiveness of elections using the XRCOMP measure in the Polity IV data set (see Marshall et al., 2017). This measure refers to the "Competitiveness of Executive Recruitment" and is defined as follows: "Competitiveness refers to the extent that prevailing modes of advancement give subordinates equal opportunities to become superordinates." (Gurr, 1974, p. 1483). Marshall et al. (2017) note that "[f]or example, selection of chief executives through popular elections matching two or more viable parties or candidates is regarded as competitive." Such a country receives a value of 3. In our sample of aid-recipients in Africa, the median equals 1.5 for the time period covered in our regression analysis, which is between 1990 and 2017.[8] A value of 1.5 refers to an in-between situation where "Chief executives are determined by hereditary succession, designation, or by a combination of both" on one hand, and a transitional arrangement "between selection (ascription and/or designation) and competitive elections" on the other. Only two countries, South Africa and Mauritius, have the perfect score of 3 for all the years covered in our analysis. We prefer this democracy measure over – for example – the composite measures such as "POLITY" or "POLITY2" or other composite measures such as Freedom House as this measure captures *one* important and tangible dimension of a democracy, namely whether the transfer of power is regulated and succeeds as a result of competitive elections.[9] It is also plausible that donors are more likely to base

[8]We focus on the post-Soviet era because this is the era that is most relevant in terms of donor aid-selectivity as discussed in the previous section and the upper bound in our time frame is dictated by data availability. The coverage of PWT 9.1 ends in 2017.

[9]Note that for countries in which the transfer of power is unregulated – i.e. for example through forceful seizures of power – XRCOMP has a value of zero. For composite measures of democracy as for example POLITY2, other components such as "constraints on the executive" enter additive in into the measure, which implies that competitive elections can be perfectly substituted with constraints on the executive, which may

their aid allocation decision on such a tangible property related to democracy.

Data on development assistance is taken from the Organization for Economic Co-operation and Development (OECD). We define total aid as the total gross flow of development aid (excluding emergency aid, food aid and debt relief) from government donors, multilateral organizations, NGOs, private foundations, and the private sector. We hereby follow Annen and Kosempel (2009) and capture development aid resulting in new cash inflows (gross aid). Aid is measured in constant USD.

All the data used in the aid allocation regressions is averaged across 4-year periods starting with the 1990–1993 period and ending with the 2014–2017 period. This is done because our main variable of interest, the share of women legislators, changes about every 4 years for the average country. In addition, averaging has the advantage of smoothing the data series, which is desirable as aid can be fairly volatile. Averaging is also beneficial for reducing the risk of a dynamic panel bias when estimating our dynamic panel model with a lagged dependent variable. Arellano-Bond tests for auto-correlation show that our estimates do not suffer from such a bias (Roodman, 2009).

The amount of aid disbursed to recipient countries is modeled as a function of the share of women legislators in parliaments and other controls. The panel model is expressed as:

$$Aid_{r,t} = \alpha Aid_{r,t-1} + \beta Sharew_{r,t-1} + \delta X^{'}_{r,t-1} + \rho_r + \gamma_t + v_{r,t} \tag{3.1}$$

where $Aid_{r,t}$ is the amount of development aid received by recipient country r in period t

not be a useful property in the context here.

from all donors. $X_{r,t}$ comprises recipient-country specific explanatory variables in period t (e.g., population size, GDP per capita). The coefficient of interest is β. If the hypothesis that donor countries are selective to policies promoting political gender equality in recipient countries is correct, then we expect the sign of β to be positive. The regressions control for one-period lag in aid through the inclusion of $Aid_{r,t-1}$. The coefficient for the share of women legislators, β, then captures the additional aid amounts that a recipient country receives from aid donors when the country demonstrates commitment to political gender equality. Using this dynamic panel model also implies that changing the gender composition in the national parliament in a given period has in addition to the immediate impact measured by β also a longer term effect if $\alpha > 0$. That effect equals $\beta \sum_{t=0}^{T} \alpha^t$ after T periods or $\beta/(1-\alpha)$ in the very long-run.

Our model includes period fixed effects denoted by γ_t. These fixed effects help to address any omitted variable bias concerns in the model caused by excluding unobserved variables that evolve over time but are constant across recipient countries. We include also aid recipient country fixed effects denoted by ρ_r. This will control for unobserved time-invariant recipient country effect. Other controls typically included in the aid literature such as colonial ties and religious beliefs of recipient countries will be absorbed by the country fixed effects. For example, the colonial past of aid recipient countries has been shown to affect the amount of aid that a country receives (Alesina and Dollar, 2000; Rajan and Subramanian, 2008).

Additional controls are included in the regression equation to make the regressions comparable with other aid allocation regressions and to address possible omitted variable bias and potential endogeneity issues. For instance, if a recipient country that is more democratic also tends to have a higher share of women legislators, and if a democratic country is also likely to attract more aid then our estimate is biased if we fail to control for democracy. As indicated earlier, we use the XRCOMP measure from Polity IV to control for democracy (Marshall et al., 2017). We also include GDP per capita and population as controls as it is standard in aid allocation regressions to include these two controls. GDP per capita captures the poverty-selectivity of aid allocations and a control of population captures a well documented donor bias in aid allocations towards smaller countries (see for example Collier and Dollar, 2001). As discussed earlier, donor's aid allocation have become more policy selective over the last three decades. Here we follow Annen and Knack (2020) and include the World Bank's Country Policy and Institutional Assessment policy index (CPIA) as a control. This measure is useful in the context here as it is used as part of the distribution key for the International Development Assistance funds (IDA) by the World Bank. We also include merchandise trade (as a percentage of GDP) to capture commercial interests, and temporary membership in the United Nations Security Council (UNSC) following Dreher and Jensen (2013) to capture geopolitical interests. The UNSC measure controls for the political motivation behind aid donations. A recipient country who is a member of the UNSC may receive more aid amounts from donors to influence the recipient country's political support in favor of donors' interests at the UNSC meetings. In an article published by Fox News on May 27, 2011, Congressman Steve Chabot, who at that time sat on the House Foreign Affairs Committee of the United States stated that "[i]f we are giving money to countries consistently voting against our interest, we ought to cut them off." The same article states that the US Ambassador to the U.N. at the time – John Bolton – proposed cutting off all aid allocations to 30 countries who consistently vote against the United States interests at the United Nations meetings.[10] Finally, we also control for the occurrence of conflicts in recipient countries. This control may be relevant in our context as wars have been observed to lead to an increase in women empowerment through the transition of women into the labor force and subsequently into decision making roles (Powley, 2004; Mpoumou, 2004). If aid donors are also likely to give more aid amounts to conflict-affected countries and regions then the omission of a control for conflicts can produce an omitted variable bias. For example, "Peace, Justice and Strong Institutions" is one of the United Nation's Sustainable Development Goals which aims at significantly reducing all forms of violence and related death rates everywhere and many aid donors have made conflict-affected countries and fragile states a high priority in recent years (Collier et al., 2003; World Bank, 2011; Ellison, 2016; Findley, 2018). Data on conflicts is taken from the Uppsala Conflict Data Program (UCDP) database (Pettersson and Öberg, 2020; Sundberg and Melander, 2013).

We control for whether there is a change in leadership in a recipient country. In our data, we observe that quotas were likely to be enacted into law following an election. Also, found that USAID was used to meddle elections.

[10]Source: https://www.foxnews.com/politics/its-all-your-money-foreign-aid-to-muslim-arab-nations

Table 3.3 presents the first regression results. The table shows that there is a positive relationship between lagged share of women in parliament and current aid for Africa in all five regressions reported in this table. Column 1 shows an overall strong correlation between these two variables without any further controls other than the period and country fixed effects. Columns 2–5 all control for lagged aid, democracy, population size, and GDP per capita and our coefficient of interest ranges between 0.10 and 0.14 depending on additional controls that are included in the regressions. We observe the estimate to be robust to the various set of controls included in the regressions. The midpoint of this interval suggests that a 10 percent increase in the share of women legislators is associated with a 1.3 percent increase in aid. For example, an increase in the share of women legislators from 15 to 20 percent is associated with a 4.3 percent increase in aid. The long-term impact of such a change holding everything else constant amounts to 7.92 percent in additional aid. The estimate on the lagged aid measure is also highly robust. We test for the presence of first-order serial correlation in the errors using the Arellano-Bond test for serial correlation. The p value from the Arellano-Bond test are reported at the bottom of each regression results table. The null hypothesis for this test is zero autocorrelation in the disturbances. The p values from the tests are very large, therefore we fail to reject the assumption of zero serial correlation in the residuals.[11] We also find a positive and significant aid effect of democracy almost in identical magnitudes for the share of women legislators. Column 3 includes an interaction term between the share of women legislators and democracy. We find a positive

[11] We run this test by running the fixed effect regressions using the Least Square Dummy Variable (LSDV) estimator and then running the "abar" test developed by Roodman (2009) in Stata.

Table 3.3: Aid selectivity in Africa: share of women legislators (1990-2017)

	(1)	(2)	(3)	(4)	(5)
Log of aid, lag		0.5035***	0.5073***	0.4507***	0.4556***
		(0.0815)	(0.0821)	(0.0812)	(0.0804)
Log of sharew., lag	0.1998***	0.1417***	0.0949	0.1274***	0.1356***
	(0.0564)	(0.0435)	(0.0597)	(0.0452)	(0.0474)
Democracy, lag		0.0968**	0.0002	0.1004***	0.0987**
		(0.0375)	(0.0929)	(0.0370)	(0.0370)
Log of GDPpc, lag		-0.0581	-0.0566	-0.0876	-0.0967
		(0.1308)	(0.1352)	(0.1353)	(0.1261)
Log of population, lag		-0.2186	-0.1305	-0.5378	-0.4902
		(0.3559)	(0.3789)	(0.3224)	(0.3828)
Log of CPIA, lag				0.4455	0.4620
				(0.2712)	(0.2898)
Conflict, lag					0.0833
					(0.1469)
UNSC, lag					0.1369
					(0.1000)
Trade, lag					-0.0007
					(0.0035)
Log of sharew × dem., lag			0.0464		
			(0.0371)		
Country FE	Yes	Yes	Yes	Yes	Yes
AR(1)		0.6366	0.5245	0.5142	0.6353
N	302	302	302	302	302
R-squared	0.32	0.55	0.55	0.56	0.57
F statistic	14.29	28.58	32.66	36.18	33.01

Note: The table presents estimates of the effect of the log of the share of women legislators on the log of developmental aid. See Table 3.7 in the Appendix for a full description and sources of all variables. The sample includes African countries between 1990 and 2017. All regressions include a constant term, period- and country fixed effect, which are not reported. Recipient-level cluster-robust standard errors are reported in parenthesis. Significance levels : $*$: 10 percent $**$: 5 percent $***$: 1 percent

but small effect, which is not statistically significant. This implies that donors seem not to tailor their gender-selective aid towards more democratic African countries. Column 4 includes the log of the CPIA as additional control presenting thereby a similar aid allocation regression as in Annen and Knack (2020). A notable difference, however, is that this regression includes only African countries and uses 4-year period averages instead of yearly data. This regression confirms the policy-selectivity of aid allocations. In addition, our coefficient of interest increases slightly when adding this control. Income and population have the expected sign in all the regressions but the coefficients are not statistically significant. Finally, Column 5 adds the final set of controls, which all do not affect our coefficient of interest.

Table 3.4: Aid selectivity in Africa: quota-systems (1990-2017)

	(1)	(2)	(3)	(4)	(5)
Log of aid, lag		0.4831***	0.4309***	0.4727***	0.4290***
		(0.0935)	(0.0912)	(0.0926)	(0.0907)
Legis. quota, lag	-0.0640	-0.0531	-0.0906	-0.1134	-0.1375
	(0.1989)	(0.1150)	(0.1051)	(0.1250)	(0.1007)
Reserved seats	0.7088***	0.4390***	0.4135***	0.3269**	0.3138**
	(0.1464)	(0.1401)	(0.1244)	(0.1510)	(0.1334)
Log of sharew., lag				0.1175**	0.1137**
				(0.0438)	(0.0471)
Democracy, lag		0.1042***	0.1053***	0.1024***	0.1028***
		(0.0378)	(0.0386)	(0.0355)	(0.0355)
Log of GDPpc, lag		-0.0614	-0.0997	-0.0473	-0.0863
		(0.1329)	(0.1245)	(0.1378)	(0.1322)
Log of population, lag		-0.2651	-0.5641	-0.1991	-0.4469
		(0.3797)	(0.3838)	(0.3664)	(0.3842)
Log of CPIA, lag			0.5040*		0.4577*
			(0.2529)		(0.2677)
Conflict, lag			0.0844		0.1158
			(0.1622)		(0.1572)
UNSC, lag			0.0817		0.1184
			(0.0923)		(0.0982)
Trade, lag			-0.0004		-0.0008
			(0.0034)		(0.0034)
Country FE	Yes	Yes	Yes	Yes	Yes
AR(1)		0.7779	0.7789	0.6229	0.6211
N	302	302	302	302	302
R-squared	0.33	0.55	0.57	0.56	0.58
F statistic	20.18	37.05	32.79	30.46	33.60

Note: The table presents estimates of the effect of a quota system on the log of developmental aid. See Table 3.7 in the Appendix for a full description and sources of all variables. The sample includes African countries between 1990 and 2017. All regressions include a constant term, period- and country fixed effect, which are not reported. Recipient-level cluster-robust standard errors are reported in parenthesis. Significance levels : $*$: 10 percent $**$: 5 percent $***$: 1 percent

Table 3.4 tests for possible channels through which gender-selective aid may work. In particular, we include a dummy variable for whether a country uses legislated candidate quotas or reserved seats in a given year. We find a strong and statistically significant aid effect for reserved seats but not for quotas. Column 1 shows an overall strong correlation between these two variables without any further controls other than the period and country fixed effects. Columns 2 and 3 include additional controls, including lagged aid. We estimate a coefficient between 0.41 and 0.44 for reserved seats depending on the set of controls. The mid-point of these estimates suggests that adopting reserved seats for women legislators increases aid by 53 percent in the following period and by 77 percent in the long-run. We observe that the other controls behave very similarly as in the Table 3.3. Columns 4 and 5 add the share of women legislators. As expected, coefficients of the share of women legislators and reserved seats fall but they both remain positive and statistically significant. This implies that changing the share of women in parliaments has positive aid effects independent of a reserved seat policy.

3.5 Robustness Checks

Table 3.5 repeats the analysis using our sample without Rwanda. We have observed that Rwanda has had the largest share of women legislators in the last ten years. All the estimates are highly robust to that change which suggests that our results are not driven by Rwanda.

All the countries in North Africa have legislated candidate quotas or reserved seats. Table 3.6 repeats the analysis using our sample with Sub-Saharan African countries only. All the estimates are highly robust to that change. The coefficients of share of women legislators and reserved seats are much higher and they remain statistically significant.

We repeat the analysis using OLS and system GMM. The results from these estimation methods are qualitatively similar to those reported in the paper. From the system GMM regression, we find that "an increase in the share of women legislators by 10 percent for a recipient country is associated with an increase of 1.2 percent in aid", and from the OLS regression, we find that "an increase in the share of women legislators by 10 percent for a recipient country is associated with an increase of 1.1 percent in aid".

Table 3.5: Aid selectivity in Africa: share of women legislators excluding Rwanda (1990-2017)

	(1)	(2)	(3)	(4)
Log of sharew., lag	0.1387***	0.1343***	0.1106**	0.1081**
	(0.0440)	(0.0470)	(0.0435)	(0.0466)
Reserved seats			0.3543**	0.3469**
			(0.1697)	(0.1501)
Country FE	Yes	Yes	Yes	Yes
Full Controls	No	Yes	No	Yes
AR(1)	0.8272	0.7864	0.5879	0.5799
N	295	295	295	295
R-squared	0.55	0.56	0.56	0.57
F statistic	27.40	30.73	26.43	26.52

Note: The table presents estimates of the effect of the log of the share of women legislators on the log of developmental aid. See Table 3.7 in the Appendix for a full description and sources of the variables. The sample includes African countries (excluding Rwanda) between 1990 and 2017. All regressions include a constant term, period- and country fixed effect, which are not reported. Recipient-level cluster-robust standard errors are reported in parenthesis. Significance levels : $*$: 10 percent $**$: 5 percent $***$: 1 percent

Table 3.6: Aid selectivity in Africa: share of women legislators—only Sub-Saharan Africa countries (1990-2017)

	(1)	(2)	(3)	(4)
Log of sharew., lag	0.1458***	0.1346**	0.1277***	0.1197**
	(0.0485)	(0.0509)	(0.0459)	(0.0487)
Reserved seats			0.3628**	0.3567***
			(0.1346)	(0.1011)
Country FE	Yes	Yes	Yes	Yes
Full Controls	No	Yes	No	Yes
AR(1)	0.7901	0.7026	0.4416	0.3875
N	274	274	274	274
R-squared	0.56	0.58	0.58	0.60
F statistic	28.28	35.01	44.45	43.79

Note: The table presents estimates of the effect of the log of the share of women legislators on the log of developmental aid. See Table 3.7 in the Appendix for a full description and sources of the variables. The sample includes African countries (excluding Rwanda) between 1990 and 2017. All regressions include a constant term, period- and country fixed effect, which are not reported. Recipient-level cluster-robust standard errors are reported in parenthesis. Significance levels : $*$: 10 percent $**$: 5 percent $***$: 1 percent

3.6 Conclusion

This study provides a rigorous quantitative analysis for the connection between policy-selective aid and women political participation in aid-recipient countries in Africa. Using data from the Inter-Parliamentary Union and Varieties of Democracy (Coppedge et al., 2020; Pemstein et al., 2020) over the period 1960 to 2018, we observe a big jump in the share of women legislators in Africa beginning in the post-Soviet era. The analysis reveals that the growth in the number of women who run in and win elections in the lower chamber of national parliaments across Africa cannot be entirely attributed to an improvement in the feminist attitude of electorates. Instead, the increase in the share of women legislators in Africa can be linked to the adoption of gender quotas. Over the past four decades, feminist actions and efforts have been taken by the United Nations and its member states to bridge the gender inequality gap that exists across the world. Further, many aid donors have taken a feminist approach to international development assistance in the last two decades. This paper investigates whether there is a relationship between women representation in parliaments across Africa and foreign aid allocations. The paper estimates that an increase in the share of women legislators by 10 percentage point for a recipient country is associated with an increase of 1.3 percent in aid on average. We also find that aid recipient countries who reserve special seats for women in parliaments receive an additional 53 percent in aid amounts on average. Although we find that democratic countries receive more aid, our results show that donors do not tailor their gender-selective aid towards more democratic African countries. Our findings provide evidence in support of the growing selectivity of aid donors to gender equality and women empowerment issues in Africa. Future research can investigate the effectiveness of "aid-influenced" quota adoption. That is, whether women have become critical leaders (like speakers of parliament, committee chairs, more women at the negotiating table) or whether the women legislators selected into parliament truly influence policies that affect women – what has been termed as substantive representation.

3.7 Way forward/Future Research

The mechanisms used to implement reserved seats vary from country to country. Different countries adopt different methods to recruit women to fill special seats in parliament.

In some countries, women are elected directly by voters using a women-only ballot list whereas in other countries, women are indirectly selected into office by leaders of political parties. For the latter, the number of women a political party is entitled to appoint depends on the proportion of votes the political party obtained during the general open seats elections. For instance, if a political party obtained 50 percent of the votes in the general election for open seats, then that political party is entitled to nominate women to fill in 50 percent of the seats reserved for women.

In Tanzania, political parties use internal party mechanisms to nominate special seats women candidates. Seats are allocated by the electoral commission to parties based on the number of popular votes (at least 5 percent) that each party obtained in the regular parliamentary elections for open seats. In Uganda and Kenya, one seat is reserved for women in each district. Each political party in a district nominates one woman onto a women-only ballot. Voters exercise their universal suffrage and the woman with the most votes gets elected into parliament from the district. In Zimbabwe and Sudan, multiple seats are reserved for women in each district. In every district, each political party defines a women-only candidate list and voters vote for a list of their choice. The women seats are then allocated among the political parties based on the relative votes each political party list receive. Suppose district 1 is given the option to elect 10 women into parliament. Each political party in district 1 will present a list of 10 women to be voted for. If 60 percent of the voters vote for party A's list of women, then party A gets to appoint 6 out of 10 women on its list in the order in which they appear on the candidate list.

We categorize the mechanisms to fill seats reserved for women as follows:

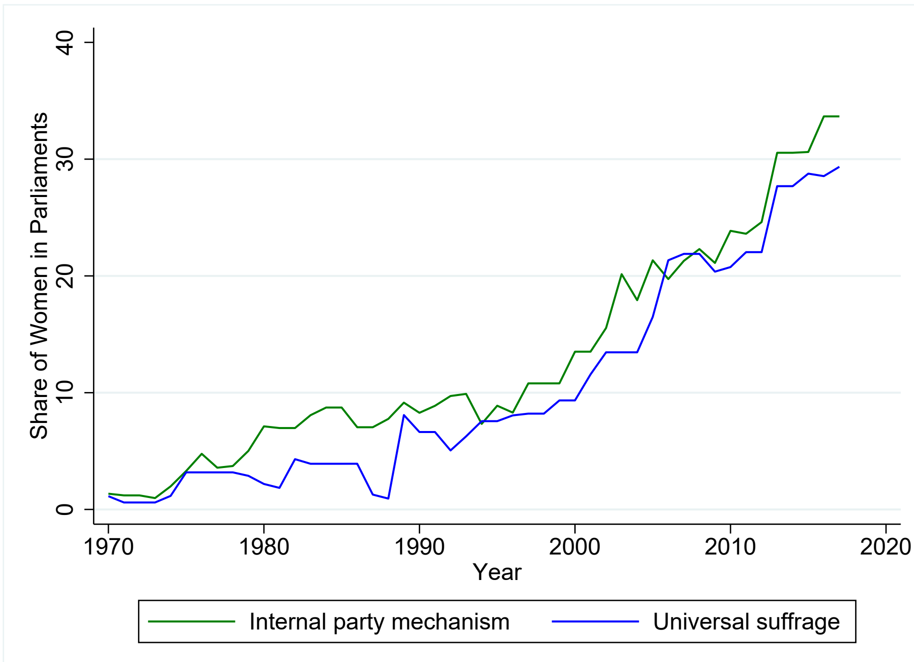
- Elected by universal suffrage: This is when special seats reserved for women in parliament are filled directly by voters using an ‘only female candidate ballot’ (e.g., Uganda, Kenya, Mauritania, Burundi, Egypt)

- Elected through party internal mechanism: This is when women are nominated by political parties or special committees and clans to fill special seats in parliaments. (e.g., Tanzania, Zimbabwe, South Sudan, Sudan, Rwanda, Morocco, Somalia)

For 7 out of 12 countries in Africa who have reserved seats, the special seats are not determined directly by voters. This paper will focus on a large unexplored aspect of gender quotas, that is, why African countries prefer reserved seats to candidate quotas. Some studies have examined the effectiveness of reserved seats as a means of reducing gender inequality in parliaments in single case studies. However, the political motivation for the adoption of reserved seat has hardly received any attention in the quota literature. Figure 3.7 displays the share of women legislators in Africa by the two reserved seats mechanisms. Over time, the share of women legislators is increasing for both groups. However, the share of women legislators in countries who fill special women seats through internal party mechanism is relatively higher compared with countries who fill special women seats through direct elections. The countries who fill special seats through party internal mechanism reserve at least 20 percent of the seats in the Lower Chamber of the Legislature for women (except Morocco with 15 percent). We assert that political parties, especially the biggest political parties, benefit from reserving more seats for women. These benefits come in the form of reduced competition for open seats and reduced electoral spending. Politicians in Africa usually offer gifts and monies to the electorates during open seat elections, i.e., vote-buying. Therefore, election costs are reduced for political parties if fewer seats are available in the general elections and more seats are reserved for women. It is also advantageous for the political party leaders to have the option to nominate women because the women nominees are more likely to "follow the leader" when decisions are made in parliament. Political party leaders may also appear democratic and modern to the rest of the world by adopting reserved seats (Tøraasen, 2017).

The goal is to design a game theoretic-model that explains the relationship between the rents that political parties enjoy and the adoption of reserved seats in developing countries. As mentioned before, the rents can come in the form of reduced electoral spending, reduced electoral competition, and women appointees who vote in favor of their leaders in parliaments.

Figure 3.7: Share of women in parliaments for African countries with reserved seats (1970-2016)



The green line graphs the average share of women legislators in African countries who fill the special seats reserved for women through internal party mechanisms. The blue line graphs the average share of women legislators for African countries who fill the special seats reserved for women through voting. Data on quotas is from the Varieties of democracy (Vdem) database. Data on the share of women in parliaments is based on a new data set created by the authors.

Appendix

Table 3.7: Variables and their definition

Variable	Definition	Source
Total Aid	Log of Total Aid disbursements from donor country d to recipient country r in year t (in constant 2011 US$)	OECD
GDP per capita	Log of Gross Domestic Product per capita (in PPP prices) for country c in year t	Penn World
Population	Population size for country c in year t	Penn World
Share of women	The percentage of members of parliament in the Lower Chamber of Legislature who are women for country c in year t	Data created by authors based on html and pdf files extracted from IPU (data.ipu.org) and Vdem
Democracy	3 if the selection of head of states is through popular elections involving two or more parties or candidates for country c in year t	Vdem
Quota	1 if country c has statutory gender quota for all political leaders with or without sanctions in year t	Vdem
Reserved seat	1 if country c has reserved seats in the legislature for women in year t	Vdem
Conflict	1 if a conflict involving government and rebel groups resulted in 1000 deaths or more for country c in year t	Uppsala Conflict Data Program (UCDP)
UNSC	1 if a country c is a member of the United Nations Security Council in year t	UN.org
Trade	The percentage of the GDP that is associated with merchandise trade (both imports and exports) for country c in year t	World Development Indicator Database
CPIA	An index that assesses the quality of country c's policies and institutions in year t	World Bank

Table 3.8: Countries who have implemented Gender Quotas/Reserved Seats in parliaments (Lower Chamber)

Country	Quota Year	Country	Quota Year	Country	Reserved Seats Year
Angola	2008	Macedonia	2002	Afghanistan	2005
Albania	2009	Mongolia	2012	Burundi	2005
Argentina	1993	Mauritania	2006	Bangladesh	1973
Armenia	2003	Mexico	2003	Egypt	1980
Algeria	2012	Montenegro	2012	Haiti	2015
Belgium	1999	Nepal	1991	Iraq	2010
Burkina Faso	2012	Nicaragua	2016	Jordan	2003
Bosnia and Herzegovina	1998	Niger	2004	Kenya	1997
Bolivia	1997	North Korea	1998	Mauritania	2013
Brazil	1998	Paraguay	1998	Morocco	2002
China	2008	Peru	2000	Pakistan	1970
DRC	2011	Poland	2011	Philippines	1987
Republic of the Congo	2012	Portugal	2009	Rwanda	2003
Colombia	2014	Romania	2004	Saudi Arabia	2014
Cape Verde	2011	Senegal	2012	Somalia	2004
Costa Rica	1998	Serbia	2007	South Sudan	2015
Croatia	2015	Slovenia	2008	Sudan	1974
Djibouti	2003	Solomon Isl.	2014	Tanzania	1975
Dominican Republic	1998	South Korea	2000	Uganda	1989
Ecuador	1998	Spain	2008	Zimbabwe	2013
Egypt	2012	Timor Leste	2007		
El Salvador	2015	Tunisia	2011		
France	2002	Uruguay	2014		
Guinea	2013	Uzbekistan	2004		
Greece	2012	Venezuela	1998		
Guyana	2001	Vietnam	2016		
Honduras	2001				
Indonesia	2004				
Ireland	2016				
Iraq	2005				
Italy	1994				
Kyrghyzstan	2007				
Lesotho	2012				
Liberia	2005				
Libya	2012				

This table presents the list of all countries (in an alphabetical order) who have implemented legislated candidate quotas or reserved seats and the year in which they are implemented.
Source: Varieties of Democracy (V-Dem) Database (2018).

Table 3.9: How reserved seat quotas are implemented (African countries)

Country	Quota %	Reserved seats/quota system
Burundi	30%	Burundi has gender quotas through party lists where 1 in 4 candidates must be a woman. If the quota does not result in 30% women representation, they make that top-up through a process of co-optation. The Electoral Administration adds, from the candidate lists that have obtained at least 5% of the votes cast, more members from the under-represented gender until the quota requirements are met.
Egypt	10%	64 seats to which the nominations was restricted to women was previously adopted. However, this has been cancelled since the 2010 election. Parties are now obliged to nominate at least one woman as part of their district candidate lists which they have to submit for the 46 districts electing 332 seats contested through a proportional system.
Kenya	13%	The constitution reserves 47 seats for women deputies to be elected from 47 counties. In every county, each political party nominates one woman onto a women-only list, voters vote and the woman who wins most of the votes in each county is selected to represent the county.
Mauritania	14%	20 women are elected on a single nation-wide women-only ballot list. Parties nominate women onto the list, voters vote and the women who win the most votes get elected into parliament.
Morocco	15%	The seats are filled by winners elected through a proportional representation system based on nation-wide closed party lists of women candidates. Each party defines a candidate list of women and voters will vote for a list. The relative vote for each list determines how many candidates from each list will be elected.
Rwanda	30%	24 women are elected from each province (2 women per province) and from the city of Kigali in a women-only ballot. They are elected by an assembly made up of various councils and committee members.
Somalia	30%	Candidates for parliamentary seats are nominated by the country's major clans and vetted by the Technical Selection Committee in compliance with the criteria outlined in the Constitution. The failure to meet the stated commitments on the 30 per cent reserved seats for women in the 2012 Federal Parliament is largely due to the lack of agreement among the clans which govern the country.
South Sudan	25%	Women members shall be elected on the basis of proportional representation at the national level from closed party lists.
Sudan	25%	Each party will present a women-only ballot list. A voter will vote for only one women's list of their choice. Seats are allocated by the Commission according to the proportional representation among these parties. The seats shall be won by the candidates in the order in which their names appear in the party list from top to bottom.
Tanzania	29%	In 1995 and 2000, the Commission allocated the women seats based on the number of constituency seats each party won. Since 2005, the Commission has allocated these seats based on the number of popular votes each party received in the parliamentary election. Parties use different internal mechanisms to nominate special-seat candidates.
Uganda	47%	Women are elected in a separate election through a simple plurality system of women-only ballots; one woman per district. Each party nominates one woman onto the list, voters vote and the woman who wins most of the votes in each district is selected to represent the district.
Zimbabwe	22%	In each of Zimbabwe's 10 provinces, 6 women are elected under a women only party-list system of proportional representation. Each party will define a candidate list of 6 women and voters will vote for a list. The relative vote for each list determines how many candidates from each party list will be elected.

Source: International Institute for Democracy and Electoral Assistance (International IDEA) (idea.int)

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