Essays in Inequality and Gender in Developing Countries

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Rodrigo Lovaton Davila

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Dedication

To my parents, Lourdes and Marco.

Abstract

This dissertation is comprised of three essays: two of which focus on the impacts of changes in maternity leave legislation on women's employment status and fertility, and the third concentrates on aggregation methods for the construction of asset-based proxy measures for household socioeconomic status in developing countries. In the first essay, I explore the effects of maternity leave on labor market outcomes in six countries in Latin America (Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela). The evidence shows that maternity leave has a positive effect on the labor force participation and unemployment-to-population ratio of women of childbearing age. In the second essay, I investigate the impact of maternity leave on fertility for the same set of six countries. Results suggest that maternity leave has small negative effects on higher order births for young adult women (18 and 30 years old), while it has small positive effects on fertility for older adult women (31 and 45 years old). If we consider these two effects, the evidence indicates that increases in maternity leave duration are associated to postponing some additional births. Finally, the third essay analyzes the performance of alternative methods to aggregate data for an asset-based wealth index using ordinal variables. Despite recommendations given by previous research, results suggest a relatively similar performance of principal components analysis on dichotomized data with respect to other methods that work with ordinal variables.

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Chapter 1: Introduction

This dissertation is comprised of three essays. Two essays focus on the impacts of changes in maternity leave legislation on women's employment status and fertility and the third analyzes aggregation methods used to construct asset-based proxy measures of household socioeconomic status in developing countries. Even though the topics examined in the first two essays on maternity leave are not closely related to the construction of asset-based indices, all three essays focus on inequalities and gender in the context of developing countries.

Maternity leave after childbirth helps parents to reconcile work and family responsibilities, but it could also raise the relative costs of hiring women. In the first essay, I explore the effects of maternity leave duration on labor force participation, employment, and unemployment in six countries in Latin America (Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela). I use census microdata from these selected Latin American countries and apply a pseudo-panel method (Deaton, 1985) to estimate the effects of these countries' maternity leave regulations for the period between 1960 and 2013. Maternity leave duration was constructed based on the review of laws for each of the countries over the period of analysis. For every country analyzed, maternity leave is paid and replacement rates are 100 percent or close to a fully paid salary. The empirical strategy exploits changes in the benefit duration at different points in time for different countries and also compares women of childbearing age (18 to 30 years old) to other groups that should not be affected by the enactment of these regulations, to obtain difference-in-differences-in-differences (DDD) estimates.

The estimates in this first essay show that maternity leave has positive effects on the labor force participation and unemployment-to-population ratio of women of childbearing age. That is, maternity leave creates incentives for women of childbearing age to participate in the labor market. Even though the employment effects for women (or mothers) of childbearing age are positive, these effects are not statistically significant. Given that increases in employment depend also on the labor demand side, it is possible that employers perceive women as more costly due to the potential need to hire temporary workers or to reorganize production during the time that new mothers are on leave. Therefore, women are not necessarily getting more jobs, despite the higher incentives for them to participate in the labor market, which seems to drive the increase in unemployment.

In the second essay, I investigate the effects of paid maternity leave on the number of children for the same six countries in Latin America (Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela). Maternity leave after childbirth is hypothesized to lower the costs of having children and thus to increase fertility. For each of the six countries analyzed, maternity leave is paid and the duration of this benefit was created from the review of national legislation. For this essay, I use a similar set of census microdata from these selected Latin American countries, covering the period between 1960 and 2011, and apply a pseudo-panel method (Deaton, 1985) to estimate the effects of these countries' maternity leave regulations. The empirical strategy also exploits changes in the benefit duration at different points in time for different countries to produce difference-in-differences (DD) estimates.

Results from the second essay show that maternity leave has a negative effect on higher order births for young adult women (between 18 and 30 years old), while it has positive effects on fertility for older adult women (between 31 and 45 years old). The estimated effects are relatively small with respect to fertility levels in the countries under analysis. In the case of higher order births, the direction of effects of maternity leave duration depends on the woman's age: there is a decrease in higher order births for women between 18 and 30 years old, while there is an increase for women between 31 and 45 years old. If one considers these two effects together, increases in maternity leave duration are associated with the postponement of some higher order births for young adult women.

The construction of wealth indices based on housing characteristics and asset ownership has been widely used when other measures of socioeconomic status are not available. A popular approach has been to apply principal components analysis (PCA) on data recoded to binary indicators (Filmer and Pritchett, 2001). However, this procedure has been criticized since standard PCA methods are not designed to handle discrete data. In the third essay, I compare alternative aggregation procedures that have been proposed to overcome this issue. The essay uses data from twelve developing countries. The evidence indicates that methods based on ordinal data have high agreement in rankings, but the PCA procedure on dichotomized data also has reasonable agreement with these measures. The alternative measures do not have striking differences in their relationship with a set of education, fertility, and mortality outcomes, both based on wealth index quintiles and on regression analysis. Finally, none of the asset-based indices outperformed the rest in terms of similarities of rankings with the logarithm of income per capita.

In this sense, despite recommendations given by previous research (Howe *et al*., 2008; Kolenikov and Angeles, 2009), results suggest a relatively similar performance of the PCA procedure on dichotomized data with respect to methods based on ordinal data. Furthermore, given the possible difficulties in the calculation of polychoric correlations, standard correlation methods for PCA applied on ordinal or binary data may be preferred. The implementation of an asset count index is not recommended, because of the limited variability of the measure (which creates problems to define household rankings) and the difficulties to include information on housing characteristics (which are mostly ordinal variables that would need to be recoded to be added in a "count").

Chapter 2: The Impact of Maternity Leave on Women's Employment Status in Latin America

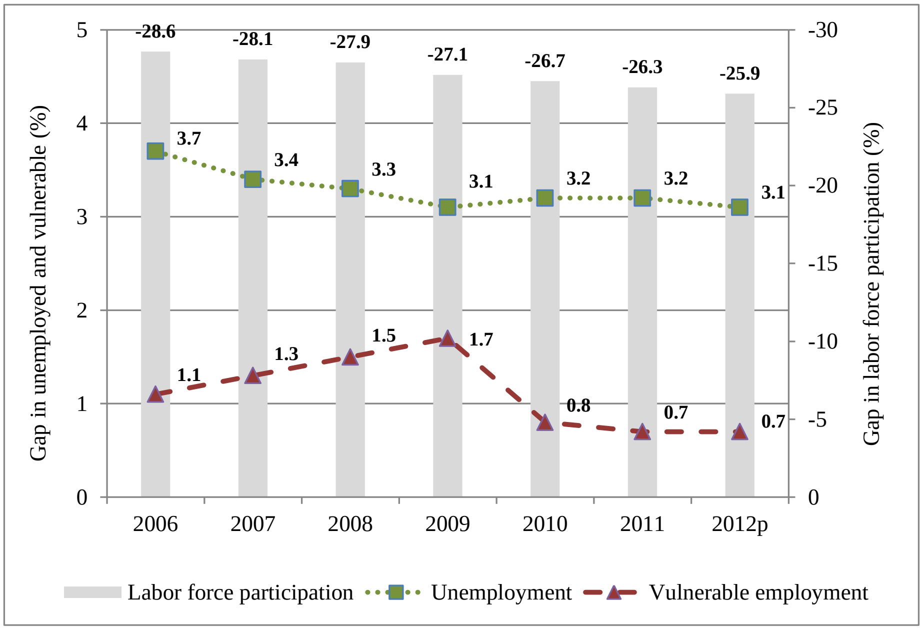
2.1. <u>Introduction</u>

The gap in labor market outcomes by gender remains large today in almost all countries. This is reflected not only in considerable differences in labor force participation, employment, and unemployment rates, but also in a higher concentration of women working in certain industries and occupations. Labor market regulations on maternity protection and childcare at the workplace help harmonize work and family responsibilities, which are usually a barrier for women to continue participating in the labor market after they have children. However, creating these benefits for women could also raise the relative costs of hiring women.

A recent paper by the International Labor Organization (ILO, 2012) examines long-term trends in gender gaps in labor force participation, employment, unemployment, wage and salary employment, and industry and occupational segregation. The report shows that aggregate differences in labor force participation by gender in the world decreased in the 1990s but stayed constant in the 2000s, with participation rates falling for both men and women during both decades (mainly due to increased education for younger cohorts and to population aging). While gaps in employment and unemployment had some convergence in the early 2000s, these gaps widened in the late 2000s (ILO, 2012, p. 4-5). The gender gaps in salaried employment and in wages, as well as industry and occupational segregation, are currently large. Women are more likely to be family workers (self-employed in a family business) than men, which typically implies lower quality of employment (ILO, 2012, p. 22-23). Recent trends show that women are increasingly working in service sectors and moving out of agriculture in developing countries. In terms of occupational segregation, women are over-represented in mid-skill occupations, such as clerks, service, and shop/sales workers (ILO, 2012, p. 24-26).

In Latin America and the Caribbean, employment gender gaps were high during the 2000s: women experienced 3-3.5 percentage points more unemployment (even higher for youth) and 25-30 percentage points less employment and labor force participation; while the gap in vulnerable employment (family and own account workers) was relatively small at about 1 percentage point. Even though some labor force indicators improved over the period, as shown in Figure 1 (such as the difference in labor force participation), the gaps in labor market outcomes for men and women are still high. For instance, the unemployment rate for women in 2011 was 9.0 percent while it was only 5.9 percent for men, implying a gap of 3.1 percentage points or 1.5 times higher unemployment, for women.

Figure 2.1: Gender Gaps (percentage points) in Selected Labor Market Outcomes, Latin America and the Caribbean (2006-2012) [1/ 2/]



Source: Global Employment Trends for Women, International Labor Organization (2012).

1. Estimates for 2012 are preliminary.

2. The gap is calculated as the difference between the indicator for women and for men.

In this chapter, I explore the effects of changes in the duration of maternity leave on selected labor market outcomes. This question has not been extensively explored in the relevant literature and, to the best of my knowledge, no prior evidence has been provided for Latin America. For this purpose, I use data from six countries in Latin America (Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela) and apply a pseudo-panel method to estimate the effects of these regulations on employment status from 1960 to 2013. This method allows me to take advantage of existing cross-sectional data by creating synthetic longitudinal data to assess the effects of these policies on the labor market. These countries were selected due to the availability of data covering a long span of time (mostly microdata from the Integrated Public Use Microdata Series - International project) and of sources to track changes in the duration of maternity leave over time. For the analysis, I tracked all changes in the duration of maternity leave, starting in the 1960s, based on the relevant labor laws for each country. For every country analyzed, maternity leave is paid and replacement rates are 100 percent or close to a fully paid salary.

The chapter is organized as follows. In the next section, I discuss the expected consequences of maternity leave and review the empirical literature on its labor market effects. In Section 2.3, I summarize the changes in the duration of maternity leave in Latin America over the relevant period, focusing on the selected countries for this study. In Section 2.4, I describe the data and the empirical strategy. Next, in Section 2.5 I present some general trends in the labor market in Latin America, from a gender perspective, and the estimated effects of changes in maternity leave regulations on labor force participation, the employment-to-population ratio, and the unemployment-to-population ratio. Finally, a discussion based on the results of the study is shown in Section 2.6.

2.2. Conceptual Framework

2.2.1. What Are the Expected Effects of Maternity Leave?

A mother's choices regarding employment and care provision after childbirth depend on several factors. The time spent taking care of the newborn child and being away from work has value for the mother, while it also has potential positive effects on the health of both the mother and the child. However, it could also imply losing wages if leave is unpaid or if the mother temporarily exits the labor market. Based on these factors, the working mother may have a desired length of leave.[1] Without regulations on this matter, the duration of leave voluntarily offered by the employer is expected to be lower than the mother's desired amount (Klerman and Leibowitz, 1997). The mother faces the decision of staying at her current job and taking leave (presumably shorter than desired), or quitting to take care of her child, thus losing wages that she could have earned and later incurring the costs of searching for a new job.

Mandated maternity leave offers job protection and, if it is paid, replacement of lost wages during the leave period. Therefore, with mandated maternity leave, mothers have an extended duration of leave available (assuming mandated leave is longer than that voluntarily provided by the employer) and are able to retain their jobs and return to them after leave (also avoiding search costs for a new job). Even if the regulated leave is shorter than the desired duration, it offers the additional benefit of returning to the pre-childbirth job. For women already in the labor market, there are two potential positive effects on employment outcomes (Ruhm, 1998; Baker and Milligan, 2008a; Baum and Ruhm, 2013). After leave regulation is in place, some women may decide to continue in their pre-childbirth jobs, instead of quitting, and use leave to spend time with their children. In addition, for this group that would have quit their jobs without mandated leave, leave duration is expected to be shorter than the time they would otherwise have been out of the labor market.[2] For women who were previously (to the leave regulation) out of the labor market, working may now be more attractive and they may decide to find

[1] The time spent taking care of the child after birth on leave could have an optimal duration if we assume, for example, given unpaid leave, that the value of staying with the newborn child is strictly decreasing as the child ages and there is a constant wage offered to the mother in the labor market (Klerman and Liebowitz, 1997).

[2] As Ruhm (1998) argues, this happens "because the gap between desired leave duration and that offered by the firm decreases, while the benefits of maintaining the employment relationship are little changed" (p. 290).

a job before having a child, so that they are also covered by job protection (Ruhm, 1998; Akgunduz and Plantenga, 2013). All these effects imply an increase in the labor supply.

There are also effects on employers due to potential additional costs associated with leave. First, it may be necessary to re-organize production processes or even replace workers on leave, which includes hiring and training temporary personnel (Baum, 2003; Akgunduz and Plantenga, 2013; Thévenon and Solaz, 2013; Das and Polacheck, 2014). If the costs to the firms are high, we may observe increased occupational segregation if employers restrict women's employment to occupations where the impact of leave is smaller (i.e. where it is relatively easier to re-organize work or to find temporary replacements) (Ruhm, 1998). Second, the employer may be required to pay a subsidy to workers taking leave. This is not always true since subsidies offered to workers on leave are often funded through some social security mechanism or with public resources, so these are not necessarily costs to the business. A decrease in labor demand is expected, but only to the extent that there are additional costs to the firms (Ruhm, 1998; Thévenon and Solaz, 2013). Thus, mandated maternity leave could also create barriers for women if employers perceive that it creates additional costs; this may lead in turn to reductions in female wages or female employment (Van der Meulen, 1999; Abramo and Todaro, 2002).

In the long term, continuation of the work relationship leads to higher accumulation of human capital specific to the firm. Thus, the mother is expected to receive a higher wage than what she would have earned in a new job given her firm-specific human capital and higher productivity. In addition, these are also benefits for the employer associated with higher worker's retention (Ruhm, 1998; Baum, 2003; Das and

Polacheck, 2014).[3] These effects would imply an outward shift of the labor demand (Ruhm, 1998; Akgunduz and Plantenga, 2013) and are more likely for skilled or specialized workers (Baum, 2003; Thévenon and Solaz, 2013). Nevertheless, regulated leave that is too long may harm the worker's human capital, productivity, or career advancement (Ruhm, 1998; Lalive and Zweimüller, 2009; Thévenon and Solaz, 2013).

This simple theoretical framework suggests that mandated maternity leave is expected to have a positive effect on the labor supply (due to the incentives for mothers to be employed) and a negative effect on the labor demand (due to the possible costs associated with leave). In the long term, there could be additional positive effects on the labor demand due to higher productivity associated to firm-specific human capital. Overall, this indicates that the effect on employment is ambiguous, while we should observe a negative net effect on wages. Empirical evidence on the direction and statistical significance of these effects is reviewed next, covering mainly the United States, Canada, and other countries in the Organization for Economic Cooperation and Development (OECD).

2.2.2. Empirical Evidence

Parental leave policies are widely available in OECD countries and some research focuses on their labor market effects. In general, these studies take advantage of variation across countries and over time in the duration of maternity leave to investigate effects on

[3] Baum (2003) argues that leave regulation "may benefit employers by preserving employer-employee relationships if permanent separations are costly" because of the firm's investment in workers and the cost of finding permanent replacements (p. 773).

different labor market outcomes. Many of them also identify treatment and control groups to produce difference-in-differences-in-differences (DDD) estimates, based mostly on comparisons of women and men of certain ages. A summary of the main characteristics of the studies discussed in this section is shown in Table 2.A1 in the Appendix to this chapter. Winegarden and Bracy (1995) estimated the effects of paid maternity leave on labor force participation of women between 20 and 34 years old in 17 OECD countries. Their results show that a longer duration of paid maternity leave is associated with increased labor force participation of young women (a 10 percent increase in leave increases labor force participation by about 2 percentage points). Ruhm (1998) used yearly data for nine European countries that underwent significant changes in paid parental leave, resulting in variation in the level of these benefits across countries and over time. He performed comparisons between men and women to obtain DDD estimates.[4] Results indicate that paid leave durations of three to nine months increase women's employment by about 3 to 4 percentage points. Even though shorter leave periods do not have substantial wage effects, paid leave reduces wages by about 3 percentage points for long leave entitlements (nine months).

Two recent papers build on the findings by Ruhm (1998) and present qualitatively similar results. Both of them produced DDD estimates based on the variation of leave duration across countries over time and comparisons between women and men. Thévenon and Solaz (2013) expanded Ruhm's analysis to 30 OECD countries (including some that more recently introduced parental leave regulations) and examined effects of paid

[4] Men are used mostly as a comparison group to identify the effects on women, but some calculations are also performed against women 45-54 years old and against men 25-34 years old. Ruhm (1998) argues that men are a good comparison group "since women use virtually all parental leave in most countries" (p. 286).

parental leave on several labor market outcomes. Their results show a widening employment gender gap for leave durations shorter than one year or longer than two years (that is, negative effects on female employment with respect to males), while the gap narrows for leave periods between one and two years. Furthermore, paid leave has a positive effect on average working hours and a negative effect on earnings for women with respect to men, for leave durations longer than eighteen weeks and shorter than two years. Akgunduz and Plantenga (2013) examined the effects of parental leave legislation on a wide set of labor market outcomes for 16 European countries. An important difference between this and other studies is that the leave variable is constructed combining paid and unpaid leave.[5] This paper also finds evidence of positive effects on women's employment and on weekly working hours, which are smaller the longer the duration of leave. Even though the authors did not find statistically significant effects on wages in the manufacturing sector, their results show negative effects on wages in the financial intermediation sector and on the share of women in high-level occupations.

In the case of the United States, state legislation on maternity leave was passed between the late eighties and early nineties in a few states and the Family and Medical Leave Act (FMLA) was created in 1993, which is federal legislation mandating unpaid maternity leave (Klerman and Leibowitz, 1997; Waldfogel, 1999; Baum, 2003). Before the enactment of this legislation, leave was essentially voluntarily offered by employers (Baker and Milligan, 2008a). The FMLA provided up to 12 weeks of job-protected leave

[5] The authors' objective is to incorporate additional periods of unpaid leave available in some countries. They use a weighted measure, in which "leave is weighted by 33% if the replacement level is between 0% and 33%, by 66% if the replacement rate is between 33% and 67%, and by 100% if the replacement rate is above 67%" (Akgunduz and Plantenga, 2013, p. 850). Even though this measure seems more comprehensive, it may complicate interpretation of results since the effects of paid and unpaid leave could be different on the outcomes of interest.

(not guaranteed to be paid) that could be used for maternity reasons for workers fulfilling certain requirements (minimum number of hours worked over the previous year) and only for firms with 50 or more workers (Waldfogel, 1999).

In the context of the creation of this legislation in the U.S., some research has examined the effects of these regulations on certain labor market outcomes and leave-taking behavior by parents. These studies have generally looked at differences across states with or without specific leave policies and have identified population groups that are used as treatment and control groups. Klerman and Leibowitz (1997) evaluated the employment effects of maternity leave state laws using microdata from the 1980 and 1990 censuses, before and after the implementation of state laws. Even though their results show some evidence of positive effects on employment, these are statistically insignificant based on their DDD estimates. Two additional studies used data from the Current Population Survey (CPS). Waldfogel (1999) examined the effects of the FMLA on employment and hourly wages. She compared states that had no job-protected leave before the passage of this federal legislation (so these were "affected" by the FMLA) to others that did have regulations on this matter (where the FMLA overlapped with existing benefits). Waldfogel (1999) argued that the FMLA could have both negative and positive effects on the labor market outcomes of interest, so the net effect was uncertain. Even though she found increased leave coverage after the FMLA, her results indicate a small positive employment effect and no impact on hourly wages. In a more recent paper, Han, Ruhm, and Waldfogel (2009) explored the impact of three parental leave policies (the FMLA, state laws, and the state temporary disability insurance) on employment and leave-taking. The authors estimated separate models for mothers and fathers, expecting differential effects on employment and leave for both groups. Leave regulations did not appear to have statistically significant effects on employment but were associated with more time on leave for both mother and father, particularly for more educated parents and married mothers (who were expected to have higher coverage).

More recently, some states in the U.S. enacted paid maternity leave legislation or had plans to create it. The California Paid Family Leave (CPFL) program was the first legislation to be implemented on this matter. The CPFL offers all private sector workers a maximum of 6 weeks of paid leave, since 2004, for different reasons including childbirth (Rossin-Slater, Ruhm, and Waldfogel, 2013; Baum and Ruhm, 2013; Das and Polachek, 2014). Several studies follow a difference-in-differences approach to estimate the effects of this policy. Rossin-Slater, Ruhm, and Waldfogel (2013) showed that the CPFL has positive effects on leave-taking (particularly for less advantaged groups), weekly working hours for mothers of children 1- to 3-years-old, and wage income (broadly corresponding to additional working hours, but with imprecise estimated effects). Baum and Ruhm (2013) also found evidence that the CPFL has a positive impact on leave-taking and medium-term effects on hours and weeks worked (but not statistically significant effects on wages), as well as on the probability that a mother has returned to work between nine to twelve months after giving birth and the likelihood of job continuity (with the pre-childbirth employer) for some women. Even though Das and Polacheck (2014) reported an increase in labor force participation for young women after this policy went into effect (about 1.5 percentage points), they also found higher unemployment (between 0.3 and 1.5 percentage points) and unemployment duration. These are considered as unanticipated effects of the CPFL.

Canada progressively expanded job-protected leave through provincial and federal regulations from zero to 12 weeks in the 1960s and to 52 or more weeks by the end of 2000 (Ten Cate, 2003; Baker and Milligan, 2008a; Hanratty and Trzcinski, 2009). Payments during leave are provided separately through the unemployment insurance, which covers work interruptions due to childbirth (Ten Cate, 2003; Baker and Milligan, 2008a). Ten Cate (2000, 2003) examined the effects of leave entitlements' variation across provinces. She observed that these benefits increased the probability of returning to work within two years after childbirth and the employment rates for females with small children (zero to two years old). Baker and Milligan (2008a) analyzed changes in leave duration in Canada over three different time periods. These corresponded to the initial expansion of these benefits around the 1960s and 1970s, the first large increase in the 1990s, and the more recent one that occurred at the end of 2000. Their results show that shorter leave durations (up to 17 or 18 weeks) increase mothers' employment but only longer ones have an effect on the mothers' time spent at home after birth. Furthermore, maternity leave has a positive effect on job continuity with the pre-childbirth employer for any benefit duration. Hanratty and Trzcinski (2009) focused on the latest policy change after 2000. Results indicate that longer leave duration is associated with a delayed return to work for mothers within one year after birth, but with convergence to the pre-expansion levels once paid leave eligibility expires. Furthermore, no evidence was found regarding a decrease in relative employment for mothers of children age one with respect to those with children age three to four.

A few studies specifically investigated the effects of maternity leave on the likelihood of returning to work for new mothers. Baum (2003) estimated the effect of maternity leave laws in the U.S. (both state and federal) on the mother's probability of returning to work and the timing of this return. This paper showed that maternity leave legislation has a positive effect on the mother's probability to return to her pre-childbirth job and negative effect on the probability of starting a new job. It finds a delayed return to this job, generally decreasing the probability of return prior to month two and increasing it after. Pronzato (2009) studied the influence of parental leave policies in nine European countries on the woman's hazard of returning to work after childbirth. She concluded that extended job-protected leave increases the likelihood of returning to work, and paid leave positively influences the mother's time spent at home during the child's first year of life. Furthermore, women with higher human capital return to work sooner while those with higher family income return later.

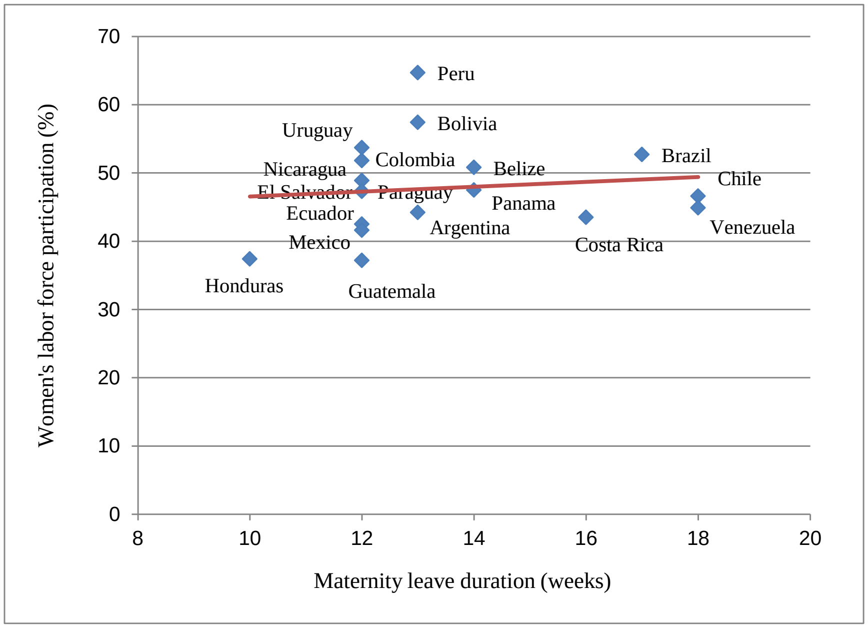
In summary, empirical studies for OECD countries suggest that maternity leave duration has positive effects on women's labor force participation, employment, and working hours, while these are negative for wages (Winegarden and Bracy, 1995; Ruhm, 1998; Akgunduz and Plantenga, 2013; Thévenon and Solaz, 2013). Some evidence also indicates increased occupational segregation (Akgunduz and Plantenga, 2013). However, results are generally conditional on leave duration, such that there might be an optimal length to maximize the (desired) effects of regulated maternity leave. Furthermore, several studies find that regulated maternity leave is associated with women's increased probability of returning to work and continuing to work with the same employer (Ten Cate, 2000; Baum, 2003; Baker and Milligan, 2008a; Pronzato, 2009; Baum and Ruhm, 2013). In the specific case of the United States, effects on women's employment are mostly not statistically significant (except small effects reported by Waldfogel, 1999), despite the fact that there is higher coverage and leave-taking, as well as longer time on leave, after the enactment of these laws (Waldfogel, 1999; Baum, 2003; Han, Ruhm, and Waldfogel, 2009). An important difference in this regard is that most studies for the U.S. are based on changes in unpaid leave (except for the case of California) while other countries examined mainly paid leave regulations. For Latin America, there are no studies focused on the effects of maternity leave, and we can only find a few on the related area of the costs of job security (such as Kugler, 1999; Heckman and Pages, 2000).

2.3. Maternity Leave in Latin America

Regulations with respect to maternity leave play an important role in the reconciliation of work and family responsibilities; they provide alternatives for female workers facing increased family responsibilities through time off during pregnancy and after childbirth, coverage of the labor income lost during that time, and increased job security. The level of protection offered by these labor market regulations varies across countries in Latin America. In the region, the duration of maternity leave benefits in 2009-2010 ranged from 10 (Honduras) to 18 weeks (Chile and Venezuela), with an average of 13.5 weeks of leave (ILO/UNDP, 2009; Pautassi and Nieves, 2011; national legislation). The duration of the benefit in some countries is extended for special circumstances such as multiple births or health issues of the mother or the newborn child. During maternity leave, female workers receive a subsidy (generally 100 percent of their salary) which is often fully financed from public resources (and most frequently, social security), although some national legislations require that employers cover some fraction of it. In those cases where social security covers the maternity leave subsidy, there may be certain requirements in terms of minimum contributions (to social security) from the benefited worker. Most countries also have some form of firing restrictions, which prohibit firing a pregnant woman and usually extend this protection even over some period of time after birth.

Figure 2 compares the duration of maternity leave (in weeks) with labor force participation (percent) for women. The labor force participation rates for women range from 37.2 percent (Honduras) to 64.7 percent (Peru) and were calculated by the ILO for the years 2009 or 2010 based on Labor Force Surveys. The duration of maternity leave is taken from national legislation or secondary sources and corresponds to similar years to the labor force participation rates. As it can be observed in Figure 2, there is a slight increase in labor force participation for women in countries with longer maternity leave duration. However, this could be affected by outliers (such as Peru or Guatemala) and may also be related to other factors, so further investigation is required.

Figure 2.2: Central and South America, Duration of Maternity Leave and Labor Force Participation (%) for Women [1/]



Source: ILOSTAT database for labor force participation and national legislation, ILO/UNDP (2009), Pautassi and Nieves (2011) for labor regulations

1. Labor force participation data corresponds to years 2009 or 2010 based mainly on Labor Force Surveys.

For this essay, I tracked the relevant maternity leave regulations over time to identify changes in the duration of leave. Tracking maternity leave regulations was necessary because existing data from secondary sources do not cover the whole period analyzed in this essay and the information is not available on a year-by-year basis (to be able to identify the exact year when a change in duration occurred). Maternity leave regulations were identified (for each country and over time) mainly through labor law studies (describing the overall regulation of the labor market for a country), databases from the International Labor Organization (ILO), and cross references found in national legislation.[6] The duration of maternity leave (year-by-year) was then constructed based on the analysis of the identified pieces of legislation. The detailed information about the specific maternity leave regulations examined (and their year of enactment) is included as a final appendix to the dissertation. A summary of changes in leave duration by country are shown in Table 2.A2 in the Appendix to this chapter. For the years starting in the 1990s, the constructed duration of maternity leave was verified by comparing it to secondary data.[7] The constructed leave duration series is consistent with the secondary sources examined.

In all the countries examined, these regulations are set at the national level, not at lower political divisions, thus, there is no variation in the duration of leave within countries. Furthermore, these regulations are implemented though labor codes that cover a broad spectrum of the labor force, with only a few exceptions targeting certain workers or sectors. Below, in Figure 3, I present the changes in leave duration weighted by each country's population starting in 1950 for the six countries that will be analyzed in this paper (Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela). As can be observed, maternity leave has been extended over the years in the countries examined and most changes occurred after the 1990's. The evolution of leave duration is associated to

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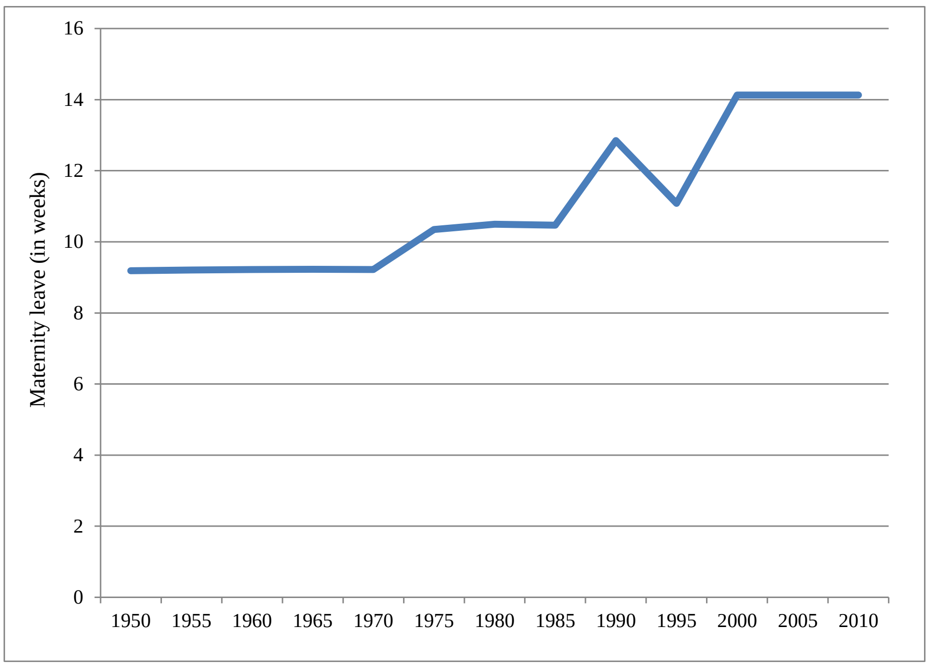
[6] The databases consulted from the ILO are: NATLEX - National Labor, Social Security, and Related Human Rights Legislation Database (retrieved from: http://www.ilo.org/dyn/natlex/natlex4.home) and TRAVAIL - Working Conditions Laws Database (retrieved from: http://www.ilo.org/dyn/travail/travmain.home). In the countries analyzed, legislation texts typically include references to previously existing regulations or to current complementary regulations. For instance, as a common practice, the text of a new legislation includes a reference to the legislation that it abolishes.

[7] The following databases were used to verify the constructed duration of leave: the TRAVAIL - Working Conditions Laws Database (retrieved from: http://www.ilo.org/dyn/travail/travmain.home), the ILO Social Security Database (retrieved from: http://www.ilo.org/dyn/sesame/IFPSES.SocialDatabase), and the World Bank Cross Country Data (retrieved from: https://www.quandl.com/WORLDBANK). These databases include a few data points for maternity leave duration, but they do not explicitly identify the exact year when duration changed. However, I was able to verify if the constructed series was at least consistent with the few available data points from these secondary sources.

recommendations established over time by the ILO conventions and other documents on the matter, which are being adopted progressively by countries around the world. In particular, ILO Conventions 3 (1919) and 103 (1952) recommended a minimum of 12 weeks of leave, while Convention 183 (2000) recommended 14 weeks, and Recommendation 191 (2000) a total of 18 weeks.

Figure 2.3: Weighted Duration of Maternity Leave (by Population) for Selected Countries in Latin America (1950-2010) [1/]



Source: World Population Prospects 2012 revision (Department of Economic and Social Affairs, United Nations) for each country's population and national legislation for labor regulations.

1. Data include the following countries: Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela. The duration of leave is weighted by each country's population.

Maternity leave regulations were instituted for most of the countries examined between the 1920s and 1940s and in all cases duration in weeks has increased at least once afterwards. For every country analyzed, maternity leave is paid and replacement rates are 100 percent or close to a fully paid salary. In Bolivia, maternity leave was regulated at 8.6 weeks (60 days) in the 1940s by the General Labor Law and it increased to 12.9 weeks in 1975. In the case of Chile, maternity leave had a duration of 9 weeks (in 1925), which was expanded to 12 weeks in 1966 and to 18 weeks in 1972. This benefit was initially only for laborers but it was extended in the 1950s to other types of workers. Colombia regulated maternity leave at 8 weeks in 1938 (and maintained the same duration in the Labor Code of 1950), increased it to 12 weeks with the reform of the labor laws in 1990, and to 14 weeks in 2011. The duration of maternity leave in Ecuador was 6 weeks in the Labor Code of 1961, extended to 8 weeks in 1978, and to 12 weeks in 1997. For Peru, maternity leave was mandated for a duration of 8.6 weeks (60 days) as far back as 1925 but became unregulated in 1995; and the benefit was reinstated only one year later, in 1996, at 12.9 weeks (90 days). In Venezuela, the Labor Law of 1936 instituted maternity leave for 12 weeks, but it increased to 18 weeks in 1990 and further to 26 weeks in 2012. These are the base leave durations, since many of these regulations include longer leave duration in special cases for mothers, such as premature, underweight, or multiple births.

2.4. Methodology

2.4.1. Data

In this paper, I use data from six countries in Latin America. These six countries were selected due to the availability of microdata covering a long span of time that includes years both before and after legislation changes. Data from the Minnesota

Population Center include a total of 41 censuses available from the region, which for some countries cover the entire period between the 1960s and 2010. This period seems to be a sufficient time span to observe the impacts of changes in labor regulations on maternity leave. Table 2.1 shows all the data years available from the Integrated Public Use Microdata Series (IPUMS) International that was used for the analysis. The datasets are 10 percent samples of the corresponding censuses, except for Chile 1960 (1 percent), Colombia 1964 (2 percent), and Ecuador 1962 (3 percent), Therefore, given the large proportion of the population included in the census samples, all the data are nationally representative and also representative for lower geographical units.

In addition, the census data are complemented by nationally representative household surveys, particularly for those countries with relatively few censuses available. In the case of Peru, for example, additional data points were obtained by using the Living Standards Measurement Study (LSMS) survey for 1985-86 and the National Household Survey (ENAHO) for 2013. The household surveys that are used in this paper are also included in Table 2.1.

Table 2.1: Availability of Census Microdata and Household Surveys for Selected Latin American Countries (Year of Data Collection)

Census round	1960-1969	1970-1979	1980-1989	1990-1999	2000-2009	2010
Bolivia	NA	1976	NA	1992	2001	EH 2009 *
Chile	1960	1970	1982	1992	2002	CASEN 2011 *
Colombia	1964	1973	1985	1993	2005	NA
Ecuador	1962	1974	1982	1990	2001	2010
Peru	NA	NA	LSMS 1985-86 *	1993	2007	ENAHO 2013 *
Venezuela	NA	1971	1981	1990	2001	NA

Source: Census microdata from the Integrated Public Use Microdata Series (IPUMS) International; the 1985-86 Living Standards Measurement Study (LSMS) and the 2013 National Household Survey (ENAHO) from the National Institute of Statistics and Informatics (INEI) - Peru; the 2009 Household Survey (EH) from the National Institute of Statistics (INE) - Bolivia; and the National Socioeconomic Characterization Survey (CASEN) from the Ministry of Social Development - Chile.

* = Household Survey. This includes the 1985-86 LSMS and 2013 ENAHO for Peru, the 2009 EH for Bolivia, and the 2011 CASEN for Chile.

NA=Not available

The following labor market outcomes will be examined: labor force participation, the employment-to-population ratio, and the unemployment-to-population ratio (shown in equations 2.1a, 2.1b, and 2.1c).[8] These outcomes are hypothesized to be affected by regulations on maternity leave, following the discussion in Section 2.2 of the conceptual framework. The labor force participation rate is defined as the ratio of the economically active population (those employed and unemployed) over the working-age population (that is, the active plus the inactive). The employment-to-population ratio is the ratio of

[8] At the individual level, the observed outcome is whether the person is employed, unemployed, or economically inactive. However, the outcomes of interest are the ratios shown in equations 2.1a, 2.1b, and 2.1c when the analysis is performed at the cohort level, as will be discussed next in the empirical strategy sub-section.

the employed over the working-age population, while the unemployment-to-population ratio is the ratio of the unemployed over the working-age-population.

$$\text{Labor Force Participation} = \frac{\text{Econ. Active population}}{\text{Econ. Active population} + \text{Econ. Inactive poulation}} \ldots (2.1a)$$

$$\text{Employment-to-population} = \frac{\text{Employed population}}{\text{Econ. Active population} + \text{Econ. Inactive poulation}} \ldots (2.1b)$$

$$\text{Unemployment-to-population} = \frac{\text{Unemployed population}}{\text{Econ. Active population} + \text{Econ. Inactive poulation}} \ldots (2.1c)$$

The three outcomes analyzed are defined using the person's labor force status: employed persons are those who performed work for pay or who were temporarily absent from their job, the unemployed individuals are those without a job that were available to work and were seeking a job, and the economically inactive population comprises those who are neither employed nor unemployed (United Nations, 2008; United Nations, 2010). The use of data from different sources raises the question whether labor force status has been consistently defined and whether the outcomes have any measurement error. Recommendations for census data collection suggest applying detailed questionnaires to help correctly classify specific cases, such as domestic services provided by paid workers (that are employed) or persons not currently seeking work but who made arrangements to start a paid job in the future (that are unemployed) (United nations, 2008). Furthermore, a short reference time period (such as week) is recommended to reduce possible recall errors (United Nations, 2008).

Based on these general recommendations, I analyzed the questionnaires from the data sources used in this essay. About half of the data sources included only a single question to determine labor force status (followed by other related questions such as the person's occupation or economic sector), while other datasets had a set of sequential questions for this purpose (which were more detailed for the household surveys than censuses). Even though the complexity of questions had some variability across data sources, all questionnaires distinguished those who were employed, unemployed, looking for work for the first time, and the different possible statuses for those economically inactive (such as students, pensioners, or persons performing unpaid household duties). In the case of unemployment, most questions referred directly to its definition criteria, that is, persons who are "not working but available and seeking a job."[9] The reference period used in the questionnaires was the previous week for most datasets (25 out of the 29), while only four of them referred to the person's "current" employment status. Therefore, the questionnaire design for the datasets used in this essay did not reveal evidence of major inconsistencies in data collection or potential measurement errors in the assessment of the person's labor force status. In addition, the influence of measurement errors in labor force status on the estimation results is expected to be small, because the effect of maternity leave duration are based on (cohort) fixed effects models (that will be described in the following sub-section).

The most important explanatory variable corresponds to the duration of maternity leave (in weeks) set by regulation for the selected countries. As previously discussed, this regulation is hypothesized to have an effect on labor market outcomes by creating incentives for working mothers to stay in the labor force and by changing the relative

[9] For instance, the Bolivia 1992 census includes options for those that "looked for work having worked before" and persons who "looked for work for the first time." For those data sources that instead asked whether a person was "unemployed," the enumeration instructions explained the criteria that needed to be met in order to classify a person under this status. For example, the Chile 1960 census asked to identify the "unemployed," which were defined in the enumeration instructions as persons that do not have an occupation but who are looking for work.

costs of employing men and women. In particular, maternity leave is expected to have a positive effect on the labor supply and a negative effect on the labor demand, which suggests a negative effect on wages and an ambiguous effect on employment. The duration of leave variable was created based on a review of the laws for each of the countries over the period of analysis (starting in the 1960's) and a summary of changes in leave duration by country is shown in Table 2.A2 in the Appendix to this chapter. Maternity leave is regulated at the national level for the selected countries, so there are no differences in duration within countries. Other variables used in the analysis include the person's educational attainment, marital status, and family size from the various microdata sources and the annual growth of Gross Domestic Product (GDP) per capita for each country from the World Development Indicators.

2.4.2. Empirical Strategy

The effects of the relevant labor market regulations will be assessed through a difference-in-differences-in-differences (DDD) approach. The identification strategy is based on: (i) the variation of the duration of maternity leave over time (at different times for different countries); and (ii) a comparison of women of childbearing age (between 18 to 30 years old) against other groups that are presumably not affected by changes in these regulations. Women of childbearing age are expected to be the most affected from increases in the duration of maternity leave. I also test whether there are differential effects on women of childbearing age who have already given birth. Two comparison groups are proposed, whose labor market outcomes should not be influenced by maternity leave. The first is women out of childbearing age (41 to 55 years old). The second is men between 18 and 30 years old, who are expected to be more similar to women of the same age range (in terms of human capital and other characteristics).

Two criteria guided the choice of cutoff ages for the control and treatment groups. The first criterion is based on the age ranges examined in previous research in this area. In the cross-country studies discussed in Section 2.2, the treatment group is often defined as women between 20 or 25 to 34 years old, while control groups generally include women out of childbearing age (45 to 54 or 60 years old) or men (Winegarden and Bracy, 1995; Ruhm, 1998; Akgunduz and Plantenga, 2013). The second criterion uses information on age-specific fertility rates for the countries under analysis. The average number of children born per woman for ages between 15 to 49 years old is shown in Table 2.A7 in the Appendix to this chapter. In the table, the number of children per woman starts increasing significantly at 18 years old (changes of 2 percentage points or more of total fertility for each additional year of age), while these increases per year of age generally peak at 30 years old (and additional changes become progressively smaller after this age). The proposed age range represents about half of the average number of children born per woman across the countries of interest. Even though a larger age range could be applied, it is reasonable to expect important decisions regarding participation in the labor market and fertility to be taken in the selected ages.

Consider the labor market outcome Y_{ijt}, where *i* denotes an individual, *j* is the group (women of childbearing age, older women, or men between 18 and 30 years old), *k* is the country, and *t* is the time period:

$$Y_{ijkt} = \alpha_j + \beta_j T_t + \varphi_j X_{ijkt} + \upsilon_j Z_{kt} + \lambda_j ML_{kt} + \theta_{ijk} + \varepsilon_{ijkt} \quad \ldots (2.2)$$

In equation (2.2), α_j is a group specific intercept, T_t is a general time effect common across all countries (corresponding to each decade or census round from Table 2.1), X_{ijkt} are time-varying controls for each individual (educational attainment and marital status dummies), Z_{kt} is a time-varying control at the country level (annual growth of GDP per capita), ML_{kt} is the duration of maternity leave (in weeks), θ_{ijk} is an individual fixed effect, and ε_{ijkt} is an error term. This empirical specification follows Ruhm (1998), Akgunduz and Plantenga (2013), and Thevenon and Solaz (2013).

Equation (2.2) is a difference-in-differences (DD) estimate given that maternity leave laws change at different times in different countries. This equation would not be a DD estimate if the analysis were based on one country. However, with multiple countries, those implementing changes in maternity leave duration at time "t" are the treatment group, while those where it remained unchanged serve as the control group. Thus, the country that extended the duration of maternity leave last is the control group for the rest. During the time period when data are available for this study, I observe one or two changes in leave duration for each country, except for Bolivia (which is used as a "control" throughout all time periods). Furthermore, most leave duration changes occur at different time periods, such that no more than two countries increase it simultaneously over any given time period.

Given that no longitudinal data are available to estimate this model, I will apply a pseudo-panel method (Deaton, 1985) to census microdata (complemented with household surveys). This approach averages the data within cohorts, which are treated as observations of the same unit over time. Consider the following simplified model (Deaton, 1985; Verbeek, 2008):

$$y_{it} = x_{it} \cdot \beta + \alpha_i + \varepsilon_{it} \quad \ldots (2.3)$$

In equation (2.3), x_{it} are controls, the subscript i refers to an individual observed in time period t (coming from independent cross-section datasets), and α_i is an individual fixed effect. We can define groups such that each individual i belongs to only one of them and this group membership is fixed over time (Deaton, 1985; Verbeek and Nijman, 1992; Verbeek, 2008). As Deaton (1985) suggests, an obvious criterion to define these groups are age cohorts, that is, all individuals born in a specific year or multiple years. Then, we can aggregate all individuals i belonging to cohort c in time t, such that we obtain a model based on the sample cohort means:[10]

$$\bar{y}_{ct} = \bar{x}_{ct} \cdot \beta + \bar{\alpha}_c + \bar{\varepsilon}_{ct} \quad \ldots (2.4)$$

This produces a pseudo or synthetic panel where the unit of observation is the cohort over time. This model yields consistent estimates for β even if $\bar{\alpha}_c$ is correlated with any of the controls, by treating $\bar{\alpha}_c$ as a parameter to estimate. In this model, $\bar{\alpha}_c$ is treated as a time-invariant fixed effect for the cohort, which is reasonable if we average across a large number of observations (Verbeek and Nijman, 1992; Verbeek, 2008).

[10] Even though the analysis uses mostly censuses, this model is still based on sample cohort means and not population estimates, given that the datasets are samples from the original census data (10 percent in most cases).

In this paper, the basic set of results is produced using single years of birth to define the cohorts, so that the relevant variables are transformed into cohort means by birth year. In addition, some results are based on cohorts by year of birth and educational attainment, identifying persons with less than primary education completed and those with primary completed or further education.[11] Since the analysis is based primarily on census data (complemented with household surveys), I obtain a reasonable number of observations within each cohort. Previous methodological research suggests that cohorts for the synthetic panel should comprise more than 100 to 200 observations (Verbeek and Nijman, 1992; Verbeek and Vella, 2005), while several empirical applications worked with minimum cohort sizes of about 100 to 500 observations (Browning, Deaton, and Irish, 1985; Banks, Blundell, and Preston, 1994; Blundell, Browning and Meghir, 1994; Blundell, Duncan, and Meghir, 1998; Propper, Rees, and Green, 2001; Warunsiri and McNown, 2010).

Returning to the full model in equation (2.2) and after taking the cohort means of each variable, the cohort fixed effects model is defined as:

$$\bar{Y}_{cjkt} = \alpha_j + \beta_j T_t + \varphi_j \bar{X}_{cjkt} + \upsilon_j Z_{kt} + \lambda_j ML_{kt} + \bar{\theta}_{cjk} + \bar{\varepsilon}_{cjkt} \quad \ldots (2.5)$$

In equation (2.5), c denotes the cohort and $\bar{\theta}_{cjk}$ is a cohort fixed effect, so we are controlling for unobserved cohort characteristics that are fixed over time, and other controls are similar to equation (2.2) but expressed as cohort means. In effect, equation

[11] Given that the pseudo-panel method requires that group membership is fixed over time, the assumption is that persons 18 years old or more with less than primary education or primary completed are unlikely to switch groups over time. Blundell, Duncan, and Meghir (1998) also construct a pseudo-panel based on year of birth and educational attainment, although the groups identified for the cohorts correspond to "those who left education at the minimum legal age and those who continued beyond the minimum" (p. 838).

(2.5) is analogous to equation (2.2) but it is based on sample cohort means data for $\overline{Y}_{cjkt}$ and $\overline{X}_{cjkt}$.

Next, consider a treatment (*j*=*f*) and a control group (*j*=*m*), such that:

$$\overline{Y}_{cfkt} = \alpha_f + \beta_f T_t + \varphi_f \overline{X}_{cfkt} + \upsilon_f Z_{kt} + \lambda_f ML_{kt} + \overline{\theta}_{cjk} + \overline{\varepsilon}_{cfkt} \quad \ldots (2.6a)$$

$$\overline{Y}_{cmkt} = \alpha_m + \beta_m T_t + \varphi_m \overline{X}_{cmkt} + \upsilon_m Z_{kt} + \lambda_m ML_{kt} + \overline{\theta}_{cjk} + \overline{\varepsilon}_{cmkt} \quad \ldots (2.6b)$$

As previously discussed, equations (2.6a) and (2.6b) are already difference-in-differences (DD) estimates. Furthermore, comparisons will be implemented between women of childbearing age (18 to 30 years old), the treatment group, against men in the same age range and against older women (41 to 55 years old), the two control groups. By defining the treatment and two comparison groups, it allows one to control for other country specific factors contemporaneous to changes in maternity leave duration that could have an effect on the relevant labor market outcomes. That is, the use of treatment and comparison groups allows controlling for possible bias if these country specific factors are correlated with increases in maternity leave duration; for example, if maternity leave benefits are extended when labor force participation or employment rates are growing (Ruhm, 1998). The intuition is that these country specific factors would affect both the control and treatment groups, so the differential in outcomes across groups eliminates the possible bias introduced by these factors in the estimates of maternity leave duration effects.

Thus, the difference-in-differences-in-differences (DDD) model is estimated as:

$$\overline{Y}_{ckt} = \alpha + \omega F + \beta_1 T_t + \beta_2 F * T_t + \varphi_1 \overline{X}_{ckt} + \varphi_2 F * \overline{X}_{ckt} + \upsilon_1 Z_{kt} + \upsilon_2 F * Z_{kt}$$

$$+ \lambda_1 ML_{kt} + \lambda_2 F * ML_{kt} + \overline{\theta}_{ck} + \overline{\varepsilon}_{ckt} \ \ldots \ (2.7)$$

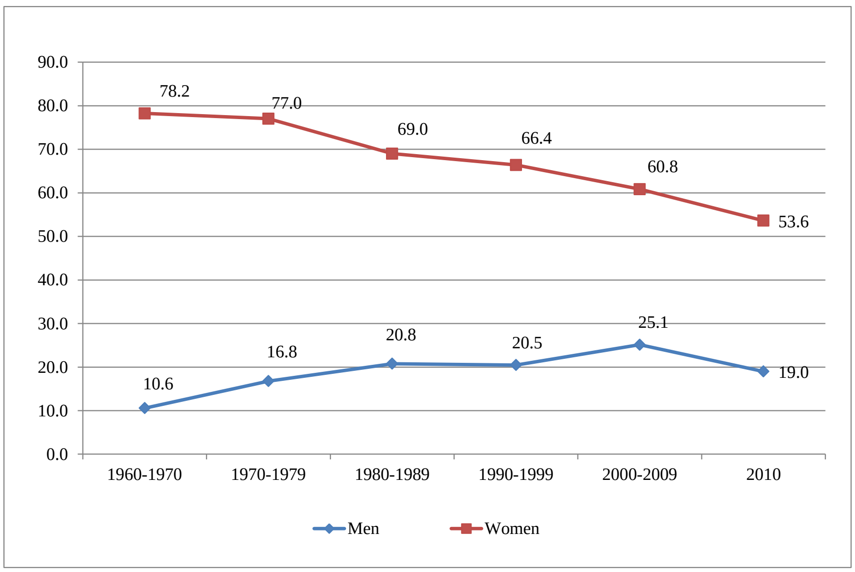
In equation (2.7), F is a dummy variable equal to 1 for women of childbearing age (18 to 30 years old) and to 0 otherwise and the model includes interactions of F with the other controls. In this equation, the differential effect between the treatment and control groups is represented by λ_2 for maternity leave. From equations (2.6a), (2.6b), and (2.7), we infer that $\lambda_1=\lambda_m$ and $\lambda_2=\lambda_f-\lambda_m$. Based on the conceptual framework discussion, it would be expected that $\lambda_m=0$, such that the effects on the labor market outcomes should be identified through λ_2. The model in equation (2.7) is a difference-in-differences-in-differences (DDD) estimate given that changes in maternity leave duration occur at different times in the countries under analysis (such that some countries are controls for those extending duration) and comparisons are also performed between a treatment (women of childbearing age) and two control groups within each country (women out of childbearing age and men between 18 and 30 years old).

2.5. <u>Results</u>

2.5.1. Labor Market Trends by Gender in Selected Countries

This subsection presents trends found in the labor markets of the six countries of this study to provide background for the analysis. The labor market gaps by gender in Latin America have generally decreased over time. The proportion of the population that is inactive has been growing for men while it has been decreasing considerably for women; therefore, the gap in labor force participation (the complement to the inactive population) is closing over time for all countries examined (Figure 4 below and Figure A1 in the Appendix to this chapter). Nevertheless, the proportion of inactive women is still very high in all years analyzed, ranging from 70-80 percent in the 1960s census round to 35-55 percent for more recent data.

Figure 2.4: Weighted Inactivity Rates (%) for Population Ages 15-64 for Selected Countries in Latin America (1960-2010) [1/]



Data source: Integrated Public Use Microdata Series (IPUMS) International.
1. Data include the following countries: Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela. The inactivity rates have been calculated using only census microdata (census years shown in Table 2.1) and the rates are weighted by the population of each country.

The proportion of employed women (with respect to all women in the labor force) by education and other demographic characteristics shows some clear patterns in the six countries examined. The results for women between 15 and 64 years old are shown in Table 2.2; they include only the 2000 census round, but similar figures were found for all census rounds available for each country. As expected, the proportion of women employed is higher at older ages, reaching usually a maximum at 35-44 years old and decreasing afterwards (upper part of Table 2.2). This proportion increased over time for all age groups and the largest growth is generally found for women between 35-54 years old. The employment rates are consistently higher for more educated women (middle part of Table 2.2), but the changes over time show that women with less than primary and primary completed tended to have higher increases. Finally, even though the number of children under age 5 living in the household is associated with lower employment rates (lower part of Table 2.2), employment rates have also increased over time for this group at a relatively high rate. Therefore, the largest growth in employment rates for women in the six countries under analysis is associated with women at older ages, women with low educational attainment, and women with more children.

Participation in unpaid work is generally considered to be an indicator of low labor productivity. The proportion of the employed population who participate in unpaid family work by gender is shown in Figure A2 in the Appendix to this chapter. This gender gap is positive (i.e. higher rates for women) but it is narrowing over time in Bolivia and Peru; however, men generally have higher proportions of unpaid work than women in Chile, Ecuador, and Venezuela. Therefore, no clear pattern is identified based on this indicator.

Table 2.2: Employed Women (%) Ages 15-64 for Selected Countries in Latin America (Census Round 2000) by Age, Education, and Number of Children Under 5 Years Old

	Bolivia 2001	Chile 2002	Colombia 2005	Ecuador 2001	Peru 2007	Venezuela 2001
Age group						
15-24 years	33.63	20.40	23.22	28.21	28.08	20.07
25-34 years	48.47	43.11	45.14	41.00	45.84	42.73
35-44 years	53.05	41.17	45.74	42.16	48.24	48.73
45-54 years	50.71	39.58	38.68	37.87	44.96	41.85
55-64 years	40.25	24.23	18.73	27.04	31.28	20.86
Obs.	*233,543*	*507,544*	*1,188,789*	*357,263*	*877,188*	*725,831*
Educational attainment						
Less than primary	43.67	20.77	20.04	28.19	30.30	21.86
Primary	38.62	25.55	27.83	29.17	28.64	28.56
Secondary	47.94	45.67	45.07	47.09	44.18	48.73
University	68.98	71.59	71.29	73.83	67.40	55.50
Obs.	*231,047*	*507,544*	*1,184,080*	*356,079*	*877,188*	*723,046*
Children under 5 years						
No children	45.07	35.43	36.72	36.98	41.37	36.60
1 child	43.25	30.47	33.47	32.76	34.85	33.35
2+ children	35.27	25.78	20.95	25.41	27.04	20.98
Obs.	*233,543*	*507,544*	*1,188,789*	*357,263*	*877,188*	*725,831*

Data source: Integrated Public Use Microdata Series (IPUMS) International.

The data on occupation and sector (industry) by gender show clear patterns of segregation. The sectors with higher proportions of women, in wage or salaried employment, in all countries analyzed correspond to service jobs (Table 2.A3 in the Appendix to this chapter): in this sector, we find about 60-80 percent of women compared to only 25-45 percent of men. In turn, men in wage or salaried employment have higher concentrations in either agriculture/mining or manufacturing (about 50-60 percent). This gender segregation by sector of employment remains relatively constant over time. The classification of workers in wage or salaried employment by occupation shows similar evidence with clear trends across gender (Table 2.A4 in the Appendix to this chapter). Women are more concentrated in clerk or service occupations, while men work in higher proportions in craft or plant/machine operator occupations. Other occupational categories have smaller differences, or have no clear difference by gender.

The Duncan Index of dissimilarity indicates the percentage of the population that would have to be moved between sectors or occupations to achieve a similar distribution across gender. The Duncan index for these countries shows moderate to high values of segregation for occupation (ranging from 0.21 to 0.64) and it shows a generally decreasing pattern over time (Table 2.A6 in the Appendix to this chapter). In turn, the index shows relatively large values for industry (ranging from 0.37 to 0.64), but also with a decreasing trend over time (Table 2.A5 in the Appendix to this chapter). The highest values of occupational segregation are found for Colombia, while the highest values for industry segregation are for Bolivia.

2.5.2. Effects of Maternity Leave on Labor Force Participation, the Employment-to-Population Ratio, and the Unemployment-to-Population Ratio

The difference-in-differences (DD) estimates of equations (2.6a) and (2.6b) for cohorts defined by year of birth are shown in Table 2.3 for labor force participation, the employment-to-population ratio, and the unemployment-to-population ratio for the treatment and control groups. If there are effects of maternity leave regulations, these effects may be stronger for young mothers (as compared to other women of childbearing age), since this policy is designed precisely to reconcile work and family responsibilities. The table shown below includes only the estimated coefficients for maternity leave, but the model controls for educational attainment, marital status, family size, annual growth of the GDP per capita, and dummy variables for each time period.

Results show that maternity leave duration has a positive effect on labor force participation, but it is statistically significant only for young women (mothers) and men. For these groups, this effect implies an increase of about 1.0 to 1.1 percentage points for an additional week of leave, with respect to an activity rate around 20-40 percent for young women and 80 percent for young men. The effects on employment are generally not statistically significant, except for a positive and marginally significant effect for older women (p=0.09). Even though the estimated effect on employment for mothers between 18 and 30 years old is larger than the effect on all women of childbearing age, neither of these is statistically significant. For unemployment, I obtain positive and statistically significant effects for all groups, except for older women. For young women, this effect implies an increase of 0.6 percentage points with respect to an unemployment-to-population ratio of about 1 to 3 percent over time. The effect for young mothers is considerably smaller, about half of the overall effect on young women. For young men, there is an increase of about 0.5 percentage points with respect to an unemployment-to-population ratio over time of about 4 to 7 percent. Therefore, these effects are relatively important with respect to the unemployment-to-population ratio for young women, while they are smaller for young men and young mothers.

Table 2.3: Difference-in-Differences (DD) Model, Pseudo-Panel (Year of Birth) Cohort Fixed-Effects Regressions for Labor Market Outcomes [1/]

Outcome	Treatment groups		Control groups	
	Women 18-30 years old	**Mothers 18-30 years old**	**Men 18-30 years old**	**Women 41-55 years old**
Labor force participation				
Mat. leave coefficient	**0.0116****	**0.0114***	**0.0104***	**0.0071**
	[0.0046]	*[0.0061]*	*[0.0063]*	*[0.0045]*
Obs.	338	335	338	389
Groups	274	272	274	288
Cohort size (average)	12,121	7,013	11,611	5,764
Employment-to-population ratio				
Mat. leave coefficient	**0.0050**	**0.0084**	**0.0055**	**0.0075***
	[0.0039]	*[0.0058]*	*[0.0058]*	*[0.0043]*
Obs.	338	335	338	389
Groups	274	272	274	288
Cohort size (average)	12,121	7,013	11,611	5,764
Unemployment-to-population ratio				
Mat. leave coefficient	**0.0066*****	**0.0029****	**0.0049****	**-0.0003**
	[0.0016]	*[0.0013]*	*[0.0024]*	*[0.0004]*
Obs.	338	335	338	389
Groups	274	272	274	288
Cohort size (average)	12,121	7,013	11,611	5,764

Data source: Census microdata from the Integrated Public Use Microdata Series (IPUMS) - International; selected household surveys from the National Institute of Statistics and Informatics (INEI) - Peru, the National Institute of Statistics (INE) - Bolivia, and the Ministry of Social Development - Chile.

Clustered standard errors (for the cohort) in brackets, *** p<.01,** p<.05, * p<.10

1. Regressions control for educational attainment (dummies for completed primary, completed secondary, and completed university), marital status, family size, annual growth of GDP per capita, and dummy variables for each time period.

The estimates for the difference-in-differences-in-differences (DDD) model from equation (2.7) for cohorts defined by year of birth are shown in Table 2.4, which include comparisons of young women (and mothers) against the two comparison groups. The table presents the main effect of maternity leave, as well as the interaction with women of childbearing age. Results show that both the main effect of maternity leave and its interaction are generally not statistically significant for labor force participation or the employment-to-population ratio. The main effects of maternity leave ("maternity leave coefficient" in the DDD estimates) suggest positive effects on labor force participation for younger men and on the employment-to-population ratio for older women, but both are only marginally statistically significant. Nevertheless, maternity leave duration does increase unemployment for women of childbearing age, as we observe in the comparisons against older women in the first two columns. This effect is smaller for young mothers with respect to all women of childbearing age. Furthermore, I also find statistically significant effects on unemployment for young men ("maternity leave coefficient" in the last two columns of DDD estimates). Based on the sign of the interactions of maternity leave with women of childbearing age, results imply that increases in unemployment are larger for young women with respect to young men and larger for young men with respect to young mothers (even though the interaction coefficients are not statistically significant in both cases).

Table 2.4: Difference-in-Difference-in-Differences (DDD) Model, Pseudo-Panel (Year of Birth) Cohort Fixed-Effects Regressions for Labor Market Outcomes (Women and Mothers Ages 18-30 against Comparison Groups) [1/]

Outcome	Comparison			
	Women 18-30 against women 41-55	**Mothers 18-30 against women 41-55**	**Women 18-30 against men 18-30**	**Mothers 18-30 against men 18-30**
Labor force participation				
Mat. leave coefficient	**0.0071**	**0.0071**	**0.0104***	**0.0104***
	[0.0045]	*[0.0045]*	*[0.0063]*	*[0.0063]*
Interaction with women in ch. age	**0.0045**	**0.0042**	**0.0011**	**0.0009**
	[0.0063]	*[0.0076]*	*[0.0078]*	*[0.0088]*
Obs.	727	724	676	673
Groups	562	560	548	546
Cohort size (average)	8,720	6,342	11,867	9,323
Employment-to-population ratio				
Mat. leave coefficient	**0.0075***	**0.0075***	**0.0055**	**0.0055**
	[0.0043]	*[0.0043]*	*[0.0058]*	*[0.0058]*
Interaction with women in ch. age	**-0.0024**	**0.0009**	**-0.0004**	**0.0028**
	[0.0057]	*[0.0072]*	*[0.0069]*	*[0.0082]*
Obs.	727	724	676	673
Groups	562	560	548	546
Cohort size (average)	8,720	6,342	11,867	9,323
Unemployment-to-population ratio				
Mat. leave coefficient	**-0.0003**	**-0.0003**	**0.0049****	**0.0049****
	[0.0003]	*[0.0004]*	*[0.0024]*	*[0.0024]*
Interaction with women in ch. age	**0.0069*****	**0.0033****	**0.0016**	**-0.0019**
	[0.0016]	*[0.0013]*	*[0.0028]*	*[0.0026]*
Obs.	727	724	676	673
Groups	562	560	548	546
Cohort size (average)	8,720	6,342	11,867	9,323

Data source: Census microdata from the Integrated Public Use Microdata Series (IPUMS) - International; selected household surveys from the National Institute of Statistics and Informatics (INEI) - Peru, the National Institute of Statistics (INE) - Bolivia, and the Ministry of Social Development - Chile.

Clustered standard errors (for the cohort) in brackets, *** p<.01,** p<.05, * p<.10

1. Regressions control for educational attainment (dummies for completed primary, completed secondary, and completed university), marital status, family size, annual growth of GDP per capita, dummy variables for each time period, and interactions of these with women of childbearing age. Results are numerically equivalent in the first, third, and fifth row (main effect of maternity leave) if they are based on the same comparison group.

The analysis was extended to cohorts defined by year of birth and educational attainment (identifying persons with less than primary and those with primary completed or further education) for young women and men. The additional set of estimates allows one to check whether results are consistent when year of birth cohorts are split by education and it also produces a larger number of cohort observations to gain statistical precision. The difference-in-differences (DD) estimates of equations (2.6a) and (2.6b) for cohorts defined by year of birth and educational attainment are shown in Table 2.5 for labor force participation, the employment-to-population ratio, and the unemployment-to-population ratio for young women and men. The number of observations in Table 2.5 is almost double the number of cases in the corresponding estimates presented in Table 2.3 due to the combination of years of birth with educational attainment.

Results show that the effects on labor force participation and the employment-to-population ratio are not statistically significant. The effects on the unemployment-to-population ratio show an increase of about 0.6 to 0.7 percentage points for an additional week of leave. The difference-in-differences-in-differences (DDD) estimates (not shown here) suggest that the effects for young women and men are not statistically different. Thus, the size and statistical significance of the unemployment effects are consistent with previous results based on year of birth cohorts.

Table 2.5: Difference-in-Differences (DD) Model, Pseudo-Panel (Year of Birth and Educational Attainment) Cohort Fixed-Effects Regressions for Labor Market Outcomes [1/]

	Women 18-30 years old	Men 18-30 years old
Labor force participation		
Mat. leave coefficient	**0.0035**	**0.0065**
	[0.0048]	*[0.0044]*
Obs.	639	639
Groups	522	526
Cohort size (average)	6,410	6,141
Employment-to-population ratio		
Mat. leave coefficient	**-0.0022**	**-0.0011**
	[0.0042]	*[0.0043]*
Obs.	639	639
Groups	522	526
Cohort size (average)	6,410	6,141
Unemployment-to-population ratio		
Mat. leave coefficient	**0.0057*** **	**0.0077*** **
	[0.0012]	*[0.0017]*
Obs.	639	639
Groups	522	526
Cohort size (average)	6,410	6,141

Data source: Census microdata from the Integrated Public Use Microdata Series (IPUMS) - International; selected household surveys from the National Institute of Statistics and Informatics (INEI) - Peru, the National Institute of Statistics (INE) - Bolivia, and the Ministry of Social Development - Chile.

Clustered standard errors (for the cohort) in brackets, *** p<.01,** p<.05, * p<.10

1. Regressions control for educational attainment (dummies for completed primary, completed secondary, and completed university), marital status, family size, growth of GDP per capita, and dummy variables for each time period.

Finally, I also verified results using a placebo or falsification test. The objective is to test whether results could be influenced by other policies that may have been implemented at similar time periods to increases in maternity leave duration or by general trends in the labor market. Similar exercises have been performed by previous research in this topic (Baker and Milligan, 2008a; Rossin-Slater *et al.*, 2013; Das and Polachek, 2014). In particular, using the same data, the changes in duration of leave were coded as if they occurred one time period earlier than the date of true enactment. If the estimated results reflect the effect of changes in maternity leave duration and not some other contemporaneous factors, the estimates of the leave duration variable for this exercise should not be statistically significant. Results for the falsification test are presented in Table 2.6. In all cases, I obtain relatively small coefficients, and none of them is statistically significant. This provides further evidence to support estimated effects shown for labor force participation, employment-to-population, and unemployment to population. Table 2.6 includes results only for the difference-in-differences estimates, but the triple difference model provides similar evidence.

Table 2.6: Falsification Test, Difference-in-Differences (DD) Model, Pseudo-Panel (Year of Birth) Cohort Fixed-Effects Regressions for Labor Market Outcomes [1/]

Outcome	Treatments		Controls	
	Women 18-30 years old	Mothers 18-30 years old	Men 18-30 years old	Women 41-55 years old
Labor force participation				
Mat. leave coefficient	**-0.0045**	**-0.0042**	**-0.0041**	**-0.0034**
	[0.0062]	*[0.0075]*	*[0.0074]*	*[0.0053]*
Obs.	338	335	338	389
Groups	274	272	274	288
Cohort size (average)	12,121	7,013	11,611	5,764
Employment-to-population ratio				
Mat. leave coefficient	**-0.0019**	**-0.0037**	**-0.0058**	**-0.0038**
	[0.0049]	*[0.0073]*	*[0.0067]*	*[0.0052]*
Obs.	338	335	338	389
Groups	274	272	274	288
Cohort size (average)	12,121	7,013	11,611	5,764
Unemployment-to-population ratio				
Mat. leave coefficient	**-0.0025**	**-0.0005**	**0.0017**	**0.0003**
	[0.0020]	*[0.0013]*	*[0.0025]*	*[0.0004]*
Obs.	338	335	338	389
Groups	274	272	274	288
Cohort size (average)	12,121	7,013	11,611	5,764

Data source: Census microdata from the Integrated Public Use Microdata Series (IPUMS) - International; selected household surveys from the National Institute of Statistics and Informatics (INEI) - Peru, the National Institute of Statistics (INE) - Bolivia, and the Ministry of Social Development - Chile.

Clustered standard errors (for the cohort) in brackets, *** p<.01,** p<.05, * p<.10

1. Regressions control for educational attainment (dummies for completed primary, completed secondary, and completed university), marital status, family size, growth of GDP per capita, and dummy variables for each time period.

2.6. Discussion

The evidence presented suggests that maternity leave creates incentives for women of childbearing age to participate in the labor market. The size of the effects is similar for both the sample of all women and the sample of those that have already given birth. Even though the employment effects for women (or mothers) in childbearing age are positive, they are not statistically significant. Given that increases in employment depend also on the labor demand side, it is possible that employers perceive women as more costly due to the potential need to hire temporary workers or to reorganize production during the time that new mothers are on leave. Therefore, women are not necessarily getting more jobs, despite the higher incentives for them to participate in the labor market, which seems to drive the increase in unemployment.

The positive effects in labor force participation and unemployment observed on young men are unexpected. Two possible factors may explain these effects. If maternity leave is indeed helping young women to make work and family responsibilities more compatible, it is possible that this also reduces the family burden for young men and creates more incentives for them to work. Furthermore, young men may be facing increased competition in the labor market if indeed maternity leave is creating incentives for more women to be economically active. This incentive may be translated not only in changes in employment status but also in the desire to work additional hours or full-time (rather than part-time) when the duration of maternity leave increases. Thus, increases in the women's economically active population could lead to higher unemployment for young men.

What is the preferred comparison group? A priori, neither group is preferred, since young men may be more similar to young women given possible cohort differences, but there could also be joint household decisions affecting the labor supply of young men that do not influence older women (Ruhm, 1998). The evidence shows some effects on labor force participation and unemployment for young men (which were not expected) but no statistically significant effects on labor market outcomes for older women (only marginally significant for employment). This suggests that the latter group may be a better comparison than the former to estimate the effects of this policy.

How does results compare to the previous research? The literature on this topic reports some positive effects on labor market outcomes associated with maternity leave, mainly on labor force participation, employment, and working hours (Winegarden and Bracy, 1995; Ruhm, 1998; Akgunduz and Plantenga, 2013; Thévenon and Solaz, 2013), but its importance may be conditional on the specific leave duration and also on whether leave is paid or unpaid. In the six countries being examined, leave duration over time increased from about 6 or 8 weeks to 12 weeks between the years 1960 to 2010. In comparison, Thévenon and Solaz (2013) calculate an average duration of maternity leave around 19 weeks in 2011 across the OECD countries, with even much longer leave duration for certain countries (such as the 52 weeks offered in United Kingdom). Moreover, these figures do not include additional (parental) leave not specifically allocated to the mother. Previous evidence shows that some effects depend on leave duration, such that there may be an optimal duration of leave. For example, Thévenon and Solaz (2013) report positive and statistically significant effects of leave duration on women's employment for periods between 1 and 2 years, while shorter leave durations showed some negative effects (although not statistically significant). Thus, it is possible that we do not observe statistically significant employment effects given the relatively short leave durations in the six countries under analysis (and the Latin American region more generally) compared to other regions previously studied.

Furthermore, there is some evidence on unemployment effects associated with maternity leave. Das and Polacheck (2014) found positive effects of the California Paid

Family Leave (CPFL) on the unemployment rate and unemployment duration among young women in California in comparison with states that did not adopt this policy. The authors argue that increases in labor force participation combined with a reduced demand for young women (due to hiring of temporary workers and depreciation of worker's skills during leave) may have led to higher unemployment rates for this group. Therefore, even though the literature on this field generally reports some positive effects on labor market outcomes, there is some evidence supporting the increases in unemployment found for the six Latin American countries included in this study.

The data available have some limitations for the analysis. Some of the data sources do not include specific instructions on how women on leave should be coded. For instance, among the Chile census samples, only the 1982 census explicitly states that "leave" should be considered as being employed but not at work, the same as sickness, vacations, and other reasons. Even though employed but not at work appears as an option in the relevant questionnaire item in most cases, it is possible that some women on leave may have been classified as unemployed, but no secondary data are available to test for this issue. Other limitations arise from information unavailable in our data sources. Some adjustment in women's labor market behavior in response to longer leave duration may occur through changes in working hours (or the proportion of women working full- versus part-time), rather than employment status. However, only a few census datasets collect information on this matter. Furthermore, countries often have some eligibility requirements expressed in terms of previous work or previous contributions to be able to take leave. This information would allow for identification of a sample of women for whom we would expect a larger effect of increases in leave duration, as it has been done by previous studies (Rossin-Slater, Ruhm, and Waldfogel, 2013; Baum and Ruhm, 2013), but work history is not available in the census data from these six countries.

Finally, the nature of the data being analyzed implies that I am able to observe only the employment status at the time of data collection (for the same cohort at different points in time), but not labor market behavior around the time of a birth or the effects on other outcomes over time. Therefore, it is not feasible to examine changes in leave-taking, time on leave, or job continuity, which are the focus of a number of studies in this area (Waldfogel, 1999; Ten Cate, 2000; Baum, 2003; Baker and Milligan, 2008a; Han, Ruhm, and Waldfogel, 2009; Hanratty and Trzcinski, 2009; Pronzato, 2009; Baum and Ruhm, 2013; Rossin-Slater, Ruhm, and Waldfogel, 2013). Even though the findings in this paper suggest that a longer duration of maternity leave is associated with increased labor force participation and higher unemployment for women of childbearing age, other outcomes for women or their children may have been positively influenced. For instance, job continuity for new mothers could lead to higher accumulation of human capital specific to the firm and better professional prospects in the long run. In addition, increased leave-taking could be associated with extended breastfeeding and positive effects on child development (Baker and Milligan, 2008b; Baker and Milligan, 2010). Thus, further investigation is required on these topics.

2.7. Appendix: Additional Tables and Figures

Table 2.A1: Summary of Studies on the Labor Market Effects of Maternity Leave

Study	Coverage		Data			Estimation strategy
	Geographic	Time	Source	Leave variable	Outcomes	
Cross-country (OECD or Europe)						
Winegarden and Bracy (1995)	17 OECD countries	Four time periods (1959, 1969, 1979, and 1989)	Various sources, country level	Weeks of paid maternity leave (maximum length)	Infant mortality, labor force participation, and general fertility rate	Fixed (country) effects, simultaneous equations for the three outcomes
Ruhm (1998)	9 European countries	1969 to 1993	Various sources, country level	Weeks of paid leave multiplied by the average wage replacement rate ("full-pay" weeks)	Employment-to-population ratio and hourly wages	Diff-in-diff, comparisons of (i) women against men and (ii) women aged 25-34 against men aged 25-34 and women aged 45-54
Pronzato (2009)	9 European countries	1994 to 2001	European Community Household Panel (ECHP)	Dummy variables to indicate whether women are eligible for job-protected leave or transfers	Time on leave (timing of return to work)	Hazard discrete models, sample of women who had a child at the time of the survey
Thevenon and Solaz (2013)	30 OECD countries	1970 to 2010	Various sources, country level	Weeks of paid maternity leave (maximum length)	Employment-to-population ratio, weekly working hours, and weekly earnings of full-time employees	Diff-in-diff, comparison of women and men aged 25 to 54
Akgunduz and Plantenga (2013)	16 European countries	1970 to 2010	Various sources, country level	Weighted leave (maternity and parental leave weighted by wage replacement rates)	Employment-to-population ratio, weekly working hours, hourly wages in manufacturing/financial intermediation, and share of women in high-level occupations	Diff-in-diff, comparisons of (i) women against men and (ii) women aged 25-34 against men aged 45-54

United States						
Klerman and Leibowitz (1997)	United States	1980 and 1990	Census microdata	Dummy variables for the periods after maternity state laws were in force	Employment rate (distinguishes on leave and at work)	Diff-in-diff, comparison of mothers of infants (under age 1) and those with older children (ages 2 to 3)
Waldfogel (1999)	United States	1992 to 1995	Current Population Survey (CPS), March	Dummy variables for maternity state laws and for the period after the implementation of the FMLA	Leave coverage, leave-taking, employment rate, and hourly wages	Diff-in-diff, comparison of women aged 19-45 with children and women aged 19-45 with infants (under age 1) against childless women aged 19-45, men aged 19-45, and women aged 46 to 60
Baum (2003)	United States	1988 to 1994	National Longitudinal Survey of Youth (NLSY)	Dummy variables for availability of leave regulation (and weeks of leave) across maternity state laws or FMLA	Leave-taking, probability of return to pre-childbirth job, and time on leave (timing of return to work)	Sample of women who gave birth during data collection
Han, Ruhm, and Waldfogel (2009)	United States	1988 to 2004	Current Population Survey (CPS), June Fertility Supplements and other months	Dummy variable for availability of leave regulation across FMLA, state temporary disability insurance (TDI) programs, and maternity state laws (or length of leave)	Employment rate and time on leave (timing of return to work)	Diff-in-diff, comparison of new mothers or fathers (during the birth month and three months after) against those who will have a birth 11 or 12 months after the survey date
Rossin-Slater, Ruhm, and Waldfogel (2013)	United States (California versus other states)	1999 to 2010	Current Population Survey (CPS), March	Dummy variable for the period after the California Paid Family Leave (CPFL) was in force, July 2004	Leave-taking (different definitions), work status and working hours last week, work status and working hours last year, and yearly wage income	Diff-in-diff, comparison of mothers with infants or young children (1- to 3-year-old) in California to women with older children (5- to 17-years-old) and childless women in California, and mothers with infants in other states
Baum and Ruhm (2013)	United States (California versus selected states)	2000 to 2010	National Longitudinal Survey of Youth (NLSY)	Dummy variable for births after the California Paid Family Leave (CPFL) was in force, July 2004 (coded as zero for births in other states or in California before that date)	Leave-taking, time on leave (timing of return to work), probability of return to pre-childbirth job, hourly wages, weekly working hours, and annual weeks of work	Diff-in-diff, comparison of California parents (mothers and fathers) of children born before/after the CPFL and against the corresponding parents in selected states (based on pre-program trends and number of observations available)

Das and Polachek (2014)	United States (California versus other states)	1996 to 2002	Current Population Survey (CPS), March	Dummy variable for the period after the California Paid Family Leave (CPFL) was in force, July 2004	Labor force participation, unemployment rate, and unemployment duration	Diff-in-diff, comparison of young women (less than 42 years old) in California to other states, as well as young men and older workers (men and women)
Canada						
Ten Cate (2000)	Canada	1993 to 1998	Survey of Labour and Income Dynamics (SLID) (longitudinal data)	Weeks of job-protected leave at the province level	Probability of return to pre-childbirth job and time on leave (timing of return to work)	Variation in duration of job-protected leave at the province level (across provinces and over time)
Ten Cate (2003)	Canada	1976 to 2000	Labor Force Survey (LFS), March and September	Weeks of job-protected leave at the province level	Employment rate	Diff-in-diff, comparison of women with young children (0- to 2-year-old) to women with older children (3- to 5-years-old)
Baker and Milligan (2008a)	Canada	1976 to 2002	Labor Force Survey (LFS), April and October	Weeks of job-protected leave at the province level // Dummy variable indicating an extension of job-protected leave	Employment rate (employed and at work, employed and on leave) and job continuity	Main results based on a sample of mothers with a child less than 1 year old and exploiting variation in duration of job-protected leave at the province level (across provinces and over time); additional results performing comparisons against married men and married childless women
Hanratty and Trzcinski (2009)	Canada	1999 to 2003	National Longitudinal Survey of Chidren and Youth (NLSCY)	Dummy variable for the period after the latest expansion in leave duration, in December 2000	Employment rate and time on leave (timing of return to work)	Diff-in-diff, comparison of women with children age 1 year old against those with children age 3 to 4 years old

Table 2.A2: Evolution of Maternity Leave Duration (in Weeks) for Selected Countries (1925-2012) [1/]

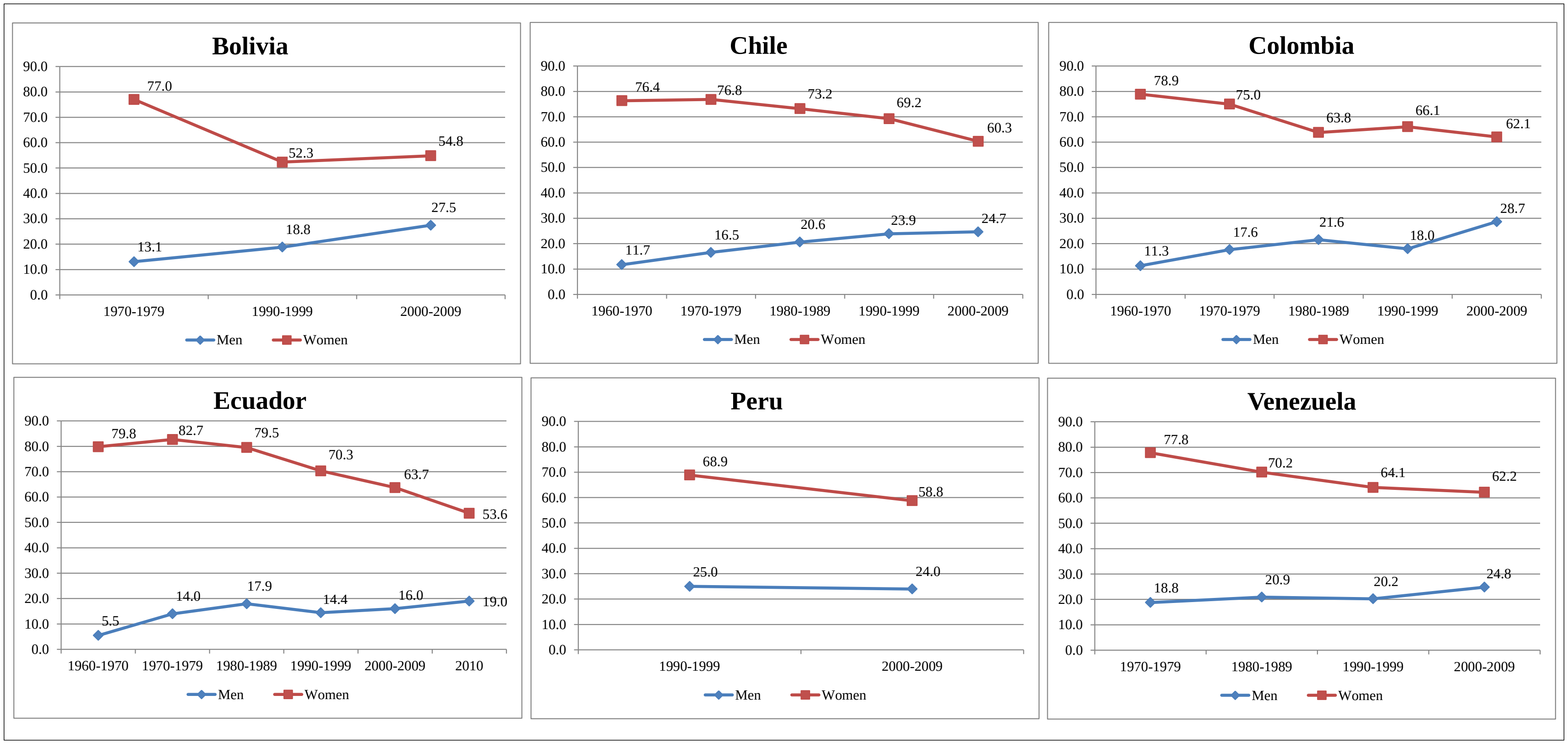
Year	Bolivia	Chile	Colombia	Ecuador	Peru	Venezuela
1925	NA	8.6	NR	NA	8.6	NR
1930	NA	8.6	NR	NA	8.6	NR
1935	NA	12	NR	NA	8.6	NR
1940	8.6	12	8	NA	8.6	12
1945	8.6	12	8	NA	8.6	12
1950	8.6	12	8	NA	8.6	12
1955	8.6	12	8	NA	8.6	12
1960	8.6	12	8	NA	8.6	12
1965	8.6	12	8	6	8.6	12
1970	8.6	12	8	6	8.6	12
1975	12.9	18	8	6	8.6	12
1980	12.9	18	8	8	8.6	12
1985	12.9	18	8	8	8.6	12
1990	12.9	18	12	8	8.6	18
1995	12.9	18	12	8	NR	18
2000	12.9	18	12	12	12.9	18
2005	12.9	18	12	12	12.9	18
2010	12.9	18	12	12	12.9	18
2012	12.9	18	14	12	12.9	26

Source: National legislation

NR = Not regulated, NA = Data not available on maternity regulations

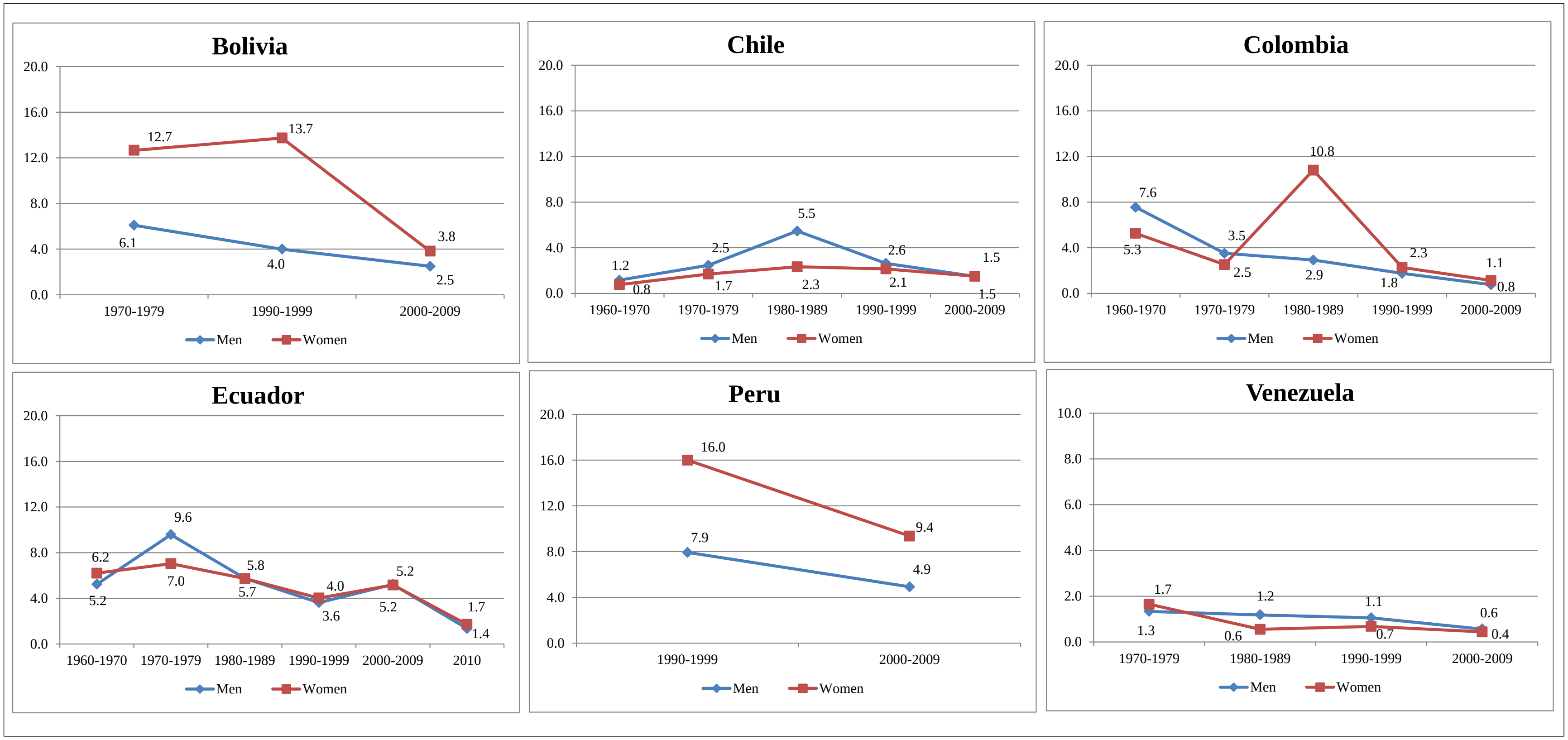
1. The duration of maternity leave obtained from national legislation was compared to the following secondary sources: the ILO TRAVAIL - Working Conditions Laws Database (retrieved from: http://www.ilo.org/dyn/travail/travmain.home), the ILO Social Security Database (retrieved from: http://www.ilo.org/dyn/sesame/IFPSES.SocialDatabase), and the World Bank Cross Country Data (retrieved from: https://www.quandl.com/WORLDBANK). Leave duration included in this table is consistent with the secondary sources examined.

Figure 2.A1: Inactive Population (%) Ages 15-64 for Selected Countries, by Gender (1960-2010)



Data source: Integrated Public Use Microdata Series (IPUMS) International.

Figure 2.A2: Unpaid Workers (%) for Employed Population Ages 15-64 for Selected Countries, by Gender (1960-2010)



Data source: Integrated Public Use Microdata Series (IPUMS) International.

Table 2.A3: Sector (%) for Employed Population Ages 15-64 for Selected Countries, Only Wage/Salary Employment (1960-2010) [1/]

Gender	Men						Women					
Data year	**1960-1970**	**1970-1979**	**1980-1989**	**1990-1999**	**2000-2009**	**2010**	**1960-1970**	**1970-1979**	**1980-1989**	**1990-1999**	**2000-2009**	**2010**
Bolivia												
Agriculture & Mining	na	30.3	na	19.4	15.6	na	na	4.4	na	4.1	3.1	na
Manft., Utilities, & Const.	na	25.8	na	32.2	32.2	na	na	6.6	na	8.8	9.2	na
Trade	na	1.6	na	8.2	11.3	na	na	3.2	na	6.4	11.1	na
Services	na	42.3	na	40.2	40.9	na	na	85.8	na	80.7	76.6	na
Obs.	*na*	*41,113*	*na*	*49,388*	*67,251*	*na*	*na*	*12,404*	*na*	*21,790*	*37,474*	*na*
Chile												
Agriculture & Mining	39.7	29.7	26.0	22.5	16.0	na	3.2	3.7	2.0	3.8	4.0	na
Manft., Utilities, & Const.	27.8	32.9	26.6	32.7	28.4	na	16.8	25.5	9.7	14.3	8.7	na
Trade	6.9	9.0	9.5	12.3	17.4	na	5.4	12.5	9.5	12.9	14.0	na
Services	25.6	28.5	37.9	32.6	38.2	na	74.5	58.4	78.9	69.0	73.2	na
Obs.	*14,482*	*131,556*	*139,329*	*206,667*	*244,809*	*na*	*4,459*	*28,376*	*64,029*	*99,020*	*144,958*	*na*
Colombia												
Agriculture & Mining	46.9	43.0	na	29.8	36.2	na	5.4	4.6	na	4.1	6.7	na
Manft., Utilities, & Const.	25.7	26.3	na	26.5	16.9	na	13.8	16.9	na	17.3	9.2	na
Trade	4.2	7.4	na	15.5	8.3	na	6.4	10.1	na	17.6	6.9	na
Services	23.2	23.3	na	28.3	38.6	na	74.4	68.4	na	61.0	77.3	na
Obs.	*39,601*	*219,972*	*na*	*331,170*	*297,299*	*na*	*13,051*	*83,812*	*na*	*174,075*	*119,695*	*na*
Ecuador												
Agriculture & Mining	56.7	na	26.5	26.9	24.4	25.7	16.7	na	3.5	6.0	7.4	7.6
Manft., Utilities, & Const.	19.6	na	29.2	23.8	25.3	25.8	10.6	na	15.3	14.4	13.0	11.4
Trade	1.8	na	6.2	7.1	13.6	13.8	3.9	na	10.2	10.8	14.2	16.2
Services	21.8	na	38.1	42.3	36.7	34.8	68.9	na	71.0	68.8	65.5	64.8
Obs.	*14,544*	*na*	*76,466*	*85,588*	*113,604*	*215,238*	*3,567*	*na*	*24,959*	*34,677*	*56,639*	*114,494*
Peru												
Agriculture & Mining	na	na	na	23.9	18.4	na	na	na	na	4.7	6.0	na
Manft., Utilities, & Const.	na	na	na	24.8	23.6	na	na	na	na	11.9	8.7	na
Trade	na	na	na	11.5	9.7	na	na	na	na	10.6	12.2	na
Services	na	na	na	39.8	48.3	na	na	na	na	72.8	73.1	na
Obs.	*na*	*na*	*na*	*204,935*	*319,071*	*na*	*na*	*na*	*na*	*89,297*	*193,829*	*na*
Venezuela												
Agriculture & Mining	na	na	11.9	14.2	14.9	na	na	na	1.5	1.8	1.8	na
Manft., Utilities, & Const.	na	na	35.6	30.2	24.5	na	na	na	15.0	14.4	8.7	na
Trade	na	na	12.5	12.6	18.3	na	na	na	11.1	11.7	13.5	na
Services	na	na	39.9	43.0	42.4	na	na	na	72.4	72.2	76.0	na
Obs.	*na*	*na*	*189,184*	*219,560*	*295,895*	*na*	*na*	*na*	*91,012*	*98,748*	*195,774*	*na*

Data source: Integrated Public Use Microdata Series (IPUMS) International. na = Not available

1. Agriculture and mining include also fishing and forestry; services include hotels and restaurants, transportation and communications, financial services and insurance, public administration and defense, real state and business services, education, health and social work, private household services and others.

Table 2.A4: Occupation (%) for Employed Population Ages 15-64 for Selected Countries, Only Wage/Salary Employment (1960-2010) [1/]

Gender	Men						Women					
Data year	**1960-1970**	**1970-1979**	**1980-1989**	**1990-1999**	**2000-2009**	**2010**	**1960-1970**	**1970-1979**	**1980-1989**	**1990-1999**	**2000-2009**	**2010**
Bolivia												
Managers	na	1.2	na	3.0	3.6	na	na	0.6	na	1.8	2.9	na
Professionals	na	11.1	na	9.9	10.3	na	na	28.1	na	22.0	20.1	na
Technicians	na	0.7	na	5.6	8.4	na	na	0.5	na	6.4	7.2	na
Clerk/Services	na	14.5	na	16.2	15.3	na	na	59.0	na	25.3	30.0	na
Craft/Operators	na	62.2	na	52.7	55.3	na	na	7.9	na	9.3	8.7	na
Elementary occupations	na	10.4	na	12.6	7.2	na	na	3.9	na	35.2	31.1	na
Obs.	*na*	*38,333*	*na*	*49,613*	*67,520*	*na*	*na*	*12,261*	*na*	*22,341*	*38,889*	*na*
Chile												
Managers	1.7	0.7	1.8	1.9	1.8	na	0.8	0.5	0.8	1.2	1.1	na
Professionals	3.5	4.1	6.4	5.4	8.3	na	10.8	11.3	15.9	12.6	15.4	na
Technicians	2.1	3.4	4.3	4.7	13.5	na	4.6	12.0	16.2	7.9	19.2	na
Clerk/Services	11.7	17.4	18.6	19.8	18.8	na	25.1	56.5	51.0	36.8	31.2	na
Craft/Operators	39.7	59.9	50.8	48.5	35.1	na	14.3	15.1	7.3	10.1	5.3	na
Elementary occupations	41.3	14.5	18.1	19.8	22.6	na	44.5	4.6	8.8	31.3	27.8	na
Obs.	*14,164*	*126,148*	*138,458*	*203,142*	*223,516*	*na*	*4,471*	*41,014*	*65,312*	*100,220*	*133,370*	*na*
Colombia												
Managers	2.8	2.0	na	na	na	na	1.0	0.6	na	na	na	na
Professionals	4.0	3.9	na	na	na	na	5.9	9.5	na	na	na	na
Technicians	1.4	1.3	na	na	na	na	8.8	1.2	na	na	na	na
Clerk/Services	14.3	17.8	na	na	na	na	68.8	68.9	na	na	na	na
Craft/Operators	75.0	64.2	na	na	na	na	15.0	14.9	na	na	na	na
Elementary occupations	2.6	10.8	na	na	na	na	0.5	4.9	na	na	na	na
Obs.	*39,106*	*226,998*	*na*	*na*	*na*	*na*	*13,258*	*86,359*	*na*	*na*	*na*	*na*
Ecuador												
Managers	0.4	1.3	1.7	1.9	3.5	2.6	0.1	0.7	0.9	1.4	3.5	3.1
Professionals	3.9	5.5	9.9	11.5	8.2	7.9	13.3	17.8	24.9	24.1	17.7	19.9
Technicians	1.1	1.6	4.3	6.7	3.7	5.0	3.0	3.9	7.5	7.8	6.9	7.5
Clerk/Services	8.4	11.0	11.9	15.7	18.9	21.5	34.5	24.2	25.0	24.2	33.3	33.4
Craft/Operators	23.6	29.3	36.3	31.2	39.2	34.0	9.9	12.5	10.0	8.5	13.9	9.7
Elementary occupations	62.6	51.3	35.8	33.0	26.5	29.0	39.2	41.0	31.8	34.0	24.8	26.4
Obs.	*14,436*	*68,822*	*71,336*	*82,236*	*114,144*	*212,391*	*3,620*	*16,117*	*24,793*	*39,679*	*58,757*	*114,565*
Peru												
Managers	na	na	na	1.4	0.5	na	na	na	na	1.1	0.4	na
Professionals	na	na	na	12.7	13.3	na	na	na	na	27.6	24.3	na
Technicians	na	na	na	7.5	8.3	na	na	na	na	9.4	10.5	na
Clerk/Services	na	na	na	18.2	19.8	na	na	na	na	27.2	27.6	na
Craft/Operators	na	na	na	29.8	26.7	na	na	na	na	7.2	5.7	na
Elementary occupations	na	na	na	30.5	31.4	na	na	na	na	27.6	31.5	na
Obs.	*na*	*na*	*na*	*208,549*	*306,302*	*na*	*na*	*na*	*na*	*91,566*	*187,807*	*na*
Venezuela												
Managers	na	na	4.4	4.1	7.1	na	na	na	1.4	2.2	6.5	na
Professionals	na	na	5.9	7.0	5.2	na	na	na	17.3	19.6	14.7	na
Technicians	na	na	13.2	14.1	10.3	na	na	na	14.0	16.3	15.6	na
Clerk/Services	na	na	18.8	17.1	15.6	na	na	na	35.3	32.3	34.4	na
Craft/Operators	na	na	38.9	37.7	36.9	na	na	na	6.8	7.0	3.9	na
Elementary occupations	na	na	18.9	20.0	24.9	na	na	na	25.2	22.7	25.0	na
Obs.	*na*	*na*	*170,612*	*194,860*	*290,612*	*na*	*na*	*na*	*88,037*	*94,597*	*197,124*	*na*

Data source: Integrated Public Use Microdata Series (IPUMS) International. na = Not available

1. Managers include legislators and senior officials; technicians include associates; craft and operators include skilled agricultural workers and plant and machine operators.

Table 2.A5: Duncan Index of Dissimilarity (Sector or Industry) for Employed Population Ages 15-64 for Selected Countries [1/]

	1960-1970	1970-1979	1980-1989	1990-1999	2000-2009	2010
Bolivia	na	0.635	na	0.536	0.484	na
Chile	0.615	0.443	0.533	0.458	0.435	na
Colombia	0.638	0.533	na	0.439	0.514	na
Ecuador	0.570	na	0.509	0.431	0.399	0.413
Peru	na	na	na	0.414	0.376	na
Venezuela	na	na	0.432	0.466	0.439	na

Data source: Integrated Public Use Microdata Series (IPUMS) International.

1. Based on major categories from the International Standard Industrial Classification of economic activities (ISIC).

Table 2.A6: Duncan Index of Dissimilarity (Occupation) for Employed Population Ages 15-64 for Selected Countries [1/]

	1960-1970	1970-1979	1980-1989	1990-1999	2000-2009	2010
Bolivia	na	0.623	na	0.439	0.486	na
Chile	0.283	0.520	0.539	0.404	0.315	na
Colombia	0.634	0.575	na	na	na	na
Ecuador	0.353	0.282	0.308	0.275	0.273	0.272
Peru	na	na	na	0.276	0.215	na
Venezuela	na	na	0.380	0.364	0.344	na

Data source: Integrated Public Use Microdata Series (IPUMS) International.

na = Not available

1. Based on major categories from the International Standard Classification of Occupations (ISCO) 1988.

Table 2.A7: Number of Children Born per Woman (average) by Age

Age	Bolivia 1976	Bolivia 1992	Bolivia 2001	Chile 1960	Chile 1970	Chile 1982	Chile 1992	Chile 2002	Colombia 1973	Colombia 1985	Colombia 1993
15	0.02	0.03	0.04	0.01	0.22	0.03	0.05	0.20	0.03	0.02	0.03
16	0.05	0.09	0.09	0.05	0.24	0.06	0.09	0.08	0.07	0.06	0.14
17	0.13	0.17	0.20	0.12	0.30	0.13	0.15	0.14	0.19	0.14	0.27
18	0.26	0.35	0.35	0.19	0.40	0.23	0.24	0.22	0.35	0.27	0.43
19	0.46	0.55	0.51	0.36	0.52	0.38	0.38	0.33	0.57	0.42	0.62
20	0.73	0.85	0.74	0.56	1.16	0.53	0.48	0.43	0.87	0.69	0.82
21	0.96	0.99	0.91	0.64	1.26	0.68	0.62	0.54	1.03	0.83	0.95
22	1.22	1.30	1.11	0.91	1.43	0.86	0.76	0.66	1.37	1.04	1.12
23	1.50	1.52	1.33	1.20	1.61	1.03	0.92	0.76	1.67	1.26	1.31
24	1.77	1.75	1.55	1.45	1.77	1.17	1.05	0.85	1.94	1.44	1.45
25	2.05	1.99	1.78	1.69	2.05	1.34	1.17	0.96	2.30	1.65	1.56
26	2.38	2.23	1.99	1.94	2.18	1.57	1.32	1.07	2.52	1.83	1.70
27	2.66	2.44	2.16	2.09	2.46	1.69	1.44	1.21	2.84	2.03	1.86
28	2.93	2.68	2.46	2.26	2.62	1.87	1.58	1.34	3.11	2.17	1.99
29	3.29	2.91	2.64	2.32	2.99	2.03	1.72	1.45	3.39	2.40	2.11
30	3.56	3.09	2.91	2.43	3.35	2.17	1.87	1.58	3.78	2.57	2.29
31	3.97	3.31	3.07	2.97	3.61	2.36	1.99	1.73	3.98	2.71	2.31
32	4.10	3.50	3.24	3.19	3.70	2.55	2.14	1.87	4.32	2.89	2.56
33	4.35	3.58	3.53	3.39	4.01	2.63	2.25	1.95	4.63	3.08	2.71
34	4.61	3.72	3.71	3.14	4.14	2.75	2.33	2.05	4.84	3.26	2.78
35	4.81	3.96	3.91	3.43	4.11	2.82	2.40	2.11	4.97	3.38	2.91
36	5.17	4.24	4.12	3.85	4.33	3.06	2.53	2.16	5.32	3.66	3.03
37	5.27	4.38	4.27	3.62	4.58	3.15	2.64	2.27	5.58	3.84	3.20
38	5.50	4.47	4.48	3.67	4.68	3.27	2.69	2.34	5.75	3.98	3.32
39	5.70	4.47	4.61	3.92	4.79	3.44	2.78	2.43	5.91	4.27	3.38
40	5.58	4.68	4.73	3.62	4.87	3.48	2.82	2.47	5.88	4.33	3.58
41	5.89	4.69	4.86	4.15	5.19	3.68	2.87	2.52	6.12	4.61	3.52
42	5.94	4.83	4.99	3.78	5.03	3.84	2.95	2.60	6.32	4.65	3.74
43	6.12	4.86	5.14	3.98	5.36	4.04	3.08	2.65	6.47	5.07	3.90
44	6.23	5.05	5.28	4.09	5.40	4.14	3.10	2.65	6.72	5.26	4.01
45	5.97	4.93	5.32	3.76	5.06	3.98	3.05	2.66	6.33	5.23	4.16
46	6.09	5.18	5.36	3.97	5.38	4.23	3.19	2.72	6.62	5.45	4.25
47	6.35	5.19	5.36	4.02	5.31	4.27	3.35	2.77	6.56	5.65	4.40
48	6.30	5.08	5.44	3.57	5.19	4.35	3.40	2.80	6.60	5.68	4.64
49	6.29	5.31	5.60	3.77	5.32	4.48	3.52	2.88	6.63	5.85	4.65

Data source: Integrated Public Use Microdata Series (IPUMS) International.

Table 2.A7 (continued): Number of Children Born per Woman (average) by Age

Age	Colombia 2005	Ecuador 1974	Ecuador 1982	Ecuador 1990	Ecuador 2001	Ecuador 2010	Peru 1993	Peru 2007	Venezuela 1971	Venezuela 1990	Venezuela 2001
15	0.04	0.04	0.10	0.18	0.04	0.04	0.04	0.02	0.06	0.08	0.04
16	0.10	0.08	0.17	0.07	0.09	0.10	0.07	0.06	0.12	0.15	0.09
17	0.19	0.17	0.30	0.16	0.18	0.19	0.13	0.12	0.23	0.27	0.18
18	0.32	0.36	0.50	0.28	0.31	0.31	0.23	0.20	0.35	0.42	0.30
19	0.46	0.56	0.74	0.44	0.46	0.44	0.34	0.30	0.53	0.60	0.44
20	0.63	0.92	1.01	0.68	0.65	0.58	0.54	0.43	0.76	0.80	0.60
21	0.77	1.05	1.20	0.80	0.83	0.73	0.67	0.54	0.92	1.00	0.75
22	0.94	1.43	1.54	1.03	0.97	0.89	0.91	0.92	1.19	1.20	0.91
23	1.07	1.68	1.73	1.23	1.17	1.04	1.13	1.06	1.43	1.40	1.07
24	1.21	2.05	1.96	1.45	1.33	1.18	1.29	1.19	1.74	1.60	1.25
25	1.35	2.43	2.24	1.65	1.48	1.33	1.52	1.29	1.99	1.79	1.38
26	1.49	2.68	2.47	1.86	1.62	1.49	1.70	1.41	2.23	1.99	1.54
27	1.66	2.98	2.72	2.07	1.79	1.64	1.95	1.56	2.46	2.19	1.68
28	1.80	3.33	2.98	2.29	1.96	1.77	2.15	1.67	2.70	2.37	1.83
29	1.92	3.54	3.11	2.53	2.04	1.88	2.29	1.79	2.95	2.55	1.96
30	2.08	4.05	3.50	2.79	2.32	2.00	2.58	1.98	3.22	2.80	2.11
31	2.18	3.99	3.55	2.92	2.40	2.13	2.66	2.03	3.31	2.96	2.28
32	2.30	4.48	3.96	3.13	2.61	2.24	2.94	2.19	3.52	3.12	2.38
33	2.40	4.62	4.12	3.36	2.66	2.37	3.18	2.30	3.75	3.35	2.53
34	2.47	4.98	4.33	3.52	2.82	2.45	3.23	2.45	3.87	3.43	2.62
35	2.57	5.28	4.60	3.67	2.91	2.52	3.50	2.54	3.93	3.61	2.69
36	2.64	5.50	4.93	3.88	3.08	2.67	3.64	2.72	4.03	3.79	2.82
37	2.70	5.88	5.11	4.08	3.16	2.78	3.85	2.83	4.10	3.89	2.95
38	2.77	5.87	5.33	4.35	3.33	2.82	4.15	2.93	4.19	4.03	3.05
39	2.87	6.19	5.64	4.40	3.38	2.91	4.23	3.09	4.18	4.22	3.08
40	2.93	6.26	5.74	4.59	3.55	2.97	4.51	3.22	4.12	4.46	3.23
41	2.97	6.39	5.91	4.65	3.61	3.06	4.47	3.26	4.20	4.49	3.30
42	3.08	6.46	6.06	4.84	3.73	3.08	4.67	3.37	4.24	4.68	3.40
43	3.16	6.53	6.40	5.08	3.85	3.17	4.83	3.46	4.25	4.81	3.51
44	3.20	6.78	6.56	5.25	3.95	3.26	4.90	3.49	4.22	4.96	3.55
45	3.28	6.57	6.39	5.30	4.01	3.28	5.10	3.69	4.06	5.08	3.56
46	3.34	6.76	6.60	5.50	4.12	3.33	5.22	3.77	4.03	5.29	3.66
47	3.39	6.72	6.71	5.63	4.24	3.43	5.23	3.81	4.02	5.36	3.71
48	3.47	6.55	6.59	5.63	4.27	3.50	5.56	3.96	4.02	5.49	3.84
49	3.53	6.74	6.74	5.83	4.44	3.56	5.51	3.99	4.09	5.61	3.88

Data source: Integrated Public Use Microdata Series (IPUMS) International.

Chapter 3: The Impact of Maternity Leave on Women's Fertility in Latin America

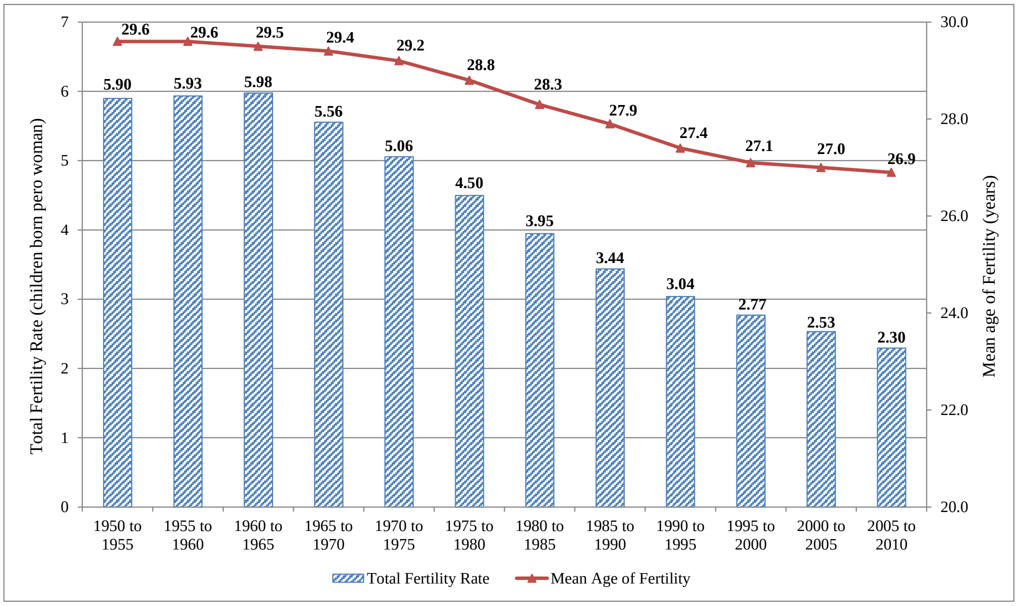
3.1. Introduction

Latin America has shown a marked trend of declining fertility over the last 60 years. This pattern is relevant for public policies due to its impact on the dependency ratio (the number of dependents divided by the working-age-population), which has implications for health and pension systems. This phenomenon coincides within a context of women becoming increasingly more economically active and acquiring further education. According to a report by the Economic Commission for Latin America and the Caribbean (2011), fertility began to decline considerably in the region by the mid-twentieth century, as Latin American countries continued to develop, and this brought further changes in reproductive behavior. The total fertility rate (Figure 1) declined in Latin America from around six children born per woman in the 1950s and 1960s to a number very close to the reproduction level after 2005, while the mean age at fertility decreased by about three years during that period.

In a context of declining fertility, maternity leave could be viewed as a possible tool for policy makers. Even though these regulations were not originally developed with explicit objectives with respect to fertility, the literature on this area has identified effects of maternity benefits on the number or timing of births. Maternity leave helps women to reconcile work and family responsibilities, providing time off during pregnancy and after childbirth, replacement of labor income lost during that time, and increased job security.

These features of maternity leave lower the costs of having children and thus are hypothesized to increase fertility.

Figure 3.1: Selected Fertility Indicators, Latin America (1950-2010) [1/]



Source: Economic Commission for Latin America and the Caribbean (ECLAC).

1. The total fertility rate (a synthetic measure) is the average number of children that would be born to a woman during her lifetime based on the current age-specific fertility rates; the mean age of fertility is the average age of women at the birth of their children (i.e. age at birth weighted by current age-specific fertility rates).

In this chapter, I explore the effects of changes in the duration of maternity leave on selected fertility outcomes. This question has not been extensively explored in the relevant literature and, to the best of my knowledge, no prior evidence has been provided for Latin America. For this purpose, I use data from six countries in Latin America (Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela) and apply a pseudo-panel method to estimate the effects of these regulations on the probability that a woman has any children and on the probabilities of higher order children from 1960 to 2011. This method allows me to take advantage of existing cross-sectional data by creating synthetic longitudinal data to assess the effects of these policies on reproductive behavior. These countries were selected due to the availability of data covering a long-term span (mostly microdata from the Integrated Public Use Microdata Series - International project) and sources to track changes in the duration of maternity leave over time. For the analysis, I tracked all changes in the duration of maternity leave, starting in the 1960s and based on the relevant labor laws for each country. For every country analyzed, maternity leave is paid and replacement rates are 100 percent or close to a fully paid salary.

The chapter is organized as follows. In the next section, I discuss the expected consequences of maternity leave and review the empirical literature on its fertility effects. In Section 3.3, I describe the data and the empirical strategy. The summary of changes in the duration of maternity leave in Latin America that are used in the analysis were described in the previous Chapter. Next, in Section 3.4 I present data on some general fertility trends in Latin America, and the estimated effects of changes in maternity leave regulations on the probability that a woman has any number of children and on the probabilities of having higher order children. Finally, a discussion of the results is presented in Section 3.5.

3.2. Conceptual Framework

3.2.1. What Are the Expected Effects of Maternity Leave?

Household decisions about fertility depend on individual preferences and on child production costs (Willis, 1973; Becker, 1981). Two main cost categories have been identified in the literature in this topic: the direct costs of raising a child (including the purchase of goods such as food or clothing plus childcare expenditures) and foregone labor market earnings of the mother or other care provider (Ermisch, 1988; Walker, 1995; Ronsen, 2004; Björklund, 2007). The latter category includes both the current loss of labor market earnings (i.e. work time sacrificed for childcare activities) and the potential impact on future earnings given the loss of human capital (work experience) associated to interruptions in labor force participation.

The hypothesis is that policies reducing child production costs will increase fertility (Hyatt and Milne, 1991; Gauthier and Hatzius, 1997; Averett and Whittington, 2001; Björklund, 2006). Two components of maternity leave could affect these production costs: the duration of leave and the subsidy paid during leave. Longer leave duration implies larger losses of labor market income and human capital, but also more time off work to recover after a birth and to take care of the newborn, accompanied by job security (Gauthier and Hatzius, 1997).[12] The monetary benefits paid during leave cover foregone labor market earnings, lowering the opportunity cost of childbearing. Therefore, more generous leave benefits (in terms of duration or pay) are expected to increase fertility. Maternity leave may also create incentives for women out of the labor market to enter the workforce and could reduce fertility for them (Zhang, Quan, and Van Meerbergen, 1994). This last effect may attenuate the expected increase in fertility, and thus the overall impact of maternity leave on fertility needs to be tested empirically.

[12] On this issue, the authors argue that even though a longer leave implies a higher opportunity cost, the other option would be to quit their job. Therefore, longer leave duration should increase fertility if "women deciding to have a child are willing to forgo earnings" (Gauthier and Hatzius, 1997, p. 296).

Different aspects of fertility decisions may be influenced by changes in these benefits. The effect of maternity leave could be more important for higher order than first births. Averett and Whittington (2001) argue that most women experience at least one birth (in a study for the United States), such that the reduction in production costs of children could have a stronger effect on second or further births. Gauthier and Hatzius (1997) suggest that the cost of higher order births may be even lower than the first one if we assume economies of scale in chil production. Following this argument, the cost reduction due to maternity leave (a "fixed" benefit) for a second or a third child is more important than for the first one in relative terms. Furthermore, even if the total number of children at the end of a woman's reproductive years remains unchanged, the timing of births may be affected through specific features in the design of maternity leave. Previous research on this topic identified some of these features. Women are often required to meet minimum work requirements (or social security contributions) to be eligible for leave or to receive larger leave benefits. Because of this, maternity leave could delay births for young workers until they are able to meet these requirements (Averett and Whittington, 2001; Björklund, 2006). In some countries the leave renewal is regulated (such as Austria or Sweden) and workers are allowed to take leave again if there is an additional birth within certain time after the end of a previous leave. The literature identifies a "speed premium" incentive if women decide to shorten the spacing for the next child to be able to renew and continue on maternity leave (Hoem, 1993; Ronsen, 2004; Björklund, 2006; Lalive and Zweimüller, 2009).

Certain characteristics of women or their families may create differential effects of maternity leave on fertility. Household socioeconomic status and the mother's occupation or career orientation are some of the factors highlighted in the literature. The impact of maternity leave on fertility is expected to be higher for low-earning parents if the wage replacement rate is regulated as a fixed amount or if it is subject to ceilings based on income (Phipps, 2000; Lalive and Zweimüller, 2009). The reason for this effect is that the leave subsidy would represent, in those cases, a larger proportion of foregone labor market income. But in a broader sense, the leave subsidy component is assumed to create higher incentives among low-earning parents since it allows them to avoid financial distress (Lalive and Zweimüller, 2009). Job protection provides women a guarantee that they can return to their previous job after the end of their leave. This feature of maternity leave is valued more by individuals with stronger career orientation or working in skilled occupations, for whom losing their jobs is more costly (Lalive and Zweimüller, 2009). Thus, fertility effects may be larger for these women who value relatively more the job protection offered by maternity leave.

3.2.2. Empirical Evidence

The fertility effects of parental leave policies in OECD countries have been examined by several papers (see Table 3.A1 in the Appendix to this chapter). A few studies analyze cross-country aggregate data, but most of the evidence corresponds to country-specific studies. Winegarden and Bracy (1995) estimated the effects of paid maternity leave on the general fertility rate of women between 15 and 44 years old in 17 OECD countries. Their results suggest that maternity leave did not have an impact on fertility[13]. Gauthier and Hatzius (1997) examine the effects of cash and maternity benefits using time series of aggregate data from a wider set of 22 industrialized countries. Even though the authors find that cash benefits from family allowances have a positive effect on fertility, the duration of maternity leave and its wage replacement rates do not have a statistically significant effect.

In the United States, state legislation on maternity leave was passed between the late eighties and early nineties in a few states and the Family and Medical Leave Act (FMLA) was created in 1993, which is federal legislation mandating maternity leave (Klerman and Leibowitz, 1997; Waldfogel, 1999; Baum, 2003). Before the enactment of this legislation, leave was essentially voluntarily offered by employers (Baker and Milligan, 2008a). Averett and Whittington (2001) focus on the period before the enactment of the FMLA to determine if there is any evidence of women sorting into firms offering maternity leave and whether the availability of this benefit influences the probability of a birth. Their data do not support the hypothesis that women choose jobs based on their leave policies. However, maternity leave increases the probability of a birth for women who already have at least one child (by about 3 to 7 percentage points), while it does not seem to have effects on first births. Cannonier (2014) analyzes the impact of the FMLA on both the probability of a birth and the spacing between births. The author uses a difference-in-differences strategy by comparing births for women eligible and ineligible to the FMLA. His results show a positive effect on the probability of a birth (1.5 percentage points more for the first child and 0.6 percentage points for the

[13] Winegarden and Bracy (1995) estimated a system of equations for fertility, labor force participation, and infant mortality. The reduced form estimates showed no net impact of maternity leave on births, although there is a positive impact based on the structural estimates. In the former set of results, the effect of maternity leave on fertility was offset by its positive effect on labor force participation and its negative effect on infant mortality, both which lower fertility.

second), while eligible women give birth earlier (about 12 months for the first child and 8.5 months for the second).

Canada expanded job-protected leave through provincial and federal regulations from zero to 12 weeks in the 1960s and to 52 or more weeks by the end of 2000 (Ten Cate, 2003; Baker and Milligan, 2008a; Hanratty and Trzcinski, 2009). Maternity leave benefits are paid through the unemployment insurance since 1971 (Ten Cate, 2003; Baker and Milligan, 2008a). Two papers work with country level data to examine the effect of paid maternity leave on fertility following the introduction of the unemployment insurance act. Hyatt and Milne (1991) conclude that benefits paid during maternity leave have a positive and statistically significant effect on fertility, translating into a small increase between 0.09 and 0.26 per cent in the total fertility rate (for a one percent increase in the real value of maternity benefits). Zhang, Quan, and Van Meerbergen (1994) analyze fertility effects of paid maternity leave in the context of other family policies in Canada, including the personal tax exemption for children, child tax credit, and family allowances. Even though they find positive and statistically significant effects of the three tax-transfer programs on fertility, paid maternity benefits do not have a statistically significant effect on fertility. Phipps (2000) uses longitudinal microdata to determine whether the availability of paid leave benefits has behavioral implications. The evidence suggests that women do not adjust their labor market supply in order to become eligible for maternity benefits. Furthermore, the author concludes that paid maternity benefits do not affect the probability of having a birth, which she argues could be explained by the relative unimportance of the monetary maternity leave benefits with respect to the overall costs of raising a child.

Most of the evidence on fertility effects of maternity leave comes from studies on the Nordic countries. Paid maternity leave was introduced in Sweden, Norway, and Finland between the 1950s and 1960s (Hoem, 1993; Ronsen, 2004; Duvander and Andersson, 2006). In all three countries, leave would be later modified to include men as beneficiaries (i.e. parental leave), thus allowing parents to have flexibility for either of them (or both) to use the paid leave period. An important policy change occurred in the 1990s in Norway and Sweden, when part of the parental leave (that could be shared by both parents) was required to be taken exclusively by each parent (Duvander and Andersson, 2006; Duvander, Lappegard, and Andersson, 2010). For Sweden, another relevant aspect in the design of the parental leave scheme was created in the 1970s, which allowed a parent to retain the compensation level from one birth to the next one (without demonstrating further work-related income) during certain limited period of time (Hoem, 1993; Björklund, 2006).

Several studies examine parental leave policies in these countries. For Sweden, Hoem (1993) reports evidence that parents increased their fertility before the end of the eligibility period to retain the compensation level from the previous birth, which is identified as a "speed premium" that creates incentives to reduce the time interval between births. Björklund (2006, 2007) describes the evolution of the parental leave scheme in Sweden along with other relevant changes in family policies, such as the provision of child care and child allowances. Based on the comparison of fertility indicators for Sweden against other European countries, he concludes that the overall policy changes contributed to a higher fertility level and a shorter interval between births. Ronsen (2004) explores the effects of the duration of parental leave in Norway and

Finland on the hazard of first, second, and third births. Results show that increases in the duration of parental leave have a positive effect on birth rates, particularly for third births and in Finland.

Three studies analyze fertility outcomes in Sweden and Norway based on longitudinal data from linked administrative registers, which include information on births and parental leave use. The policy variables of interest for these papers are the parental leave uptake and how leave is allocated between mother and father. Even though these studies have potential endogeneity issues acknowledged by the authors, they still present some evidence on decisions on the use of parental leave and fertility. Duvander and Andersson (2006) find a positive correlation between the father's uptake of leave and the probability of a second or third birth in Sweden, while a very high or very low mother's leave uptake is negatively correlated with the probability of a second birth. Duvander, Lappegard, and Andersson (2010) show that the father's uptake of leave is positively correlated with the probability of a second or a third birth in both countries, while a long leave period for the mother is positively correlated with a third birth (which they argue could be related to the mother's weaker work orientation). Lappegard (2010) analyzes parental leave along with other related policies in Norway, including the provision of childcare and childcare cash benefits. Results from this study also suggest that parental leave has a positive correlation with second births (particularly when leave is shared with the father), but a negative correlation with third births (that the author speculates could be related to couples with more traditional family roles and higher fertility).

Finally, Lalive and Zweimüller (2009) analyze the effects of changes in the duration of parental leave in Austria, which was extended from 1 to 2 years in 1990 (for births after July 1st) and then reduced by 6 months in 1996. Before and after these reforms, parents were exempted from the work requirement to renew parental leave for a subsequent birth happening within 3.5 months from a previous leave granted. Lalive and Zweimüller (2009) argue that the initial grace period to renew parental leave (12 plus 3.5 months) was relatively short for parents to have a new birth. However, the leave duration increase in 1990 made a new birth (biologically) more feasible within the longer grace period for leave renewal (24 plus 3.5 months). The authors compare women who had a first born just before and after the 1990 reform, to determine the effects of the extended leave for the current birth and the longer grace period for leave renewal. Their results show that the proportion of women with second births is higher for post-reform than pre-reform mothers (in about 5 percentage points). Furthermore, they compare women who had a first birth just before the 1990 reform (but who would be eligible for the new leave duration for additional births) to those who had it three years before that, to assess the effect of the extended leave available for future births. Similarly, they find evidence of (strong) positive effects on second births (in about 7 percentage points). These fertility effects are persistent over time, as they are observed even ten years after the first birth. The 1996 reform basically reduced the time interval between births but it did not have effects on the number of second births.

Overall, empirical studies on maternity leave reveal some effects on fertility. The cross-country studies available do not find evidence of fertility effects of maternity leave (Winegarden and Bracy, 1995; Gauthier and Hatzius, 1997), similar to the small or not statistically significant effects reported for Canada (Hyatt and Milne, 1991; Zhang, Quan, and Van Meerbergen, 1994; Phipps, 2000). Most of the evidence presented by the literature suggests positive effects of parental leave on higher order births (Averett and Whittington, 2001; Ronsen, 2004; Duvander and Andersson, 2006; Lalive and Zweimüller, 2009; Duvander, Lappegard, and Andersson, 2010; Lappegard, 2010; Cannonier, 2014), while a few studies find also positive effects on first births (Cannonier, 2014) or in the (shorter) spacing between births (Hoem, 1993; Lalive and Zweimüller, 2009; Cannonier, 2014). In the specific case of the Nordic countries, the allocation of leave between parents also shows that the father's uptake is positively correlated with higher order births (Duvander and Andersson, 2006; Duvander, Lappegard, and Andersson, 2010; Lappegard, 2010). For Latin America, there are no studies focused on fertility effects of leave.

3.3. Methodology

3.3.1. Data

In this paper, I use data from six countries in Latin America. These six countries were selected due to the availability of microdata covering a long span of time that includes years both before and after legislation changes. Data from the Minnesota Population Center include a total of 41 censuses available from the region, which for some countries cover the entire period between the 1960's and 2010. This seems to be a sufficient time span to observe the impacts of changes in labor regulations on maternity leave. Table 3.1 shows all the data years available from the Integrated Public Use

Microdata Series (IPUMS) International that will be used for the analysis. The datasets are 10 percent samples of the corresponding censuses, except for Chile 1960 (1 percent), Therefore, given the large proportion of the population included in the census samples, all the data are nationally representative and also for lower geographical units.

In addition, the census data are complemented by nationally representative household surveys, particularly for those countries with relatively few censuses available. In the case of Peru, for example, an additional data point was obtained by using the Living Standards Measurement Study (LSMS) survey for 1985-86. The household surveys that are used in this paper are also included in Table 3.1 and all are nationally representative.

Table 3.1: Availability of Census Microdata and Household Surveys for Selected Latin American Countries (Year of Data Collection)

Census round	1960-1969	1970-1979	1980-1989	1990-1999	2000-2009	2010
Bolivia	NA	1976	NA	1992	2001	EH 2009 *
Chile	1960	1970	1982	1992	2002	CASEN 2011 *
Colombia	NA	1973	1985	1993	2005	NA
Ecuador	NA	1974	1982	1990	2001	2010
Peru	NA	NA	LSMS 1985-86 *	1993	2007	NA
Venezuela	NA	1971	NA	1990	2001	NA

Source: Census microdata from the Integrated Public Use Microdata Series (IPUMS) International; the 1985-86 Living Standards Measurement Study (LSMS) from the National Institute of Statistics and Informatics (INEI) - Peru; the 2009 Household Survey (EH) from the National Institute of Statistics (INE) - Bolivia; and the National Socioeconomic Characterization Survey (CASEN) from the Ministry of Social Development - Chile.
* = Household Survey. This includes the 1985-86 LSMS for Peru, the 2009 EH for Bolivia, and the 2011 CASEN for Chile.
NA=Not available

The outcomes to be analyzed are based on the number of children ever born reported by women of reproductive age. Three outcomes are defined based on this variable: having one or more, two or more, or three or more children (shown in equations 3.1a, 3.1b, and 3.1c). This definition allows one to distinguish between the decision of having children from that of having multiple children, as potential differential effects of maternity leave on higher order fertility have been identified in the literature (Averett and Whittington, 2001; Ronsen, 2004; Duvander and Andersson, 2006; Lalive and Zweimüller, 2009; Duvander, Lappegard, and Andersson, 2010; Lappegard, 2010; Cannonier, 2014). Furthermore, this distinction is relevant given that the most significant changes in fertility over the last decades in Latin America refer to higher order children (as will be discussed in Section 3.5).

$$\text{Children } 1+ = \mathrm{I}\left[\text{Children ever born} \geq 1\right] \ldots \text{ (3.1a)}$$

$$\text{Children } 2+ = \mathrm{I}\left[\text{Children ever born} \geq 2\right] \ldots \text{ (3.1b)}$$

$$\text{Children } 3+ = \mathrm{I}\left[\text{Children ever born} \geq 3\right] \ldots \text{ (3.1c)}$$

Information on children ever born is an indicator of lifetime fertility that is assessed retrospectively at the time of data collection, and it excludes stillbirths, miscarriages, and abortions (United Nations, 1983; United Nations, 2010). Therefore, this measure may be affected by recall bias. Some common problems in data collection are the under-report of children living elsewhere or those who had died, which is considered to be relatively more important for women in their late reproductive years (ages 35 or more); in addition, but to a lesser extent, there could be over-report due to the inclusion of stillbirths, fetal deaths, or adopted children (United Nations, 1983; United Nations, 1989). Because of these reporting issues, recommendations for census enumeration suggest including detailed fertility questions to assess data quality and to reduce recall bias (United Nations, 1983; United Nations, 2008; United Nations, 2010). That is, questionnaires should comprise not only the total number of children ever born (as a single question), but also the breakdown by gender and place of residence (living at home or elsewhere), and the number of children still alive and who have died. The examination of census (and household survey) questionnaires reveals that most data sources included follow up questions intended to reduce recall bias (such as the number of surviving children or children living elsewhere) and many of them also had a breakdown of children ever born by gender.[14] Thus, the design of the questionnaires should have attenuated reporting errors of children ever born, but some under-report could be still present in the data. Further investigation on this issue is required, but it seems reasonable to expect that under-reporting will at least not bias the estimated effects of maternity leave regulations, because questionnaires for the same country have been relatively consistent over time and no major differences in the questions used for data collection have been identified across countries.[15]

The most important explanatory variable corresponds to the duration of maternity leave (in weeks) set by regulation for the selected countries. As previously discussed, this

[14] In the data used for this essay only the Chile 1960 and Venezuela 1971 censuses had a single question inquiring about the number of children ever born. Other data sources included at least one follow up question, which most often referred to the number of surviving children. Furthermore, six out of the 26 datasets had the breakdown of children ever born by gender.

[15] For example, bias in the estimates may exist if improvements in data collection over time reduced significantly under-report of children ever born simultaneously to increases in the duration of maternity leave. Nevertheless, besides the Chile 1960 and Venezuela 1971 censuses, other data sources included comparable questions over time and across countries. In the case of Colombia, for instance, all four censuses asked about surviving children and the timing of the last birth, and some breakdown of questions by gender was considered in the 1985, 1993, and 2005 censuses.

regulation is hypothesized to reduce household production costs of children and, in turn, increase fertility. Increases in the duration of leave imply longer time out of the labor market for women (i.e. larger losses in terms of foregone wages and human capital depreciation), but they also allow them to recover after a birth and take care of the newborn. The subsidy paid during leave covers foregone labor market earnings and reduces the opportunity cost of childbearing. In this sense, increases in maternity leave are expected to have a positive effect on fertility. Nevertheless, maternity leave could also create incentives for women to enter the workforce and this may reduce fertility for them.

The duration of leave variable was created based on a review of the laws for each of the countries over the period of analysis (starting in the 1960's) and a summary of changes in leave duration by country is shown in Table 3.A2 in the Appendix to this chapter. Maternity leave is regulated at the national level for the selected countries, so there are no differences in duration within countries. This benefit is paid in all countries examined and the replacement rates are almost always the whole wage or salary. Other variables used in the analysis include the woman's educational attainment, marital status, and family size from the various microdata sources, and the proportion of rural population and infant mortality rates (children under one year old) for each country from the World Development Indicators.

3.3.2. Empirical Strategy

The effects of maternity leave will be assessed through a difference-in-differences (DD) approach. The identification strategy is based on the variation of the duration of maternity leave over time (at different times for different countries). Women during their reproductive years are expected to be affected by changes in these regulations. Reductions in the production cost of children through maternity leave should primarily affect economically active women (possibly increasing fertility), but this may also create incentives for inactive women to participate in the labor market (possibly decreasing fertility). Furthermore, there could be changes in the timing of births even if there are no effects on the final number of children.

Therefore, two age groups are used for the analysis: women between 18 to 30 years old and women between 31 to 45 years old. The criterion to choose the cutoff ages is based on the age-specific fertility rates for the countries under analysis. The average number of children born per woman for ages between 15 and 49 years old is shown in Table 3.A5 in the Appendix to this chapter. In particular, the number of children per woman starts increasing significantly at 18 years old (changes of 2 percentage points or more of total fertility for each additional year of age), while these increases per year of age generally peak at 30 years old (and additional changes become progressively smaller after this age). Thus, the age range for the first group represents about half of the average number of children born per woman across the countries of interest. The second age group represents a smaller share in the total fertility observed, but it is a stage in the women's reproductive years when a larger proportion of higher order births are occurring.

The definition of these groups would also allow identification of possible changes in the timing of births.

Consider the fertility outcome Y_{ijt}, where *i* denotes an individual, *j* is the group (women between 18 and 30 years old or women between 31 and 45 years old), *k* is the country, and *t* is the time period:

$$Y_{ijkt} = \alpha_j + \beta_j T_t + \varphi_j X_{ijkt} + \upsilon_j Z_{kt} + \lambda_j ML_{kt} + \theta_{ijk} + \varepsilon_{ijkt} \quad \ldots (3.2)$$

In equation (3.2), α_j is a group specific intercept, T_t is a general time effect common across all countries (corresponding to each decade or census round from Table 3.1), X_{ijkt} are time-varying controls for each individual (educational attainment and marital status dummies), Z_{kt} are time-varying controls at the country level (proportion of rural population and infant mortality rates), ML_{kt} is the duration of maternity leave (in weeks), θ_{ijk} is an individual fixed effect, and ε_{ijkt} is an error term. This empirical specification follows Zhang, Quan, and Van Meerbergen (1994), Winegarden and Bracy (1995), and Averett and Whittington (2001).

Equation (3.2) is a difference-in-differences (DD) estimate given that maternity leave laws change at different times in different countries. This equation would not be a DD estimate if the analysis were based on one country. However, with multiple countries, those implementing changes in maternity leave duration at time "t" are the treatment group, while those where it remained unchanged serve as the control group. Thus, the country that extended the duration of maternity leave last is the control group for the rest. During the time period when data are available for this study, I observe one or two changes in leave duration for each country, except for Bolivia (which is used as a "control" throughout all time periods). Furthermore, most leave duration changes occur at different time periods, such that no more than two countries increase it simultaneously over any given time period.

Given that no longitudinal data are available to estimate this model, I will apply a pseudo-panel method (Deaton, 1985) to census microdata (complemented with household surveys). This approach averages the data within cohorts, which are treated as observations of the same unit over time. Consider the following simplified model (Deaton, 1985; Verbeek, 2008):

$$y_{it} = x_{it} \cdot \beta + \alpha_i + \varepsilon_{it} \;\ldots\; (3.3)$$

In equation (3.3), x_{it} are controls, the subscript i refers to an individual observed in time period t (coming from independent cross-section datasets), and α_i is an individual fixed effect. We can define groups such that each individual i belongs to only one of them and this group membership is fixed over time (Deaton, 1985; Verbeek and Nijman, 1992; Verbeek, 2008). As Deaton (1985) suggests, an obvious criterion to define these groups are age cohorts, that is, all individuals born in a specific year or multiple years. Then, we can aggregate all individuals i belonging to cohort c in time t, such that we obtain a model based on the sample cohort means:[16]

$$\bar{y}_{ct} = \bar{x}_{ct} \cdot \beta + \bar{\alpha}_c + \bar{\varepsilon}_{ct} \;\ldots\; (3.4)$$

This produces a pseudo or synthetic panel where the unit of observation is the cohort over time. This model yields consistent estimates for β even if $\bar{\alpha}_c$ is correlated

[16] Even though the analysis uses mostly censuses, this model is still based on sample cohort means and not population estimates, given that the datasets are samples from the original census data (10 percent in most cases).

with any of the controls, by treating $\overline{\alpha}_c$ as a parameter to estimate. In this model, $\overline{\alpha}_c$ is treated as a time invariant fixed effect for the cohort, which is reasonable if we average across a large number of observations (Verbeek and Nijman, 1992; Verbeek, 2008).

In this paper, the basic set of results is produced using single years of birth to define the cohorts, so that the relevant variables are transformed into cohort means by birth year. In addition, results are also estimated based on cohorts by year of birth and educational attainment, identifying persons with less than primary education completed and those with primary completed or further education.[17] Since the analysis is based primarily on census data (complemented with household surveys), I obtain a reasonable number of observations within each cohort. Previous methodological research suggests that cohorts for the synthetic panel should comprise more than 100 to 200 observations (Verbeek and Nijman, 1992; Verbeek and Vella, 2005), while several empirical applications worked with minimum cohort sizes of about 100 to 500 observations (Browning, Deaton, and Irish, 1985; Banks, Blundell, and Preston, 1994; Blundell, Browning and Meghir, 1994; Blundell, Duncan, and Meghir, 1998; Propper, Rees, and Green, 2001; Warunsiri and McNown, 2010).

Returning to the full model in equation (3.2) and after taking the cohort means of each variable, the cohort fixed effects model is defined as:

$$\overline{Y}_{cjkt} = \alpha_j + \beta_j T_t + \varphi_j \overline{X}_{cjkt} + \upsilon_j Z_{kt} + \lambda_j ML_{kt} + \overline{\theta}_{cjk} + \overline{\varepsilon}_{cjkt} \quad \ldots (3.5)$$

[17] Given that the pseudo-panel method requires that group membership is fixed over time, the assumption is that persons 18 years old or more with less than primary education or primary completed are unlikely to switch groups over time. Blundell, Duncan, and Meghir (1998) also construct a pseudo-panel based on year of birth and educational attainment, although the groups identified for the cohorts correspond to "those who left education at the minimum legal age and those who continued beyond the minimum" (p. 838).

In equation (3.5), c denotes the cohort and $\bar{\theta}_{cjk}$ is a cohort fixed effect (so we are controlling for unobserved cohort characteristics that are fixed over time), and similarly to equation (3.2), α_j is a group specific intercept, T_t is a time effect (corresponding to each decade or census round from Table 3.1), $\bar{X}_{cjkt}$ are time-varying controls for each cohort (educational attainment and marital status dummies), Z_{kt} are time-varying controls at the country level (proportion of rural population and infant mortality rates), ML_{kt} is the duration of maternity leave (in weeks), and $\bar{\varepsilon}_{cjkt}$ is an error term. In effect, equation (3.5) is analogous to equation (3.2) but it is based on sample cohort means data for $\bar{Y}_{cjkt}$ and $\bar{X}_{cjkt}$.

The specification shown in equation (3.5) is the difference-in-differences (DD) model to be estimated. The empirical strategy does not include a comparison group, as changes in maternity leave duration could affect both women between 18 to 30 years old and women between 31 to 45 years old. Even though the model controls for cohort fixed-effects and time period variables, bias may still be introduced in the estimated effects of maternity leave if time-varying country specific factors are correlated to changes in leave duration. The availability of an appropriate comparison group would allow estimating a difference-in-differences-in-differences (DDD) model to eliminate this possible source of bias. This limitation in the analysis is associated to the fertility outcomes being observed only for women in their reproductive years (who are expected to be "treated"), but needs to be acknowledged for the interpretation of results.

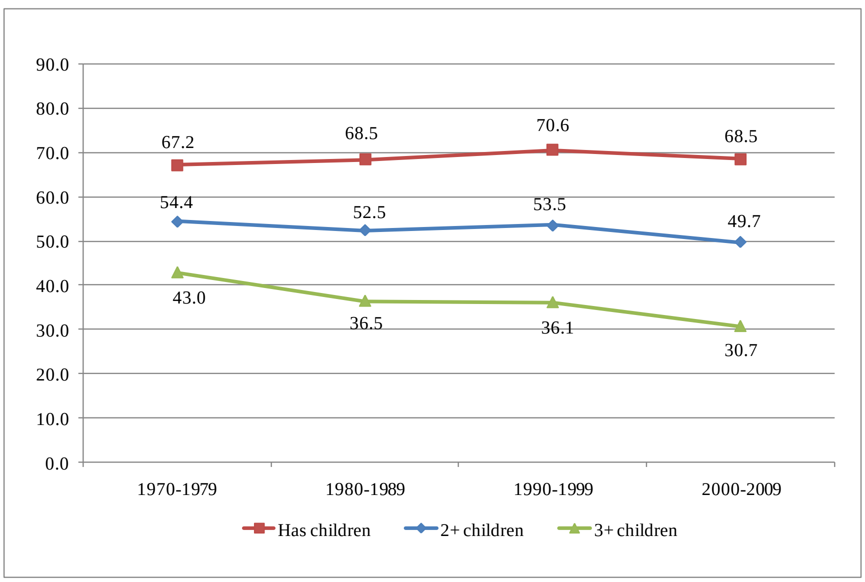
3.4. Results

3.4.1. Fertility Trends by Gender in Selected Countries

This subsection presents some fertility trends for the six countries of this study to provide background for the analysis. The average number of children born per woman has been declining from about six in the 1950s and 1960s to a level close to population reproduction after 2005, as previously shown in Figure 1. Even though the proportion of women in reproductive ages who have children (mothers) has not changed considerably, this proportion has been declining over time for those giving birth to multiple children. As we observe in Figure 2, the percentage of women with children remained relatively stable at around 70 percent between the 1970s and 2000s. In contrast, women with two or more children decreased by 5 percentage points and those with three or more decreased by 13 percentage points over the same time period. This pattern also implies that women having only one child increased from about 12 percent in the 1970s to 18 percent in the 2000s. Given that these changes may reflect timing of births (rather than completed fertility) combined with the age structure of the population, fertility rates by women's age are also examined.

The conclusions are similar if we disaggregate these fertility indicators by age groups (Figure A1 in the Appendix to this chapter). The proportion of women with any number of children (mothers) remained stable across age groups over time: about 35 percent for women of ages 15 to 24 years old, 80 percent for women of ages 25 to 34 years old, and 90 percent for women of older ages. The figures also show a decrease in the proportion of women with two or more children and women with three or more children across all age groups, although this difference is more pronounced for the two intermediate age groups (between 25 and 44 years old). The evidence underscores two inter-related adjustments in fertility behavior over time. First, the younger cohorts have postponed some higher order births. In effect, even though the proportion of women with two or more children decreased considerably over time for the age group between 25 and 34 years old (by about 15 percentage points), this reduction is smaller for older women (about 1 percentage point for women between 44 and 49 years old). That is, a similar proportion of women give birth to two or more children by the end of their reproductive ages, but younger cohorts delayed the second birth. A similar pattern is observed for women with three or more children. Second, the changes over time for the oldest age group (between 45 and 49 years old) confirm that there has been an actual decrease in higher order births, but it is larger for women with three or more births.

Figure 3.2: Women by Number of Children (%) for Population Ages 15 to 49 for Selected Countries in Latin America (1960-2010) 1/



Data source: Integrated Public Use Microdata Series (IPUMS) International.

1. Data include the following countries: Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela. The proportions of women by number of children have been calculated using only census microdata (census years shown in Table 3.1) and the rates are weighted by the population of each country.

The proportion of women with children by educational attainment and other demographic characteristics shows some patterns in the six countries examined. The results for women between 15 and 49 years old are shown below in Table 3.2 and include only the 2000 census round, but similar patterns are found in data for other census rounds available for each country. For the younger age group (15 to 24 years old), only about a third of women have any children, yet this number increases to about four out of five for the next older age group (25 to 34 years old). Furthermore, more than 90 percent of women in the two older age groups (at the end of their reproductive age) have at least one child in the six countries under analysis. The proportion of women with any children is considerably higher for those with the lowest educational attainment (less than primary completed), while it tends to be smaller for women with secondary or tertiary education. The proportion of women with children is considerably smaller for those that are single (about 20 to 30 percent across countries) with respect to other marital statuses (married, in union, divorced, or widowed). Finally, the proportion of mothers among inactive women is larger than among those unemployed (possibly due to a stronger family orientation) but smaller than those employed.[18]

[18] The table describes correlations of fertility with other person characteristics to identify the possible direction of their effects on fertility. Nevertheless, some of these correlations may, in fact, hide multivariate relationships. For example, single women are younger on average and this should explain part of the differences observed in the data.

Table 3.2: Women with Children (%) Ages 15 to 49 for Selected Countries in Latin America (Census Round 2000) by Age, Education, Marital Status, and Employment

	Bolivia 2001	Chile 2002	Colombia 2005	Ecuador 2001	Peru 2007	Venezuela 2001
Age group						
15-24 years	40.4	32.5	35.3	36.1	30.6	33.3
25-34 years	85.6	78.9	79.6	82.4	84.9	78.9
35-44 years	94.5	91.4	88.6	91.6	93.9	91.2
45-49 years	95.2	92.5	88.7	92.5	95.1	93.3
Obs.	*181,460*	*380,170*	*976,753*	*305,173*	*676,979*	*605,299*
Educational attainment						
Less than primary	87.0	87.1	82.2	77.4	87.6	81.7
Primary	67.4	73.7	67.1	66.1	61.3	72.0
Secondary	61.9	67.2	57.9	61.3	63.7	57.0
University	62.9	59.1	55.7	64.9	74.1	49.0
Obs.	*179,424*	*380,170*	*966,048*	*304,327*	*676,979*	*603,312*
Marital status						
Single	27.9	33.5	28.8	18.2	17.7	24.0
Married or in union	94.6	93.8	89.3	93.4	93.3	91.3
Other	97.4	96.3	93.5	96.4	97.8	95.7
Obs.	*181,460*	*380,170*	*966,221*	*304,719*	*676,979*	*604,080*
Employment status						
Employed	75.8	73.6	69.9	67.8	72.9	73.1
Unemployed	65.5	62.6	54.7	58.6	53.7	51.4
Inactive	69.2	71.5	66.9	68.0	66.6	64.5
Obs.	*180,036*	*380,170*	*962,832*	*296,339*	*676,979*	*604,140*

Data source: Integrated Public Use Microdata Series (IPUMS) International.

Results from Table 3.2 were also produced for women with two or more or three or more children, to determine if similar patterns are observed for women with higher order children (see Tables 3.A3 and 3.A4 in the Appendix to this chapter). For these two outcomes, comparable fertility behavior is found by age, educational attainment, and marital status. The proportion of women with higher order children is about 80 percent for two or more and 60 percent for three or more for the last two age groups, while it is the highest for women with less than primary completed and for non-single women. Therefore, the size of the differences for higher order children by these characteristics may be larger in some cases, but the direction of the differences is consistent with the proportion of motherhood analyzed in Table 3.2.

However, there are some changes in the pattern by employment status. Even though unemployed women still have the lowest proportion of higher order children compared to those employed or inactive, this proportion tends to be larger for those inactive with respect to employed women (particularly for three or more children). That is, despite the fact that employed women tend to report being mothers slightly more often than women who are inactive, inactive women tend to have higher order children slightly more often. This difference may be explained due to a stronger family orientation from those women, who have a relatively weaker attachment to the labor market.

3.4.2. Effects of Maternity Leave on Fertility

The difference-in-differences (DD) estimates of equation (3.5) for cohorts defined by year of birth are shown in Table 3.3 for any number of children, two or more children, or three or more children for the two age groups previously defined. If there are effects of maternity leave regulations, these could happen at different stages over the women's reproductive years. These effects could be translated into changes in the number or in the timing of births. The table shown below includes only the estimated coefficients for maternity leave, but the model controls for the woman's educational attainment, marital status, and family size, the country level proportions of rural population and infant mortality rates, and dummy variables for each time period.

Results show that maternity leave duration does not have a statistically significant effect on motherhood (one or more children ever born) for either age group. This is not surprising given the high proportion of women who are mothers in these countries (about

70 percent of women). However, there are some effects for higher order births. For the younger age group, the effect is zero for two or more births and negative for three or more births, although not statistically significant for either outcome. In contrast, the effect of maternity leave on fertility for the older group of women is positive and statistically significant for higher order births, both two or more or three or more children. In particular, this effect implies an increase of about 0.4 to 0.5 percentage points for an additional week of leave, with respect to a proportion of women with two or more births around 50 percent and three of more births around 40 percent. Therefore, there is a relatively small positive effect for higher order births and only for women between 31 and 45 years old, based on cohorts defined by year of birth.

Table 3.3: Difference-in-Differences (DD), Pseudo-Panel (Year of Birth) Cohort Fixed-Effects Regressions for Fertility Outcomes [1/]

Outcome	Women 18-30 years old	Women 31-45 years old
Has children		
Mat. leave coefficient	**-0.0017**	**0.0001**
	[0.0047]	*[0.0012]*
Obs.	312	358
Groups	262	277
Cohort size (average)	11,331	8,247
Two or more children		
Mat. leave coefficient	**0.0001**	**0.0041****
	[0.0044]	*[0.0018]*
Obs.	312	358
Groups	262	277
Cohort size (average)	11,331	8,247
Three or more children		
Mat. leave coefficient	**-0.0069**	**0.0048***
	[0.0049]	*[0.0025]*
Obs.	312	358
Groups	262	277
Cohort size (average)	11,331	8,247

Data source: Census microdata from the Integrated Public Use Microdata Series (IPUMS) - International; selected household surveys from the National Institute of Statistics and Informatics (INEI) - Peru, the National Institute of Statistics (INE) - Bolivia, and the Ministry of Social Development - Chile.
Clustered standard errors (for the cohort) in brackets, *** p<.01,** p<.05, * p<.10
1. Regressions control for educational attainment (dummies for completed primary, completed secondary, and completed university), marital status, family size, labor force participation and employment, the proportion of rural population, the mortality rate for children under one year old, and dummy variables for each time period.

The analysis was extended to cohorts defined by year of birth and educational attainment (identifying persons with less than primary and those with primary completed or further education) for the two age groups. The additional set of estimates allow one to check whether the results are consistent when year of birth cohorts are split by education and it also produces a larger number of cohort observations to gain statistical precision. The difference-in-differences (DD) estimates of equation (3.5) for cohorts defined by year of birth and educational attainment are shown in Table 3.4 for any number of children, two or more children, or three or more children for both age groups of women. The number of observations in Table 3.5 is almost double the number of cases in the corresponding estimates presented in Table 3.4 due to the combination of years of birth with educational attainment. The regressions use a similar set of control variables.

Results show that there is a positive and statistically significant effect of maternity leave duration on motherhood, but only for women between 31 and 45 years old. The effect on motherhood represents an increase of about 0.4 percentage points for an additional week of leave for this group, which is relatively small with respect to around 70 percent of women who become mothers in the countries under analysis. Furthermore, the effect on women between 18 and 30 years old has similar size (0.26 percentage points), but the difference in precision (larger standard errors) implies that it is not statistically significant. The coefficients for maternity leave for higher order births show opposite effects across the two age groups. For the younger group, we observe a negative effect of maternity leave on the probability of two or more or three or more children, which is statistically significant for the latter outcome. For the older group, we observe a positive and statistically significant effect for both higher order births outcomes. Again, these effects are relatively small compared to the proportion of women with multiple children.

Even though the estimated effects are slightly larger for the cohorts defined by year of birth and educational attainment (with respect to cohorts defined only by year of birth), the evidence seems to be consistent with previous results. Furthermore, the estimates suggest that some women may postpone additional births until a later age in life (given the opposite signs across age groups), while there is a small net effect on motherhood for women between 31 and 45 years old.

Table 3.4: Difference-in-Differences (DD), Pseudo-Panel (Year of Birth and Educational Attainment) Cohort Fixed-Effects Regressions for Fertility Outcomes [1/]

Outcome	Women 18-30 years old	Women 31-45 years old
Has children		
Mat. leave coefficient	**0.0026**	**0.0044*** **
	[0.0036]	*[0.0016]*
Obs.	589	679
Groups	500	533
Cohort size (average)	6,000	4,346
Two or more children		
Mat. leave coefficient	**-0.0057**	**0.0076*** **
	[0.0036]	*[0.0021]*
Obs.	589	679
Groups	500	533
Cohort size (average)	6,000	4,346
Three or more children		
Mat. leave coefficient	**-0.0092** **	**0.0063** **
	[0.0037]	*[0.0025]*
Obs.	589	679
Groups	500	533
Cohort size (average)	6,000	4,346

Data source: Census microdata from the Integrated Public Use Microdata Series (IPUMS) - International; selected household surveys from the National Institute of Statistics and Informatics (INEI) - Peru, the National Institute of Statistics (INE) - Bolivia, and the Ministry of Social Development - Chile.
Clustered standard errors (for the cohort) in brackets, *** p<.01,** p<.05, * p<.10
1. Regressions control for educational attainment (dummies for completed primary, completed secondary, and completed university), marital status, family size, labor force participation and employment, the proportion of rural population, the mortality rate for children under one year old, and dummy variables for each time period.

Finally, I also verified results using a placebo or falsification test. The objective is to test whether results could be influenced by other policies that may have been implemented at similar time periods to increases in maternity leave duration or by general fertility trends. Similar exercises have been performed by previous researchers (Baker and Milligan, 2008a; Rossin-Slater *et al*., 2013; Das and Polachek, 2014). In particular, using the same data, the changes in duration of leave were coded as if they occurred one time period earlier than the date of true enactment. If the estimated results reflect the effect of changes in maternity leave duration and not some other contemporaneous factors, the estimates of the leave duration variable for this exercise should not be statistically significant.

Results for the falsification test are presented in Table 3.5. In almost all cases, I obtain relatively small coefficients, and only one of them is marginally statistically significant (motherhood for women between 18 and 30 years old). This provides further evidence to support estimated effects shown for having any number of children, two or more children, or three or more children for both age groups of women.

Table 3.5: Falsification Test, Difference-in-Differences (DD), Pseudo-Panel (Year of Birth) Cohort Fixed-Effects Regressions for Fertility Outcomes [1/]

Outcome	Women 18-30 years old	Women 31-45 years old
Has children		
Mat. leave coefficient	**0.0095***	**0.0013**
	[0.0053]	*[0.0014]*
Obs.	312	358
Groups	262	277
Cohort size (average)	11,331	8,247
Two or more children		
Mat. leave coefficient	**-0.0047**	**0.0008**
	[0.0061]	*[0.0024]*
Obs.	312	358
Groups	262	277
Cohort size (average)	11,331	8,247
Three or more children		
Mat. leave coefficient	**-0.0014**	**0.0016**
	[0.0067]	*[0.0030]*
Obs.	312	258
Groups	262	277
Cohort size (average)	11,331	8,247

Data source: Census microdata from the Integrated Public Use Microdata Series (IPUMS) - International; selected household surveys from the National Institute of Statistics and Informatics (INEI) - Peru, the National Institute of Statistics (INE) - Bolivia, and the Ministry of Social Development - Chile.
Clustered standard errors (for the cohort) in brackets, *** p<.01,** p<.05, * p<.10
1. Regressions control for educational attainment (dummies for completed primary, completed secondary, and completed university), marital status, family size, labor force participation and employment, the proportion of rural population, the mortality rate for children under one year old, and dummy variables for each time period.

3.5. Discussion

In this paper, I explore the effects of maternity leave on the probability that a woman has any children and on the probabilities that she has two or more children and three or more children in six countries in Latin America (Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela). The empirical strategy exploits changes in the benefit duration at different points in time for different countries. The effects of the selected countries' maternity leave regulations are estimated using census microdata and applying pseudo-panel methods, which were first proposed by Deaton (1985). Even though previous research has examined the fertility effects of maternity leave for developed countries, no previous studies analyze data for Latin America. Therefore, this paper contributes to the existing literature by providing new evidence for this region.

The evidence presented in this paper indicates that maternity leave has only a small effect on fertility, and mainly for higher parity births; the main effect is that it creates incentives to postpone some births until a later stage in life. Maternity leave duration has a positive and statistically significant effect on the proportion of women with any number of children, but only for the group of women between 31 and 45 years old. In particular, the estimates reveal an increase in motherhood for this group by approximately 0.4 percentage points for each additional week of leave, which is relatively small with respect to the 70 percent of women who are currently mothers (on average) in the countries under analysis. In the case of higher order births, the direction of the effect of maternity leave duration depends on the woman's age: there is a decrease in higher order births for women between 18 and 30 years old (by about 0.6 to 0.9 percentage points per week of maternity leave), while there is an increase for women between 31 and 45 years old (by about 0.4 to 0.8 percentage points per week of maternity leave). If we consider these two effects together, increases in maternity leave duration are associated with the postponement of some higher order births for young adult women.

Results based on the pseudo-panel data for the countries under analysis are consistent with previous findings on this topic. Even though most evidence using country level data shows statistically insignificant effects of maternity leave on fertility (Winegarden and Bracy, 1995; Zhang, Quan, and Van Meerbergen, 1994; Gauthier and Hatzius, 1997), a study for Canada (Hyatt and Milne, 1991) found a small increase between 0.09 and 0.26 percent in the total fertility rate for a one percent increase in the real value of maternity benefits. Empirical evidence using longitudinal data or administrative records reports positive effects of maternity leave on first births (Cannonier, 2014) but the main effects are on higher order births (Averett and Whittington, 2001; Ronsen, 2004; Lalive and Zweimüller, 2009; Cannonier, 2014; among others). For instance, in the case of the United States, Averett and Whittington (2001) concluded that the availability of employer provided leave increases the probability of an additional child by about 3 to 7 percentage points. In addition, Cannonnier (2014) found that the Family and Medical Leave Act (FMLA) increased the probability of a first birth by 1.5 percentage points and the probability of a second birth by 0.6 percentage points. Results of this study using pseudo-panel data show effects that are smaller but are generally consistent with those previously identified in the literature.

As a final note, there are some limitations for the data analysis. Previous studies on this topic analyze the total fertility rate for a population (Hyatt and Milne, 1991; Zhang, Quan, and Van Meerbergen, 1994; Gauthier and Hatzius, 1997; among others) or the probability of a birth (Phipps, 2000; Ronsen, 2004; Lalive and Zweimüller, 2009; Cannonier, 2014; among others). These outcomes refer to observed fertility over a certain time period (usually a year),[19] which offer the advantage that individual behavior refers to

[19] The total fertility rate (a synthetic measure) is the average number of children that would be born to a woman during her lifetime and it is obtained by adding up all the current age-specific fertility rates at a given point in time (i.e. for a certain year).

the same time period where changes in the duration of maternity leave are identified. The microdata available for this study most often do not include information about births over a pre-determined time period (i.e. births last year), but rather the woman's number of children ever born. Therefore, instead of any shifts in individual behavior, the indicators used are the result of the cohorts' accumulated fertility over time.

3.6. Appendix: Additional Tables and Figures

Table 3.A1: Summary of Studies on the Fertility Effects of Maternity Leave

Study	Coverage		Data			Estimation strategy
	Geographic	Time	Source	Leave variable	Outcomes	
Cross-country (OECD or Europe)						
Winegarden and Bracy (1995)	17 OECD countries	Four time periods (1959, 1969, 1979, and 1989)	Various sources, country level	Weeks of paid maternity leave (maximum length)	Infant mortality, labor force participation, and general fertility rate	Fixed (country) effects, simultaneous equations for the three outcomes
Gauthier and Hatzius (1997)	22 OECD countries	1970 to 1990	Various sources, country level	Weeks of maternity leave (paid or unpaid) and payment received during this period (as a percentage of regular earnings)	Total fertility rate	Fixed (country) effects, time series of aggregate data for multiple countries
United States						
Averett and Whittington (2001)	United States	1985 to 1992	National Longitudinal Survey of Youth (NLSY)	Availability of employer provided leave (reported by survey respondent)	Desired fertility and probability of a birth	Hazard discrete model. Sample of working women, both including all ages and restricted to 22 to 26 years old.
Cannonier (2014)	United States	1989 to 2010	National Longitudinal Survey of Youth (NLSY)	Dummy variable for the time period after the implementation of the FMLA, 1993	Probability of a first or second birth	Diff-in-diff, comparison of eligible and non-eligible women from a sample in which all have participated in the labor force. Hazard discrete model.

Canada						
Hyatt and Milne (1991)	Canada	1948 to 1986	Various sources, country level	Ratio of the weekly female wage rate plus maternity benefits (weighted) to the female weekly wage rate	Total fertility rate	Time series of aggregate data. Leave is assumed to be endogenous (and the model uses IV) and its effect is interacted with the proportion of employed females to the total number of females.
Zhang, Quan, and Van Meerbergen (1994)	Canada	1921 to 1988	Various sources, country level	Dummy variable for the period after maternity benefits were available through the Unemployment Insurance program, in 1971	Total fertility rate	Time series of aggregate data
Phipps (2000)	Canada	1988 to 1990	Labor Market Activities Survey (LMAS) (longitudinal data)	Dummy variable for women eligible for maternity benefits (based on employment history in preceding year) and the replacement rate (considering ceiling on benefits for higher-income women)	Probability of a birth	Sample of women 16 to 64 years old
Nordic countries						
Hoem (1993)	Sweden	1961 to 1990	Country level fertility indicators	Dummy variable for periods after the increase in the grace period to retain benefits level from previous birth ("speed premium")	Birth rates (for the first, second, third, and fourth child)	Analysis of standardized birth rates trends
Ronsen (2004)	Norway and Finland	1988 and 1989	Norwegian Family and Occupation Survey (1988) and Finnish Population Survey (1989), administrative register data for Norway, and other sources	Duration of parental leave (in weeks for Norway and months for Finland)	Probability of a first, second, or third birth	Hazard discrete model. Data includes retrospective life histories on childbearing, employment, and demographic characteristics.

Bjorklund (2006)	Sweden against other 10 European countries	1925 to 1958	Various sources, country level	Analysis of changes in overall family policies, including leave duration and wage replacement rates	Completed fertility, age at first birth, time interval (in years) between first and last child, and total fertility rate	Diff-in-diff (means), comparison of Sweden against other countries that experienced less important changes in family policies
Duvander and Andersson (2006)	Sweden	1988 to 1999	Administrative register data (fertility, taxes, and education) and other sources	Paid parental leave benefits and their shares for mother and father in the 2 years following a childbirth	Probability of a second or third birth	Hazard discrete model. Sample of co-residing couples that had their first or second common child in the period of analysis, restricted to January births (to have complete income data).
Duvander, Lappegard, and Andersson (2010)	Norway and Sweden	1988 to 2003 (differs for each country)	Administrative register data (fertility, taxes, and education) and other sources	Paid parental leave benefits and their shares for mother and father in the 2 years following a childbirth	Probability of a second or third birth	Hazard discrete model. Sample of co-residing couples that had their first or second common child in the period of analysis, restricted to January births (to have complete income data) and couples eligible for parental benefits.
Lappegard (2010)	Norway	1995 to 2004	Administrative register data (fertility, taxes, and education) and other sources	Parental leave use and its shares for mother and father during their first or second child's first year	Probability of a second or third birth	Hazard discrete model with fixed (municipality) effects. Sample of co-residing couples that had their first or second common child in the period of analysis.
Other countries						
Lalive and Zweimüller (2009)	Austria	1985 to 2000	Austrian Social Security Database (ASSD)	Dummy variables for each period after parental leave duration changed, in July 1990 and in July 1996	Probability of an additional birth	Sample of women ages 15 to 45 who are eligible for parental leave and who gave birth in the month before or after each reform.

Table 3.A2: Evolution of Maternity Leave Duration (in Weeks) for Selected Countries (1925-2012) [1/]

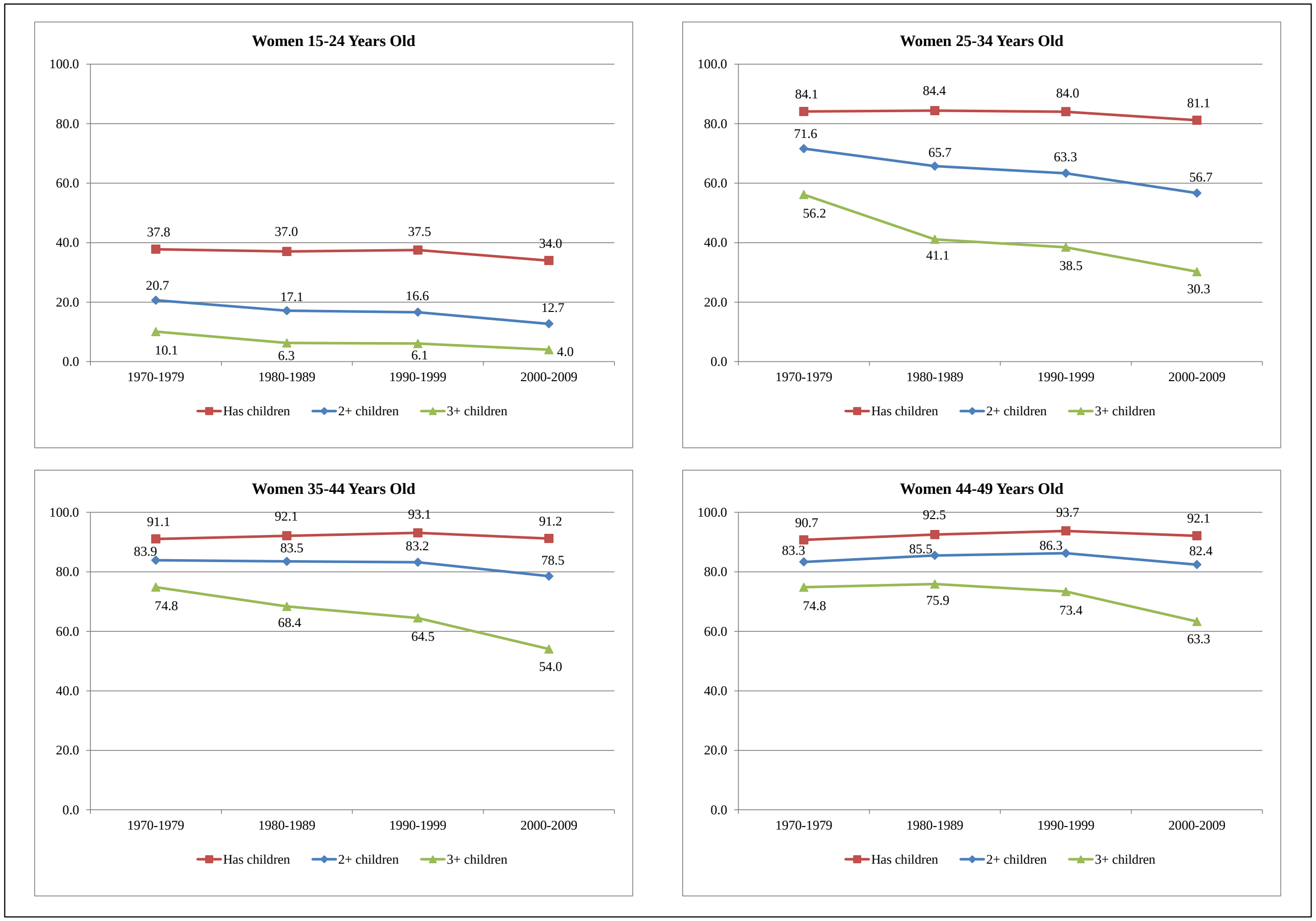
Year	Bolivia	Chile	Colombia	Ecuador	Peru	Venezuela
1925	NA	8.6	NR	NA	8.6	NR
1930	NA	8.6	NR	NA	8.6	NR
1935	NA	12	NR	NA	8.6	NR
1940	8.6	12	8	NA	8.6	12
1945	8.6	12	8	NA	8.6	12
1950	8.6	12	8	NA	8.6	12
1955	8.6	12	8	NA	8.6	12
1960	8.6	12	8	NA	8.6	12
1965	8.6	12	8	6	8.6	12
1970	8.6	12	8	6	8.6	12
1975	12.9	18	8	6	8.6	12
1980	12.9	18	8	8	8.6	12
1985	12.9	18	8	8	8.6	12
1990	12.9	18	12	8	8.6	18
1995	12.9	18	12	8	NR	18
2000	12.9	18	12	12	12.9	18
2005	12.9	18	12	12	12.9	18
2010	12.9	18	12	12	12.9	18
2012	12.9	18	14	12	12.9	26

Source: National legislation

NR = Not regulated, NA = Data not available on maternity regulations

1. The duration of maternity leave obtained from national legislation was compared to the following secondary sources: the ILO TRAVAIL - Working Conditions Laws Database (retrieved from: http://www.ilo.org/dyn/travail/travmain.home), the ILO Social Security Database (retrieved from: http://www.ilo.org/dyn/sesame/IFPSES.SocialDatabase), and the World Bank Cross Country Data (retrieved from: https://www.quandl.com/WORLDBANK). Leave duration included in this table is consistent with the secondary sources examined.

Figure 3.A1: Women by Number of Children (%) for Population Ages 15 to 49 for Selected Countries in Latin America (1960-2010) [1/]



Data source: Integrated Public Use Microdata Series (IPUMS) International.

1. Data include the following countries: Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela. The proportions of women by number of children have been calculated using only census microdata (census years shown in Table 3.1) and the rates are weighted by the population of each country.

Table 3.A3: Women with Two or More Children (%) Ages 15 to 49 for Selected Countries in Latin America (Census Round 2000) by Age, Education, Marital Status, and Employment

	Bolivia 2001	Chile 2002	Colombia 2005	Ecuador 2001	Peru 2007	Venezuela 2001
Age group						
15-24 years	18.8	7.6	13.6	15.5	9.3	14.3
25-34 years	68.1	48.0	57.0	61.6	55.1	57.4
35-44 years	87.0	76.2	75.9	82.0	79.8	79.1
45-49 years	88.9	80.5	78.9	85.1	85.2	83.9
Obs.	*181,460*	*380,170*	*976,753*	*305,173*	*676,979*	*605,299*
Educational attainment						
Less than primary	76.5	71.6	71.1	65.4	75.6	70.8
Primary	52.3	52.5	50.2	49.7	44.4	56.5
Secondary	40.5	42.1	32.3	39.9	38.5	36.0
University	39.4	38.3	31.7	42.5	43.8	29.7
Obs.	*179,424*	*380,170*	*966,048*	*304,327*	*676,979*	*603,312*
Marital status						
Single	12.4	11.7	13.2	8.1	5.9	12.8
Married or in union	79.1	71.1	69.5	73.4	69.1	71.4
Other	80.8	74.2	72.1	71.1	64.1	71.9
Obs.	*181,460*	*380,170*	*966,221*	*304,719*	*676,979*	*604,080*
Employment status						
Employed	58.8	46.5	47.1	48.3	49.1	52.9
Unemployed	43.1	35.5	31.7	38.2	31.6	31.1
Inactive	54.9	52.5	50.5	52.5	47.9	49.5
Obs.	*180,036*	*380,170*	*962,832*	*296,339*	*676,979*	*604,140*

Data source: Integrated Public Use Microdata Series (IPUMS) International.

Table 3.A4: Women with Three or More Children (%) Ages 15 to 49 for Selected Countries in Latin America (Census Round 2000) by Age, Education, Marital Status, and Employment

	Bolivia 2001	Chile 2002	Colombia 2005	Ecuador 2001	Peru 2007	Venezuela 2001
Age group						
15-24 years	7.6	1.5	4.2	5.0	2.3	5.1
25-34 years	47.5	17.6	31.3	36.1	26.2	32.9
35-44 years	73.9	43.4	52.1	62.5	54.0	55.9
45-49 years	78.8	54.0	60.2	71.3	65.8	65.0
Obs.	*181,460*	*380,170*	*976,753*	*305,173*	*676,979*	*605,299*
Educational attainment						
Less than primary	65.3	49.0	55.1	51.8	59.6	57.1
Primary	38.1	28.3	30.3	32.8	27.5	38.1
Secondary	22.7	17.6	12.9	19.6	17.6	16.4
University	17.4	17.0	11.0	19.1	16.5	12.1
Obs.	*179,424*	*380,170*	*966,048*	*304,327*	*676,979*	*603,312*
Marital status						
Single	7.2	4.7	6.6	4.2	2.2	7.4
Married or in union	60.1	36.8	43.7	49.0	42.1	46.0
Other	60.2	43.0	45.2	46.8	37.8	45.8
Obs.	*181,460*	*380,170*	*966,221*	*304,719*	*676,979*	*604,080*
Employment status						
Employed	43.3	21.7	25.5	29.9	27.6	31.2
Unemployed	26.9	17.7	16.9	21.6	16.4	17.0
Inactive	41.5	28.3	32.6	35.7	29.8	33.1
Obs.	*180,036*	*380,170*	*962,832*	*296,339*	*676,979*	*604,140*

Data source: Integrated Public Use Microdata Series (IPUMS) International.

Table 3.A5: Number of Children Born per Woman (average) by Age

Age	Bolivia 1976	Bolivia 1992	Bolivia 2001	Chile 1960	Chile 1970	Chile 1982	Chile 1992	Chile 2002	Colombia 1973	Colombia 1985	Colombia 1993
15	0.02	0.03	0.04	0.01	0.22	0.03	0.05	0.20	0.03	0.02	0.03
16	0.05	0.09	0.09	0.05	0.24	0.06	0.09	0.08	0.07	0.06	0.14
17	0.13	0.17	0.20	0.12	0.30	0.13	0.15	0.14	0.19	0.14	0.27
18	0.26	0.35	0.35	0.19	0.40	0.23	0.24	0.22	0.35	0.27	0.43
19	0.46	0.55	0.51	0.36	0.52	0.38	0.38	0.33	0.57	0.42	0.62
20	0.73	0.85	0.74	0.56	1.16	0.53	0.48	0.43	0.87	0.69	0.82
21	0.96	0.99	0.91	0.64	1.26	0.68	0.62	0.54	1.03	0.83	0.95
22	1.22	1.30	1.11	0.91	1.43	0.86	0.76	0.66	1.37	1.04	1.12
23	1.50	1.52	1.33	1.20	1.61	1.03	0.92	0.76	1.67	1.26	1.31
24	1.77	1.75	1.55	1.45	1.77	1.17	1.05	0.85	1.94	1.44	1.45
25	2.05	1.99	1.78	1.69	2.05	1.34	1.17	0.96	2.30	1.65	1.56
26	2.38	2.23	1.99	1.94	2.18	1.57	1.32	1.07	2.52	1.83	1.70
27	2.66	2.44	2.16	2.09	2.46	1.69	1.44	1.21	2.84	2.03	1.86
28	2.93	2.68	2.46	2.26	2.62	1.87	1.58	1.34	3.11	2.17	1.99
29	3.29	2.91	2.64	2.32	2.99	2.03	1.72	1.45	3.39	2.40	2.11
30	3.56	3.09	2.91	2.43	3.35	2.17	1.87	1.58	3.78	2.57	2.29
31	3.97	3.31	3.07	2.97	3.61	2.36	1.99	1.73	3.98	2.71	2.31
32	4.10	3.50	3.24	3.19	3.70	2.55	2.14	1.87	4.32	2.89	2.56
33	4.35	3.58	3.53	3.39	4.01	2.63	2.25	1.95	4.63	3.08	2.71
34	4.61	3.72	3.71	3.14	4.14	2.75	2.33	2.05	4.84	3.26	2.78
35	4.81	3.96	3.91	3.43	4.11	2.82	2.40	2.11	4.97	3.38	2.91
36	5.17	4.24	4.12	3.85	4.33	3.06	2.53	2.16	5.32	3.66	3.03
37	5.27	4.38	4.27	3.62	4.58	3.15	2.64	2.27	5.58	3.84	3.20
38	5.50	4.47	4.48	3.67	4.68	3.27	2.69	2.34	5.75	3.98	3.32
39	5.70	4.47	4.61	3.92	4.79	3.44	2.78	2.43	5.91	4.27	3.38
40	5.58	4.68	4.73	3.62	4.87	3.48	2.82	2.47	5.88	4.33	3.58
41	5.89	4.69	4.86	4.15	5.19	3.68	2.87	2.52	6.12	4.61	3.52
42	5.94	4.83	4.99	3.78	5.03	3.84	2.95	2.60	6.32	4.65	3.74
43	6.12	4.86	5.14	3.98	5.36	4.04	3.08	2.65	6.47	5.07	3.90
44	6.23	5.05	5.28	4.09	5.40	4.14	3.10	2.65	6.72	5.26	4.01
45	5.97	4.93	5.32	3.76	5.06	3.98	3.05	2.66	6.33	5.23	4.16
46	6.09	5.18	5.36	3.97	5.38	4.23	3.19	2.72	6.62	5.45	4.25
47	6.35	5.19	5.36	4.02	5.31	4.27	3.35	2.77	6.56	5.65	4.40
48	6.30	5.08	5.44	3.57	5.19	4.35	3.40	2.80	6.60	5.68	4.64
49	6.29	5.31	5.60	3.77	5.32	4.48	3.52	2.88	6.63	5.85	4.65

Data source: Integrated Public Use Microdata Series (IPUMS) International.

Table 3.A5 (continued): Number of Children Born per Woman (average) by Age

Age	Colombia 2005	Ecuador 1974	Ecuador 1982	Ecuador 1990	Ecuador 2001	Ecuador 2010	Peru 1993	Peru 2007	Venezuela 1971	Venezuela 1990	Venezuela 2001
15	0.04	0.04	0.10	0.18	0.04	0.04	0.04	0.02	0.06	0.08	0.04
16	0.10	0.08	0.17	0.07	0.09	0.10	0.07	0.06	0.12	0.15	0.09
17	0.19	0.17	0.30	0.16	0.18	0.19	0.13	0.12	0.23	0.27	0.18
18	0.32	0.36	0.50	0.28	0.31	0.31	0.23	0.20	0.35	0.42	0.30
19	0.46	0.56	0.74	0.44	0.46	0.44	0.34	0.30	0.53	0.60	0.44
20	0.63	0.92	1.01	0.68	0.65	0.58	0.54	0.43	0.76	0.80	0.60
21	0.77	1.05	1.20	0.80	0.83	0.73	0.67	0.54	0.92	1.00	0.75
22	0.94	1.43	1.54	1.03	0.97	0.89	0.91	0.92	1.19	1.20	0.91
23	1.07	1.68	1.73	1.23	1.17	1.04	1.13	1.06	1.43	1.40	1.07
24	1.21	2.05	1.96	1.45	1.33	1.18	1.29	1.19	1.74	1.60	1.25
25	1.35	2.43	2.24	1.65	1.48	1.33	1.52	1.29	1.99	1.79	1.38
26	1.49	2.68	2.47	1.86	1.62	1.49	1.70	1.41	2.23	1.99	1.54
27	1.66	2.98	2.72	2.07	1.79	1.64	1.95	1.56	2.46	2.19	1.68
28	1.80	3.33	2.98	2.29	1.96	1.77	2.15	1.67	2.70	2.37	1.83
29	1.92	3.54	3.11	2.53	2.04	1.88	2.29	1.79	2.95	2.55	1.96
30	2.08	4.05	3.50	2.79	2.32	2.00	2.58	1.98	3.22	2.80	2.11
31	2.18	3.99	3.55	2.92	2.40	2.13	2.66	2.03	3.31	2.96	2.28
32	2.30	4.48	3.96	3.13	2.61	2.24	2.94	2.19	3.52	3.12	2.38
33	2.40	4.62	4.12	3.36	2.66	2.37	3.18	2.30	3.75	3.35	2.53
34	2.47	4.98	4.33	3.52	2.82	2.45	3.23	2.45	3.87	3.43	2.62
35	2.57	5.28	4.60	3.67	2.91	2.52	3.50	2.54	3.93	3.61	2.69
36	2.64	5.50	4.93	3.88	3.08	2.67	3.64	2.72	4.03	3.79	2.82
37	2.70	5.88	5.11	4.08	3.16	2.78	3.85	2.83	4.10	3.89	2.95
38	2.77	5.87	5.33	4.35	3.33	2.82	4.15	2.93	4.19	4.03	3.05
39	2.87	6.19	5.64	4.40	3.38	2.91	4.23	3.09	4.18	4.22	3.08
40	2.93	6.26	5.74	4.59	3.55	2.97	4.51	3.22	4.12	4.46	3.23
41	2.97	6.39	5.91	4.65	3.61	3.06	4.47	3.26	4.20	4.49	3.30
42	3.08	6.46	6.06	4.84	3.73	3.08	4.67	3.37	4.24	4.68	3.40
43	3.16	6.53	6.40	5.08	3.85	3.17	4.83	3.46	4.25	4.81	3.51
44	3.20	6.78	6.56	5.25	3.95	3.26	4.90	3.49	4.22	4.96	3.55
45	3.28	6.57	6.39	5.30	4.01	3.28	5.10	3.69	4.06	5.08	3.56
46	3.34	6.76	6.60	5.50	4.12	3.33	5.22	3.77	4.03	5.29	3.66
47	3.39	6.72	6.71	5.63	4.24	3.43	5.23	3.81	4.02	5.36	3.71
48	3.47	6.55	6.59	5.63	4.27	3.50	5.56	3.96	4.02	5.49	3.84
49	3.53	6.74	6.74	5.83	4.44	3.56	5.51	3.99	4.09	5.61	3.88

Data source: Integrated Public Use Microdata Series (IPUMS) International.

Chapter 4: Testing Alternative Aggregation Methods Using Ordinal Data for a Census Asset-Based Wealth Index

4.1. Introduction

The asset-based index approach to measure socioeconomic status has been widely used as an alternative measure of that status when income and expenditure data are not available. Principal components analysis (PCA) on data recoded to binary indicators (Filmer and Pritchett, 2001) is one of the most frequently used procedures to construct such an index. However, this approach has been subject to criticism, given that the standard PCA method does not consider that many asset variables are in fact categorical or ordinal. Furthermore, the variable dichotomization procedure not only generates spurious negative correlations (across binary indicators derived from the same categorical or ordinal variable) but also neglects the ordering of categories that may contribute additional information to define the index (Howe *et al*., 2008; Kolenikov and Angeles, 2009).

The use of ordinal data and polychoric correlations has been proposed as an alternative to overcome these criticisms of the commonly used approach that applies PCA to binary data (Kolenikov and Angeles, 2009). The performance of aggregation procedures based on asset ordinal data has not been extensively tested. Howe *et al*. (2008) found that the choice of categorical versus binary data had a strong influence on the agreement between alternative indices defined from living conditions variables. Kolenikov and Angeles (2009) compared PCA applied to binary indicators to other methods using ordinal variables. Their results show better performance of indices based on ordinal data according to different criteria, including the proportion of data variability explained by the index and its statistical significance in explaining women's fertility. Thus, they do not recommend working with binary indicators unless there is no information at all regarding the ordering of categories. Other research on this topic has examined the question on aggregation procedures for asset variables, but not through methods appropriate to deal with discrete asset data (Montgomery *et al*., 2000; Bollen *et al*., 2002; Filmer and Scott, 2012).

In this chapter, I use census data to test alternative aggregation procedures to define a asset-based wealth index based on information of housing characteristics and assets.[20] The type and number of variables available vary widely in census microdata, in comparison to the more standard asset information typically included in household surveys (such as in the case of the Demographic and Household Surveys). This data variability provides an appropriate setting to test the relative performance of asset-based indices produced by alternative PCA methods, some of which are designed to handle ordinal variables. In particular, I explore whether these alternative methods generate similar household rankings and whether there are differences in their relation with selected education, fertility, and mortality outcomes. Results are also compared against the logarithm of income per capita for those datasets with this information available.

The chapter is organized as follows. In the next section, I discuss previous research on methods to aggregate data on housing characteristics and asset ownership to define a

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[20] Throughout the paper, I will refer to indices constructed from information on housing characteristics and assets simply as asset-based indices or asset indices (which they are frequently called).

proxy measure of socioeconomic status. In Section 4.3, I describe the data and the methods to construct the indices that are analyzed, including principal component analysis and the use of polychoric correlations. Next, in Section 4.4, I show the results of the study. Section 4.5 has a discussion of the main findings of the study. The appendix to this chapter includes additional tables.

4.2. Literature Review

Filmer and Pritchett (2001) examined the use of housing characteristics and asset ownership to define an alternative measure of household socioeconomic status. This practical approach is motivated by the fact that income or expenditures are not always available in microdata. In their application, categorical variables are transformed into binary indicators (where each category is recoded as a separate variable) and principal components analysis (PCA) is used to assign weights to each indicator to construct an index. The authors found not only comparable rankings of households based on asset or expenditure data but also that these measures had similar predictive power to explain school enrollment using microdata from India, Indonesia, Nepal, and Pakistan. The method proposed by Filmer and Pritchett (2001), which applied PCA to dichotomized asset data, has been widely used as a control for household socioeconomic status in other studies that examine a variety of outcomes (see, for example, Bollen *et al*., 2002; Minujin and Bang, 2002; Houweling *et al*., 2003; Rutstein and Johnson, 2004; McKenzie, 2005; Lindelow, 2006; Bollen *et al*., 2007; Filmer and Scott, 2012; Wagstaff and Watanabe, 2003).

The use of information on housing characteristics and assets to define a proxy measure for household socioeconomic status leads to the question about the methods used to aggregate (i.e. produce weights for) the data. This question has been previously explored by several studies in this field. Montgomery *et al*. (2000) analyzed the use of individual living conditions variables against an index represented by the simple sum of these indicators. Their evidence indicates that either of these alternatives had limited explanatory power for consumption expenditures per adult, but they were useful proxies in regressions explaining fertility, child mortality, or children's schooling. Bollen *et al*. (2002) applied four different aggregation methods on information from consumer durable goods, including the number of assets, their current and median value, and PCA on binary indicators. The authors conclude that the number of durable assets and the binary PCA have stronger effects on children ever born than the current or median value of assets. Howe *et al*. (2008) worked with several methods to calculate weights, including PCA on categorical and dichotomized data. The study suggests that the choice of data (categorical versus dichotomized variables) had more influence on the agreement of indices than the different methods that were used to weight the data, while all the aggregation procedures had similar moderate agreement with consumption expenditures. Filmer and Scott (2012) compared a variety of approaches to measure welfare based on living conditions data, which include indices derived from an asset count, the traditional PCA on binary indicators, item response theory (IRT), and predicted per capita household expenditures. Their results show that household rankings are not identical and they depend on which measure is used, but differences in outcomes across these rankings are robust to this choice. Overall, conclusions regarding the relative performance of these methods do not strongly advocate for the use of one of them before the rest.

The Filmer and Pritchett (2001) approach to produce the index has been subject to criticism by more recent research, given that many asset variables are ordinal (such as dwelling ownership, type of water supply, or predominant walls material). Some specific issues have been identified (Howe *et al*., 2008; Kolenikov and Angeles, 2009). PCA relies largely on the calculation of the variance of the data --as it will be later discussed-- to produce the weights for the index. However, the methods frequently used to calculate the variance-covariance matrix for PCA neglect the fact the asset data are primarily discrete. In fact, PCA is based on the assumption that the data follow a multivariate (joint) normal distribution, which is clearly violated when working with asset data (Kolenikov and Angeles, 2009).[21] Furthermore, the dichotomization process generates spurious negative correlations between binary indicators derived from the same categorical variable, because by construction these come from categories that are mutually exclusive.[22] Therefore, the variance of the data used for PCA is based not only on the (positive) correlations between asset variables (hypothesized to depend on unobserved household wealth), but also on artificial negative correlations between indicators defined from the same categorical variable. The index could then reflect variability associated to these spurious correlations rather than that of unobserved

[21] Kolenikov and Angeles (2009) argue that normality should be at least a reasonable distribution approximation, given that non-normal distributions of indicators entail that "some of the properties of the principal components no longer hold or need to be revised" (p. 161). Other authors suggest that normality is not always required but imply that it may be necessary for certain properties of PCA; for example, Jolliffe (2002), p. 19 or Timm (2002), p. 447.

[22] For example, having a dirt floor implies that a household cannot declare any other alternative as predominant construction material of the dwelling floors. Thus, the binary indicators will have some negative correlation just for being defined based on categories from the same variable.

household wealth.[23] Finally, the ordering of categories implied in ordinal variables contributes additional information, but this information is lost when these variables are transformed into binary indicators. PCA on binary indicators is considered assumption-free, both in terms of the order of categories (which may be incorrect) and the scale given to the distance between categories. However, ignoring the ordering may exclude useful information to rank households if the options clearly follow a certain order.

Kolenikov and Angeles (2009) examined methods to define a wealth index based on ordinal data, to address the issues of the popular approach proposed by Filmer and Pritchett (2001). In particular, they compared PCA on dichotomized variables to PCA using polychoric correlations (on ordinal data) and PCA on ordinal variables using the standard methods to calculate the variance of the data.[24] The authors implemented a simulation study based on artificial data, where PCA based on binary indicators is compared against the two methods based on ordinal variables. Results showed that transforming categorical variables into binary indicators leads to lower performance, mostly in terms of the proportion of variance of the data explained by the index. In addition, even if the ordering of categories is incorrect, there is no evidence that PCA on binary indicators produces better results. The study also includes an empirical example to illustrate the differences between these procedures using data from the Bangladesh 2000

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[23] On this matter, Kolenikov and Angeles (2009, p. 138) argue that the "PCA method then needs to take into account both the fundamental (usually positive) correlations between observed variables and the spurious (negative) correlations between the dummy variables produced from a single factor," such that the "PCA procedure may not be able to recover the SES from the data, as the directions of greater variability may now correspond to those spurious correlations."

[24] Polychoric correlations are designed to handle discrete data unlike the standard correlations used in PCA. They yield a different estimate of the variance-covariance matrix when the latent variables of interest are continuous but we observe ordinal data. The details about the calculation of PCA with polychoric correlations are presented in Section 4.3 on the methodology.

Demographic and Health Survey. This empirical application showed a relatively comparable ranking of households by wealth quintiles for the standard PCA on ordinal variables and the polychoric PCA, but higher disagreements of both indices with that derived from PCA on binary indicators. Moreover, when using each index as a control variable to explain fertility, results are very similar for the methods based on ordinal variables, while the binary PCA has lower significance and smaller coefficients. Based on this evidence, the authors to conclude that PCA on binary indicators is "not recommended unless there is absolutely no information about the ordering of categories" (Kolenikov and Angeles, 2009, p. 162). This recommendation is very strong and I examine it in the rest of this chapter.

4.3. Methodology

4.3.1. Principal Components Analysis (PCA) and Polychoric Correlations

Before describing the data and aggregation procedures that are applied in this study, I present a brief summary of principal components analysis (PCA) and the calculation of polychoric correlations, as these concepts are used throughout the analysis. The objective of PCA is to find a subset of principal components that represents most of the variation in a set of x_j variables. Therefore, PCA is a data dimensionality-reduction technique. Given a set of x_j variables (j=1, ...,p), PCA produces a linear combination, denoted by PC_1, that maximizes the variance of a weighted sum of the x_j variables by using weights w_{ij}, as follows:

$$PC_1 = W_{\boldsymbol{1}}' \boldsymbol{X} = w_{11} \cdot x_1 + w_{12} \cdot x_2 + \ldots + w_{1p} \cdot x_p \ldots \text{ (4.1)}$$

A second linear combination, PC_2, can likewise be defined, that also maximizes the remaining variance of the x_j variables that was not captured by PC_1 and is uncorrelated with PC_1. More generally, PC_k can be defined as the *kth* linear combination or principal component that maximizes the remaining variance of the x_j variables and is uncorrelated with $PC_1, PC_2, ..., PC_{k-1}$. The total number of uncorrelated linear combinations or principal components that can be constructed is equal to the number of x_j variables in the data.

In order to determine the weights, consider the optimization problem where the objective function is the variance of the principal component $VAR\left(W_j^{'}X\right)$ subject to the constraint of $W_j^{'}W_j = 1$ (imposed to find a single maximum) (Jolliffe, 2002; Timm, 2002):

$$L = VAR\left(W_j^{'}X\right) - \eta\left(W_j^{'}W_j - 1\right) = W_j^{'}\sum W_j - \eta\left(W_j^{'}W_j - 1\right) \ldots \text{ (4.2)}$$

In equation (4.2), $\sum$ is the variance-covariance matrix of the x_j variables in the data. Differencing equation (4.2) with respect to the weights W_j gives the first order condition for this (constrained) maximization problem:

$$\frac{\partial L}{\partial W_j} = \sum W_j - \eta W_j = 0 \ldots \text{ (4.3a)}$$

$$\left(\sum - \eta I\right)W_j = 0 \ldots \text{ (4.3b)}$$

This optimization problem is equivalent to finding the eigenvalues of the matrix $\sum$, where λ_j is an eigenvalue and the weight W_j is its corresponding eigenvector (Jolliffe, 2002; Timm, 2002). From equation (4.3b), it is also possible to infer that the eigenvalues are equal to the variance of the corresponding principal component or $VAR\left(W_j^{'}X\right) = \lambda_j$ (Jolliffe, 2002; Kolenikov and Angeles, 2009). A common measure to assess the performance of PCA is the proportion of total variance explained by each principal component. Following the previous result, this proportion can be calculated as:

$$\frac{VAR\left(W_j^{'}X\right)}{\sum_{i=1}^{p} VAR\left(W_i^{'}X\right)} = \frac{\lambda_j}{\sum_{i=1}^{p} \lambda_i} \ldots (4.4)$$

In the standard calculation of PCA, the covariance matrix $\sum$ is estimated using the sample variance and covariance formulas:

$$S = \frac{1}{n-1} X'X \ldots (4.5)$$

In equation (4.5), X is a matrix with each element derived from the original x_j variables after subtracting their sample means: $x_{ij} = \left(x_{ij} - \bar{x}_j\right)$. Given that the x_j variables could have different scales or measurement units, a common procedure is to work with the standardized data (zero mean and unit variance), which leads to an optimization problem analogous to equation (4.2) but based on the correlation matrix of the data (Jolliffe, 2002; Kolenikov and Angeles, 2009).

Principal components analysis (PCA) was not originally designed to handle discrete variables, but continuous (and normally-distributed) data. Polychoric correlations have been proposed as an alternative to calculate the correlation matrix necessary to perform PCA on discrete asset data (Kolenikov and Angeles, 2009). Consider ordinal variables y_k with d_k categories derived from an underlying continuous variable y_k* using a set of thresholds $\gamma_1^k, ..., \gamma_{d-1}^k$, such that (Drasgow, 2006; Kolenikov and Angeles, 2009):

$$y_k = y_{ki} \text{ } if \text{ } \gamma_{i-1}^k \leq y_k{*} < \gamma_i^k \ldots (4.6)$$

In equation (4.6), i=1, 2, …, d_k and the first and last thresholds are defined as $\gamma_0^k = -\infty$ and $\gamma_d^k = \infty$, respectively. Furthermore, assuming underlying continuous variables y_1* and y_2* that are distributed following a bivariate normal distribution, the probability of an observation (y_{1i}, y_{2j}) is given by (Olsson, 1979; Drasgow, 2006; Holgado-Tello *et al.*, 2010):

$$Pr(y_{1i}, y_{2j}) = \int_{\gamma_{i-1}^1}^{\gamma_i^1} \int_{\gamma_{j-1}^2}^{\gamma_j^2} \phi(y_1, y_2; \rho) dy_1 dy_2 \ldots (4.7)$$

In equation (4.7), $\phi(Y_1, Y_2; \rho)$is the bivariate normal distribution function with correlation ρ. Based on this probability, the likelihood function for a sample with n_{ij} observations of values (y_{1i}, y_{2j}) can be defined as:

$$L = \prod_{i=1}^{d_1} \prod_{j=1}^{d_2} \Pr(y_{1i}, y_{2j})^{n_{ij}} \ldots (4.8a)$$

$$\ln L = \sum_{i=1}^{d_1} \sum_{j=1}^{d_2} n_{ij} \ln \Pr(y_{1i}, y_{2j}) \ldots (4.8b)$$

In equations (4.8a) and (4.8b), d_1 and d_2 are the number of categories for y_1 and y_2. The likelihood function is then maximized with respect to ρ and the set of thresholds $(\gamma_1^1, ..., \gamma_{d-1}^1; \gamma_1^2, ..., \gamma_{d-1}^2)$ to obtain maximum likelihood estimates of the model parameters (Drasgow, 2006). The solution for ρ from the resulting system of equations is then used to create each element of the correlation matrix for PCA. Therefore, polychoric correlations yield an estimate derived from the underlying (unobserved) continuous variables, which is argued to be the "true" measurement of the correlation structure. In practical terms, standard correlation methods (such as the sample variance and covariance) seem to produce smaller-sized estimates, when applied to discrete data, relative to polychoric correlations (Kolenikov and Angeles, 2009).

4.3.2. Data

The analysis in this study was performed using selected census samples from the IPUMS-International project. The data from IPUMS-International offers the main advantage that most of the variables necessary for the analysis have been previously harmonized, which produces more comparable results. The selected census samples are shown in Table 4.1. The main criterion for the selection of datasets was the availability of income data, given only a small proportion of censuses collect this information.[25] Further, I excluded census samples that did not have data for all the outcomes of interest for this study, particularly those without questions on children ever born and on children

[25] In particular, only 36 of the 277 census samples available (at the time this essay was written) through the IPUMS-International project included income data. However, given that the asset-based approach is primarily relevant for developing countries, then Canada, Puerto Rico, and the United States were excluded from the possible datasets for the analysis.

surviving. Three additional census datasets were included to have more variability in the total number of assets available, which include Cambodia 1998 (only 6 items), Colombia 2005 (35 items), and Peru 1993 (30 items).

The countries represented in Table 4.1 are mainly from Latin America (Brazil, Colombia, Dominican Republic, Mexico, Panama, and Peru), while two census samples correspond to other regions (Cambodia and South Africa). The datasets are 10 percent samples of the corresponding censuses, except for Mexico 1970 (1 percent), Brazil 1970 (5 percent), and Brazil 2000 (6 percent). Therefore, given the large proportion of the population included, all the data are nationally representative and are also representative of lower geographical units.

Table 4.1: Asset Variables Available for Selected Census Samples

Census sample	Variables available	Type of variables available					Household income
		Housing characteristics	Assets	Binary	Ordinal	Count	
Brazil 2000	21	11	10	8	7	6	Yes
Brazil 2010	20	11	9	10	8	2	Yes
Cambodia 1998	6	6	0	1	4	1	No
Colombia 1973	13	13	0	2	9	2	Yes
Colombia 2005	35	18	17	18	10	7	No
Dominican Republic 2002	26	13	13	14	10	2	Yes
Mexico 1970	13	11	2	8	4	1	Yes
Mexico 2000	24	14	10	12	10	2	Yes
Panama 1980	19	13	6	7	10	2	Yes
Panama 2010	26	14	12	7	10	9	Yes
Peru 1993	30	11	19	20	9	1	No
South Africa 1996	10	9	1	1	8	1	Yes

Data source: Integrated Public Use Microdata Series (IPUMS) - International.

The set of selected samples include varied information on housing characteristics and asset ownership at the household level (Table 4.1). The detailed list of variables available in each census sample is shown in Table 4.A1 in the Appendix to this chapter.

The datasets have a wide number of variables available, ranging from 6 for Cambodia 1998 to 35 for Colombia 2005. The vast majority of data are discrete, given about half of the variables are binary and a slightly smaller proportion are ordinal. The specific variables included depend on the dataset, but cover aspects such as the predominant construction materials of the roof, dwelling ownership, type of water source, number of rooms in the dwelling, ownership of diverse durable assets, among others.

Household income information was collected in nine of the selected census samples datasets. The availability of income data allows me to compare the indices based on housing characteristics and assets against a more traditional measure of household socioeconomic status. Income is not a perfect measure of socioeconomic status because of issues such as measurement error (given the complexity of data collection) and year-to-year variability (Montgomery *et al.*, 2000; Bollen *et al.*, 2002). However, acknowledging these possible limitations, the income data provide an additional criterion in this analysis, to examine the relative performance of the asset-based indices. Income is adjusted by the number of household members to obtain a per capita variable, and it is also log-transformed due to its highly skewed distribution.

4.3.3. How are the indices defined?

Data on housing characteristics and assets are summarized to define a proxy measure of socioeconomic status. In general, this household measure is defined as:

$$WI_i = a_1 \cdot x_{1i} + a_2 \cdot x_{2i} + \ldots + a_n \cdot x_{ni} \quad \ldots \ (4.9)$$

In equation (4.9), WI_i is the index for household i, x_j are variables representing housing characteristics and assets, and a_j is the weight assigned to variable x_j. All the variables available in the data, which were described in Table 4.1, are used for the construction of the indices.

I tested four alternative aggregation methods that have been proposed in the literature (Filmer and Pritchett, 2001; Kolenikov and Angeles, 2009). Two methods work with the original unrecoded data (including binary, ordinal, and count variables), while the other two only use binary indicators:

(i) PCA based on the standard calculation of correlations and applied to the unrecoded (discrete) asset data, which is called "ordinal PCA" in the rest of the chapter.

(ii) PCA based on polychoric correlations on the ordinal data or "polychoric PCA."

(iii) PCA on the recoded data, where all ordinal variables have been transformed into binary indicators, and using the standard calculation of the correlation matrix. This is the common method performed in previous studies and that is identified in this paper as "binary PCA."

(iv) A count index, as a simple aggregation procedure to compare against the more sophisticated PCA-based methods. For this last index, each household variable receives a weight equal to 1, such that we count the number of asset characteristics that the household "owns."[26] The count index was defined only

[26] Certain variables available in the data registered the number of assets of a certain type. For example, the Brazil 2000 census sample included not only a binary response (yes/no) for televisions, cars, and air conditioning units, but also the number of such assets owned by the household. In these cases, the count index considered only ownership, such that

using durable asset variables (i.e. it excludes housing characteristics), so it is not fully comparable with the other methods.

The inclusion of the three PCA-based aggregation methods allows examination of changes in household rankings due to the use of polychoric correlations (by comparing the polychoric PCA against the ordinal PCA) and to the use of ordinal versus binary data (by comparing the ordinal PCA against the binary PCA). For all the PCA-based methods, only the first component is retained. The first component represents by definition the maximum variance extracted from the asset data. Therefore, I will follow the usual assumption in the literature that the first component "represents" (or is a proxy for) household socioeconomic status based on the various asset indicators used to produce it (Filmer and Scott, 2012).[27] Other components may be related to other dimensions of "household socioeconomic status," but were not used in the analysis.

The ordering of categories may be an important input to define the indices, as it conveys additional information to rank households. The asset data had an implicit ordering in most cases and it was just necessary to assign the lowest value of each variable to the "worst" option and the highest to the "best" option. An example is shown in Table 4.2 for the main construction material of the roof from the Dominican Republic 2002 census sample, including both the ordinal and binary versions of the data. Some basic data manipulation was implemented to obtain ordinal versions of categorical variables. The original ordering of categories was modified only in a few cases where certain categories were clearly misclassified in the scale; for example, if dirt was second

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multiple assets of the same kind were not double counted. The PCA-based indices do include the number of assets and not only its ownership, if this information is available.

[27] For some further discussion on the use of higher order components, see Filmer and Scott (2012).

(or penultimate) in the scale rather than in one of the ends in the flooring material categories. Furthermore, some categories were dropped from the analysis (i.e. transformed to missing) if their position in the scale was unclear, which most often happened for "other" (residual) responses. For instance, "other" was excluded from the roof main construction material for Dominican Republic 2002, as it was unclear how to rank the construction material represented by this option with respect to the rest of the alternatives. In addition, a few categories were combined if they seemingly represented similar positions in a scale. For example, "palm leaves" and "cane" were combined for the roof main construction material for Dominican Republic 2002, given they did not appear to be qualitatively different, as shown in Table 4.2. In order to have comparable information to define the indices using the alternative aggregation methods, any categories that were dropped or combined were treated the same way for the dichotomized version of the data.

Table 4.2: Data Recoding, Main Construction Material of Roof, Dominican Republic 2002

Original variable	Ordinal	Binary
0 = Unoccupied households	*(dropped)*	*(dropped)*
1 = Concrete	4 = Concrete	1 = Concrete, 0 = No
2 = Zinc	3 = Zinc	1 = Zinc, 0 = No
3 = Asbestos	2 = Asbestos	1 = Asbestos, 0 = No
4 = Palm leaves	1 = Palm leaves or cane	1 = Palm leaves or cane, 0 = No
5 = Cane		
6 = Other	*(dropped)*	*(dropped)*

Data source: Integrated Public Use Microdata Series (IPUMS) - International.

Based on these aggregation methods, I examine in the next section whether they produce similar household rankings and whether there are differences in their relation to selected education, fertility, and mortality outcomes. Comparisons with income data are included throughout the analysis. Household rankings produced by each measure are compared through Spearman rank correlations and cross-classification by wealth quintiles. The relationships with the selected outcomes were observed both through differences across wealth index quintiles and in regression analysis using each index as a control variable. The objective is to test whether the direction and strength of the relation of each proxy for socioeconomic status are consistent across diverse outcomes that are typically examined in social and economic research. The specific outcomes analyzed include school enrollment, literacy, completion of primary school, completion of secondary school, having any children, and having experienced any child death. For higher comparability of results, the same set of controls (i.e. available across datasets) was used for regressions estimated for the same outcome. The implicit assumption behind the proposed analysis is that a stronger relation with the outcome is considered to show a better performance of the socioeconomic status measure (Bollen *et al.*, 2007; Kolenikov and Angeles, 2009). This assumption is also followed by other research in this area. The analysis does not intend to estimate a causal relationship between socioeconomic status and the selected outcomes, but rather verify that they are indeed correlated. In the study by Kolenikov and Angeles (2009) this is distinguished as a "weaker requirement of internal validity" from the causal effect that needs to be verified for external validity of the measure (p. 159).

4.4. Results

4.4.1. Calculation of weights using PCA

The data on housing characteristics and asset ownership were used to define the proposed indices, where three of them had weights calculated by applying PCA. For the PCA-based indices, a first step was to examine the sign and size of the weights assigned to each item. Kolenikov and Angeles (2009) refer to a "natural ordering" of categories that follow a monotone relation among them in the case of ordinal variables. For example, it is expected that dirt be the worst and concrete be the best floor materials and that they would be assigned the lowest and highest value in the correspondent ordinal variable, while other flooring materials would be assigned a number within that range. If this ordering is meaningful in terms of the relation between the asset variable and (unobserved) household socioeconomic status, then weights assigned by PCA are expected to follow the ordering, assigning larger positive values to the most "desirable" household characteristics and larger negative values to the least "desirable" ones. Nevertheless, it should also be noted that in the case of binary PCA, weights tend to be larger for assets more unequally distributed across households (McKenzie, 2005; Vyas and Kumaranayake, 2006). That is, assets that are owned by all or by very few households will receive relatively smaller weights, as they do not vary much across households (and PCA is defined from the variability of the data).

An example of weights obtained by the three PCA-based methods is shown in Table 4.3 for the main flooring materials in the Colombia 2005 census sample. Categories in the table are ordered from best to worst. In the example, the weights assigned by the polychoric PCA method follow the ordering of categories, where carpet flooring ("best" option) is assigned the largest positive weight, dirt or sand ("worst" option) are assigned the largest negative weight, and the rest of categories receive intermediate values as weights. The binary PCA weights do not satisfy this monotonicity condition, given these are not strictly increasing as we move from the "worst" to the "best" flooring material. In particular, the weight assigned to tile flooring (0.153) is larger than the corresponding to carpet (0.064), while a similar issue occurs for cement or gravel flooring (-0.070) with respect to rough wood (-0.051). The disagreement between the implicit ordering of the variable and the weights calculated by the binary PCA seems to be driven by the larger frequencies associated to tile and cement (reported by about 30-40 percent of households) with respect to carpet and rough wood (reported only by 5-7 percent of households). In general, weights assigned by binary PCA to other ordinal variables do not necessarily follow their order of categories, similarly to this example.

Table 4.3: Weights for Flooring Material by Alternative Aggregation Methods, Colombia 2005 Census Sample

Main flooring material (best to worst)	Household proportion (%)	PCA aggregation method		
		Binary	Ordinal	Polychoric
Carpet, marble, parquet, or polished wood	6.8	0.064		0.388
Tile, vinyl, clay tile, or brick	44.8	0.153		0.119
Cement or gravel	33.9	-0.070	0.247	-0.103
Rough wood, board, plywood, or other vegetable	4.4	-0.051		-0.235
Dirt or sand	10.2	-0.166		-0.352

Data source: Integrated Public Use Microdata Series (IPUMS) - International.

The PCA based methods can also be compared by examining the proportion of variance of the data explained by each of them. The proportion of variance explained by the first principal component is calculated as the ratio of the first eigenvalue to the sum of all eigenvalues from the variance-covariance matrix, similarly for any of the PCA-based measures. Results are included in Table 4.4. As we observe, this criterion shows that the indices based on the calculation of polychoric correlations explain a larger proportion of the overall data variability, followed by ordinal PCA, and lastly by the indices based on dichotomized variables. The differences in proportion of variance explained by each of the indices are consistent and large across all datasets analyzed. For example, while the asset-based wealth index using binary PCA only explains 16 percent of the data variability for Peru 1993, PCA on ordinal variables achieves 26 percent, and the polychoric PCA 46 percent of the overall variability. In the bottom row of the table, on average, about 18 percent of variance is explained by the first component of binary PCA, 33 percent for ordinal PCA, and 49 percent for polychoric PCA. This evidence is consistent with previous findings by Kolenikov and Angeles (2009).

Table 4.4: Proportion of Variance Explained (%) by the First Principal Component, for Alternative Aggregation Methods

Census sample	PCA aggregation method		
	Binary	Ordinal	Polychoric
Brazil 2000	17.83	29.05	44.76
Brazil 2010	12.93	23.36	42.18
Cambodia 1998	20.15	35.45	45.81
Colombia 1973	15.62	37.46	45.30
Colombia 2005	15.51	27.18	45.14
Dominican Republic 2002	15.12	25.07	46.69
Mexico 1970	29.70	42.64	57.10
Mexico 2000	17.32	32.86	51.34
Panama 1980	17.68	35.16	49.51
Panama 2010	13.59	29.64	49.61
Peru 1993	15.56	25.88	46.15
South Africa 1996	23.29	52.28	60.91
Simple average	***17.86***	***33.00***	***48.71***

Data source: Integrated Public Use Microdata Series (IPUMS) - International.

4.4.2. Household Rankings

The household classifications resulting from each of the wealth indices can be compared to assess the consistency of rankings. As a first step to compare rankings, I calculated Spearman rank correlations between indices, both with respect to the index based on polychoric correlations (Table 4.5a) and to the logarithm of income per capita (Table 4.5b). The correlations show very high correspondence of the polychoric PCA index with other asset-based wealth indices (correlations larger than 0.9 in most cases), being the largest for the ordinal PCA and binary PCA, followed by the asset count index. Similarly, we observe a high congruence of rankings (around 0.5 to 0.6) if we compare any of these alternative aggregation methods against rankings based on the logarithm of income per capita. This evidence suggests not only that household rankings are (reasonably) similar to rankings based on the (logarithm of) household income per capita, but also that these rankings are highly consistent across the alternative aggregation methods for the asset-based measures analyzed.

Table 4.5a: Correlation Coefficients between Wealth Indices and Polychoric PCA Index

	Asset count	Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	0.920	0.977	0.998	1.000	0.716
Brazil 2010	0.821	0.956	0.990	1.000	0.594
Cambodia 1998	NA	0.886	0.983	1.000	NA
Colombia 1973	NA	0.983	0.998	1.000	0.531
Colombia 2005	0.858	0.981	0.997	1.000	NA
Dominican Republic 2002	0.901	0.975	0.999	1.000	0.487
Mexico 1970	0.714	0.989	0.999	1.000	0.555
Mexico 2000	0.930	0.988	0.997	1.000	0.628
Panama 1980	0.830	0.980	0.997	1.000	0.659
Panama 2010	0.910	0.973	0.995	1.000	0.375
Peru 1993	0.900	0.987	0.998	1.000	NA
South Africa 1996	NA	0.989	0.997	1.000	0.615
Simple average	***0.865***	***0.972***	***0.996***	***1.000***	***0.573***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

Table 4.5b: Correlation Coefficients between Wealth Indices and Log of Income per Capita

	Asset count	Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	0.677	0.693	0.721	0.716	1.000
Brazil 2010	0.523	0.524	0.567	0.594	1.000
Cambodia 1998	NA	NA	NA	NA	NA
Colombia 1973	NA	0.521	0.530	0.531	1.000
Colombia 2005	NA	NA	NA	NA	NA
Dominican Republic 2002	0.412	0.485	0.485	0.487	1.000
Mexico 1970	0.420	0.553	0.554	0.555	1.000
Mexico 2000	0.583	0.635	0.621	0.628	1.000
Panama 1980	0.518	0.649	0.665	0.659	1.000
Panama 2010	0.323	0.364	0.371	0.375	1.000
Peru 1993	NA	NA	NA	NA	NA
South Africa 1996	NA	0.606	0.605	0.615	1.000
Simple average	***0.494***	***0.559***	***0.569***	***0.573***	***1.000***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

Households were classified into quintiles based on the wealth index measures to further examine consistency of household rankings. In order to compare classifications, I calculated the proportion of households that were classified in the same, in a higher, or in a lower quintile across pairs of indices. The detailed classifications by quintile are available, but are not presented here. In particular, this procedure was performed to compare the classifications based on the polychoric PCA index and the log of income per capita against other measures. Results for the proportion of households classified in the same quintile are shown below in Tables 4.6a and 4.6b, while those for lower and higher quintiles are included in the Appendix to this chapter.

The discrepancies in classifications by quintiles reveal larger differences than the Spearman rank correlation coefficients, but still a sizable overlap. Household classification by quintiles is highly consistent between the two PCA methods that use ordinal variables (polychoric and ordinal), relative to other measures. In fact, more than 90 percent of total households are classified in the same wealth quintile if we use the original unrecoded data (ordinal variables) for any of the census samples examined, disregarding the method applied to calculate the correlations matrix for PCA. The household classification by quintiles for the polychoric PCA also has a relatively large overlap with binary PCA for most of the datasets analyzed (around 75-85 percent), except for Cambodia 1998, which may be explained by the limited number of variables available for this census sample. The correspondence of wealth quintiles based on the polychoric PCA index is smaller with the asset count index (55 percent of households were classified in the same quintile on average) and the logarithm of income per capita (40 percent of households were classified in the same quintile on average).

In terms of cross-classifications into lower or higher quintiles with respect to the polychoric PCA index (Tables 4.A2a and 4.A3a in the appendix to this chapter), results show that a similar proportion of households move up or down for other measures, except for the asset count. For example, if we compare quintiles based on the polychoric PCA index for Brazil 2000, about 2.5 percent of households are classified in a higher or lower quintile for the ordinal PCA index, 9 percent for the binary PCA index, and 28 percent for the log of income per capita. However, 35.9 percent of households are relatively less "wealthy" using the asset count index but only 6.9 percent of households appear to be "wealthier" in this case. The larger discrepancies in classification into lower quintiles are mainly explained by the relatively lower variability of the count index, given this method only produces discrete values that range between zero and the total number of asset variables available in the dataset. This lower variability implies that a larger number of households are assigned the same number for the index, thus creating issues to calculate the cutoff points to define the quintiles.

Table 4.6a: Comparison of Classification by Quintiles
Households Classified in the Same Quintile with respect to Polychoric PCA Index (%)

	Asset count	Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	57.2	82.4	94.7	100.0	43.3
Brazil 2010	50.1	74.9	95.2	100.0	38.5
Cambodia 1998	NA	32.9	90.5	100.0	NA
Colombia 1973	NA	79.5	94.6	100.0	36.2
Colombia 2005	59.9	84.7	96.2	100.0	NA
Dominican Republic 2002	56.9	78.4	95.2	100.0	34.3
Mexico 1970	35.1	83.7	96.1	100.0	38.3
Mexico 2000	67.8	82.3	94.5	100.0	42.5
Panama 1980	53.4	79.2	93.1	100.0	41.7
Panama 2010	60.5	79.8	93.4	100.0	38.4
Peru 1993	49.0	83.3	96.0	100.0	NA
South Africa 1996	NA	85.0	92.4	100.0	38.4
Simple average	***54.4***	***77.2***	***94.3***	***100.0***	***39.1***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

The overlap in classifications by quintiles in Table 4.6b show that all indices based on housing characteristics and assets have a relatively similar performance when compared to quintiles based on the logarithm of household income per capita. In all cases, about 35 to 45 percent of households coincide in the same wealth quintile defined by income, even following the simple approach of counting household durable assets. Furthermore, analogous to previous results, the asset count tends to classify households more often into lower quintiles than higher quintiles when compared to income, as it can be observed in Tables 4.A2b and 4.A3b in the Appendix to this chapter.

Overall, the consistency of household classification by quintiles is more moderate (but sizable) between income and all the indices based on housing characteristics. The existing discrepancies could be explained by the concept of household wealth that each measure is capturing: while housing characteristics and assets reflect accumulation of material well-being for a household (a stock), income reflects monetary gains from household members over some specified period of time (a flow). Additionally, it is possible that income is a noisy measure due to measurement error and year-to-year fluctuations, issues that are not expected to affect asset information. Nevertheless, we do not observe a relatively better or worse performance by any of the material wealth measures based on alternative aggregation procedures when compared to income per capita. All of them classify about the same proportion of households into the same wealth quintile as income.

Table 4.6b: Comparison of Classification by Quintiles
Households Classified in the Same Quintile with respect to Log of Income per Capita (%)

	Asset count	Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	40.4	42.8	43.9	43.3	100.0
Brazil 2010	35.7	36.8	38.4	38.5	100.0
Cambodia 1998	NA	NA	NA	NA	NA
Colombia 1973	NA	35.9	36.1	36.2	100.0
Colombia 2005	NA	NA	NA	NA	NA
Dominican Republic 2002	32.5	34.3	34.3	34.3	100.0
Mexico 1970	28.3	38.6	38.3	38.3	100.0
Mexico 2000	40.9	42.7	42.4	42.5	100.0
Panama 1980	35.4	41.0	42.3	41.7	100.0
Panama 2010	35.5	38.4	38.5	38.4	100.0
Peru 1993	NA	NA	NA	NA	NA
South Africa 1996	NA	39.4	38.7	38.4	100.0
Simple average	***35.5***	***38.9***	***39.2***	***39.1***	***100.0***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

Finally, graphical analysis of kernel density estimates for each of the asset-based indices was used to identify whether there are any differences along the resulting distribution of wealth (see Figure 4.A1 in the Appendix to this chapter). The purpose is to visually inspect the resulting distributions based on the alternative aggregation methods applied. In order to have comparable scales, all indices were standardized (i.e. they have zero mean and unit variance). Even though alternative aggregation procedures were implemented to calculate the variable weights, all the PCA-based indices appear to produce very similar distributions, either using the dichotomized or the original unrecoded data. This result was surprising given that the binary PCA works with transformed (dichotomized) data. For these three indices, the only noticeable discrepancy happens for Cambodia 1998, where there seems to be larger disagreement in left tail of the distribution, possibly explained by the small number of variables available for this census sample.

The most important differences correspond to the comparisons between the asset count indices against the PCA-based indices. For all the graphs, as it would be expected, the density mass is concentrated around a more limited set of values for the asset count (which produces a less smooth distribution). As previously discussed, this distribution can be explained by the definition of the index, which has discrete values that range from zero up to the maximum number of assets available in the data. For example, the Mexico 1970 sample has only two asset variables available to define the asset count; therefore, the resulting index could have only values of zero, one, or two. The three possible values can be clearly observed in the kernel density estimate for this dataset. However, the asset count does produce a comparable (but not smooth) distribution with respect to the PCA-based indices in the case of Panama 2010. The reason for these similarities is that PCA-based methods assigned weights of similar size to many items included in this dataset, such that the equal (unit) weights used in the count index resemble the weights produced by PCA. For instance, ordinal PCA assigned weights of 0.231 to the type of lighting, 0.224 to the main method of garbage disposal, 0.207 to the fuel used for cooking, 0.204 to ownership of a stove, and so forth.

4.4.3. How are outcomes changing across the indices?

Measures of socioeconomic status are often used in economic research to examine differences across outcomes of interest and as a control in regression analysis. In this sub-section, I examine the extent to which the proposed indices produce similar-sized differences across quintiles of household wealth and have comparable coefficients in regression analysis for a set of six selected education, fertility, and mortality outcomes. The selected outcomes are some of those frequently analyzed in social and economic research.

The differences for school enrollment (for children ages 7 to 14) by quintile of wealth for each of the alternative proxies of socioeconomic status are shown in Table 4.7a. As expected, across almost all datasets and measures, I obtained increasing proportions of children enrolled in school when moving from the bottom ("poorest") to the top ("richest") quintile. The detailed proportions of children enrolled in school are not shown, but are available upon request. In Table 4.7a, I present a summary figure: the average difference in school enrollment across quintiles, which was calculated as the difference between the top and bottom quintile divided by four.[28] This number represents the average change in the outcome of interest when we move along wealth quintiles. In Table 4.7a, the differences in school enrollment across quintiles are strikingly similar for

[28] The average difference presented is equivalent to calculating the difference between second and bottom, third and second, fourth and third, and top and fourth quintile, and averaging across these numbers.

all the PCA-based measures. For example, the change in school enrollment is about 2.8 percent across quintiles for Mexico 2000 and 3.6 percent for Panama 1980 for the binary PCA, ordinal PCA, or polychoric PCA. If these measures are compared to quintiles based on the logarithm of income per capita, the numbers for the average difference across quintiles are also similar but are consistently smaller than the rest. Across datasets and measures, the largest differences in school enrollment between quintiles correspond to the two census samples from the 1970s (Mexico 1970 and Colombia 1973), which also coincide with the two lowest average school enrollment rates.

Table 4.7a: Average Change in School Enrollment (children ages 7-14) across Wealth Quintiles, for Alternative Aggregation Methods [1/]

	Mean	Average difference across quintiles			
		Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	94.5	2.29	2.33	2.34	1.75
Brazil 2010	96.9	0.59	0.64	0.64	0.43
Cambodia 1998	64.5	5.66	4.53	4.17	NA
Colombia 1973	62.2	11.49	11.42	11.30	8.10
Colombia 2005	90.8	3.96	4.02	4.10	NA
Dominican Republic 2002	87.4	0.79	0.85	0.88	0.20
Mexico 1970	69.4	7.93	8.11	8.10	6.67
Mexico 2000	92.8	2.84	2.88	2.87	1.87
Panama 1980	87.7	3.56	3.54	3.59	3.06
Panama 2010	97.0	1.00	1.02	1.00	0.09
Peru 1993	87.0	3.36	3.43	3.42	NA
South Africa 1996	88.9	3.15	3.20	3.19	2.45
Simple average	***84.9***	***3.89***	***3.83***	***3.80***	***2.73***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

1. The average difference is calculated as the difference between top and bottom quintiles divided by four.

The differences in school enrollment by household socioeconomic status are also examined through regressions controlling for each of the measures based on the alternative aggregation methods. The regression results for school enrollment are shown in Table 4.7b. The table shows the odds-ratio for each census sample and socioeconomic status measure. As it is expected, almost all measures show a positive (odds-ratio larger than one) and a statistically significant coefficient for SES in explaining school enrollment. The only notable exception is observed for the log of income per capita for Panama 2010, which is marginally statistically significant and implies a negative effect on school enrollment.

The coefficients for the asset count and the three PCA-based methods are very similar in size, and all of them are larger than the effect of the logarithm of income per capita. If we further examine the (small) differences in the effects across samples, none of the measures is consistently larger than the rest. However, it is possible to identify a pattern based on the (small) differences in coefficients: the polychoric PCA tends to have the first or second largest coefficient (for ten out of twelve samples), followed by the ordinal PCA (for six out of twelve samples). But differences are small on average.

Table 4.7b: Logit model for School Enrollment (Children Ages 7-14)
Wealth Index Coefficient (Odds-ratio) 1/

	Asset count	Binary PCA	Polychoric PCA	Ordinal PCA	Log income per capita
Brazil 2000	**2.224**	**2.000**	**2.184**	**2.100**	**1.444**
	0.088	*0.059*	*0.074*	*0.060*	*0.034*
Brazil 2010	**1.405**	**1.378**	**1.434**	**1.393**	**1.088**
	0.050	*0.033*	*0.039*	*0.030*	*0.024*
Cambodia 1998	**NA**	**1.489**	**1.363**	**1.442**	**NA**
		0.040	*0.033*	*0.031*	
Colombia 1973	**NA**	**1.883**	**1.792**	**1.804**	**1.249**
		0.030	*0.026*	*0.026*	*0.025*
Colombia 2005	**2.078**	**1.927**	**2.059**	**1.963**	**NA**
	0.044	*0.096*	*0.097*	*0.089*	
Dominican Republic 2002	**1.131**	**1.126**	**1.142**	**1.145**	**1.012** #
	0.023	*0.025*	*0.028*	*0.029*	*0.018*
Mexico 1970	**1.314**	**1.543**	**1.553**	**1.558**	**1.321**
	0.037	*0.073*	*0.071*	*0.071*	*0.040*
Mexico 2000	**1.780**	**1.775**	**1.820**	**1.774**	**1.123**
	0.093	*0.105*	*0.099*	*0.088*	*0.034*
Panama 1980	**1.647**	**1.816**	**1.863**	**1.842**	**1.283**
	0.071	*0.094*	*0.090*	*0.096*	*0.055*
Panama 2010	**1.709**	**1.552**	**1.689**	**1.611**	**0.943** ##
	0.078	*0.075*	*0.091*	*0.079*	*0.027*
Peru 1993	**1.330**	**1.337**	**1.399**	**1.386**	**NA**
	0.074	*0.062*	*0.093*	*0.101*	
South Africa 1996	**NA**	**1.786**	**1.775**	**1.729**	**1.300**
		0.046	*0.046*	*0.046*	*0.019*

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

1. All estimates are statistically significant at the 1 percent level unless otherwise noted (### p>.01, ## p>.05, # p>.10). The table includes odds-ratio coefficients in bold and clustered standard errors in italic. Standard errors are clustered using mesoregions for Brazil, districts for Cambodia and Panama, municipalities for Colombia and Dominican Republic, states for Mexico, provinces for Peru, and magisterial districts for South Africa. The estimation sample is restricted to persons 7 to 14 years old that are not household heads.

Control variables: sex, age, and age-squared of the child; sex, age, age-squared, and educational attainment of household head (dummies for primary, secondary, and university); urban/rural.

The differences across wealth quintiles were also measured for child mortality for women who ever gave birth, aged between 18 and 30 years old at the time of data collection. The information on child mortality was not directly available in the data, but it was approximated using children ever born and children surviving. Therefore, it does not refer to deaths of children within certain ages as it is typically reported (under one or under five years old), but to any child death implicitly declared in the data.

In almost all cases, I obtained decreasing proportions of child deaths when moving from the bottom ("poorest") to the top ("richest") quintile. That is, child mortality decreases with higher household socioeconomic status, as one would expect. The detailed results by quintile are not shown, but are also available upon request. In Table 4.8a, I present the same summary figure calculated before: the average change in child mortality across quintiles. The evidence is similar to results for school enrollment: the average change across quintiles is highly similar for all the PCA-based measures and it is larger (in absolute value) than the corresponding number for the logarithm of income per capita. For instance, the decrease in the child mortality rate across wealth quintiles using asset-based indices is about 4.6 percent for Colombia 1973 and 2.6 percent for South Africa 1996, while slightly smaller numbers (in absolute value) are found for quintiles based on the logarithm of income per capita.

Table 4.8a: Average Change in Child Mortality (Women Who Ever Gave Birth, Ages 18-30) across Wealth Quintiles, for Alternative Aggregation Methods [1/]

	Mean	Average difference across quintiles			
		Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	5.1	-1.67	-1.70	-1.72	-1.29
Brazil 2010	2.4	-0.60	-0.62	-0.62	-0.46
Cambodia 1998	16.5	-2.10	-1.87	-1.70	NA
Colombia 1973	17.5	-4.61	-4.64	-4.61	-4.11
Colombia 2005	2.8	-0.83	-0.83	-0.86	NA
Dominican Republic 2002	11.6	-0.78	-0.85	-0.84	-0.58
Mexico 1970	NA	NA	NA	NA	NA
Mexico 2000	6.6	-2.06	-2.04	-2.06	-1.64
Panama 1980	6.0	-1.50	-1.45	-1.44	-1.43
Panama 2010	3.3	-1.02	-1.02	-1.04	-0.34
Peru 1993	11.0	-4.46	-4.33	-4.36	NA
South Africa 1996	9.5	-2.57	-2.58	-2.58	-1.83
Simple average	***8.4***	***-2.02***	***-1.99***	***-1.98***	***-1.46***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

1. The average difference is calculated as the difference between top and bottom quintiles divided by four. Child mortality is derived from children ever born and children surviving, as declared by women who ever gave birth.

The regressions for child mortality are shown in Table 4.8b. The coefficients for household socioeconomic status as a control in a regression to explain child mortality appear to be statistically significant and consistently negative (except for only one statistically insignificant coefficient for the log of income per capita for Panama 2010). Similarly to school enrollment, the coefficients of all the asset-based indices are of similar size and tend to be stronger than the corresponding coefficient for the logarithm of income per capita. None of the measures is consistently larger than the rest if we examine the small differences across measures. The polychoric and ordinal PCA are again generally the first or second largest (negative) coefficients across samples, but differences are small on average.

Table 4.8b: Logit Model for Child Mortality (Women Who Ever Gave Birth, Ages 18-30) Wealth Index Coefficient (Odds-ratio) 1/

	Asset count	Binary PCA	Polychoric PCA	Ordinal PCA	Log income per capita
Brazil 2000	**0.568**	**0.579**	**0.541**	**0.555**	**0.676**
	0.015	*0.011*	*0.012*	*0.012*	*0.012*
Brazil 2010	**0.739**	**0.777**	**0.731**	**0.762**	**0.816**
	0.010	*0.013*	*0.012*	*0.012*	*0.012*
Cambodia 1998	**NA**	**0.747**	**0.797**	**0.769**	**NA**
		0.013	*0.017*	*0.016*	
Colombia 1973	**NA**	**0.770**	**0.763**	**0.757**	**0.826**
		0.022	*0.018*	*0.018*	*0.019*
Colombia 2005	**0.754**	**0.716**	**0.680**	**0.706**	**NA**
	0.017	*0.016*	*0.016*	*0.016*	
Dominican Republic 2002	**0.906**	**0.921**	**0.900**	**0.903**	**0.919**
	0.022	*0.029*	*0.027*	*0.028*	*0.025*
Mexico 1970	**NA**	**NA**	**NA**	**NA**	**NA**
Mexico 2000	**0.718**	**0.693**	**0.696**	**0.708**	**0.864**
	0.010	*0.011*	*0.010*	*0.010*	*0.010*
Panama 1980	**0.884**	**0.780**	**0.780**	**0.785**	**0.863**
	0.039	*0.059*	*0.062*	*0.061*	*0.036*
Panama 2010	**0.606**	**0.574**	**0.548**	**0.571**	**1.005 #**
	0.051	*0.042*	*0.048*	*0.571*	*0.030*
Peru 1993	**0.781**	**0.594**	**0.623**	**0.611**	**NA**
	0.031	*0.011*	*0.012*	*0.012*	
South Africa 1996	**NA**	**0.614**	**0.616**	**0.628**	**0.815**
		0.014	*0.015*	*0.015*	*0.014*

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

1. All estimates are statistically significant at the 1 percent level unless otherwise noted (### p>.01, ## p>.05, # p>.10). The table includes odds-ratio coefficients in bold and clustered standard errors in italic. Standard errors are clustered using mesoregions for Brazil, districts for Cambodia and Panama, municipalities for Colombia and Dominican Republic, states for Mexico, provinces for Peru, and magisterial districts for South Africa. The estimation sample is restricted to women who ever gave birth between 18 to 30 years old.

Control variables: age and age-squared, marital status, educational attainment (dummies for primary, secondary, and university), family size, and urban/rural.

The analysis was applied to four other outcomes: motherhood (having any children), primary school completion, secondary school completion, and literacy, all for persons between 18 and 30 years old. Results are shown in Tables 4.A4a to 4.A7b in the Appendix to this chapter. The conclusions are qualitatively similar to the previous discussion for school enrollment and child mortality. The selected outcomes do change across the four measures of household socioeconomic status based on housing characteristics and assets, as it was expected. The estimated regression coefficients tend to be slightly larger for the asset-based measures than the logarithm of income per capita, but all are aligned in direction (positive or negative). The sign of the coefficients follows the hypothesized relation of the outcomes with household socioeconomic status (positive for the education ones and negative for having any children) and the largest coefficients are found for secondary school completion (the outcome with the largest variability in average rates across countries).

The (small) differences in coefficients between the four indices produced using the alternative aggregation methods also reveal patterns similar to those previously identified. The size of the coefficients is most comparable for the three PCA-based measures and the asset count. Even though none of the indices has consistently larger effects across samples and outcomes, the polychoric and the ordinal PCA are either the largest or second largest (or both) coefficients in almost all cases for any specific outcome. But differences in coefficients for the PCA-based measures are rather small in general.

4.5. Discussion

The construction of indices based on housing characteristics and asset ownership has been widely used when other measures of socioeconomic status are not available. This approach allows one to examine differences by socioeconomic status in outcomes of interest and to control for household socioeconomic status in regression analysis when the data do not include income or expenditures. Moreover, the use of asset-based indices also offers some advantages, such as smaller reporting errors and lower relative data collection cost than income or expenditures (Kolenikov and Angeles, 2009; Filmer and Scott, 2012). Principal components analysis (PCA) on data recoded to binary indicators (Filmer and Pritchett, 2001) is one of the most frequently used procedures to construct such an index. In this study, I compared the relative performance of the commonly used binary PCA to aggregation methods that are designed to handle discrete data.

The evidence presented in this paper indicates that methods based on ordinal asset data have higher agreement in rankings, but the common procedure of PCA on dichotomized data (and even the relatively simple method of counting assets) also has reasonable agreement with the other measures. More importantly, differences in variables of interest by wealth quintiles and the estimated coefficients on indicators of wealth in regression analysis are similar in size for all asset-based measures across a wide set of education, fertility, and mortality outcomes. Even though the asset-based indices show only moderate agreement with rankings based on the logarithm of income per capita, none of the asset-based indices outperformed the rest on this aspect. Furthermore, larger differences by wealth quintiles and in regression analysis are observed for the asset-based indices with respect to income, possibly due to noise in the income variable. Below, I discuss some specific aspects on the relative performance of each aggregation method to propose more specific recommendations.

The asset count is a relatively simple method to implement, given that it requires only adding the number of positive responses for a set of items. The coefficients for socioeconomic status measured through this proxy are of similar size but tend to be slightly smaller than the PCA-based methods. However, this method produces a more limited set of values that make more difficult to do a more fine classification of households, which created problems in the definition of wealth quintiles. Furthermore, in order to include other variables and not only durable assets, it would be necessary to recode them into binary indicators. For instance, the predominant type of flooring (an ordinal variable) cannot be added directly to an item count, but it could be recoded into "good flooring materials" (including cement and finished flooring types) and "bad flooring materials" (including dirt and similar quality materials). This recoding process may require to impose additional assumptions to determine which are "good" or "bad" categories in an ordinal variable. Therefore, the use of this measure is not recommended.

The PCA-based methods showed varying results regarding the different criteria applied to assess their performance. Similar to previous studies, weights assigned through binary PCA do not show the expected monotonicity property (from "worst" to "best" option in ordinal variables), in contrast to those derived from polychoric correlations. In addition, the proportion of variance explained had considerable differences across methods, being the largest for the polychoric PCA, followed by the ordinal PCA and then by the binary PCA. Nevertheless, all three PCA-based measures had reasonable (but not identical) agreement in terms of the household rankings, while results by quintiles or in regression analysis for the selected outcomes were strikingly similar. Therefore, if the objective is to examine differences by household socioeconomic status or the use of the measure as a control variable in a regression, any of the PCA-based measures appears to have a similar performance. In this sense, despite recommendations given by previous research (Howe *et al*., 2008; Kolenikov and Angeles, 2009), results suggest a relatively similar performance of the PCA procedure on dichotomized data with respect to methods based on ordinal data. However, given the discrepancies of rankings against income, the researcher should use the asset-based measures with more caution if the objective is different, such as identifying the poor within a certain population.

Polychoric correlations are designed to handle discrete data, in contrast to the standard correlation methods applied for PCA. However, there is one minor disadvantage of polychoric correlations. Leaving aside that they are computationally more intensive than standard correlation methods, the minor problem concerns their calculation for categories with small (or zero) frequencies. As discussed in the methodology section, the polychoric correlations are calculated from a likelihood function (equations 4.7, 4.8a, and 4.8b), which requires having non-zero frequencies for the combinations of values of the categorical variables. For example, the pairwise polychoric correlation may not be defined between variables such as electricity and ownership of an electric appliance, since at least one combination of categories may have zero frequencies. This is a minor disadvantage of polychoric correlations that does not recommend against their use, but that may require some data manipulation to work around the issue. For instance, a possible solution is to add a trivial small number to the frequencies associated to the different combinations of values of the categorical variables, so that none of these combinations has small (or zero) frequencies. In the few cases found in the data used in this study, alternatively, I combined categories with small frequencies with those that appear to represent similar positions in the scale. Based on the evidence shown in this chapter on household rankings and the relation with selected outcomes, any of the PCA-based measures seem to produce similar results. However, given the possible difficulties in the calculation of polychoric correlations, a practical recommendation would be to implement either the binary PCA or the ordinal PCA. No striking differences in performance have been found between these two measures.

4.6. Appendix: Additional Tables and Figures

Table 4.A1: Detailed Asset Variables Available for Selected Census Samples

Brazil 2000	Brazil 2010	Cambodia 1998
1 Dwelling type	1 Dwelling type	1 Ownership of dwelling
2 Rooms (number)	2 Dwelling ownership	2 Lighting
3 Sleeping rooms (number)	3 External walls material	3 Cooking fuel
4 Ownership of dwelling	4 Rooms (number)	4 Toilet
5 Ownership of land	5 Bedrooms (number)	5 Water supply
6 Water source	6 Toilet	6 Rooms (number)
7 Piped water	7 Sewage	
8 Bathrooms (number)	8 Water supply	
9 Toilet	9 Piped water	
10 Waste water	10 Garbage destination	
11 Trash	11 Electricity	
12 Radio	12 Radio	
13 Refrigerator	13 TV	
14 VCR	14 Washer	
15 Washing machine	15 Refrigerator	
16 Microwave	16 Cell phone	
17 Telephone	17 Phone	
18 Computer	18 Computer	
19 TV (number)	19 Motorcycle	
20 Private car (number)	20 Auto	
21 Air conditioner (number)		

Data source: Integrated Public Use Microdata Series (IPUMS) - International.

Table 4.A1 (continued): Detailed Asset Variables Available for Selected Census Samples

Colombia 1973	Colombia 2005	Dominican Republic 2002
1 Dwelling type	1 Dwelling type	1 Dwelling access
2 Predominant roof material	2 Wall material	2 Outer walls material
3 Outside wall material	3 Floor material	3 Roof material
4 Floor material	4 Trash removal	4 Floor material
5 Rooms (number)	5 Electricity	5 Rooms (number)
6 Bedrooms (number)	6 Sewage drains	6 Kitchen
7 Room for cooking	7 Running water	7 Tenancy
8 Water source	8 Natural gas	8 Bedrooms (number)
9 Toilet	9 Telephone	9 Cooking fuel
10 Use of toilet	10 Toilet type	10 Lighting
11 Location of toilet	11 Location of water service	11 Water source
12 Lighting	12 Bathrooms (number)	12 Toilet
13 Ownership of dwelling	13 Kitchen	13 Waste removal
	14 Ownership of dwelling	14 Refrigerator
	15 Rooms (number)	15 Stove
	16 Bedrooms (number)	16 Washing machine
	17 Source of water for cooking	17 Television
	18 Kitchen	18 Air conditioning
	19 Fuel for cooking	19 Radio/stereo
	20 Fridge	20 Car
	21 Washing machine	21 Cistern
	22 Stereo	22 Computer
	23 Water heater	23 Converter
	24 Shower	24 Generator
	25 Blender	25 Landline or cellphone
	26 Electric or gas oven	26 Internet
	27 Air conditioner	
	28 Fan	
	29 TV color	
	30 Computer	
	31 Microwave	
	32 Bike (number)	
	33 Motorcycle (number)	
	34 Ships, sailboats, boats (number)	
	35 Autos (number)	

Data source: Integrated Public Use Microdata Series (IPUMS) - International.

Table 4.A1 (continued): Detailed Asset Variables Available for Selected Census Samples

Mexico 1970	Mexico 2000	Panama 1980
1 Bath with running water	1 Dwelling type	1 Dwelling type
2 Kitchen	2 Walls	2 Rooms (number)
3 Rooms (number)	3 Roof	3 Bedrooms (number)
4 Ownership	4 Floors	4 Kitchen
5 Wall material	5 Kitchen	5 Ownership
6 Floor material	6 Bedrooms (number)	6 Exterior walls material
7 Roof material	7 Rooms (number)	7 Roof material
8 Piped water	8 Water	8 Floor material
9 Sewer connection	9 Toilet	9 Drinking water source
10 Fuel for cooking	10 Sewer	10 Sewer facilities
11 Electricity	11 Electricity	11 Bathroom
12 Radio	12 Fuel for cooking	12 Lighting
13 TV	13 Dwelling ownership	13 Fuel for cooking
	14 Radio	14 Television
	15 Television	15 Radio
	16 Videocassette player	16 Telephone
	17 Blender	17 Refrigerator
	18 Refrigerator	18 Washing machine
	19 Washing machine	19 Sewing machine
	20 Telephone	
	21 Hot water heater	
	22 Car, van, or light truck	
	23 Computer	
	24 Trash disposal	

Data source: Integrated Public Use Microdata Series (IPUMS) - International.

Table 4.A1 (continued): Detailed Asset Variables Available for Selected Census Samples

Panama 2010	Peru 1993	South Africa 1996
1 Dwelling type	1 Water source	1 Dwelling type
2 Dwelling ownership	2 Lighting (electricity)	2 Rooms (number)
3 Wall material	3 Rooms (number)	3 Dwelling ownership
4 Roof material	4 Walls	4 Fuel for cooking
5 Floor material	5 Floor	5 Fuel for heating
6 Rooms (number)	6 Sewage	6 Fuel for lighting
7 Bedrooms (number)	7 Roof	7 Water supply
8 Water supply	8 Dwelling ownership	8 Toilet
9 Toilet	9 Dwelling type	9 Refuse disposal
10 Lighting	10 Vacuum cleaner	10 Telephone
11 Garbage disposal	11 Car for private use	
12 Fuel for cooking	12 Car for work use	
13 Cook stove	13 Toilet	
14 Refrigerator	14 Bicycle	
15 Washing machine	15 Light truck for work	
16 Sewing machine	16 Room for cooking	
17 Residential phone	17 Computer	
18 Radio (number)	18 Washer	
19 Electric fan (number)	19 Floor polisher	
20 Air conditioner (number)	20 Sewing machine	
21 Cell phone (number)	21 Knitting machine	
22 Automobile (number)	22 Motorcycle	
23 TV (number)	23 Radio	
24 Cable TV	24 Refrigerator	
25 Computer (number)	25 Stereo	
26 Internet	26 Phone	
	27 Tricycle	
	28 TV black/white	
	29 TV color	
	30 Video camera	

Data source: Integrated Public Use Microdata Series (IPUMS) - International.

Table 4.A2a: Comparison of Classification by Quintiles
Households Classified in a Higher Quintile with respect to Polychoric PCA Index [1/]

	Asset count	Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	6.9	8.3	2.7	0.0	28.6
Brazil 2010	14.3	12.0	2.4	0.0	33.2
Cambodia 1998	NA	27.6	3.1	0.0	NA
Colombia 1973	NA	10.2	2.7	0.0	32.3
Colombia 2005	13.8	6.1	1.6	0.0	NA
Dominican Republic 2002	12.4	10.8	2.4	0.0	32.2
Mexico 1970	13.0	8.1	1.9	0.0	30.6
Mexico 2000	9.4	7.8	2.7	0.0	30.8
Panama 1980	12.1	10.4	3.5	0.0	32.4
Panama 2010	19.7	10.2	3.3	0.0	30.1
Peru 1993	7.7	8.3	2.0	0.0	NA
South Africa 1996	NA	7.2	3.0	0.0	29.8
Simple average	***12.1***	***10.6***	***2.6***	***0.0***	***31.1***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

1. Classified in a higher quintile by the alternative measure being compared to the polychoric PCA index.

Table 4.A2b: Comparison of Classification by Quintiles
Households Classified in a Higher Quintile with respect to Log of Income per Capita [1/]

	Asset count	Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	19.1	28.4	27.9	28.1	0.0
Brazil 2010	23.2	29.1	28.5	28.3	0.0
Cambodia 1998	NA	NA	NA	NA	NA
Colombia 1973	NA	31.7	31.6	31.5	0.0
Colombia 2005	NA	NA	NA	NA	NA
Dominican Republic 2002	29.4	33.3	33.7	33.6	0.0
Mexico 1970	20.5	30.5	31.1	31.0	0.0
Mexico 2000	23.8	26.1	26.8	26.7	0.0
Panama 1980	19.1	26.0	25.5	25.9	0.0
Panama 2010	30.7	31.4	31.4	31.5	0.0
Peru 1993	NA	NA	NA	NA	NA
South Africa 1996	NA	31.4	31.3	31.8	0.0
Simple average	***23.7***	***29.8***	***29.7***	***29.8***	***0.0***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

1. Classified in a higher quintile by the alternative measure being compared to the logarithm of income per capita.

Table 4.A3a: Comparison of Classification by Quintiles

Households Classified in a Lower Quintile with respect to Polychoric PCA Index [1/]

	Asset count	Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	35.9	9.3	2.5	0.0	28.1
Brazil 2010	35.6	13.1	2.4	0.0	28.3
Cambodia 1998	NA	39.6	6.4	0.0	NA
Colombia 1973	NA	10.3	2.7	0.0	31.5
Colombia 2005	26.3	9.2	2.2	0.0	NA
Dominican Republic 2002	30.6	10.8	2.4	0.0	33.6
Mexico 1970	51.9	8.2	2.0	0.0	31.0
Mexico 2000	22.8	9.9	2.7	0.0	26.7
Panama 1980	34.5	10.3	3.5	0.0	25.9
Panama 2010	19.8	10.1	3.3	0.0	31.5
Peru 1993	43.3	8.4	2.0	0.0	NA
South Africa 1996	NA	7.8	4.6	0.0	31.8
Simple average	***33.4***	***12.2***	***3.1***	***0.0***	***29.8***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

1. Classified in a lower quintile by the alternative measure being compared to the polychoric PCA index.

Table 4.A3b: Comparison of Classification by Quintiles

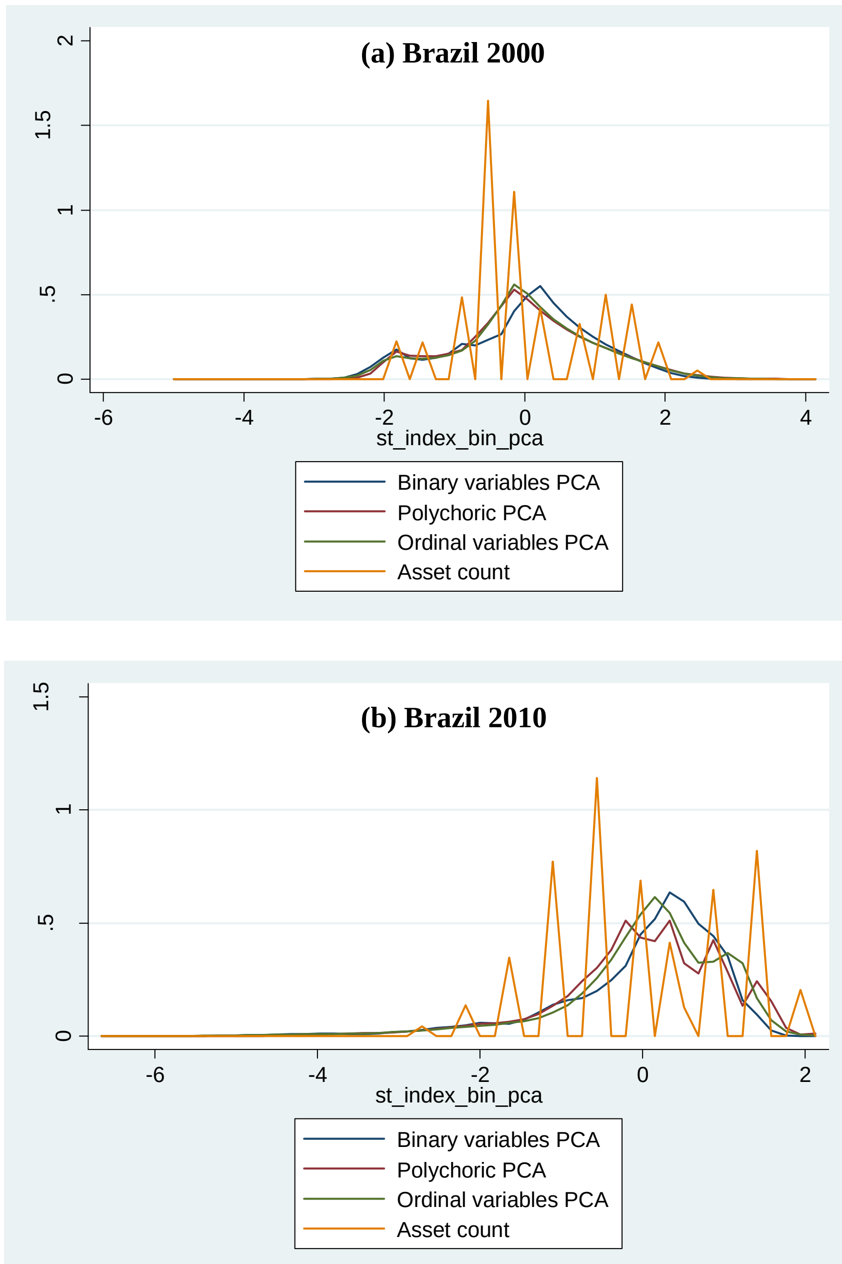
Households Classified in a Lower Quintile with respect to Log of Income per Capita [1/]

	Asset count	Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	40.5	28.8	28.2	28.6	0.0
Brazil 2010	41.2	34.1	33.1	33.2	0.0
Cambodia 1998	NA	NA	NA	NA	NA
Colombia 1973	NA	32.3	32.3	32.3	0.0
Colombia 2005	NA	NA	NA	NA	NA
Dominican Republic 2002	38.1	32.3	32.1	32.2	0.0
Mexico 1970	51.2	30.9	30.6	30.6	0.0
Mexico 2000	35.3	31.2	30.8	30.8	0.0
Panama 1980	45.5	33.1	32.2	32.4	0.0
Panama 2010	33.8	30.2	30.1	30.1	0.0
Peru 1993	NA	NA	NA	NA	NA
South Africa 1996	NA	29.2	30.0	29.8	0.0
Simple average	***40.8***	***31.3***	***31.1***	***31.1***	***0.0***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

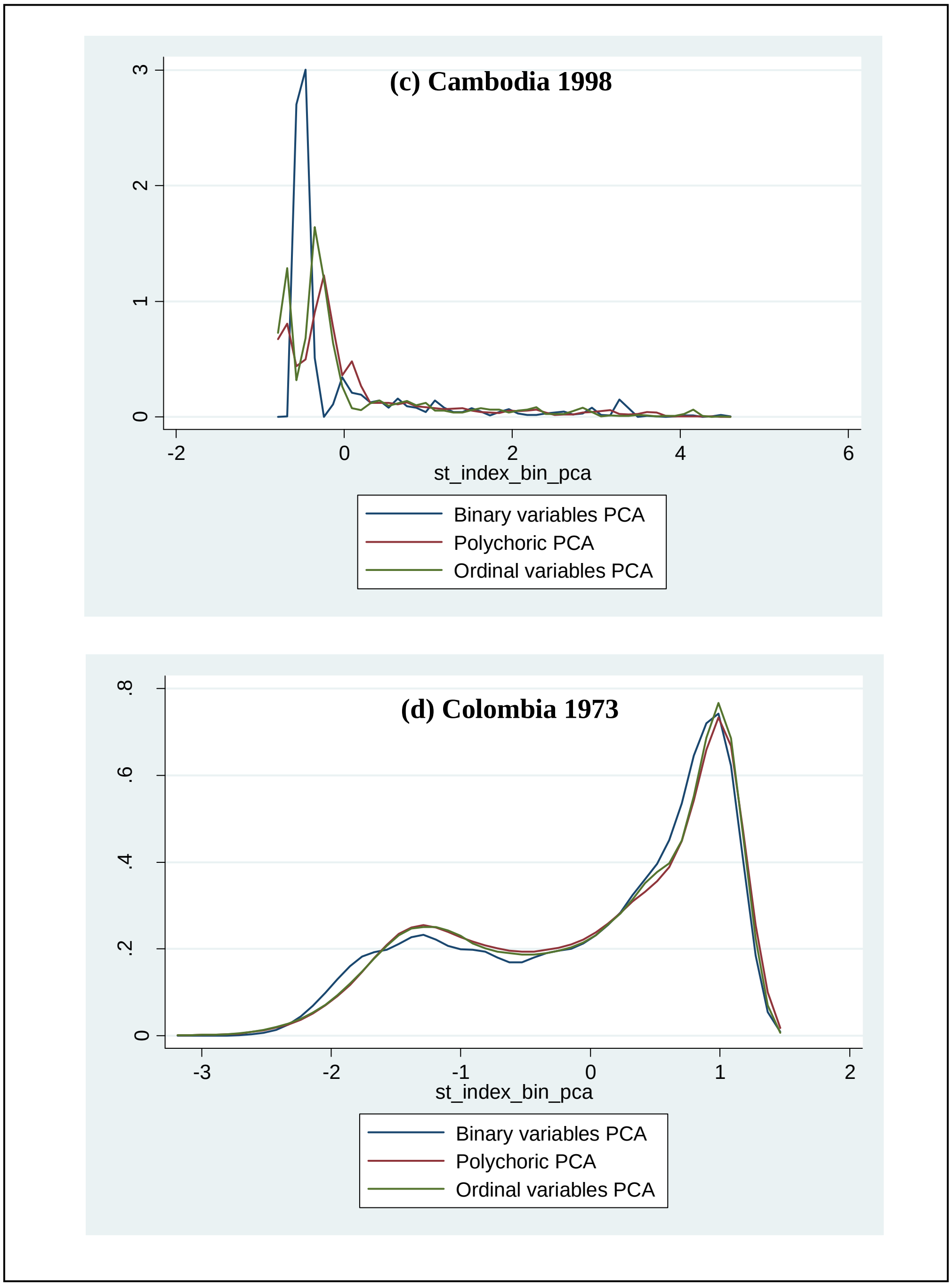
1. Classified in a lower quintile by the alternative measure being compared to the logarithm of income per capita.

Figure 4.A1: Kernel Density Estimates for Wealth Indices Based on Alternative Aggregation Methods



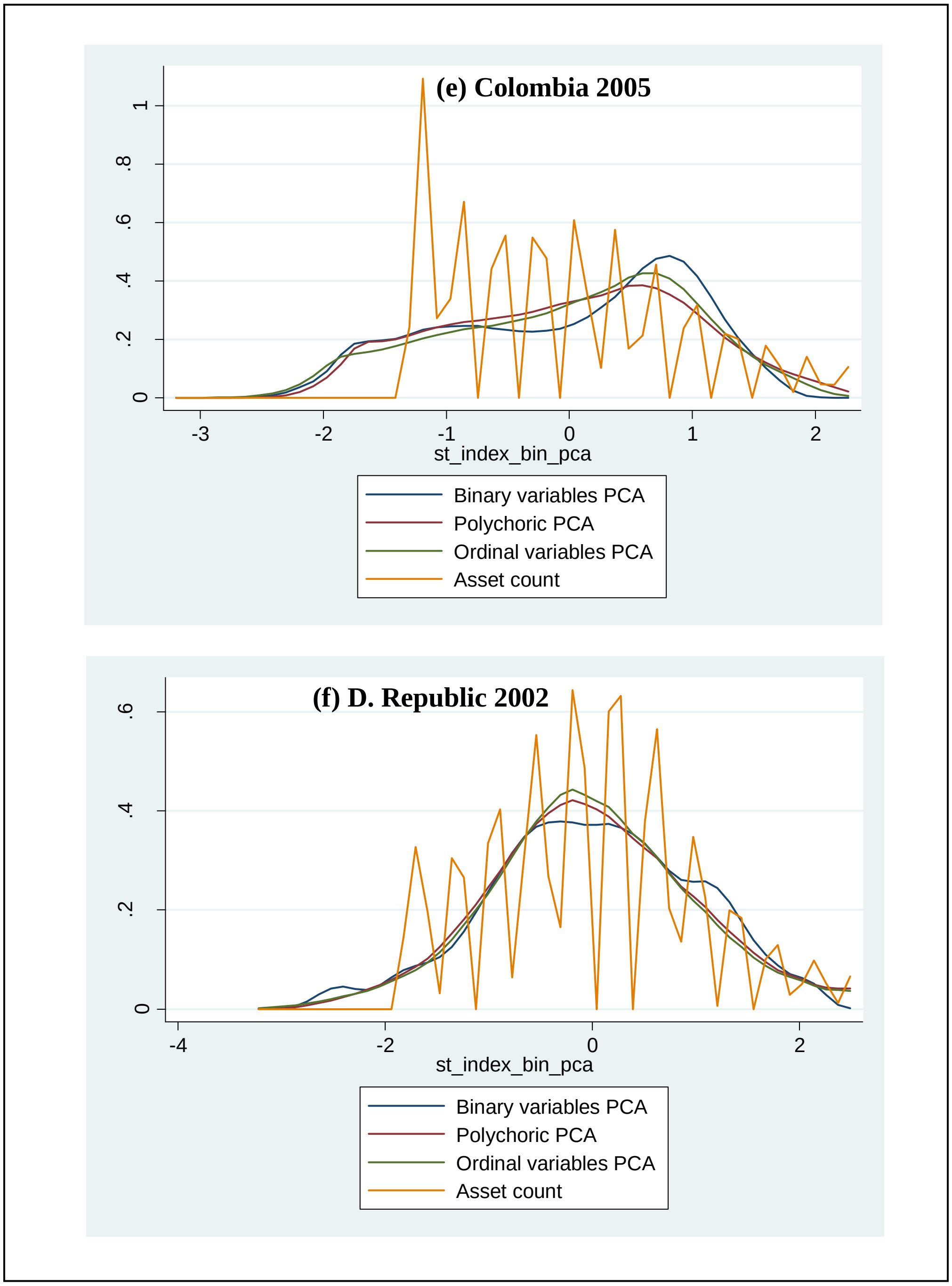
Data source: Integrated Public Use Microdata Series (IPUMS) - International.

Figure 4.A1 (continued): Kernel Density Estimates for Wealth Indices Based on Alternative Aggregation Methods



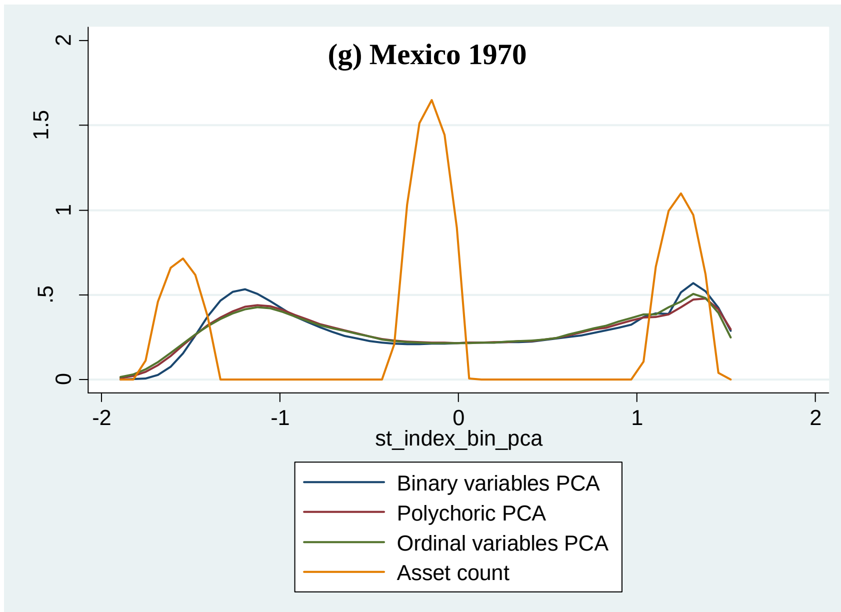
Data source: Integrated Public Use Microdata Series (IPUMS) - International.

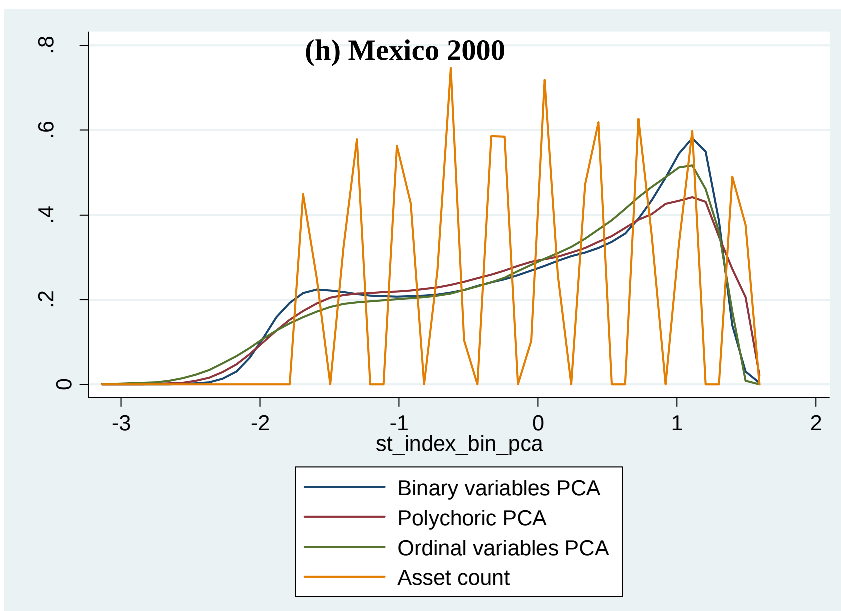
Figure 4.A1 (continued): Kernel Density Estimates for Wealth Indices Based on Alternative Aggregation Methods



Data source: Integrated Public Use Microdata Series (IPUMS) - International.

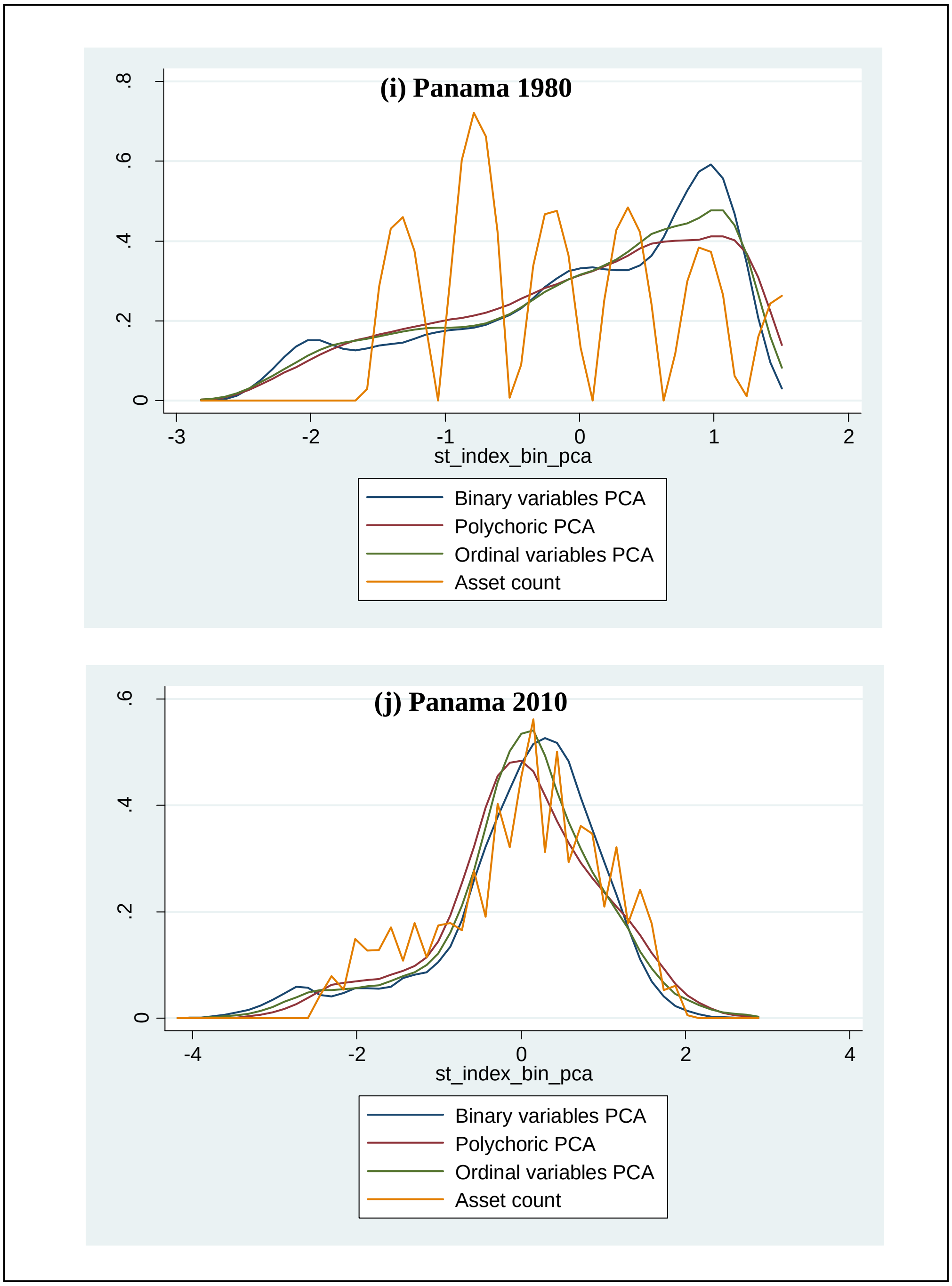
Figure 4.A1 (continued): Kernel Density Estimates for Wealth Indices Based on Alternative Aggregation Methods





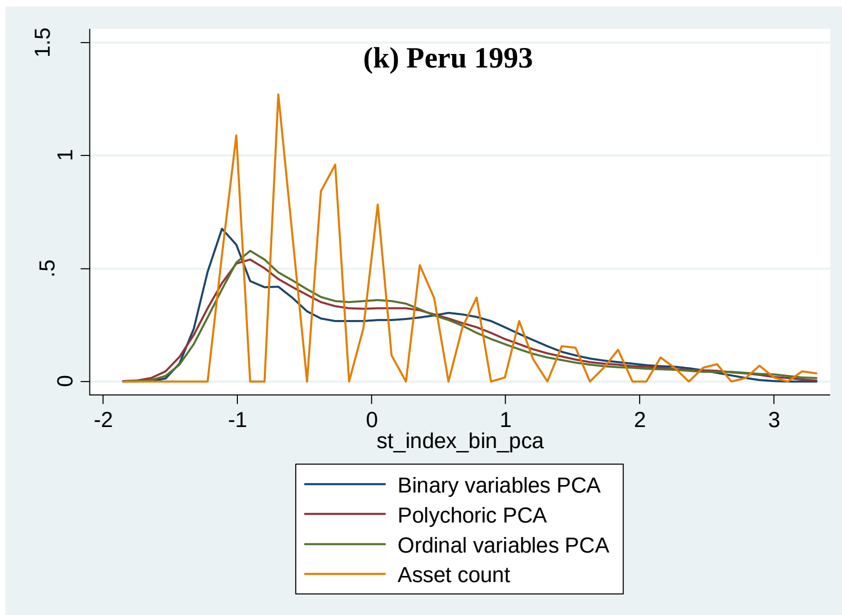
Data source: Integrated Public Use Microdata Series (IPUMS) - International.

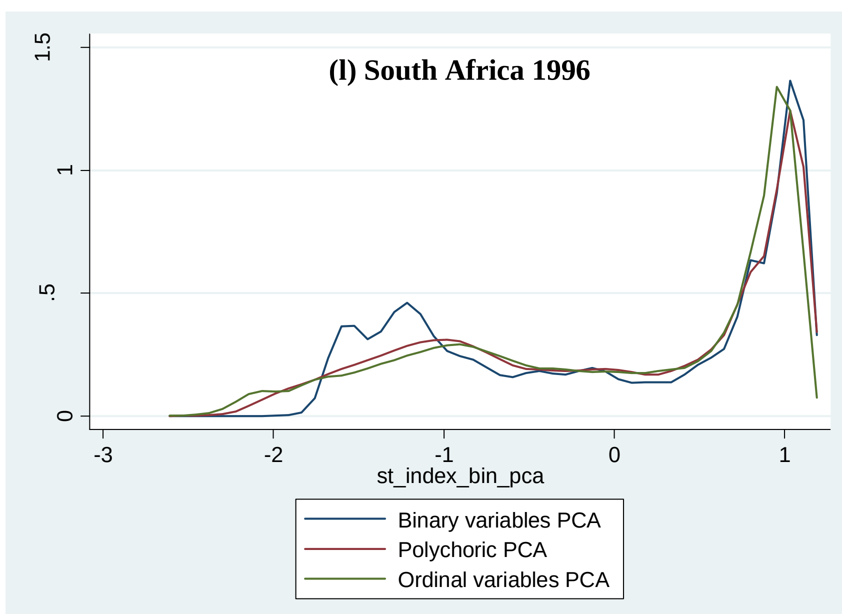
Figure 4.A1 (continued): Kernel Density Estimates for Wealth Indices Based on Alternative Aggregation Methods



Data source: Integrated Public Use Microdata Series (IPUMS) - International.

Figure 4.A1 (continued): Kernel Density Estimates for Wealth Indices Based on Alternative Aggregation Methods





Data source: Integrated Public Use Microdata Series (IPUMS) - International.

Table 4.A4a: Average Change in Literacy (persons ages 18-30) across Wealth Quintiles, for Alternative Aggregation Methods [1/]

	Mean	Average difference across quintiles			
		Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	93.4	4.21	4.30	4.33	3.73
Brazil 2010	96.8	1.88	1.98	1.98	1.71
Cambodia 1998	75.3	5.23	3.60	3.29	NA
Colombia 1973	87.7	7.01	6.75	6.77	4.56
Colombia 2005	95.0	2.71	2.74	2.87	NA
Dominican Republic 2002	92.5	4.03	4.00	4.02	2.64
Mexico 1970	79.8	9.28	9.33	9.38	8.43
Mexico 2000	95.9	3.10	3.11	3.13	2.57
Panama 1980	92.1	2.03	2.10	2.10	2.27
Panama 2010	97.1	1.42	1.42	1.44	1.24
Peru 1993	94.1	4.43	4.17	4.17	NA
South Africa 1996	NA	NA	NA	NA	NA
Simple average	***90.9***	***4.12***	***3.95***	***3.95***	***3.39***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

1. The average difference is calculated as the difference between top and bottom quintiles divided by four.

Table 4.A4b: Logit Model for Literacy (Persons Ages 18-30) Wealth Index Coefficient (Odds-ratio) [1/]

	Asset count	Binary PCA	Polychoric PCA	Ordinal PCA	Log income per capita
Brazil 2000	**3.744**	**3.069**	**3.663**	**3.408**	**2.638**
	0.108	*0.126*	*0.140*	*0.134*	*0.081*
Brazil 2010	**2.410**	**1.769**	**2.050**	**1.831**	**1.886**
	0.047	*0.054*	*0.068*	*0.055*	*0.045*
Cambodia 1998	**NA**	**1.406**	**1.310**	**1.377**	**NA**
		0.037	*0.029*	*0.027*	
Colombia 1973	**NA**	**2.228**	**2.042**	**2.038**	**1.366**
		0.054	*0.041*	*0.042*	*0.031*
Colombia 2005	**2.475**	**2.949**	**3.104**	**2.882**	**NA**
	0.108	*0.236*	*0.235*	*0.206*	
Dominican Republic 2002	**2.260**	**2.667**	**2.862**	**2.800**	**1.545**
	0.082	*0.077*	*0.096*	*0.089*	*0.035*
Mexico 1970	**1.636**	**2.022**	**2.057**	**2.063**	**1.654**
	0.043	*0.079*	*0.083*	*0.084*	*0.065*
Mexico 2000	**3.401**	**3.493**	**3.419**	**3.135**	**1.590**
	0.187	*0.154*	*0.131*	*0.097*	*0.066*
Panama 1980	**1.812**	**1.701**	**1.813**	**1.760**	**1.428**
	0.307	*0.204*	*0.227*	*0.223*	*0.101*
Panama 2010	**2.432**	**1.957**	**2.259**	**2.065**	**1.142**
	0.209	*0.119*	*0.157*	*0.119*	*0.059*
Peru 1993	**1.625**	**2.011**	**1.911**	**1.877**	**NA**
	0.165	*0.218*	*0.222*	*0.244*	
South Africa 1996	**NA**	**NA**	**NA**	**NA**	**NA**

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available

1. All estimates are statistically significant at the 1 percent level unless otherwise noted (### p>.01, ## p>.05, # p>.10). The table includes odds-ratio coefficients in bold and clustered standard errors in italic. Standard errors are clustered using mesoregions for Brazil, districts for Cambodia and Panama, municipalities for Colombia and Dominican Republic, states for Mexico, provinces for Peru, and magisterial districts for South Africa. The estimation sample is restricted to persons 18 to 30 years old that are not household heads.

Control variables: sex, age, and age-squared of the person; sex, age, age-squared, and educational attainment of household head (dummies for primary, secondary, and university); urban/rural.

Table 4.A5a: Average Change in Primary School Completion (Persons Ages 18-30) across Wealth Quintiles, for Alternative Aggregation Methods [1/]

	Mean	Average difference across quintiles			
		Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	63.3	15.90	16.01	16.02	14.26
Brazil 2010	84.9	6.96	7.28	7.29	6.69
Cambodia 1998	37.1	9.19	7.44	6.92	NA
Colombia 1973	47.1	16.76	16.59	16.56	13.90
Colombia 2005	86.4	9.20	9.30	9.44	NA
Dominican Republic 2002	75.6	10.21	10.11	10.15	7.21
Mexico 1970	35.6	16.76	17.40	17.36	15.70
Mexico 2000	87.0	8.40	8.42	8.44	6.52
Panama 1980	78.3	8.28	8.37	8.37	8.36
Panama 2010	92.3	3.97	3.99	4.04	3.31
Peru 1993	77.2	13.94	13.48	13.51	NA
South Africa 1996	84.5	6.45	6.68	6.64	5.21
Simple average	***70.8***	***10.50***	***10.42***	***10.39***	***9.02***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available
1. The average difference is calculated as the difference between top and bottom quintiles divided by four.

Table 4.A5b: Logit Model for Primary School Completion (Persons Ages 18-30) Wealth Index Coefficient (Odds-ratio) [1/]

	Asset count	Binary PCA	Polychoric PCA	Ordinal PCA	Log income per capita
Brazil 2000	**2.866**	**2.881**	**3.100**	**3.073**	**2.546**
	0.068	*0.114*	*0.102*	*0.106*	*0.059*
Brazil 2010	**1.913**	**1.714**	**1.915**	**1.782**	**1.740**
	0.028	*0.032*	*0.037*	*0.035*	*0.026*
Cambodia 1998	**NA**	**1.464**	**1.410**	**1.462**	**NA**
		0.036	*0.030*	*0.032*	
Colombia 1973	**NA**	**2.264**	**2.126**	**2.165**	**1.574**
		0.037	*0.035*	*0.034*	*0.028*
Colombia 2005	**2.514**	**2.450**	**2.643**	**2.493**	**NA**
	0.048	*0.068*	*0.072*	*0.064*	
Dominican Republic 2002	**1.851**	**2.268**	**2.354**	**2.341**	**1.550**
	0.046	*0.043*	*0.052*	*0.051*	*0.030*
Mexico 1970	**1.722**	**2.468**	**2.540**	**2.568**	**1.776**
	0.041	*0.119*	*0.106*	*0.109*	*0.078*
Mexico 2000	**2.576**	**2.778**	**2.770**	**2.641**	**1.490**
	0.113	*0.138*	*0.125*	*0.110*	*0.055*
Panama 1980	**1.921**	**2.065**	**2.161**	**2.125**	**1.689**
	0.188	*0.134*	*0.153*	*0.148*	*0.061*
Panama 2010	**2.336**	**1.989**	**2.299**	**2.120**	**1.169**
	0.131	*0.088*	*0.110*	*0.086*	*0.043*
Peru 1993	**1.848**	**2.450**	**2.411**	**2.400**	**NA**
	0.161	*0.203*	*0.254*	*0.295*	
South Africa 1996	**NA**	**1.945**	**1.996**	**1.955**	**1.463**
		0.043	*0.042*	*0.040*	*0.022*

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available
1. All estimates are statistically significant at the 1 percent level unless otherwise noted (### p>.01, ## p>.05, # p>.10). The table includes odds-ratio coefficients in bold and clustered standard errors in italic. Standard errors are clustered using mesoregions for Brazil, districts for Cambodia and Panama, municipalities for Colombia and Dominican Republic, states for Mexico, provinces for Peru, and magisterial districts for South Africa. The estimation sample is restricted to persons 18 to 30 years old that are not household heads.
Control variables: sex, age, and age-squared of the person; sex, age, age-squared, and educational attainment of household head (dummies for primary, secondary, and university); urban/rural.

Table 4.A6a: Average Change in Secondary School Completion (Persons Ages 18-30) across Wealth Quintiles, for Alternative Aggregation Methods [1/]

	Mean	Average difference across quintiles			
		Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	29.4	16.52	16.58	16.46	16.32
Brazil 2010	50.1	14.80	15.46	15.45	15.54
Cambodia 1998	4.3	3.14	3.08	3.01	NA
Colombia 1973	9.2	6.30	6.38	6.39	7.22
Colombia 2005	53.4	17.82	18.09	18.15	NA
Dominican Republic 2002	29.0	13.91	13.92	14.01	11.29
Mexico 1970	4.2	3.37	3.46	3.48	3.26
Mexico 2000	29.2	15.19	15.19	15.17	13.18
Panama 1980	27.4	14.55	15.23	15.07	14.76
Panama 2010	54.0	16.18	16.15	16.21	10.88
Peru 1993	52.4	18.56	18.26	18.31	NA
South Africa 1996	30.2	15.75	15.60	15.04	15.82
Simple average	***31.1***	***13.01***	***13.12***	***13.06***	***12.03***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available
1. The average difference is calculated as the difference between top and bottom quintiles divided by four.

Table 4.A6b: Logit Model for Secondary School Completion (Persons Ages 18-30) Wealth Index Coefficient (Odds-ratio) [1/]

	Asset count	Binary PCA	Polychoric PCA	Ordinal PCA	Log income per capita
Brazil 2000	**2.637**	**3.406**	**3.139**	**3.246**	**3.011**
	0.044	*0.097*	*0.068*	*0.072*	*0.051*
Brazil 2010	**1.991**	**2.167**	**2.372**	**2.403**	**2.147**
	0.031	*0.062*	*0.058*	*0.073*	*0.037*
Cambodia 1998	**NA**	**1.700**	**1.767**	**1.750**	**NA**
		0.058	*0.055*	*0.055*	
Colombia 1973	**NA**	**3.481**	**3.189**	**3.321**	**2.066**
		0.311	*0.210*	*0.243*	*0.050*
Colombia 2005	**2.230**	**2.495**	**2.629**	**2.630**	**NA**
	0.024	*0.055*	*0.045*	*0.053*	
Dominican Republic 2002	**1.792**	**2.327**	**2.301**	**2.290**	**1.777**
	0.031	*0.058*	*0.043*	*0.048*	*0.033*
Mexico 1970	**1.762**	**2.685**	**2.782**	**2.816**	**1.519**
	0.066	*0.313*	*0.293*	*0.306*	*0.067*
Mexico 2000	**2.223**	**2.900**	**2.887**	**3.008**	**2.000**
	0.064	*0.193*	*0.162*	*0.192*	*0.036*
Panama 1980	**1.864**	**2.641**	**2.609**	**2.750**	**2.564**
	0.121	*0.148*	*0.173*	*0.151*	*0.155*
Panama 2010	**2.480**	**2.627**	**2.666**	**2.650**	**1.241**
	0.080	*0.135*	*0.103*	*0.114*	*0.023*
Peru 1993	**1.750**	**2.396**	**2.368**	**2.303**	**NA**
	0.131	*0.178*	*0.242*	*2.303*	
South Africa 1996	**NA**	**2.308**	**2.392**	**2.338**	**2.047**
		0.060	*0.063*	*0.061*	*0.047*

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available
1. All estimates are statistically significant at the 1 percent level unless otherwise noted (### p>.01, ## p>.05, # p>.10). The table includes odds-ratio coefficients in bold and clustered standard errors in italic. Standard errors are clustered using mesoregions for Brazil, districts for Cambodia and Panama, municipalities for Colombia and Dominican Republic, states for Mexico, provinces for Peru, and magisterial districts for South Africa. The estimation sample is restricted to persons 18 to 30 years old that are not household heads.
Control variables: sex, age, and age-squared of the person; sex, age, age-squared, and educational attainment of household head (dummies for primary, secondary, and university); urban/rural.

Table 4.A7a: Average Change in Motherhood (Women Ages 18-30) across Wealth Quintiles, for Alternative Aggregation Methods [1/]

	Mean	Average difference across quintiles			
		Binary PCA	Ordinal PCA	Polychoric PCA	Log income per capita
Brazil 2000	53.0	-9.93	-10.30	-10.18	-11.21
Brazil 2010	46.9	-9.83	-10.42	-10.47	-12.53
Cambodia 1998	55.1	-0.74	-3.07	-2.71	NA
Colombia 1973	64.7	-7.16	-7.99	-7.90	-7.14
Colombia 2005	55.6	-9.10	-9.63	-9.59	NA
Dominican Republic 2002	67.7	-8.68	-9.06	-9.11	-6.71
Mexico 1970	58.4	-6.71	-7.18	-7.12	-7.55
Mexico 2000	56.1	-7.90	-8.63	-8.51	-7.59
Panama 1980	69.2	-7.03	-8.48	-8.33	-7.94
Panama 2010	56.9	-9.76	-10.42	-10.29	-7.97
Peru 1993	57.5	-11.57	-11.83	-11.87	NA
South Africa 1996	63.5	-6.40	-6.03	-5.75	-5.82
Simple average	***58.7***	***-7.90***	***-8.59***	***-8.49***	***-8.27***

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available
1. The average difference is calculated as the difference between top and bottom quintiles divided by four.

Table 4.A7b: Logit Model for Motherhood (Women Ages 18-30) Wealth Index Coefficient (Odds-ratio) [1/]

	Asset count	Binary PCA	Polychoric PCA	Ordinal PCA	Log income per capita
Brazil 2000	**0.787**	**0.762**	**0.758**	**0.751**	**0.525**
	0.009	*0.009*	*0.009*	*0.009*	*0.010*
Brazil 2010	**0.692**	**0.754**	**0.693**	**0.720**	**0.472**
	0.009	*0.014*	*0.013*	*0.015*	*0.012*
Cambodia 1998	**NA**	**0.810**	**0.831**	**0.826**	**NA**
		0.015	*0.017*	*0.015*	
Colombia 1973	**NA**	**0.710**	**0.679**	**0.678**	**0.641**
		0.024	*0.023*	*0.024*	*0.016*
Colombia 2005	**0.803**	**0.923**	**0.851**	**0.871**	**NA**
	0.012	*0.028*	*0.025*	*0.027*	
Dominican Republic 2002	**0.775**	**0.739**	**0.718**	**0.723**	**0.744**
	0.014	*0.021*	*0.014*	*0.014*	*0.016*
Mexico 1970	**0.935**	**0.927**	**0.911**	**0.912**	**0.768**
	0.029	*0.031*	*0.031*	*0.031*	*0.029*
Mexico 2000	**0.809**	**0.803**	**0.776**	**0.776**	**0.675**
	0.020	*0.021*	*0.022*	*0.022*	*0.035*
Panama 1980	**0.806**	**0.791**	**0.765**	**0.770**	**0.552**
	0.035	*0.051*	*0.049*	*0.054*	*0.035*
Panama 2010	**0.754**	**0.757**	**0.715**	**0.722**	**0.796**
	0.043	*0.059*	*0.048*	*0.053*	*0.020*
Peru 1993	**0.723**	**0.638**	**0.646**	**0.658**	**NA**
	0.023	*0.026*	*0.033*	*0.036*	
South Africa 1996	**NA**	**0.797**	**0.805**	**0.830**	**0.715**
		0.016	*0.016*	*0.016*	*0.018*

Data source: Integrated Public Use Microdata Series (IPUMS) - International. NA = Not available
1. All estimates are statistically significant at the 1 percent level unless otherwise noted (### p>.01, ## p>.05, # p>.10). The table includes odds-ratio coefficients in bold and clustered standard errors in italic. Standard errors are clustered using mesoregions for Brazil, districts for Cambodia and Panama, municipalities for Colombia and Dominican Republic, states for Mexico, provinces for Peru, and magisterial districts for South Africa.
Control variables: age and age-squared, marital status, educational attainment (dummies for primary, secondary, and university), family size, and urban/rural.

Conclusions

The first two essays contribute useful evidence concerning the effects of maternity leave on women's employment status and fertility in Latin America, a region for which no previous studies on this topic are available. Results from the first essay show that maternity leave has positive effects on the labor force participation and unemployment-to-population ratio for women of childbearing age. However, effects on women's employment are statistically insignificant. The evidence presented in the second essay indicates that maternity leave has a small effect on fertility, and mainly for higher order births, which implies that it creates incentives to postpone births until a later stage in life.

The empirical results are based on data from six Latin American countries (Bolivia, Chile, Colombia, Ecuador, Peru, and Venezuela), but are useful evidence for policymakers in other countries in the region, particularly in contexts where women are increasingly participating in the labor market, acquiring more education, and reducing their lifetime fertility. Future research could focus on the effects of changes in other family benefits (such as paternity leave, sick child leave, childcare facilities, among others) in the context of overall labor market regulations in Latin America. Moreover, the availability of additional datasets, particularly with higher frequency of data collection (yearly or similar), may be used to perform country-specific analyses of the impact of changes in maternity leave duration.

Two specific contributions arise from the two chapters on the impact of changes in maternity leave regulations. The empirical strategy takes advantage of existing census microdata that spans a long time period by applying a pseudo-panel method (Deaton, 1985). To the best of my knowledge, no previous study has used this method to produce synthetic longitudinal data from censuses to assess the effects of any policy. Therefore, this opens an opportunity for further research that can take advantage of existing large collections of census microdata. In addition, the data on the duration of maternity leave was not previously available for the long span covered in this study. The creation of yearly time series data for this variable is a valuable contribution of this research.

The construction of wealth indices based on housing characteristics and asset ownership has been widely used as a proxy for socioeconomic status. Results from the third paper contribute to the discussion on which methods should be used to aggregate asset data and could be useful for any studies that require proxy measures of households' socioeconomic status. Despite recent papers that recommend not using the popular approach of PCA on binary indicators, the results of this essay suggest that this method has a similar performance to that of other methods, which are based on ordinal data. The difficulties in the calculation of polychoric correlations lead to the recommendation that standard correlation methods may be preferred to implement PCA.

Finally, the census microdata used to compare these alternative methods currently do not include any summary socioeconomic status measure based on housing characteristics and ownership of assets, despite the fact that these variables are generally available in censuses. Therefore, as a practical implication of the third essay, the results suggest that an asset-based wealth index should be added to census microdata. Two possible avenues for future research are related to the development of such measure. The theory behind the type of household wealth that is represented by the asset-based approach (with respect to income or expenditures) has not been extensively investigated.

Additionally, the availability of such asset-based wealth measure would allow analyzing changes over time in household wealth accumulation. This research could take advantage of the wide coverage in time and countries of existing census microdata collections.

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Appendix: List of National Legislation Reviewed

Bolivia

Law of 05/24/1939, General Labor Law

Supreme Decree 224 of 08/23/1943, Regulations on the General Labor Law

Law of 12/06/1949, modifies the General Labor Law

Law of 12/14/1956, Social Security Code

Supreme Decree 5315 of 09/20/1959, Regulations on the Social Security Code

Law Decree 13214 of 12/24/1975, modifies the General Labor Law

Supreme Decree 25749 of 04/24/2000, Regulations on the Law 2027 of Public Servants

Supreme Decree 1212 of 05/01/2012, modifies the Supreme Decree 25749

Chile

Law Decree 442 of 04/06/1925, Protection of Laborer Mothers

Law Decree 178 of 05/28/1931, Labor Code

Law 10383 of 08/08/1952, Law of the Laborer's Insurance

Law 11462 of 12/29/1953, modifies the Labor Code

Law 16434 of 02/26/1966, modifies the Labor Code

Law 17928 of 05/10/1973, modifies the Labor Code

Law Decree 2200 of 06/15/1978, Regulations concerning the Work Contract and Workers' Protection

Law 18620 of 07/06/1987, Labor Code

Law 19250 of 09/30/1993, modifies the Labor Code

Law 19272 of 12/09/1993, modifies the Labor Code

Law 19670 of 04/15/2000, modifies the Labor Code

Law Decree 1 of 01/19/2003, Systematized version of the Labor Code

Law 20047 of 09/02/2005, modifies the Labor Code

Law 20137 of 12/16/2006, modifies the Labor Code

Law 20545 of 10/17/2011, modifies the Labor Code

Colombia

Law 53 of 04/22/1938, Protection of Female Workers

Decree 1632 of 09/10/1938, Regulations on the Law 53

Law 90 of 12/26/1946, Social Insurance

Decree 721 of 04/01/1949, Regulations on the Compulsory Insurance on Sickness-Maternity

Decree 2663 of 09/09/1950, Labor Code

Decree 3135 of 12/26/1968, Social Security

Decree 1848 of 11/04/1969, Regulations on the Decree 3135

Law 50 of 12/28/1990, modifies the Labor Code

Decree 956 of 05/29/1996, Regulations on the

Law 755 of 07/23/2002, modifies the Labor Code

Constitutional Court verdict C-273 of 2003, interprets articles of the Labor Code

Constitutional Court verdict C-174 of 2009, interprets articles of the Labor Code

Law 1468 of 06/30/2011, modifies the Labor Code

Constitutional Court verdict C-383 of 2012, interprets articles of the Labor Code

Ecuador

Labor Code, codification of 05/18/1961

Labor Code, codification of 06/07/1971

Labor Code, codification of 08/16/1978

Decree 3159-A of 03/11/1992, modifies the General Regulations on the Law of the Civil Service and Administration Career

Labor Code, codification of 09/29/1997

Law 17 of 09/25/2003, Civil Service and Administration Career and Unification of Salaries in the Public Sector

Decree 2474 of 01/17/2005, Regulations on the Law of the Civil Service and Administration Career

Labor Code, codification of 12/16/2005

Law of 02/13/2009, modifies the Law on Civil Service and Administration Career and Unification of Salaries in the Public Sector and the Labor Code

Peru

Law 2851 of 11/19/1918, Children's and women's work

Law 26513 of 07/28/1995, Promotion of Employment

Law 26644 of 06/25/1996, Pre- and Post-Childbirth Leave for the Pregnant Worker

Law 27402 of 01/19/2001, modifies the Law 26644

Law 27606 of 12/21/2001, modifies the Law 26644

Law 28308 of 07/29/2004, Pre- and Post-Childbirth Leave for the Armed Forces and Police

Law 29409 of 09/19/2009, Paternity Leave

Decree 014-2010-TR of 12/15/2010, Regulations on the Law 29409

Decree 005-2011-TR of 05/15/2011, Regulations on the Law 26644

Law 29992 of 01/18/2013, modifies the Law 26644

Venezuela

Labor Law of 07/16/1936

Decree 1563 of 12/31/1973, Regulations on the Labor Law

Organic Labor Law of 12/20/1990

Organic Labor Law of 06/10/1997

Law to Protect Families, Maternity, and Paternity of 09/20/2007

Organic Labor Law of 05/07/2012