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Social and Demographic Determinants of Low Fertility in Brazil

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Abstract

Brazil has experienced a rapid demographic transition, going from an estimated TFR of 6.28 children per woman in 1960 to 1.90 in 2010 (IBGE 2012). In this dissertation, I explore the social and demographic determinants of low fertility in the country. The first study analyzes the association of human development, gender equality, and work-childrearing support environment with the likelihood of having a birth in 1991, 2000, and 2010. The results show a declining relative importance of the Human Development Index (HDIm) with time, while the gender equality and work-childrearing support environment increase its relative importance. The emergence of these associations in 2000 and 2010 is likely to be driven by a rise in the share of professional and managerial females in the labor market. In the second study, I explore the extent in which childbearing is a constraint on the employment of mothers with young children in Brazil. Specifically, I test whether an exogenous increase in public child care centers between 2007 and 2009 is associated with an increase in labor force participation and hours worked of mothers with the youngest child aged four or less. The results indicate that the increase in child care availability had a positive impact on female labor force participation and hours worked, specially for married mothers with lower levels of education attainment. The third study assess the plausibility of a demographic technique, the P/F Brass method, used to estimate Brazil's official TFR. The method increased the TFRs observed in Census Surveys by approximately 12% in 1991, 10% in 2000, and 19% in 2010. Because of the dramatic fertility decline in Brazil in the age groups of 15-19 and 20-24, specially in 2010, and the violation of the stable population assumption, I conclude that the Brass method should no longer be used when estimating the official TFR in the country.

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SOCIAL AND DEMOGRAPHIC DETERMINANTS OF LOW FERTILITY IN BRAZIL

To my parents, Maria Betania Castanheira and
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ABSTRACT

SOCIAL AND DEMOGRAPHIC DETERMINANTS OF LOW FERTILITY IN BRAZIL

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Brazil has experienced a rapid demographic transition, going from an estimated TFR of 6.28 children per woman in 1960 to 1.90 in 2010 (IBGE 2012). In this dissertation, I explore the social and demographic determinants of low fertility in the country. The first study analyzes the association of human development, gender equality, and work-childrearing support environment with the likelihood of having a birth in 1991, 2000, and 2010. The results show a declining relative importance of the Human Development Index (HDIm) with time, while the gender equality and work-childrearing support environment increase its relative importance. The emergence of these associations in 2000 and 2010 is likely to be driven by a rise in the share of professional and managerial females in the labor market. In the second study, I explore the extent in which childbearing is a constraint on the employment of mothers with young children in Brazil. Specifically, I test whether an exogenous increase in public child care centers between 2007 and 2009 is associated with an increase in labor force participation and hours worked of mothers with the youngest child aged four or less. The results indicate that the increase in child care availability had a positive impact on female labor force participation and hours worked, specially for married mothers with lower levels of education attainment. The third study assess the plausibility of a demographic technique, the P/F Brass method, used to estimate Brazil's official TFR. The method increased the TFRs observed in Census Surveys by approximately 12% in 1991, 10% in 2000, and 19% in 2010. Because of the dramatic fertility decline in Brazil in the age groups of 15-19 and 20-24, specially in 2010, and the violation of the stable population assumption, I conclude that the Brass method should no longer be used when estimating the official TFR in the country.

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Abstract An increasing number of *developing* countries are experiencing below replacement fertility levels. Although the factors associated with low fertility in *developed* countries have been extensively explored in the literature, analyses of the emerging low fertility in middle and low income countries continue to be rare. To help fill this gap, we use Brazil as a case study to assess whether gender equality and work-childrearing support environment, at the municipality level, are associated with the odds of having a child. Both of these factors have been identified as key determinants of low fertility and fertility variation in *developed* countries, but their association with low fertility in *developing* countries is less clear. We find that gender equality and supportive work-childrearing environments are positively associated with the odds of having higher order births in Brazil in 2000 and 2010. We also observe that in 1991 these variables were not associated with fertility. The positive association found in 2000 and 2010 is therefore in all likelihood a reversal of a relationship that may have prevailed earlier in the demographic transition when gender-equality and the ability for mothers to work were not correlated, or were negatively correlated, with fertility.

1. Introduction

More than thirty *developing* countries[1] are currently experiencing below replacement fertility levels[2] (United Nations Population Division 2013). Although the factors associated with low fertility in *developed* countries are broadly explored in the literature (for recent reviews, see Balbo, Billari, and Mills 2013 or Caldwell and Schindlmayr 2010), it remains an open question whether the rise of low fertility, and the variation of fertility within countries, in *developing* countries are driven by similar factors. For example, it is conceivable that economic factors play a larger role in explaining low fertility in *developing* countries as compared to the more affluent *developed* countries. Or, does the often relatively swift transition from high fertility to low fertility, which is characteristic of many *developing* countries, imply that the drivers of national fertility trends and within-country variation are distinct from those of more developed context, where low fertility has been commonplace for many decades.

In the last decades, a great deal of research has focused on understanding fertility differences across *developed* countries, especially between low and lowest-low fertility regimes (Kohler, Billari, and Ortega 2002, Myrskylä, Kohler, Billari 2009). Fertility at later ages has an important role in this heterogeneity (Myrskylä, Kohler, and Billari 2011; Myrskylä, Goldstein, and Cheng 2013), and a recent study has documented a recovery in the TFR of lowest-low fertility countries (Goldstein, Sobotka, and Jasilioniene 2009), which is driven by a slowing down of fertility postponement and a rise of fertility rates at relatively older ages. Nevertheless, despite the recoveries from earlier fertility troughs, Europe continues to be characterized by very low fertility

[1]We use the UN Population Division's classification (United Nations Population Division 2013), in which *developing* countries are the countries in Africa, the Caribbean, Central America, South America, Asia excluding Japan, Oceania excluding Australia and New Zealand. *Developed* countries are the countries in Europe and the United States, Canada, Australia, New Zealand, and Japan.

[2]The *developing* countries with below replacement fertility (Net Reproduction Rate < 1) are: Armenia, Aruba, Azerbaijan, Bahamas, Barbados, Brazil, Chile, China, Hong Kong SAR, Macao SAR, Costa Rica, Cuba, Curaçao, Cyprus, Dem. People's Republic of Korea, Georgia, Iran (Islamic Republic of), Lebanon, Malaysia, Martinique, Mauritius, Myanmar, Puerto Rico, Republic of Korea, Saint Lucia, Singapore, Thailand, Trinidad and Tobago, Tunisia, United Arab Emirates, Vietnam, and other non-specified areas of Eastern Asia (United Nations Population Division 2013).

and a striking fertility divide between Northern and Western Europe[3] on the one, and Southern and Eastern Europe on the other hand. Supportive work-childrearing environments and gender equity values are thought to be among the main explanations for greater fertility levels in Northern and Western Europe (McDonald 2000).

To shed light on whether and how supportive work-childrearing environments and gender equity are associated with low fertility in *developing* countries, we analyze Brazil, a heterogeneous country that reached below replacement levels in the 2000s. This institutional heterogeneity in Brazil prevails because labor policies do not harmonize work and childrearing effectively in the country (Sorj, Fontes, and Machado 2007), and the access to child care, especially low-cost and high quality child care, is likely to be one important dimension in which substantial regional heterogeneity exists. Gender equality is also at low levels, with Brazil being in the lower range of the Global Gender Gap Index (85th position among 134 countries; Hausmann, Tyson, and Zahidi 2010). However, in a heterogeneous country, the national average gender equality is not necessarily meaningful and regional variation is likely to be substantial. And arguably, regional variation in gender equality is an important determinant of regional variation in demographic behaviors, and vice versa. Specifically, gender equality and work-childrearing support environments in countries like Brazil are also more likely to depend on local institutional and socioeconomic contexts than countries with more well-established national welfare states and more harmonized policies and institutions.

In this paper, we explore Brazil's regional variation and estimate the likelihood of having a birth in Brazil using a multilevel logit model with gender equality, work-childrearing support environment, human development, and female labor force participation at the municipality level, controlling for socioeconomic characteristics at the individual level. Our key findings reinforce the importance of gender equality and supportive work-childrearing environments to fertility. In consonance with previous results (Harknett, Billari, Medalia 2014; Bauernschuster, Hener, and

[3] Western Europe with exception of Germany, Austria, and Switzerland.

Rainer 2014; Brodmann, Esping-Andersen, and Güell 2007; among others), gender equality and work-childrearing support are only associated with the probably of having higher order births, and not first births. By comparing 2010 with 2000 and 1991 (the two earliest census years for which these analyses are possible), we observe that in 1991, these variables were not associated with the probability of having a birth. In 2000, a pattern that is mostly "in between" the 1991 and 2010 results prevails. Since we observe a substantial increase in the share of professional and managerial women in this period, we hypothesize that gender equality and work-childrearing support environment become more important to and positively associated with fertility as women have better occupations in the labor market.

In addition, by comparing 2010 with the previous years, we find that the relative importance of the Human Development Index (HDIm) for higher order births *decreases* with time. Specifically, the HDIm's negative association with higher order fertility is fairly modest in 2010, in sharp contrast to a strong negative association in 1991 and 2000. For first order births, the HDIm's association with fertility is similar in 1991 and 2010, but presented no association in 2000. The reason why the HDIm is not statistically significant for first births in 2000 remains an open question. We hypothesize that a monetary crisis in 1999, with a large devaluation of the Brazilian currency, may have produced a postponement effect of first births in all individuals, independent of the Human Development Index of their municipalities. Previous research has shown that the increase in uncertainty during economic crisis can reduce fertility nine months later (Sobotka, Skirbekk, and Philipov 2011).

Overall, this paper contributes to the scarce literature on low fertility in *developing* countries. Because of the swift fertility decline in Brazil, we were able to observe the continuation of its fertility transition to post-transitional fertility levels by analyzing the three most recent Demographic Censuses. A key finding is the decrease over time of the negative association between the Human Development Index and the probability of having a higher order birth, while other variables increased its importance. It is the case for gender equality and work-childrearing support environment that started being significant in the 2000s constituting a reverse of the relationship that most likely prevailed earlier in the Demographic Transition (Mason 2001). We hypothesize that the emergence of these two variables are related to the increasing importance of females in the labor market. This is the first study that shows this changing relationship for within-country variation in a middle income *developing* country.

By distinguishing first from higher order births, we also highlight the importance of considering birth order in the fertility analysis. We conclude that, when considering fertility in a middle income *developing* country, it is important to contemplate the heterogeneity across its territory and to recognize that there is not a single factor responsible for its fertility levels, but a plethora of factors that might include first and second demographic transition elements in the same country.

2. Background

2.1 The Fertility Decline in Brazil

Socially, economically, and culturally Brazil is a very heterogeneous country. It has the fifth largest territory in the world, encompassing diverse populations and biomes. Alongside this socioeconomic heterogeneity, different demographic regimes coexist with distinctly different fertility patterns (Carvalho and Wood 1994). For example, fertility started to decline in the South and Southeast regions in the 1960s, roughly at the same time as Chile and Costa Rica, but only fifteen years later it started to decline in the North and Northeast regions of the country (Potter, Schmertmann, Assunção, and Cavenaghi 2010). The fertility heterogeneity across Brazil's territory remains, with the states of the North and Northeast having the highest fertility levels, alongside with the lowest indicators of social and economic development. The fertility transition from high to replacement levels is mostly attributed to socioeconomic development, urbanization and mortality. Potter, Schmertmann, and Cavenaghi (2002), for instance, found that a decrease in child mortality, and an increase in electrification, schooling, and female labor force participation are strongly associated with the onset and pace of the fertility decline in Brazil, while a decrease in the percentage of Catholic individuals is only weakly associated with the decline.

Because of the strong relationship between fertility and development in Brazil, many scholars also emphasize the role of Federal Government policies in the decline of fertility, despite the absence of a targeted public policy to reduce fertility (Carvalho, Sawyer, and Paiva 1981, Martine 1996, Perpétuo and Wong 2006). In particular, these studies emphasize government's role, since the 1950s, in industrialization and technology development, sanitary reform and electrification, improvements in public education and health services, and the solidification of the social security system. Additionally, some authors emphasize the role of mass communication (Faria and Potter 1999; Ferrara, Chong and Duryea 2012), hypothesized to affect fertility through changes in the perception of social roles, favoring nontraditional social norms (Faria and Potter 1999). A recent paper using the 1991 Brazilian Demographic Census shows that the expansion of Globo's coverage, the main soap opera producer in Brazil, is associated with lower birth probability between 1979 and 1991 (Ferrara, Chong and Duryea 2012). The association was more prominent in women with low socioeconomic conditions and in the middle or later ages of their reproductive life.

Contemporary below replacement fertility levels in Brazil occur together with a postponement of fertility (Rosero-Bixby, Castro-Martín, and Martín-García 2009) and an increase in cohabitation (Esteve, Lesthaeghe, R., and López-Gay 2012), suggesting that Brazil is acquiring characteristics of the second demographic transition (SDT) as defined by Lesthaeghe (2010). Three studies, so far, have explored low fertility in Brazil (Yazaki 2003, Wong and Bonifácio 2008, Alves and Cavenaghi 2009). They analyze changes in the proximate determinants of fertility, more specifically the type of contraception and age at marriage. Our paper contributes to this literature by considering contextual factors associated with fertility across Brazilian's municipalities, focusing on the regional variation in work-childrearing support environments and gender equality in small levels of geographical aggregations.

2.2 Brazilian Fertility in Context

This section briefly situates Brazilian fertility in the global context and presents its regional heterogeneity. The left panel of figure 1 shows the inverted “J” shape relation between the total fertility rate (TFR) and the human development index (HDI) as proposed by Myrskylä, Kohler, and Billari (2009). We can observe that between 1975 and 2010, Brazil had a substantial improvement in its HDI and the TFR decreased considerably. However, despite its progress, Brazil’s HDI remains relatively low (0.76) when compared to OECD countries. The right panel of Figure 1 shows the TFR and the Global Gender Gap Index in 2010 (Hausmann, Tyson, and Zahidi 2010) – which is a combination of education attainment, economic participation and opportunity, health and survival, and political empowerment of women relative to men for 134 countries. The graph shows no clear association between TFR and GGG. Brazil is located in the center of the cloud with a GGG of 0.67, ranking 85th across in the set of countries for which the GGG is available.

Figure 2 shows the total fertility rate (TFR) and the Human Development Index in Brazilian municipalities (HDIm) in 1991, 2000, and 2010. The TFR and the HDIm are calculated by the United Nations Development Programme for Brazil (http://www.pnud.org.br/IDH/). The HDIm has three dimensions: life expectancy, *per capita* income, and education. We can observe a strong correlation between TFR and HDIm in the twenty years span, with fertility decreasing as the human development increases. We can also observe the significant variation in HDIm levels across Brazilian’s municipalities, in 2010 the HDIm ranged from 0.42 in the state of Pará (North of Brazil) to 0.82 in the state of São Paulo (Southwest of Brazil). This heterogeneity explains the importance of including the HDIm in our estimations. In general, we expect a negative association between fertility and development in Brazil because the country has not attained advanced levels of development (HDI>0.9), where studies have documented the emergence of a positive HDI-TFR association (Myrskylä, Kohler, Billari 2009).

Finally, Figure 3 shows the association between regional TFR levels in Brazil and the contextual variables used in the statistical models. The municipalities' Human Development Index (HDIm) in the top left of the figure shows a strong negative correlation with TFR in 2010, as does the municipality-level ratio of working females aged 18 to 54 years old in the bottom left of the figure. The ratio of working mothers with the youngest child aged 1-5 years old, in the bottom right, has more variation and a lower negative correlation with fertility than FLFP. To measure gender equality, we compute a municipality-level version of the Global Gender Gap Index, which we denote as GGm, described in more detail in Section 3. Greater levels of GGm indicate more gender-equal societies. The municipalities' GGm, in the top right panel of Figure 3, presents a small positive correlation with TFR. This positive pattern is striking, as it almost certainly constitutes a reversal from patterns earlier in the fertility transition when gender equality was in all likelihood not associated, or negatively associated, with fertility.

2.3 Fertility, work-childrearing support environments, and gender equality

In our subsequent analyses, we use the heterogeneity in Brazilian municipalities (Figure 3) to explore the association of regional gender equality and development level with the probability of having a child at the individual level, by birth order. In these analyses, we also control for levels of female labor force participation and the ratio of working mothers with the youngest child aged 1-5 years old, as a proxy for the work-childrearing support environment in the municipality. These analyses are possible because the Brazilian Demographic Censuses provide a unique opportunity to study contextual factors associated with fertility using a small level of geographic aggregation.

Work-childrearing support environments and gender equality are among the main explanations for differences in low fertility levels. At the macro level, an important evidence of this association is a change in the correlation between fertility and female labor force participation (FLFP) across low fertility countries, going from a negative to a positive correlation in the 1980s.

Rindfuss, Guzzo, and Morgan (2003) explained that this reverse occurred by an easing of the work and motherhood role incompatibility, with some countries having their institutional and normative contexts adapted to more family-friendly environments. Rindfuss and Brewster (1996) complement that child care *availability, acceptability, accessibility, cost, and quality* are central components for improving the work-childrearing support environment. Flexibility in hours worked and family adjustments to lessen mothers' unpaid housework are also important features in this framework. Similarly, the share of mothers' unpaid housework and its association with fertility is highlighted in McDonald (2000), which explains that very low fertility levels in *developed* countries are driven by lower gender equity in family institutions.

At the micro level, the association between gender equality and work-childrearing support environments with fertility differs by birth order, and it is usually not associated with first births. Bulatao (1981) explains that motherhood transition is associated with values like establishing a family, bringing the spouse closer, loving and caring for a child, and carrying the family name. Second and third order births are associated with sibling's companionship, wanting a boy or a girl, and the pleasure to watch children grow. In general, studies have shown that individuals are more likely to forgo the later because of practical difficulties in daily life incurred by an extra child. Therefore, greater gender equality in the share of housework and better work-childrearing support environments are likely to increase the probability of having higher order births.

Although there is evidence that in Norway child care availability is associated with the probability of having first births (Rindfuss, Guilkey, Morgan, Kravdal, and Guzzo 2007), published research shows more frequently an association with second and higher order births. In this regard, a multilevel analysis across 20 European countries find a positive association between higher order births and favorable family support environments, but no association with first order births (Harknett, Billari, and Medalia 2014). In Germany, the expansion of public child care centers had a significant positive effect on second and third order births (Bauernschuster, Hener, and Rainer 2014). In Italy, the availability of publicly provided child care centers has a positive association with fertility in general (Del Boca 2002).

Studies have also compared countries with different work-childrearing support environments. Brodmann, Esping-Andersen, and Güell (2007), for instance, compare Denmark and Spain finding that women are more likely to have second order births in the former than in the latter. The two main explanations are Denmark's work-childrearing support policies, and a greater amount of child care provision from Danish fathers. An additional study compares a German-speaking region in Belgium with Western Germany (Klüsener, Neels, and Kreyenfeld 2013). The authors find that, despite having Germany's social norms, fertility is higher in the Belgium region because of its favorable institutional context, with better family and labor market policies for reconciling work and childrearing.

In relation to gender equality, recent research finds a greater probability of second and third order births when fathers take parental leave in previous births in Sweden (Duvander and Andersson 2006; Duvander, Lappegård, and Andersson 2010) and in Norway (Duvander, Lappegård, and Andersson 2010; Lappegård 2010). A greater share of family responsibilities within working couples is associated with greater probability of second births in Sweden and Hungary (Oláh 2003), in Italy (Cooke 2003), and in the United States (Torr and Short 2004). Cross-country studies using OECD countries show a positive association between TFR, work-family policies (Luci and Thévenon 2010, Thévenon 2011, Mills, Rindfuss, McDonald, and Te Velde 2011), and gender equality (Myrskylä, Kohler, and Billari 2011).

In general, the literature indicates a positive association between second and third order births with work-childrearing support environments and gender equality. But these studies have mostly been conducted in high-income countries. Our analysis offers additional insights into the relationship of fertility, gender equality, and work-childrearing support environment in the context of a *developing* middle income country, that is, an important set of countries in which low fertility becomes increasingly widespread.

3. Data and Methods

3.1 Data

We use the Brazilian Demographic Census of 2010, which provides a unique data set among middle-income countries both in terms of sample size and regional coverage. The survey was applied to 11% of Brazilians' households and has more than 20 million individuals in its sample, which statistically represents 5,565 municipalities. The percentage of households included in the sample varies by municipality size, ranging from 50% of households in municipalities with less than 2,500 inhabitants to 5% of households in municipalities with more than 500,000 inhabitants. In order to contextualize the results obtained in 2010, we also use the Brazilian Demographic Censuses of 1991 and 2000, which are equally impressive samples in terms of size and representativeness at the municipality level. The 1991 sample has 17 million individuals across 4,491 municipalities and the 2000 sample has 20 million individuals across 5,507 municipalities.

In our statistical analysis, we restrict the analytical sample to women aged 16 to 49 years old at the survey's reference date. In an attempt to capture only intended fertility, we also restrict the sample to married or cohabiting women[4]. In the 2010 Demographic Census, almost 80% of all births were born from married or cohabiting women aged 16 to 49 years old. Our final analytical sample has 2,535,219 married or cohabiting women aged 16 to 49 years old in 1991, 3,107,852 in 2000, and 3,207,816 in 2010.

Table 1 shows the descriptive statistics for the 2010 sample. We can observe that approximately 7% of married or cohabiting women aged 16 to 49 years old experienced a birth in the 12 months prior to the survey. Among these births, 40% were first births, and 60% higher order births. The average age in the sample is 34 years old and most of the women are Catholic

[4] The 2006 DHS, for instance, shows that 93% of intended pregnancies in Brazil occurred among married or cohabiting women (Brazilian DHS 2006).

(66,5%), 20% are Pentecostal protestants, 4.5% mainline protestants, 3.9% had other religion, and 5.6% had no religion affiliation. Only 9.4% of these women have completed college, 28.2% completed high school, 19.1% completed primary, and 43% have not completed primary education. Almost 40% of them lives in the North and Northeast regions and 23% in rural areas. The monthly income per capita is R$ 636.00[5], approximately US$ 356.00 in 2010.

3.2 Dependent Variable

The dependent variable is one if the person had a birth in the 12 months prior to the survey, and zero otherwise. The variable is calculated based on the following question: “What is the month and year of birth of your last live birth?”, asked to all women who answered to have had at least one live birth. In the models that control for birth order, the variable used to measure birth order is the total number of children ever born, which is based on the question “How many live births did you have?”. These questions do not depend on the children’s life status at the time of the survey.

3.3 Independent Variables

Previous research on the association of fertility with work-childrearing support environments and gender equality has mostly measured it at the country level. This national approach is reasonable in contexts where within-country variation is small, which is the case in high-income countries with strong welfare states. In Brazil, however, there is considerable variation within-country, and any understanding of the nexus between gender equality and work-

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[5] In the population, the monthly per capita income of women aged 16 to 49 years old is R$764.2. When considering all individuals in the population, official estimates of the 2010 Census show a monthly per capita income of R$767.02 (IBGE - http://www.sidra.ibge.gov.br/ , table 3578), approximately US$ 430.00 in a month and US$5,160 in a year, differing significantly from Brazil’s US$ 11,978.3 GDP per capita in 2010 (http://data.worldbank.org/indicator/NY.GDP.PCAP.CD).

childrearing support environments with fertility demands a regional analysis. For this purpose, we use municipalities as our unit of analysis, which is the smallest level of socio-political aggregation available in the Demographic Censuses of 1991, 2000, and 2010.

3.3.1 The Gender Equality Index

A gender equality assessment should reflect women's position in society relative to men. Mills (2010) have estimated the association between fertility and six internationally recognized gender equality indexes. Among the indexes analyzed by the author, we have selected the Global Gender Gap Index (Hausmann, Tyson, and Zahidi 2010) because it estimates women's relative position to men in ways that is (relatively) independent of human development levels. In addition, it is the most multidimensional of the indexes, having four dimensions: economic participation, education attainment, health and survival, and political empowerment. While the original GGG is calculated at the country level, we estimate our index at the municipality level. To do so, we calculate a municipality-level gender-equality index, denoted GGm, which is a version of the GGG adapted to Brazilian municipalities. In some domains our index has fewer variables than the original GGG, but we have at least one variable of the original index in each dimension.

Table 2 shows the variables used to estimate the Gender Gap Index for Brazilian municipalities (GGm). The first dimension is economic participation, which have the following variables: female labor force participation rate (ages 18 to 54) over male value; female legislators, senior officials, directors and managers over male value (ages 18 to 54); and female professional and technical workers over male value (ages 18 to 54). The education attainment dimension has the following variables: female literacy rate over male value; female net primary enrollment rate over male value (ages 6 to 14); female net secondary enrollment rate over male value (ages 15 to 17); and female gross tertiary enrollment rate over male value. The health and survival dimension has the sex ratio at birth converted to female over male value. Finally, the political empowerment subindex has the number of females that were elected councilmembers in the municipality over male value. We estimate the index as in Hausmann, Tyson, and Zahidi (2010), in which we truncate all the variables at the equality benchmark and provide different weights based on the standard deviations of the variables presented in Table 2. The final GGm score is the average score of the four dimensions.

In the original GGG, the economic participation subindex has two additional variables: wage equality between women and men for similar work, and ratio of female estimated earned income over male value. The original health and survival subindex has additionally a ratio of female healthy life expectancy over male value. The political empowerment subindex includes the additional variables: ratio of females with seats in parliament over male value, ratio of females at ministerial level over male value, and the ratio of the years a female was a head of state in the last 50 years over male value. Finally, the education attainment subindex has the same variables as the GGm. The political empowerment GGm subindex is the dimension with the fewer number of variables from the original GGG index, because of differences in the levels of geographic aggregation.

The variables in the GGm economic and education subindexes are calculated from the Demographic Census Survey, the health and survival subindex is calculated from the National Civil Registry, and the political empowerment subindex from the Brazilian Electoral Court. In 2010, the final GGm score had an average of 0.73, ranging from 0.49 in the state of Minas Gerais to 0.98 in the state of Rio Grande do Sul. GGm values closer to one are probably resultant of its smaller number of variables and its greater variability. The original GGG index ranged from 0.46 in Yemen to 0.85 in Iceland, and Brazil’s final score was 0.67 in 2010 (Hausmann, Tyson, and Zahidi 2010). The correlation between the GGm and the Human Development Index in Brazilian municipalities (HDIm) – an index that combines life expectancy, income *per capita*, and education attainment – is small, as expected, so greater levels of human development do not necessarily imply greater gender equality.

3.3.2 The Human Development Index for Brazilian municipalities (HDIm)

The association between socioeconomic development and fertility has long been recognized in the literature. Although there is a theoretical debate about its relative importance in relation to the diffusion of new ideas, its relevance is well established (Bryant 2007; Myrskylä, Kohler, and Billari 2009). In Brazil, in particular, socioeconomic development was crucial for the onset and spread of the fertility decline (Potter, Schmertmann, and Cavenaghi 2002). Improvements in electrification, female education, female labor force participation, and in the survivorship of children under five years old are emphasized by the authors. Because of the relevance of socioeconomic development in Brazil, and its heterogeneity in the country, it is necessary to control for municipalities' development when measuring the association of fertility with gender equality and work-childrearing support environments.

We use the Human Development Index for Brazilian municipalities (HDIm), publicly available on the website of the United Nations Development Programme for Brazil (http://www.pnud.org.br/IDH/). The HDIm is an index with three dimensions: longevity, income, and education. Longevity is measured with the life expectancy at birth and income with the municipalities' *per capita* income. The education subindex is constituted by the education attainment of the adult population and the school flow of the youth population. The first is measured by the percentage of individuals aged 18 and older with primary education. The school flow is measured by the percentage of children aged 5 and 6 years old going to school, percentage of individuals aged 11 to 13 years old in the final years of elementary school, percentage of individuals aged 15 to 17 years old that completed elementary school, and percentage of individuals aged 18 to 20 years old that completed high school. The greater the HDIm, the greater the human development in a given municipality.

3.3.3 Work-childrearing support environments

A comprehensive measure of the work-childrearing support environment in a municipality should encompass, among other things, the characteristics of public and private child care centers, the norms regarding working mothers, and the family support environment for working mothers. Unfortunately, this type of information is not available in the Brazilian Demographic Censuses. Instead of measuring work-childrearing support environment directly, we use the employment ratio of mothers with the youngest child aged 1-5 years old as a proxy for it. It is a comprehensive variable that includes the institutional and the normative dimension of the issue. We assume that the greater the ratio, the lower the work and childrearing role incompatibility and the greater the work-childrearing support environment in the municipality.

Because of the wide variation in female employment across Brazilian's municipalities, the share of employed mothers with the youngest child aged 1 to 5 years old can reflect the demand for labor, and not necessarily better work-childrearing support environments. For instance, a municipality with low levels of working mothers with the youngest child aged 1 to 5 years old, can reflect the low overall demand for female labor in the municipality, rather than the unmet need for work-childrearing support. Consequently, we control for the variation in labor demand by including the ratio of employed females aged 18 to 54 years old in the municipality.

3.3.4 Individual-Level Variables

The individual-level variables are education attainment, household income *per capita* (natural log), race, age, age squared (to allow for nonlinearities between fertility and age), and religion affiliation. In the analysis that compares 2010 with 1991 and 2000, the religion variable is not included in the models because it is not available in the 1991 Demographic Census's cd data (IBGE 1996). In addition, we include an indicator variable for rural areas and North and Northeast regions to control for regional inequalities that may affect fertility and that are not specified in the regressions.

3.4 Methodological Approach

We use multilevel logistic regressions, with individuals (level 1) nested in municipalities (level 2). Level 1 variables are education, age, religion affiliation, race, household income *per capita*, and indicators variables for individuals living in rural areas and in the North and Northeast regions of Brazil. Level 2 variables are the Gender Gap Index (GGm), the Human Development Index (HDIm), the employment ratio of mothers with the youngest child aged 1 to 5 years old, and the ratio of employed females aged 18 to 54 years old (all measured at the municipality level). This model is preferred over a one-level model with average variables at the municipality level because it accounts for differences in the composition of the population in municipalities. Moreover, individuals are the ultimate decision makers in the fertility process and, consequently, their socioeconomic characteristics are important in this process[6].

The models were executed in Stata with the “xtlogit” command allowing for random intercept. In this regard, the fertility response of the *i*th women in the *k*th municipality is $\mathbf{ln(Y)_{ik} = \beta_0 + \beta_1 X_{i,k} + \beta_2 Z_k + \alpha_k + \varepsilon_{ik}}$ with level 1 variables represented by “$X_{i,k}$” and level 2 variables by “Z_k“. The random intercept at the municipality level (α_k) reflects municipality characteristics not considered in the model, thus accounting for intra-municipality correlation. The random error (ε_{ik}) reflects individual characteristics not specified in the model. The likelihood-ratio test of rho [rho = $(var(\alpha_k))/(var(\varepsilon_{ik})+ var(\alpha_k))$] indicates whether a two-level model with random intercept is applicable. In all models, rho was statistically significant and the two-level model is justified over a one level model.

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[6] See Smith (1989) and Philipov, Thévenon, Klobas, Bernardi, and Liefbroer (2009) for further discussion.

The models that control for birth order, consider only women at risk of having birth in the correspondent birth order in the 12 months prior the survey's reference date. Thus, for example, for first order births in the 2010 Census, our dependent variable is one for women that had their first birth during August 1st 2009 and July 31st 2010, and zero for childless women in the same period. In the models that consider higher order births, our dependent variable is one for women that had their second order birth or higher during August 1st 2009 and July 31st 2010, and zero for women with at least one live birth that did not have a birth in the reference period.

4. Results

4.1 Work-childrearing support environments, Gender Equality, Human Development and the Probability of Having a Birth in 2010

Table 3 shows the estimated relationships between fertility and the level 2 variables from the multilevel logit models. The models estimate the likelihood of having a birth in Brazil during August 1st 2009 and July 31st 2010 and the table displays the odds ratio correspondent to one unit increase in each variable. All models control for race, education attainment, religion affiliation, age, age squared, log income per capita, rural locality, and North and Northeast regions. In an attempt to capture intended fertility, we consider only married and cohabiting women aged 16 to 49 years old (although our results are not sensitive to this restriction).

Model 1 in Table 3 includes the individual level variables, the Human Development Index (HDIm), and the Gender Gap Index (GGm). Model 2 includes the individual level variables, the female employment ratio, and the employment ratio of mothers with the youngest child aged 1 to 5 years old. Model 3 combines the individual level variables with the municipality-level variables of Models 1 and 2. The results are consistent across models, usually presenting the same levels of statistical significance at the $p < 0.01$ level. The statistics of fit, measured by the Akaike information criterion (AIC), have better outcomes in Model 3. The likelihood-ratio test of rho, which indicates whether a two-level model with random intercept is applicable, is statistically significant in all models and a two level model is justified over a one level model.

The literature shows that socioeconomic development, in our case measured by the Human Development Index of Brazilian's municipalities (HDIm), is negatively associated with fertility (Bryant 2007; Potter, Schmertmann, and Cavenaghi 2002), especially in countries with HDI below 0.9 (Myrskylä, Kohler, and Billari 2009). Table 3 shows the expected negative association for first and higher order births (Models 1 and 3). By estimating the predicted probabilities at certain HDIm levels, we can observe that an increase in one standard deviation in HDIm, presents a 2.4% lower first birth probability (Model 3).

Gender equality is presented in the literature as having a positive association with fertility (Myrskylä, Kohler, and Billari 2011), especially for second or higher order births (Oláh 2003; Cooke 2003; Torr and Short 2004). In this regard, our results agree with previous research from high-income countries that has shown a positive association of gender with the probability of higher order births, but no association with first order births. The predicted probabilities of Model 3 show that an increase in one standard deviations in the GGm, increases the probability of higher order births by 1.2%. So, the greater the gender equality in the municipality, the greater the probability of having higher order births.

The proxy measuring work-childrearing support environment (ratio of working mothers with the youngest child aged 1-5 years old) presents a positive association with the odds of having a child. When considering birth order, it is not statistically significant for first births, at the $p < 0.01$ level, being statistically significant only for higher order births. The predicted probabilities show that one standard deviation increase in the ratio of working mothers with the youngest child aged 1-5 years old is associated with a 6.6% increase in the likelihood of having higher order births (Table 3 - Model 3). The result is in consonance with the literature, which finds that work-childrearing support environment is positively associated with the probability of having higher order births (Harknett, Billari, Medalia 2014; Bauernschuster, Hener, and Rainer 2014;

Brodmann, Esping-Andersen, and Güell 2007; among others). We innovate by showing this positive association for the first time within a *developing* country.

Finally, the association between fertility and female labor force participation in Brazil, a country with low levels of family policies to reconcile work and childrearing, is expected to be negative. Table 3 shows that the coefficient is negative and statistically significant, so the greater the ratio of women in the labor force, the lower the probability of having higher order births. The predicted probabilities show that an increase in one standard deviations in women's employment ratio decreases the probability of a higher order birth by 4.7% (Model 3). It can be observed that, in general, the coefficients' magnitudes of the work-childrearing support environment ratio and the female employment ratio are greater than the coefficients of the Human Development Index (HDIm) and the Gender Gap Index (GGm) suggesting a stronger influence of these variables in the probability of having higher order births in the twelve months prior to the survey.

In summary, the results show that gender equality and work-childrearing support environment are associated with the probability of having higher order births in Brazil, and are not associated with the probability of having first births. These findings are consistent with previous research. The Human Development Index (HDIm), an index that combines education, income, and life expectancy, is negatively associated with fertility in all birth orders. Consequently, the greater the human development index, the lower the birth probability, showing that, in Brazil, there is no indication of a reversal of the HDI-TFR relationship shown at higher levels of HDI (HDI>0.9) by Myrskylä, Kohler, and Billari (2009). Finally, the ratio of females in the labor force is negatively associated with the probability of having higher order births in Brazil.

4. 2 Work-childrearing support and Gender Equality 10-20 years earlier

To place the above results into context, and to investigate changes in the relationships between fertility, gender equality and development over time, in this section we re-estimate the multilevel regressions presented in section 4.1 using the 1991 and 2000 Brazilian Demographic Censuses.

In the literature, there is often a clear difference between the factors that studies tend to associate with fertility before countries attain replacement levels (Net Reproduction Rate > 1), and after attaining replacement levels (Net Reproduction Rate <= 1). Work-childrearing support environments and gender equality are generally included in analysis of the latter, and are expected to influence fertility at greater levels of female employment. So arguable, the associations found in the previous section regarding gender equality and work-childrearing support environments are not expected in other fertility context. In order to test whether these variables are significant in different fertility scenarios, we estimate its relationship with the probability of having a birth in 1991 and 2000. For comparison reasons, we do not include the religion affiliation variable in the models, because it is not included in the 1991 Demographic Census's cd data (IBGE 1996), so the 2010 results are re-estimated accordingly for our comparative analyses below.

The fertility decline in Brazil was very swift, as the country transitioned from high fertility levels in the 1960s to below replacement levels in the 2000s. Official estimates show a total fertility rate of 6.28 children per woman in 1960, 5.76 in 1970, 4.35 in 1980, 2.89 in 1991, 2.38 in 2000, and 1.90 in 2010 (IBGE 2012, page 73)[7]. By comparing 1991, 2000, and 2010 we are able to observe the association of the municipality-level variables discussed in the previous section with the probability of having a child in Brazil in the fertility context that prevailed 10-20 years earlier. Unfortunately, we are unable to estimate the multilevel models for the 1960s, 1970s and 1980s because of insufficient data to estimate the GGm and the HDIm indexes.

Table 4 shows the mean and standard deviation of the municipality-level variables used in the model. We also show the subindexes of the Human Development Index (HDIm) and the

[7] Official estimates calculate the Total Fertility Rate using the P/F Brass method (IBGE 2000), which overestimates the fertility rate in low fertility contexts. We estimate that, the total fertility rate in 1991, 2000, and 2010 calculated from the Demographic Censuses are respectively 2.50, 2.16, and 1.61 children per women.

Gender Gap Index (GGm), in order to have a complete perspective of the indexes' change in time, and the total fertility rate of municipalities provided in the UNDP's database (http://www.pnud.org.br/IDH/). Overall, table 4 shows a considerable change in the variables over time. In 1991 Brazilians' municipalities had an average of 3.65 children per women, following by 2.87 in 2000, and 2.19 in 2010. TFR's cross-country variation decreased considerable in twenty years, going from a standard deviation of 1.2 children per women in 1991 to 0.5 children per women in 2010. The coefficient of variation of the TFR has also decreased, presenting a dispersion of 32.7% in 1991, 25.6% in 2000, and 23% in 2010.

The Human Development Index of Brazilian municipalities (HDIm) had a 65% increase in its average in twenty years, with a significant contribution of the three subindexes (income, education, and longevity). The education subindex ("HDIm Education" in Table 4) had the largest increase going from 0.19 in 1991 to 0.56 in 2010, indicating not only greater schooling levels of the adult population, but also greater education attainment of the youth population. The Gender Gap index of Brazilian municipalities (GGm) has also increased, with its average improving 26% in twenty years. The dimension with the greater improvement was Economic Participation ("GGm Economic Participation" in Table 4), more than doubling its average in twenty years, indicating an increase in the ratio of females in the labor force and in the number of female managers, officials, professionals, and technical workers. The female employment ratio and the ratio of employed mothers with the youngest child aged 1 to 5 years old increased approximately 10% per decade.

Table A1 of the supplemental material shows the odds-ratios of the multilevel logit models for first births. In general, the results were very similar in 1991, 2000, and 2010. The HDIm is the only municipality-level variable that was statistically significant at the $p<0.01$ level to first order births in our analysis. It was negative and statistically significant in 1991 and 2010, however it was not statistically significant in 2000. So, married or cohabiting women in municipalities with advanced human development are having lower first birth probabilities in 1991 and 2010 than similar women in worst-off municipalities. The size of the coefficient is very similar in 1991 and 2010. The reasons for a different result in 2000 remain open. We hypothesize that a monetary crisis in 1999, with a large devaluation of Brazilian currency, may have produced a postponement effect of first births in all individuals, independent of the Human Development Index of their municipalities. Previous research has shown that the increase in uncertainty during economic crisis can reduce fertility nine months later (Sobotka, Skirbekk, and Philipov 2011).

Table 5 shows the odds-ratios of the multilevel logit models estimating the likelihood of having a higher order birth in Brazil in 1991, 2000, and 2010. We can observe that the HDIms' odds-ratios are statistically significant at the $p<0.01$ level, and become closer to one in most recent years, going from 0.14 in 1991 to 0.74 in 2010. These results are expected, as previous research has argued that the negative relationship between fertility and HDI declines and potentially even reverses as societies increase their development levels (Myrskylä, Kohler, Billari 2009). The above analyses for 1990—2010 document this changing relationship for the first time for within-country fertility variation in a middle-income *developing* country.

In relation to gender equality, the theory suggests that at the beginning of the fertility transition the greater the gender equality, the lower the fertility. However, as gender equality expands to other spheres of society, the association becomes positive (McDonald 2000). We are able to document this reversal within a fairly short period – two decades – for within-country variation in fertility in Brazil. In Table 5 we can observe that GGm has a negative coefficient in 1991, but it is only statistically significant at $p=.09$. While we cannot conduct our analyses for earlier years, based on the extensive literature on gender equality and fertility in medium-to-high fertility contexts (Mason 2001), we expect that this negative relationship was even stronger in earlier years when Brazilian fertility was higher. In contrast, in 2000 and 2010 we estimate a statistically significant ***positive*** association between gender equality and fertility in Brazil. Similarly to the GGm index, the employment ratio of mothers with the youngest child aged 1 to 5 years old, our proxy measuring work-childrearing support environment, also has a significant positive association with higher order births in 2000 and 2010, but no association in 1991. The ratio of employed women aged 18 to 54 years is not statistically significant in 1991 and 2000, and it is negative and statistically significant in 2010.

The differences found in each of the three periods (1991, 2000, 2010) are expected, but this expectation is derived almost exclusively based on previous research in high-income countries that had experienced low fertility for a substantial period of time. Our estimates document the changes in the relationship between fertility, gender equality and development for the first time in a *developing* country as it transitions into below-replacement fertility. To shed light on the mechanisms that might underlie this change in the relationships between fertility, gender equality and development during 1990—2010 in Brazil, it is useful to emphasize some aspects of the descriptive statistics shown in Table 4. We can observe that the two most important changes in 1991, 2000 and 2010 were in the GGm Economic Participation subindex and the HDIm Education subindex. These indexes more than doubled in two decades, representing a considerable increase in the general education attainment of the population and in the average share of professional, legislators, senior officials, directors and managers' women in municipalities. Based on this evidence, we hypothesize that, across the decades, the increase in women's education and rank in the labor market augmented the importance of the work-childrearing support environment and gender equality in the decision of having an extra child. Consequently, these are possibly one of the reasons for the different coefficients regarding gender equality and work-childrearing support environment in Brazil in 1991 and 2000-2010.

5. Discussion

An increasing number of *developing* countries are experiencing below replacement fertility levels. In South America, official estimates show that Chile and Brazil have reached below replacement fertility in the 2000s. Because of the scarcity of published research about low fertility in South America, and in *developing* countries in general, we have analyzed Brazilian fertility in the Brazilian Demographic Censuses of 1991, 2000, and 2010. Due to its large sample size and comprehensive geographic coverage, these surveys provide an exceptional opportunity to investigate the relationship of fertility with human development, gender equality, and work-childrearing support environments in a middle income country. We focus on within-country variation in fertility as Brazil is a country with considerable variation at the municipality level in both fertility and institutional/socioeconomic contexts.

Our analyses show that the factors associated with contemporary low fertility in Brazil are similar to those in *developed* countries. In particular, gender equality and work-childrearing support environment are associated with greater probabilities of higher order births in 2000 and 2010. Our analyses are the first study that has been able to show such a relationship for within-country variation in a middle-income *developing* country. Moreover, using 1991 and 2000 census data, we are able to show that the above positive associations between fertility and gender-equality and work-childrearing support environment are a recent phenomenon that has *emerged* as Brazil's fertility transition continued and Brazil reached below-replacement fertility. The contemporary positive association is likely to constitute a reversal of a relationship that in all likelihood prevailed earlier in the demographic transition when gender-equality and the ability for mothers to work were not correlated, or were negatively correlated, with fertility.

Further research is needed for a greater understanding of the mechanisms in which work and childrearing are associated with fertility in Brazilian municipalities. Specially, analyzing childcare availability, accessibility, price, and quality could be enlightening. In addition, studying the variation in norms and attitudes about childrearing (Rindfuss and Brewster 1996), length of paid parental leaves (Duvander, Lappegård, and Andersson 2010), withdrawal and return restrictions of the labor market (Desai and Waite 1991) or workplace flexibility could likewise help for a better understanding of this association. Moreover, analyzing the variation in gender equality in Brazil's territory and its association with public policies at the local level might be helpful. Further research on gender equality and work-childrearing support environments could advance the understanding of fertility levels in Brazil, its expected directions, variation across municipalities, and possible policy interventions.

TABLES

Table 1. Descriptive Statistics of Married or Cohabiting Women aged 16 to 49 years old in the 2010 Brazilian Demographic Census Survey Sample

	Mean or Percentage	Std. Dev.
Had a birth in the 12 months prior to the survey (%)	7.1	
Had a first birth in the 12 months prior to the survey (%)	2.9	
Had a higher order birth in the 12 months prior to the survey (%)	4.2	
Black (%)	6.6	
Age	33.9	8.7
Educ Attain (ref: Less than Primary) (%)		
Primary Completed (%)	19.1	
High School (%)	28.2	
College or more (%)	9.4	
Religion (ref: Catholic) (%)		
Petencostal Protestants (%)	19.5	
Mainline Protestants (%)	4.5	
Other religion (%)	3.9	
No religion (%)	5.6	
HH Income per capita in brazilian currency (R$)	636	1,783
Rural Area (%)	23.2	
North and Northeast region	36.7	
HDIm	0.70	0.08
GGm	0.73	0.04
Mothers y1-5 Employment Ratio	0.52	0.13
Female Employment Ratio	0.57	0.11

Data Source: IBGE, Brazilian Demographic Census of 2010, and PNUD (2013).
Note: There was a total of 3,207,816 married or cohabiting women aged 16 to 49 years old in the sample of the 2010 Brazilian Demographic Census.

Table 2. Variables and Sub indexes of the Gender Gap Index (GGm) for Brazilian Municipalities in 2010

Gender Gap Index for Brazilian' Municipalities (GGm)	Mean	S. Dv.	Weight
Economic Participation Subindex	0.85	0.05	
1 Ratio: Female labor force participation over male value	0.68	0.10	0.30
2 Ratio: Female legislators, senior officials, directors and managers over male value	0.57	0.27	0.11
3 Ratio: Female professional and technical workers over male value	0.99	0.05	0.59
Education Attainment Subindex	0.99	0.01	
1 Ratio: female literacy rate over male value	1.00	0.01	0.67
2 Ratio: female net primary enrolment rate over male (Ages 6 to 14)	0.98	0.03	0.16
3 Ratio: female net secondary enrolment rate over male value (Ages 15 to 17)	0.99	0.05	0.10
4 Ratio: female gross tertiary enrolment ratio over male value	0.99	0.07	0.07
Health and Survival Subindex			
1 Sex ratio at birth (converted to female-over-male ratio)	0.91	0.12	1.00
Political Empowerment Subindex			
1 Ratio: Female Elected Councilmembers over male values	0.18	0.18	1.00
GGm Final Score	**0.73**	**0.05**	

Data Source: IBGE, Brazilian Demographic Census of 2010, National Civil Registry, Superior Electoral Court, and PNUD (2013).
Note 1: The original Global Gender Gap Index (Hausmann, Tyson, and Zahidi 2010) is calculated at the country level based on national data, our analyses require estimation based on the regional levels and, therefore, it was not possible to consider all the indicators of the original index. The original GGG economic participation subindex had two additional variables: wage equality between women and men for similar work and ratio of female estimated earned income over male value. The original GGG education attainment subindex had the same variables as the GGm. The original GGG health and survival subindex included the ratio of female healthy life expectancy over male value. Finally, the original political empowerment subindex included the additional variables: ratio of females with seats in parliament over male value, ratio of females at ministerial level over male value, and the ratio of the years a female was a head of state, in the last 50 years, over male value.
Note 2: The variable measuring the number of elected councilmembers in the municipalities depends on the election cycle and, therefore, it does not always coincide with the Demographic Census's years. In this regard, for the 2010 GGm political empowerment subindex we use the 2012 councilmembers' election. In 2000, the Census coincided with the election year. In 1991, we used the elections of 1996 because of data availability. Likewise, because of data availability, in the 1991 GGm Health and Survival Subindex we have used the sex ratio from births from 1995. The remaining variables correspond to the Census years.

Table 3. Multilevel Logit Models with random intercept at the municipality level estimating the probability of married or cohabiting women aged 16 to 49 years old having a birth in Brazil in 2010 (odds ratios)

	Any Birth Order	First Birth	Higher Order Birth
Model 1			
Human Develoment Index (HDIm)	0.75 **	0.52 **	0.73 **
Gender Gap Index (GGm)	1.12	0.83	1.29 **
Model 2			
Mothers y1-5 Employment Ratio	1.33 **	0.80	1.69 **
Female Employment Ratio	0.71 **	0.80	0.59 **
Model 3			
Human Develoment Index (HDIm)	0.77 **	0.67 **	0.71 **
Gender Gap Index (GGm)	1.11	0.91	1.25 **
Mothers y1-5 Employment Ratio	1.26 **	0.74 *	1.57 **
Female Employment Ratio	0.79 *	0.95	0.69 **
N Municipalities	5,565	5,565	5,565
N Women	3,207,816	569,051	2,638,765

Data Source: IBGE, Brazilian Demographic Census of 2010, and PNUD (2013).
Note: Table reports odds ratios. *Significant at p<.05; **p<.01. All models include controls for age, age squared, race, education attainment, log income per capita, religion affiliation, rural area, and North and Northeast region.

Table 4. Mean and Standard Deviation of Selected Variables for Brazilian Municipalities in 1991, 200 and 2010

	1991		2000		2010	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
TFR	3.65	1.19	2.87	0.74	2.19	0.50
HDIm	0.40	0.10	0.52	0.10	0.66	0.07
HDIm Income	0.52	0.09	0.58	0.09	0.64	0.08
HDIm Education	0.19	0.09	0.35	0.13	0.56	0.09
HDIm Longevity	0.65	0.08	0.72	0.07	0.80	0.04
GGm	0.58	0.07	0.66	0.06	0.73	0.05
GGm Economic Participation	0.34	0.12	0.60	0.12	0.85	0.05
GGm Education Attainment	0.97	0.03	0.98	0.02	0.99	0.01
GGm Health and Survival	0.83	0.30	0.92	0.12	0.91	0.12
GGm Political Empowerment	0.13	0.14	0.15	0.16	0.18	0.18
Mothers y1-5 Employment Ratio	0.29	0.13	0.41	0.15	0.51	0.15
Female Employment Ratio	0.33	0.13	0.45	0.13	0.55	0.13

Data Source: IBGE, Brazilian Demographic Censuses of 1991, 2000, and 2010, PNUD (2013), National Civil Registry, and Superior Electoral Court.
Note: The Brazilian Demographic Censuses had 4,491 municipalities in 1991, 5,507 in 2000, and 5,565 in 2010.

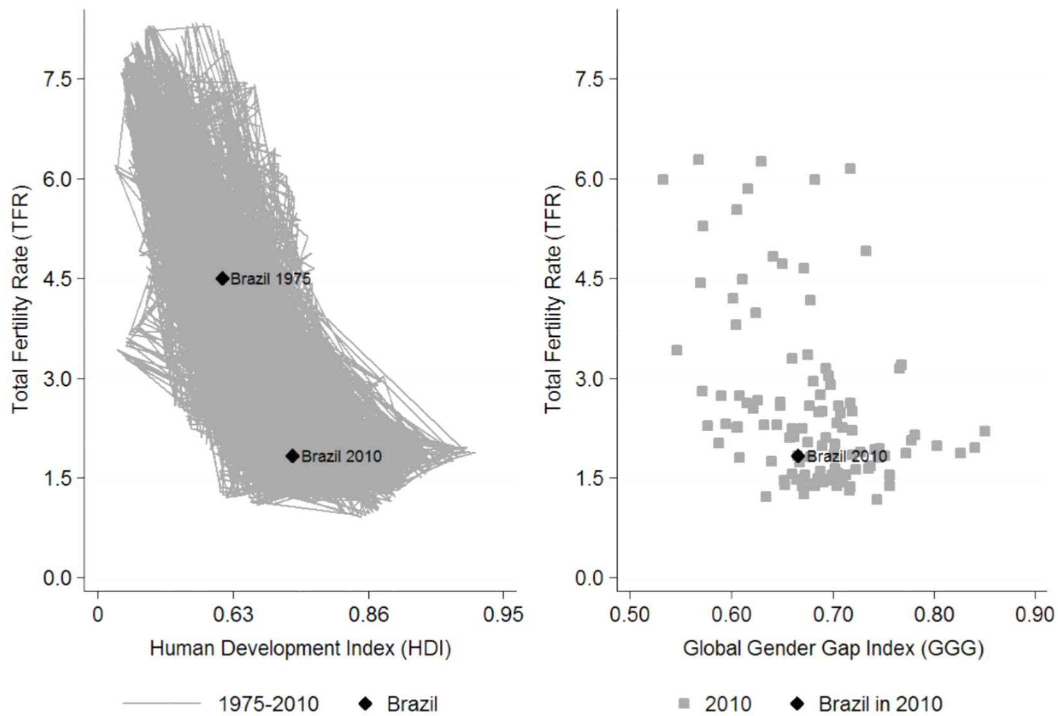
Table 5. Multilevel Logit Models with random intercept at the municipality level estimating the probability of married or cohabiting women aged 16 to 49 years old having a higher order birth in Brazil in 1991, 2000 and 2010 (odds ratios)

	1991	2000	2010
Black	1.32 **	1.21 **	1.10 **
Age	0.88 **	0.86 **	0.85 **
Age Squared	1.00 **	1.00 **	1.00 **
Educ Attainment (ref: Less than Primary)			
Primary Completed	0.81 **	0.83 **	0.89 **
High School	0.82 **	0.82 **	0.84 **
College or more	1.02	1.22 **	1.24 **
Log Income per capita (HH)	0.98 **	0.93 **	0.93 **
Rural Area	1.30 **	1.16 **	1.01
North and Northeast region	1.34 **	1.04 **	1.08 **
HDIm	0.14 **	0.28 **	0.74 **
GGm	0.87	1.35 **	1.24 **
Female Employment Ratio	1.08	0.81	0.68 **
Mothers y1-5 Employment Ratio	1.03	1.42 **	1.56 **
Constant	0.26 **	0.13 **	0.06 **
Rho	0.01 **	0.02 **	0.01 **
Sigma u	0.22	0.23	0.17
Obs	2,118,658	2,682,355	2,638,765

Data Source: IBGE, Brazilian Demographic Censuses of 1991, 2000, and 2010, PNUD (2013), National Civil Registry, and Superior Electoral Court.

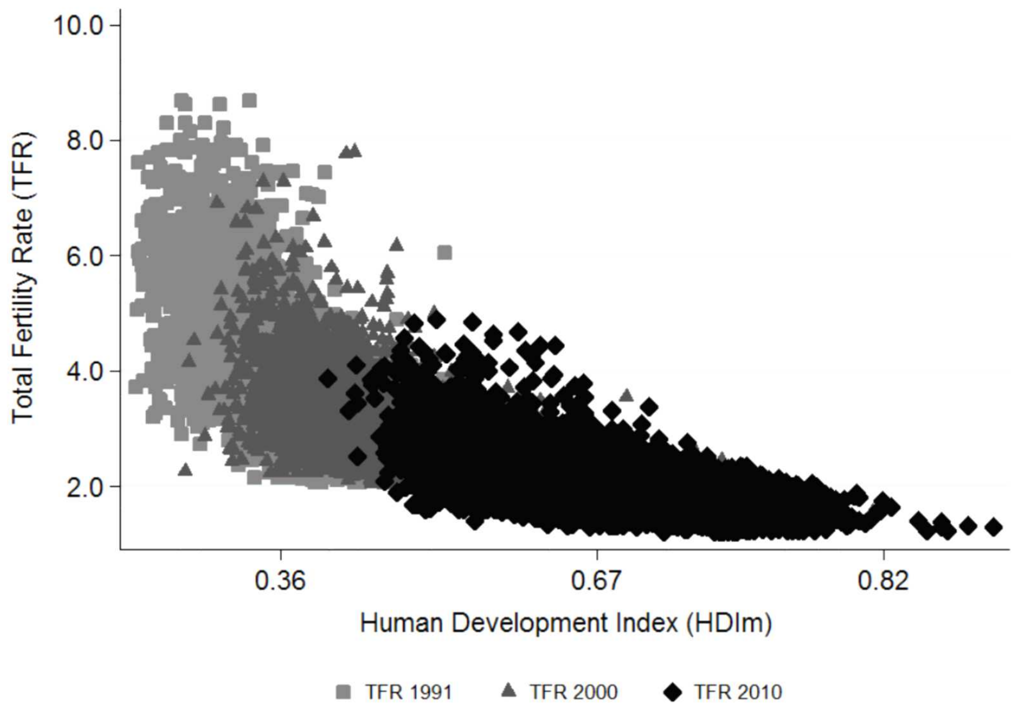
FIGURES

Figure 1. The association between total Fertility Rate and the Human Development Index for 135 countries from 1975 to 2010 and between the Total Fertility Rate and the Global Gender Gap Index in 2010



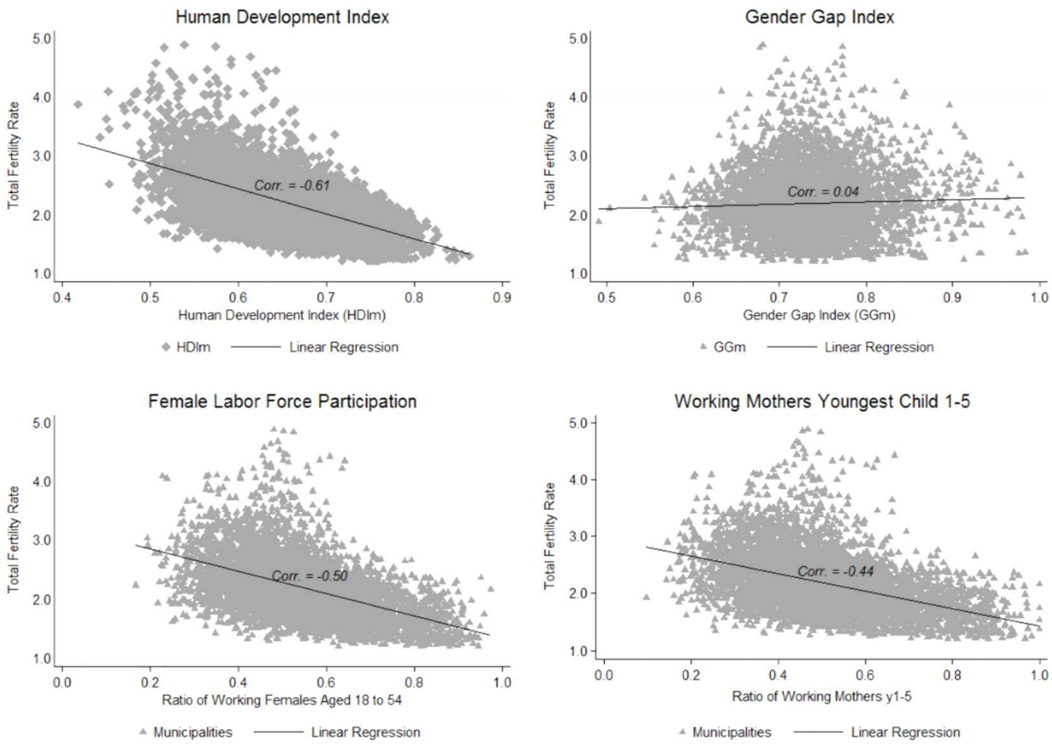
Data Source: Myrskylä, Kohler and Billari (2011) and Hausmann, Tyson, and Zahidi (2010).
Note: HDI values are in log scale [-log(1- HDI)].

Figure 2. Total Fertility Rate in relation to the Human Development Index of Brazilian's Municipalities in 1991, 2000, and 2010



Data Source: PNUD (2013).
Note: HDIm values are in log scale [-log(1- HDIm)].

Figure 3. Total Fertility Rate and Municipalities' variables (Level 2) in Brazil 2010



Data Source: IBGE, Brazilian Demographic Census of 2010 (IBGE), and PNUD (2013).

SUPLEMENTAL MATERIAL

Table A1. Multilevel Logit Models with random intercept at the municipality level estimating the probability of married or cohabiting women aged 16 to 49 years old having a first birth in Brazil in 1991, 2000 and 2010 (odds ratios)

	1991	2000	2010
Black	1.00	0.98	0.95 **
Age	0.95 **	0.92 **	0.93 **
Age Squared	1.00 **	1.00 **	1.00 **
Educ Attain (ref: Less than Primary)			
Primary Completed	0.95 **	0.96 **	0.98
High School	0.86 **	0.89 **	0.88 **
College or more	0.88 **	1.00	1.04 *
Log Income per capita (HH)	0.99 **	0.91 **	0.94 **
Rural Area	0.98	0.97 **	0.93 **
North and Northeast region	1.02	1.02	1.03
HDIm	0.77 **	1.21 *	0.62 **
GGm	0.79 *	0.97	0.91
Female Employment Ratio	1.32	1.01	0.97
Mothers y1-5 Employment Ratio	0.68 *	0.81	0.75 *
Constant	0.76 **	0.54 **	0.53 **
Rho	0.01 **	0.01 **	0.01 **
Sigma_u	0.15	0.14	0.18
Obs	325,669	425,497	569,051

Data Source: IBGE, Brazilian Demographic Censuses of 1991, 2000, and 2010, PNUD (2013), National Civil Registry, and Superior Electoral Court.

Chapter 2 Female Labor Force Participation and the Availability of Child Care Centers in Brazil

Abstract Female employment has a critical role in contributing to family income and coping with poverty in Latin American Countries. Despite the increase in FLFP in the last decade, women still face significant constraints to work. In this study, I test whether the availability of child care centers is one of these constraints. A reform in Brazil at the Federal level in 2006/2007 promoted the increase in slots availability in public child care centers. A quasi-experimental difference-in-differences design is used to assess if this exogenous increase affects the employment of mothers with the youngest child aged four or below. The results indicate that the increase in public child care availability had a positive association with female labor force participation. The impact was greater to married mothers with less than a high school degree.

1. Introduction

Female labor force participation (FLFP) has grown significantly in the last fifty years. Increases in FLFP rates are frequently related to greater female empowerment and gender equality. In developing countries, female employment has an essential role in increasing family income and coping with poverty. A study of Latin America, for instance, shows that thirty percent of the region's decrease in extreme poverty in the 2000s are attributed to increases in female labor force participation (World Bank 2012a). Despite the increases in FLFP in the last decades, women in developing countries still face significant lower participation rates than males, specially married women, and there are ongoing demands for policy designs to increase their economic activity.

An increase in the availability of child care centers is often considered a policy alternative to rise mothers' participation rates. However, estimating its effects is a challenge since the availability of child care centers in a region is usually responsive to its demand, making endogenous the association between slots' availability and mother's employment. And so, when analyzing whether an increase in slots' availability in child care centers affects mother's labor supply, it is necessary to control for the endogenous relation between the two variables, ensuring that school availability is not a direct result of mothers' demands. Few studies have considered this endogeneity, and researchers agree that more evidence is necessary in order to accumulate knowledge about the magnitude of the effect of this type of intervention (Han and Waldfogel 2001; Blau and Currie 2006). For *developing* countries, as shown by Todd (2012) in a recent assessment, the evidence is even scarcer with only a handful of studies analyzing the effect of this type of intervention on women's employment.

In this paper, the endogenous relation between slots' availability and mothers' demands is controlled by analyzing a Federal level intervention in Brazil in 2006/2007. A set of law changes in the period increased significantly the availability of slots in public provided child care centers. In 2009, Brazil had 37% more children enrolled in public child cares than in 2006, representing an increase in 5% in overall public child care attendance, and in 2013 it was 89% greater - from a total of 917,460 in 2006 to 1,730,877 in 2013 (INEP, Brazilian School Census 2006-2013). This increase occurred in all regions and states, with the highest increase in the Southeast region, which more than doubled its capacity in seven years. This extraordinary growth in enrollments is attributed to efforts at the Federal level to expand education to young children. First, the Law nº 11,274 in February 2006 changed the minimum entrance age in elementary school from seven to six years old and a constitutional amendment in December of 2006 (Constitutional Amendment nº 53/2006) limited the entrance age in preschools to a maximum of five years old. These two changes generated an increase in the availability of child preschool's slots for children five years old or younger. Another law in 2007 (Law nº 11,494/2007) extended public transfers from the National Education Fund to public and non-profit child care centers, and the program *ProInfância* (Resolution nº 6/2007) provided federal resources to municipalities in order to finance the reform and construction of child care centers and preschools - this program financed 1,721 new child care centers and preschools between 2007 and 2009.

This extraordinary increase in child care availability in Brazil can be, to a large extent, considered exogenous to mothers' demands for child care slots at the local level, because the laws were implemented at the Federal level. Therefore, this massive reform that targeted young children provide a unique opportunity to analyze the effect of increases in slots' availability in free child care centers on mothers' labor force participation using a quasi-experimental design in Brazil. Because it was a universal policy intervention, difference-in-differences is estimated to compare the change before (2004-2006) and after the reform (2007-2009) in the labor supply and hours worked of mothers with the youngest child aged four and below (treatment group) with the change in labor supply and hours worked of childless women (control group). So, the incremental change in employment outcomes of one group is measured compared to the other group. The control group is used to isolate the impact of the reform on mothers from labor market factors that affect the entire population.

This study finds that an increase in public child care attendance of approximately 5% in 2007-2009 led to an increase in the labor force participation of mothers with the youngest child in child care ages of 2.16% for married and cohabiting women and 1.2% for single women. There was also an increase of 1.78 weekly hours worked for married or cohabiting mothers. The greatest impacts were for mothers with lower education attainment, as expected since they are more likely to use public provided child care centers. Placebo estimates confirm that this results are not driven by temporal trends in labor force participation and hours worked in the treatment and control groups.

2. Background

2.1 Theoretical Framework

In order to formulate theoretical hypothesis of the increase in the availability of free child care on mothers' LFP it is convenient to separate the expectations on labor-supply at the extensive margin (participation) and at the intensive margin (hours worked). A simple static economic model of one-person labor supply is useful to make the theoretical predictions in each margin. Based on Blau (2003), I assume that the mother is the only care taker of the children and for every hour of work, one hour of child care is necessary. The budget constraint is $c = I = y + (w - p)h$, in which *c* is consumption other than child care, *I* is income net of child care expenditure, *w* is hourly wage, *p* is the hourly price of childcare, *h* hours worked, and *y* nonwage income. At the extensive margin, mothers would work when the hourly net market wage $(w - p)$ is greater than the reservation wage (the lowest wage rate in which she is willing to accept a job offer). When a free child care slot is used, *p* is zero and there is an increase in the net market wage. So, with free child care, mothers are more likely to meet their reservation wage at the same market wage *w* and, consequently, more likely to work.

At the intensive margin, the predictions are ambiguous depending of mothers' work status before and after the reform. For those not working before the reform, an increase in hours worked is expected, while for mothers working before the reform, the effect of an increase in the availability of public child care slots depend on their previous child care arrangement. No effect is expected if they are not paying for child care before the reform, and a positive, negative or no effect is expected if they are paying. A negative effect is observed if they opt to maintain their income levels *I*, and increase leisure hours, resulting in a substitution effect. A positive effect is expected if working mothers paying for child care before the reform are willing to work more hours because of the increase in the net wage rage, an income effect. Lastly, no change is expected if mothers prefer to maintain the same number of hours and present no response to changes in net income.

In sum, the availability of free of cost child care slots are likely to increase female labor force participation at the extensive margin. At the intensive margin, an increase, decrease or no effect in hours worked is expected depending on leisure preferences and child care arrangements before the reform.

2.2 Previous Empirical Research

The association between motherhood and labor force participation has long been recognized in the literature (Stycos and Weller 1967). Motherhood requires a pause in labor that can range from weeks, to months or years. Generally, the longer the pause, the higher the skill depreciation and the loss of professional connections. The difficulty in reconciling work and family is usually more dramatic in the early ages of the offspring, since young children need constant supervision and child care centers require greater worker per child ratio, and so are typically more expensive and scarcer than preschools or regular schools. The difficulty in reconciling work and the care of children is greater to mothers with lower levels of education, because their jobs usually require longer journeys (for example in factories, retail stores, and restaurants), have less flexibility in hours worked (Swanbeg et al. 2005), and constant supervision. Informal work generates greater flexibility in the hours worked, however they are usually lower paid and do not include fringe benefits.

In the last decades, a great amount of research has analyzed the associations between female labor force participation and child care availability. However, few studies have considered the endogeneity of this association, controlling for the fact that greater demands for child care centers can generate its supply. In this review, I first assess the research evidence in *developed* countries, exploring the literature that controls for this reverse causality. Then, I review the literature in *developing* countries, and in Brazil. The main objective here is to analyze the types of programs, methods of analysis, and effects on different groups of the population. The effects regarding subpopulations are mixed and they depend in a great extent on the population in which unmet child care needs are greater and on the design of the policy or subsidy.

In the United States, three authors have assessed the impact of public kindergartens on female labor force participation and they find mixed results. Gelbach (2002) uses the 1980 US Census and quarter of births (QOB) as an instrumental variable, finding a 10% increase in baseline employment and hours worked for single mothers whose youngest child is five years old, the effect is similar for married mothers. Cascio (2009), uses the 1950-90 US Censuses, but instead of using QOB, she explores the variation in the introduction of kindergarten funding across states. The author finds that single mothers with the youngest child at age five have a 7.5% greater likelihood of working and additional 2.78 weekly hours of work compared to states where the policy was not available. But, differently from Gelbach (2002), she encounters no effect on labor force outcomes of married mothers. In a recent assessment, Fitzpatrick (2012) uses the 2000 Census with a regression discontinuity based on the age of the child relative to the kindergarten's eligibility cutoff date. The author finds that public provided kindergartens increase the employment of single mothers with the youngest child aged five years old by 12.2%, but no effect in encountered for married mothers or hours worked. Although the results are consistent for single mothers, for married mothers they are mixed. The different findings across studies is explained by the period analyzed and the corresponding changes in female employment's elasticity.

Authors have also analyzed the impact of child care subsidies in the United States. Starting in 1996 as part of the welfare reform, low income mothers can receive subsidies for child care if employed or in education and training. Bainbridge et al (2003) use 1992-1997 CPS data to all unmarried women aged 16 to 44 years old that were not attending school at the time of the survey. The authors find that an increase in a $1,000 in the annual child care subsidy per single mother with a child under 13 years old, increase in 11% the employment probability. When comparing with Earned Income Tax Credit (EITC) benefit, a dollar spent in child care subsidy generated greater increases in employment than a dollar spent on tax benefits. Similarly, Han et al (2009) estimate difference-in-differences comparing families in which parents have no high school diploma (treated) to families with more education (control) interacted with the expenditures of the Child Care Development Fund (CCDF) per children younger than six in 2000. They find that an extra thousand dollars of CCDF expenditures is associated with 4% increase in employment.

Blau and Tekins (2007) using a sample of single mothers in the 1999 National Survey of America's Families and two-stage least squares, estimates that mothers receiving child care subsidies are 5% more likely to be working than single mothers not receiving subsidies, when controlling for family characteristics. Finally, Meyers et al. (2002) focusing on child care subsidies in California in 1995 estimates the probability that a women would receive subsidy and measure its impact on employment. They find that if 10% of mothers were subsidized we would observe an increase in their employment of 30%, and if 50% were subsidized and increase of approximately 75%, drawing attention to the non-linearity of this relation. Studies in the United States have also analyzed the price elasticity of labor for participation in relation to child care prices, with elasticities of -0.3 to -0.4 for married women and -0.5 to -.73 for unmarried women (Han and Waldfogel 2001).

In Canada, a child care subsidy that started in 1997 reduced child care prices to C$ 5.00 a day per child in the state of Quebec. This policy provided a great opportunity to measure with a quasi-experimental design the effect of child care subsidies on labor force participation. Lefebvre and Merrigan (2008) using the data of the 1993-2002 Survey on Labor Income Dynamics (SLID) and difference-in-differences estimates, compared mothers with at least one child aged 1 to 5 years old in Quebec with the rest of Canada finding a 13% increase in participation and a 22% increase in hours worked in 1999-2002 resultant from the policy implementation. Using another database, the Canadian National Longitudinal Survey of Children and Youth (NLST), Baker et al. (2008) found similar impact with the subsidies increasing the participation of women in two-parent families by 14.5% compared to baseline participation.

Germany also had a significant expansion in public subsidized child care centers in the 1990s, with children older than three years being entitled to attend half-day child cares. Bauernschuster and Schlotter (2013) analyze this policy implementation with date of birth instrumental variables and difference-in-differences method comparing the pre-treatment (1996) and post-treatment years (2001), with both estimates providing consistent results. When using instrumental variables, the authors find that child care subsidies were responsible for an increase of 6.5% in participation and 2.5 in weekly hours worked. In the difference-in-difference estimates there was an increase of 7% in employment when comparing treated mothers before and after the policy implementation with childless women aged 18 to 60 years old.

In Spain in 1991 there was also an important expansion in subsidized child care for children 3 years old, with an increase in the public enrollment of 3 years old from 8% in 1990 to 47% seven years later. Nollenberger and Rodriguez-Planas (2011) analyze this expansion comparing mothers whose youngest child is three years old (treatment group) from mothers whose the youngest is two years old (control group). A *dummy* variable indicating the treatment and control group is interacted with a variable from the post-implementation period providing the difference-in-differences estimate. The authors find an increase in employment of 8% and in hours worked of 9%, with the effect being greater for mothers with lower education. There was no effect of the Spain's child care reform on college educated mothers, because of their greater possibility to pay for private care and, consequently, their smaller unmet needs for subsidized care.

Havnes and Mogstad (2011) analyze the implementation of public child care in Norway after the passage of the Norway Child care Act in 1975. Using a difference-in-differences approach in the period of 1976 to 1979 they compare married mothers with the youngest child aged 3-6 years old with married mothers with the youngest child aged 7-10 years old in municipalities that had an expressive expansion in child care centers with municipalities that had no expansion. The authors found that the increase in 17.85% in child care coverage in the treated area led to an increase in 1.1% in employment rates, with the most educated mothers being more benefited from the policy. They authors conclude that this effect was negligent for the size of the expansion. Schlosser (2011) analyzes the introduction of free public preschool for children aged 3 to 4 years old in Israel. The author estimates a difference-in-differences comparing mothers of children aged 2-4 years old in towns that had the expansion in child care with towns that did not have the expansion. Mothers in 'treated' towns had an increase in 8.1% in employment and 2.83 in weekly hours worked. The effect was greater for more educated mothers as well.

Most of the studies about the impact of child care provision on labor force participation are from *developed* countries or from small scale policy interventions in *developing* countries. The only study from a *developing* country with country-level pre-school policy intervention is Berlinski and Galiani (2007) which analyzes the impact of a large-scale construction of preschool facilities in Argentina between 1994 and 2000. This policy increased preschool enrollment by 18% , employment levels by 7 to 14% and hours worked by 2.24 to 4.5 hours per week. The effects on mothers' employment were greater for women with spouse present in the household. These results can be enlighten to Brazil, however the levels of female labor force participation in Argentina are very different from Brazil, with for instance, Argentina having 48% of its female population employed in 2010-2014 and Brazil 59% in the same period (World Bank 2015). And so, it remains an open question whether the child care reform in Brazil has similar impacts as the reform in Argentina and, which are the education subgroups that are going to be most affected by the policy change.

In Brazil, there is evidence from a random lottery experiment among low income mothers in the city of Rio de Janeiro in 2008. A total of 24,000 mothers registered to receive public provided child care and 2,174 were randomly selected to have access to the program. Barros et al. (2011) compared treated and control groups and found an increase in 28% in mother's employment in the group that was able to register the children in public child care centers of Rio de Janeiro. Given the role of mothers' income in increasing family living standards and coping with poverty and misery (World Bank, 2012) this is an important result. The large fraction of mothers that remained in the queue (approximately 91% of the registered mothers) suggests a large unsatisfied demand for child care among low income mothers in city of Rio de Janeiro. This lottery experiment was focused on low income mothers in the city of Rio de Janeiro. The present study contributes to the literature by using a national representative dataset of Brazil and measuring the country-level evidence of the Child Care Reform.

In general, we can observe sizeable positive effects of policies targeted to increase the availability of subsidized child care centers and kindergartens on mothers' LFP and hours worked. The characteristics of the population most affected by these changes vary considerably across studies, with some authors finding greater effect on single mothers with lower education levels, while others encounter larger effects on married mothers with higher education levels. The subpopulations affected in these type of policies largely depend on the labor market conditions of the population analyzed, the availability of informal and private care, initial employment levels across groups, and the quality and location of the new slots in child care centers.

This article focuses on preschoolers' mothers in the context of a *developing* country. Brazil is a good case of study not only because of the law change that provides the opportunity to perform a quasi-experimental analysis, but also because of its highly unsatisfied demand for child care slots. In this study, it is hypothesized that the growth in child care availability, produced by a change in the law, increased the employed and average weekly hours worked by mothers. In the next section, more details about the Federal law change are provided and its impact on enrollments in public child care centers is analyzed.

2.3 The Child Care Reform in Brazil

In 2006/2007 the Brazilian government had a set of unprecedented actions to expand the availability of publicly provided child care's slots to children in preschool ages. In Brazil, the provision of public child care centers is a competency of the municipality, however, in this circumstance, the Federal government influenced and promoted its expansion with three main actions. First, the Law nº 11,274 in February 2006 changed the minimum entrance age in elementary school from seven to six years old. Then, a constitutional amendment in December of 2006 (Constitutional Amendment nº 53/2006) limited the entrance age in preschools to a maximum of five years old. These two changes generated an increase in the availability of child preschool's slots for children five years old or younger.

In 2007, another Federal action was implemented with a law that extended the transfers from the National Education Fund to public and non-profit child care centers (Law nº 11,494/2007). Until 2006, the Fund transferred resources to schools based only on its enrollments in elementary education, but with the 2007 law it started transferring resources to public child care centers and preschools. The last Federal measure to be highlighted is the implementation of the program *ProInfância* in 2007 (Resolution nº 6/2007). This program provides resources to finance the renovation and construction of public child care centers and preschools. It also offers resources to the purchase of furniture and equipment, such as tables, cribs, and stoves. Between 2007 and 2009 *ProInfância* financed 1,721 new public child care centers and preschools in Brazil. In this paper I focus on the changes at the level of child care centers, for children with four years and below completed in the school year. These political changes in 2006 and 2007 at the Federal level will be denominated hereafter the "Child Care Reform".

Figure 1 shows the enrollments in public child care centers in Brazil between 2004 and 2009. A steep increase resultant of the child care reform can be observed in 2007. Enrollment in public child care centers increased 37%, from a total of 917,460 in 2006 to 1,252,765 in 2009. This value, however, still far beyond the total number of children in child care ages in 2009. The total enrollment in public child care centers of children that completed four years old or less during the school year in relation to the total number of children in this age range, increased from 10.6% in 2006 to 14.9% in 2009, a total of 4.3% increase. The policies implemented with the child care reform in 2007, promoted a continued expansion in public child care enrollments in Brazil and in 2013 there was a total of 1,730,877 children enrolled in the system. This analysis is restricted to the first period of reform implementation (from 2007 to 2009) since it presents the highest growth in enrollments.

3. Data, Variables, and Method

3.1 Data

The following analysis uses the 2001-2009 National Household Sample Survey (PNAD) which is an open access survey of the National Institute of Geography and Statistics (IBGE). It is a stratified probability sample of Brazilian households with approximately 150,500 households interviewed every year. The data is cross-sectional, so it does not follow the same individual over time. In 2004 the survey included in its sample the rural areas of the Northern region of Brazil which represented two percent of the total population in the country. To maintain historical compatibility between 2001-2003 and 2004-2009, this Northern' rural population will not be considered in this analysis. Consequently, the results presented here are representative of Brazil, with the exception of the rural areas in the states of Rondônia, Acre, Amazonas, Roraima, Pará, and Amapá.

I restrict the analytical sample to women aged 16 to 44 years old at the survey's reference date, which consists of 98% of the mothers with the youngest child aged four and below. In addition, because our main goal is to observe whether there is an increase in the labor force participation of mothers that were neither working nor studying, I restrict the sample to women not attending school, which consists of 90.5% of mothers with the youngest child aged four and below. PNAD has a sizeable sample with a total of 191,155 non-studying mothers aged 16 to 44 years old with the youngest child aged four and below, and approximately 21,000 observations per year. These numbers have decreased in the 2000s because of the fertility decline. In 2001, there was approximately 22,000 non-studying mothers with the youngest child aged four and below, and nine years later this number was 14% smaller. Likewise, the number of childless non-studying women aged 16 to 44 years old increased, going from 17,256 in 2001 to 21,013 in 2009, an increase of 22%.

Married and single women differ considerably in their employment and child care choices, with married and cohabiting having greater possibility to choose between working in the labor market and staying at home. Consequently, the statistical analysis is performed separately in a sample of women living with a partner (married or cohabiting) and in a sample of women not living with a partner (single). In order to measure the impact of the increase in childcare availability on labor force participation and hours worked controlling for general labor market trends, I compare mothers with the youngest child aged four years old and below (treatment) with a control group. A good control should be comparable to the treatment group in the extent that they respond similarly to labor market shocks, but they do not receive the treatment. In this paper, I follow Eissa and Liebman (1996), and others, and use childless women as the control group for a universal policy implementation regarding mothers.

In 2004-2006, the total non-studying married women aged 16-44 years old in the treatment and control group is 131,397 and single is 115,360. Table 1 provides the size (*n*) of these two samples in the control and treatment group. We can observe that there is a total of 121,803 childless non-studying women aged 16 to 44 years old and 124,954 non-studying mothers aged 16 to 44 years old with the youngest child aged four and below. The two groups have, in average, similar ages with the mothers having a mean of 28 years and the childless women 26. However, they differ considerably by levels of education attainment, with childless women being two times more likely to have college degree than the treatment group, with the first group having 15% of women with college degree and the latter 6%. The household income from sources other than the women's own labor supply is also greater to childless women, with them having a mean of R$1,520 and mothers having a mean of R$856,00. This may be related to the fact that our treatment are more likely to have left the parent's house and started a new household, as can be observed by the share of women not living with a partner in the two samples.

The dependent variable intended to measure labor force participation at the extensive margin is one if the person has worked in the week of reference or was on vacation or on leave, and zero otherwise. At the intensive margin, the number of weekly hours usually worked is used. In relation to the control variables, there is a *dummy* indicating whether the person classified themselves as black or *pardo* (mixed-race), age centered at the mean, age squared to control for nonlinearities between employment and age, a *dummy* variable indicating whether the person has moved to other municipality in the last four years, a dummy variable indicating whether the person lived in rural areas, the log of the household income from sources other than the women's own labor supply, state fixed effects, and year fixed effects.

I also control for whether the household has received conditional cash transfers from the government, using a proxy based on the value received from non-wage transfers. Following Foguel and Barros (2010), this variable is estimated using a question asking if the household receives transfers from social programs or if it receives dividends from financial assets in the month of reference (September of every year), and, if it receives, the value of the transfer. The *Bolsa Família* program was created in October 2003, and it gradually substituted pre-existing government programs (*Auxílio Gás, Bolsa Escola, Bolsa Alimentação, and Cartão Alimentação*)[8]. From 2001 to 2003, the sum of the maximum value of these pre-existing programs are used to identify whether the household received conditional cash transfers. For the subsequent years, the values are based on the maximum *Bolsa Família* values in the period, which are R$ 58.00 in 2001-2003, R$95.00 in 2004-2006, R$ 172.00 in 2007, R$182.00 in 2008, and R$200.00 in 2009.

3.2 Methodological Approach

To test whether the child care reform affected the labor outcome of non-studying mothers with the youngest child aged 0 to 4 years old (treatment group), I use a difference-in-differences estimate. The treatment group is compared with non-studying childless women (control group) before (2004-2006) and after the child care reform (2007-2009). The control group is used to isolate the effect of the reform from other labor market factors affecting the entire population. The incremental change in employment outcomes from the treated group in the period after the reform, compared with the control group, provides the effect of the reform.

The first difference-in-differences performed is on descriptive participation results in which we observe, without controlling for demographic covariates, whether there was an incremental increase in labor force participation and hours worked for the treatment group compared with the control group in the period after the reform. The participation rates and mean hours worked (including women with zero hours worked) are obtained using the sample design and the standard error of each estimate is reported in the table. Because the differences observed can be generated by changes in the demographic composition of the treatment and control groups before and after the reform, it is important to control for demographic characteristics, and time and state dummies.

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[8] The Program of Child Labor Eradication (PETI) is not considered in this estimation because its values are usually the same as *Bolsa Família* and a family can receive one or the other, they are not cumulative.

After the basic descriptive approach, I control for covariates using the difference-in-differences method in a regression model to estimate, first, the effect of the child care reform in the probability of working (extensive margin) and, then in the average hours worked (intensive margin). When analyzing effects of the program in the extensive margin I use probit models, and for the intensive margin, tobit models[9]. Tobit models are preferred over Ordinary Least Square Models (OLS) because the distribution of hours worked is truncated at zero, and we can't assume normality. In addition, if we eliminate women that do not work from our sample (zero hours) and estimate an OLS regression we would incur in the omitted variable bias (Heckman 1979), since women are not randomly selected in nonworking and working categories. In the tobit model, a change in the independent variable xi increases the conditional mean of hours worked and the probability of the observation falling in the positive part of the distribution (Greene 2002). The standard errors of the probit and tobit models are estimated with the Huber-White standard errors correction, adjusting for heteroskedacity.

4. Results

Table 2 displays the descriptive results of the difference-in-differences estimates for labor force participation and hours worked of non-studying women aged 16 to 44 years old in Brazil for women living with a partner (married or cohabiting) and women not living with a partner (single). The first difference consists in subtracting hours worked or LFP after (2007-2009) and before (2004-2006) the child care reform. Difference-in-differences are obtained by subtracting the

[9] OLS models were estimated for weekly hours worked, and the variables presented similar signs and statistical representativeness. The exception was for rural areas in the regression of married or cohabiting mothers, which was negative in the OLS regression and positive in the tobit regression with both being statistically significant. In the regression for single mothers, the variable indicating whether the household receives conditional cash transfer was different in the tobit and OLS analysis for single mothers with a negative sign and a p-value of 0.157 in the first and negative and with a p-value of 0.006 in the OLS.

change in 2006-2007 (first difference) of the treatment group (mothers with the youngest child aged four years and below) by the change of the control group (childless women).

We can observe that there was an increase in female labor force participation of mothers with the youngest child aged four and below (treatment group) of 2% and 0.818 in hours usually worked in a week relative to difference of childless mothers in the period, a remarkable increase since the overall increase in child care attendance for this age group was 4.5% during 2007-2009. When the sample is disaggregated by education levels, we can observe that the increase in hours worked and LFP was greater to married mothers with less than high school, with a relative increase in 2007-2009 of 0.974 hours and 3.4% in participation. Mothers with high school degree also observed an increase of 1% in participation and 0.48 in hours. For mothers with college degree we do not observe an increase in participation and hours worked in 2007-2009, but rather a decrease. These results are expected given that mothers with lower education are more likely to use public child care centers. In relation to single non-studying mothers aged 16 to 44 years old we can observe an increase in participation in 2007-2009 of 0.6% for mothers with less than high school and 1% for mothers with completed high school, compared to the difference of childless non-studying women aged 16 to 44 years old in the same period. Despite the increase in labor force participation, there was no increase in hours worked by single mothers compared to single childless women.

Table 3 presents the difference-in-differences estimates for the probit and tobit model estimating the labor force participation and hours worked of non-studying mothers with the youngest child aged four years old or younger compared to childless women before and after the Child Care Reform in Brazil by companionship status. The difference-in-differences estimate is given by the interaction between a *dummy* that is one for the treatment group and zero for childless women and an indicator variable for the period of the intervention (Child04 x Pos2007). This interaction provides the incremental change in employment outcomes in 2007-2009 of one group compared to the other.

When controlling for education, race, age, migration status, other household income, conditional cash transfer status, state, and rural areas, Table 3 shows a 2.16% increase in the probability of working for non-studying married mothers with the youngest child aged four years old or younger compared to childless women in 2007-2009. For single women in the treatment group, the increase in LFP with expansion of public child cares in 2007-2009 is 1.2%, statistically significant only when $p<0.10$. Table 4 provides the probit estimates on LFP by education samples. We can observe that this increase was mainly driven by married mothers with less than high school degree, in which the average increase in 2007-2009, compared to childless women, was 3.8% . Married mothers with high school degree, had a 0.9% increase but it was not statistically significant at $p < .10$. For single women, the average increase in LFP was of 2.1% for women with less than high school degree and 2.3% for women with high school degree, both are statistically significant at $p < .05$ only.

Table 3 shows the estimation of hours worked from the tobit models by companionship status. There is an increase of 1.8 weekly worked hours for married or cohabiting mothers and an increase of 0.4 for single mothers, with the effect being statistically significant at $p<0.001$ for married or cohabiting mothers, and not statistically significant at $p<0.1$ for single mothers. Table 5 presents the estimates of hours worked by education level. The increase in hours worked in the treatment group compared to the control group was mainly driven by the sample with less than high school degree, in which the number of hours worked in a week increased by almost three hours. For single mothers, the number of hours worked increased by 1.6 hours for mothers with less than high school degree and by 1.2 hours for mothers with high school degree, but the coefficients are statistically significant only at $p<.10$.

Finally, in order to observe whether the results above are driven by underlying temporal trends, a placebo model is estimated comparing the treatment and control group in 2004-2006 and 2001-2003, previous to the reform. I assume that the intervention happened during 2004-2006. Table 6 present the estimates of the placebo intervention. There is no statistical significant change in labor force participation for married or single mothers. However, for the numbers of hours usually worked in a week, there is a statistical significant increase of 1.05 hours for mothers compared to childless women in 2004-2006, which is statistically significant at *p<.05* . It is a smaller and less significant effect compared to the observed in 2007-2009. When the regressions are estimated by education levels, the group of married women with the lowest p-value in hours worked was college educated women, which had an average increase of 1.38 hours compared to childless women, but the p-value was 0.100, and so not statistically significant at *p<.10*. And so, the effect observed at the aggregate level, with the increase in hours worked of married or cohabiting women, is mainly driven by women with the highest education attainment rather than with the lowest education attainment as observed in 2007-2009. And so, these results provide confidence that the estimates found in the 2004-2009 period are not driven by different underlying temporal trends in labor force participation and hours worked in the treatment and control groups.

In summary, I find that the child care reform in Brazil had a positive impact in the extensive and intensive margins of labor force participation for mothers with the youngest child in child care ages. The greatest impacts were for married mothers with lower education attainment. For single mothers, we could also observe an impact, however they were less significant than the impact observed for married mothers. The placebo estimates confirm that this results are not driven by a different temporal trend in labor force participation and hours worked in the treatment and control groups. Compared to the estimates reviewed in the background section, the increase in 2.16% for married or cohabiting mothers and 1.2% for single mothers and the increase of 1.78 in weekly hours worked for married or cohabiting mothers are sizeable given that the total increase in public child care enrollment rate for children aged 0 to 4 years old was 5%.

5. Discussion

Female labor force participation (FLFP) has increased in *developing* countries; however, mothers still face several constraints. The literature shows that preschool availability can be one of these constraints since a shortage of preschools reduces the possibility of mothers having access to formal child care (Gelbach, 2002; Barros et al., 2011). In this paper, I assess whether an exogenous increase in child care availability in Brazil is associated with the employment of mothers with the youngest child in child care ages (age four and below). The increase in the availability of public child care centers in 2007/2009 was motivated by a national-level reform and can be considered, to a large extent, exogenous to mothers' demand for child care slots at the local level.

The database used is the annual household survey (PNAD) with pooled cross-sectional data from 2001 to 2009. Difference-in-differences were estimated comparing mothers with the youngest child age four and below with childless women before (2004-2006) and after (2007-2009) the Brazilian's child care reform, by companionship status. The increase in the likelihood of working in 2007-2009 was estimated using a probit regression and the increase in weekly worked hours using a tobit regression. These regressions were also estimated for education samples. For married or cohabiting mothers, the increase in slots availability in public child care centers was positively associated with the likelihood of working in 2007-2009, increasing the number of working mothers in the labor force by 2.16% and 1.8 weekly worked hours. While school availability affected the probability of working and hours worked of married and cohabiting mothers, the effect was smaller on single mothers, with an increase of 1.2% in labor force participation. For single, married, and cohabiting mothers, the effect of the policy was greater to mothers with lower education attainment, specially for those with less than high school degree. Finally, placebo estimates show that the effect is not driven by temporal trends in the employment of mothers with the youngest child in child care ages.

TABLES

Table 1. Labor force participation rates and usual weekly hours worked for non-studying childless women (control) and non-studying mothers with the youngest child aged 0 to 4 years old (treatment) and its difference and difference-in-differences before (2004-2006) and after (2007-2009) the childcare reform by companionship status and education levels

	2004-2006		2007-2009		Difference		Difference-in-differences	
	LFP	Hours	LFP	Hours	LFP	Hours	LFP	Hours
Married or Cohabiting								
Mothers 0-4 [n=96,962]	0.479	16.163	0.501	17.353	0.022	1.190	0.020	0.818
	(0.004)	(0.188)	(0.004)	(0.206)	(0.004)	(0.279)	(0.007)	(0.467)
Childless [n=34,435]	0.672	26.616	0.674	26.988	0.002	0.371		
	(0.005)	(0.265)	(0.005)	(0.265)	(0.006)	(0.375)		
Married or Cohabiting, by education levels								
Less than High School, mothers 0-4 [n=61,109]	0.420	12.941	0.432	13.554	0.012	0.613	0.034	0.974
	(0.006)	(0.206)	(0.006)	(0.235)	(0.005)	(0.184)	(0.011)	(0.455)
Less than High School, childless [n=12,680]	0.546	19.929	0.524	19.568	-0.022	-0.361		
	(0.008)	(0.395)	(0.009)	(0.413)	(0.010)	(0.428)		
High School Completed, mothers 0-4 [n=29,428]	0.541	20.441	0.546	20.755	0.005	0.314	0.010	0.480
	(0.005)	(0.233)	(0.005)	(0.222)	(0.007)	(0.290)	(0.010)	(0.471)
High School Completed, childless [n=16,330]	0.707	29.200	0.703	29.034	-0.004	-0.165		
	(0.006)	(0.313)	(0.006)	(0.295)	(0.008)	(0.379)		
College or more, mothers 0-4 [n=6,425]	0.839	32.235	0.826	31.441	-0.013	-0.793	-0.010	-0.435
	(0.008)	(0.381)	(0.007)	(0.342)	(0.010)	(0.496)	(0.014)	(0.722)
College or more, childless [n=5,425]	0.895	36.296	0.891	35.937	-0.003	-0.359		
	(0.007)	(0.409)	(0.006)	(0.349)	(0.009)	(0.527)		
Single								
Mothers 0-4 [n=27,992]	0.547	21.165	0.556	21.192	0.010	0.027	-0.001	-0.514
	(0.006)	(0.244)	(0.006)	(0.257)	(0.007)	(0.354)	(0.008)	(0.434)
Childless [n= 87,368]	0.651	26.313	0.662	26.854	0.011	0.541		
	(0.004)	(0.178)	(0.004)	(0.177)	(0.004)	(0.251)		
Single, by education levels								
Less than High School, mothers 0-4 [n=18,040]	0.497	18.634	0.493	18.008	-0.004	-0.626	0.006	-0.364
	(0.007)	(0.282)	(0.007)	(0.322)	(0.009)	(0.375)	(0.011)	(0.497)
Less than High School, childless [n=24,842]	0.485	18.903	0.475	18.641	-0.010	-0.262		
	(0.006)	(0.274)	(0.006)	(0.276)	(0.007)	(0.323)		
High School Completed, mothers 0-4 [n=8,855]	0.625	25.426	0.634	25.354	0.009	-0.072	0.010	-0.032
	(0.009)	(0.411)	(0.008)	(0.376)	(0.012)	(0.546)	(0.013)	(0.579)
High School Completed, childless [n=49,088]	0.687	28.294	0.686	28.254	0.000	-0.040		
	(0.004)	(0.190)	(0.004)	(0.201)	(0.005)	(0.227)		
College or more, mothers 0-4 [n=1,097]	0.854	34.748	0.814	32.086	-0.040	-2.663	-0.035	-2.807
	(0.018)	(0.990)	(0.017)	(0.817)	(0.025)	(1.257)	(0.026)	(1.322)
College or more, childless [n=13,438]	0.870	34.526	0.866	34.670	-0.005	0.144		
	(0.005)	(0.276)	(0.004)	(0.240)	(0.007)	(0.343)		

Data Source: Brazilian National Household Survey (Pnad) from 2004 to 2009.
Note: Standard errors in parentheses. "Difference" is the difference in LFP or hours worked after (2007-2009) and before (2004-2006) the child care reform. And "Difference-in-differences" is the subtraction of the "Difference" between the treatment (mothers 0-4) and the control group (childless).

Table 2. Estimating labor force participation and usual weekly hours worked of non-studying mothers with the youngest child aged four years old and below (Child04) compared to childless women before and after the education reform (Pos2007) in Brazil by marital status

	Labor Force Participation *(Probit Estimates: Marginal Effects)*		Hours Worked Week *(Tobit Estimates)*	
	Married	Single	Married	Single
Black or Pardo	-0.009 ***	-0.012 ***	-1.075 ***	-0.997 ***
Educ Attain (ref. Less than High School)				
High School Completed	0.133 ***	0.171 ***	11.642 ***	12.532 ***
College or More	0.362 ***	0.319 ***	23.726 ***	18.257 ***
Age	0.014 ***	0.012 ***	1.164 ***	0.890 ***
Age squared	-0.001 ***	-0.001 ***	-0.070 ***	-0.072 ***
Moved in the Last Four Years	-0.053 ***	-0.010 *	-4.061 ***	0.126
Other HH Income (log)	-0.009 ***	-0.024 ***	-0.630 ***	-1.638 ***
Conditional Cash Transfer	-0.016 ***	0.008 **	-3.558 ***	-0.437
Rural Areas	0.134 ***	0.016 ***	4.039 ***	-2.677 ***
Child04	-0.165 ***	-0.115 ***	-14.007 ***	-8.989 ***
Pos2007	-0.019 ***	0.002	-1.119 **	0.142
Child04 x Pos2007	0.022 ***	0.012 *	1.785 ***	0.434
Year 2004	-0.006	0.007	-0.217	0.606 *
Year 2005	-0.001	0.007	-0.159	0.459
Year 2008	0.009 **	0.012 **	0.710 **	0.896 ***
Year 2009	0.002	-0.001	0.166	0.054
Log pseudolikelihood	-82079.2	-69463.4	-386546.3	-383794.5
Pseudo R-squared	0.10	0.09	0.02	0.02

Data Source: Brazilian National Household Survey (Pnad) from 2004 to 2009.
Note: Robust Standard errors reported. All regressions include state fixed effects.
*p < .10; **p < .05; ***p < .01

Table 3. Probit Results: Estimating labor force participation of non-studying mothers with the youngest child aged four years old and below (Child04) compared to childless women before and after the education reform (Pos2007) in Brazil by marital status and education sample

	Labor Force Participation *(Probit Estimates: Marginal Effects)*					
	Less than High School		*High School Completed*		*College or more*	
	Married	Single	Married	Single	Married	Single
Black or Pardo	-0.010 **	0.006	-0.013 **	-0.028 ***	0.016 **	0.001
Age	0.015 ***	0.011 ***	0.014 ***	0.014 ***	0.011 ***	0.011 ***
Age squared	-0.001 ***	-0.001 ***	-0.001 ***	-0.001 ***	-0.001 ***	-0.001 ***
Moved in the Last Four Years	-0.044 ***	0.016 *	-0.062 ***	-0.028 ***	-0.057 ***	-0.035 ***
Other HH Income (log)	-0.006 ***	-0.026 ***	-0.013 ***	-0.022 ***	-0.017 ***	-0.018 ***
Conditional Cash Transfer	-0.008 *	0.021 ***	-0.056 ***	-0.012 *	0.022	0.029 *
Rural Areas	0.167 ***	0.049 ***	0.028 ***	-0.030 ***	-0.016	0.032
Child04	-0.165 ***	-0.096 ***	-0.184 ***	-0.165 ***	-0.068 ***	-0.085 ***
Pos2007	-0.032 ***	0.009	-0.011	-0.001	-0.001	-0.018 *
Child04 x Pos2007	0.038 ***	0.021 **	0.009	0.023 **	-0.003	-0.022
Year 2004	-0.005	0.016 **	-0.011	0.001	-0.002	-0.005
Year 2005	-0.002	0.006	-0.003	0.010	0.009	-0.004
Year 2008	0.014 **	-0.006	0.007	0.024 ***	-0.008	0.012
Year 2009	0.008	-0.021 **	-0.003	0.011 *	-0.004	0.011
Log pseudolikelihood	-47625.2	-28473.4	-29523.4	-35080.4	-4634.4	-5587.9
Pseudo R-squared	0.06	0.04	0.04	0.05	0.03	0.05

Data Source: Brazilian National Household Survey (Pnad) from 2004 to 2009.
Note: Robust Standard errors reported. All regressions include state fixed effects.
*p < .10; **p < .05; ***p < .01

Table 4. Tobit Results: Estimating the number of usual weekly hours worked of non-studying mothers with the youngest child aged four years old and below (Child04) compared to childless women before and after the education reform (Pos2007) in Brazil by marital status and education sample

	Hours Worked Week (*Tobit Estimates*)											
	Less than High School				*High School Completed*				*College or more*			
	Married		Single		Married		Single		Married		Single	
Black or Pardo	-1.211	***	0.156		-1.230	***	-1.981	***	1.345	***	0.010	
Age	1.320	***	0.994	***	0.999	***	0.927	***	0.773	***	0.811	***
Age squared	-0.082	***	-0.088	***	-0.060	***	-0.075	***	-0.042	***	-0.050	***
Moved in the Last Four Years	-3.899	***	2.471	***	-4.453	***	-1.186	**	-3.289	***	-0.295	
Other HH Income (log)	-0.534	***	-2.485	***	-0.833	***	-1.416	***	-0.703	***	-1.027	***
Conditional Cash Transfer	-2.887	***	1.253	**	-5.995	***	-1.652	***	1.039		0.197	
Rural Areas	7.002	***	-0.083		-3.082	***	-5.443	***	-2.316	**	-1.097	
Child04	-16.101	***	-10.492	***	-14.664	***	-11.365	***	-5.530	***	-4.315	***
Pos2007	-1.936	**	0.736		-0.655		-0.173		-0.485		-0.844	
Child04 x Pos2007	2.976	***	1.593	*	0.728		1.241	*	-0.148		-2.260	*
Year 2004	0.032		1.698	**	-0.728		0.106		-0.351		-0.815	
Year 2005	-0.263		0.332		-0.190		0.781	*	0.282		-0.539	
Year 2008	1.068	*	-0.577		0.717		1.868	***	-0.287		0.709	
Year 2009	0.733		-1.686	**	-0.129		0.877	**	-0.360		0.822	
Constant	16.785	***	29.362	***	32.619	***	41.482	***	39.443	***	42.427	***
Log pseudolikelihood	-187923.0		-119511.7		149133.7		-203590.7		-47084.7		57552.3	
Pseudo R-squared	0.01		0.01		0.01		0.01		0.00		0.01	

Data Source: Brazilian National Household Survey (Pnad) from 2004 to 2009.
Note: Robust Standard errors reported. All regressions include state fixed effects.
*p < .10; **p < .05; ***p < .01

Table 5. Placebo Test: Estimating labor force participation and usual weekly hours worked of non-studying mothers with the youngest child aged four years old and below (Child04) compared to childless women in 2001-2003 and 2004-2006 in Brazil by marital status and education attainment

	Labor Force Participation (*Probit Estimates: Marginal Effects)*		Hours Worked Week *(Tobit Estimates)*	
	Married	Single	Married	Single
Total				
Child04	-0.168 ***	-0.042 ***	-15.484 ***	-4.122 ***
Pos2004	0.009	0.017 ***	0.392	1.084 ***
Child04 x Pos2004	0.002	-0.003	1.051 **	-0.307
Less than High School				
Child04	-0.161 ***	-0.044 ***	-17.520 ***	-5.866 ***
Pos2004	0.023 **	0.016 *	1.688 **	1.325 *
Child04 x Pos2004	-0.007	0.005	0.329	0.183
High School Completed				
Child04	-0.189 ***	-0.132 ***	-15.219 ***	-9.549 ***
Pos2004	0.007	0.017 **	0.441	1.130 **
Child04 x Pos2004	-0.001	-0.007	-0.012	-0.513
College or More				
Child04	-0.080 ***	-0.046 **	-6.920 ***	-2.286 **
Pos2004	-0.028 *	0.015	-1.821 **	0.066
Child04 x Pos2004	0.014	-0.039	1.383	-1.804

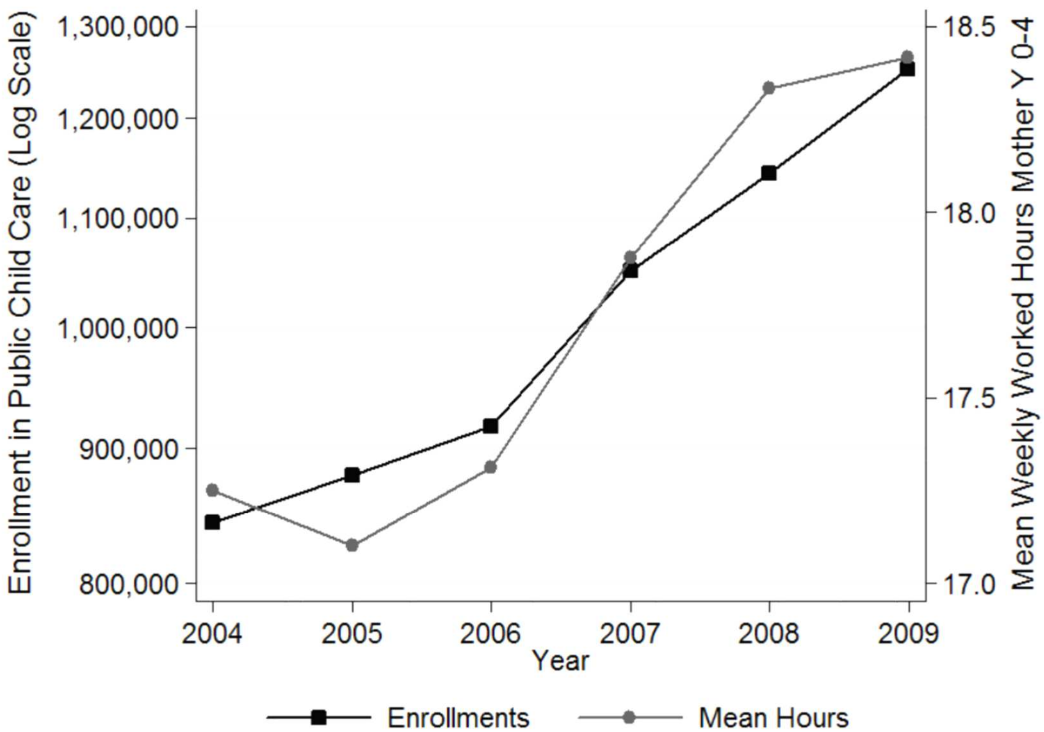
Data Source: Brazilian National Household Survey (Pnad) from 2004 to 2009.

Note: Robust Standard errors reported. All regressions include dummy for black or pardo (mixed race), age, age squared, log of other household income, dummy for women that moved in the last four years, conditional cash transfer, rural areas, and years 2004, 2005, 2008, 2009. All regressions include state fixed effects.

*p < .10; **p < .05; ***p < .01

FIGURES

Figure 1. Total enrollment in public child care centers in Brazil and mean weekly worked hours by non-studying mothers with the youngest child aged 0 to 4 years old from 2004 to 2009



Data Source: Brazilian National Household Survey (Pnad) from 2004 to 2009 and INEP Brazilian School Census 2004-2009.

Appendix 1: Exploring Additional Explanations to the Increase in FLFP in Brazil in the 2000s

In the 2000s, Brazil had other two important social changes that could affect the employment of mothers of children in child care ages: the implementation of *Bolsa Família* and a significant increase in GDP per capita. *Bolsa Família* was implemented in October 2003 and the program consisted of providing cash transfers to households in misery and in poverty. A fixed benefit was provided for any household below the misery line together with a variable benefit that depended on the number of children in the household, which is applicable for households below the poverty line. This program gradually substituted pre-existing government programs (*Auxílio Gás, Bolsa Escola, Bolsa Alimentação, Cartão Alimentação*, and *PETI* in 2006) and imposes health and educational conditionalities for its users in order to remain in the program. Before 2012, there was no education conditionality for children under age five[10], the education conditionality existed only for children aged six to seventeen years old. So, we can affirm that the increase of *Bolsa Família* users between 2001 and 2009 is not driving the increase in enrollment in public childcare centers.

The percentage of families benefited from *Bolsa Família* has expanded considerably during 2004-2009, and researches may hypothesize whether the increase in female labor force participation in the period is driven by this increase. In general, we expect that an increase in household income through *Bolsa Família* to affect the probability of working negatively, rather than positively. In addition, the maximum value of the benefit is more than half of the minimum wage, so it is not likely to affect employment search or current employment status. Previous research has shown no effect or very small negative effects of the program on employment (Foguel and Barros 2010). Consequently, if there is any effect of the *Bolsa Família* in the

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[10] In 2012, it was implemented the program *Brasil Carinhoso* which has the objective to increase the enrollment in public childcare centers of children aged four and below participating in the *Bolsa Família* program. It provides a larger FNDE's transfer per pupil to the school if the pupil is enrolled in the *Bolsa Família* program.

employment of mothers with the youngest child aged four and below, it is likely to be a negative effect, the opposite signal that we expect with the increase of slots in public childcare centers. Therefore, not affecting our estimates. Nevertheless, a control is included in our model to indicate whether the household receives *Bolsa Família* transfers. The construction of this variable is explained in Section 3.

An additional hypothesis for the increase in labor supply is related to increases in GDP per capita, which could lead to greater demand of workers and, consequently, increases in female employment. Between 2004 and 2009, Brazil’s GDP per capita increased 13.4 %, with an average annual increase of 2.6% per year. This considerable increase in GDP per capita it is likely to affect both, the treatment (mothers with the youngest child aged zero to four years old) and the control group (childless women), and its effects are captured by the year fixed effects included in the regression. Finally, previous research has shown that increases in male unemployment rates, especially in economic crises, can lead to greater employment of their wives in order to protect for household losses (Parrado and Zenteno 2001). So, it is important to observe whether there are recessions or unemployment crisis in the period observed. Data from the World Bank Indicators (2015) show that male unemployment did not have an expressive variation in Brazil, going from 6.8% in 2004 to 6.09% in 2009, with an average of 6.3%. So, it is unlikely that this factor has affected the employment of mothers with the youngest child in childcare ages.

Appendix 2: Descriptive Statistics

Table A1. Summary Statistics of Non-Studying Women Aged 16 to 44 years old in the Pnad samples from 2004 to 2009

Variables	**Childless**		**With Youngest Children Aged 0-4**	
	Mean	Std. Dev.	Mean	Std. Dev.
Usual Hours Worked in a week (for those working)	40.24	12.96	35.24	15.82
Labor Force Participation	0.65		0.50	
Black or Pardo	0.50		0.58	
Educ Attain (ref. Primary or less)	0.31		0.63	
High School	0.54		0.31	
College or More	0.15		0.06	
Age	26.3	7.1	28.0	6.36
Moved in the Last Four Years	0.09		0.11	
Other HH Income (Reais)	1,520	2,471	856	1,569
Rural Areas	0.11		0.18	
Not living with a partner	0.72		0.22	
Pos 2007 (2007, 2008, 2009)	0.51		0.48	
Sample Size	*121,803*		*124,954*	

Data Source: Brazilian National Household Survey (Pnad) from 2004 to 2009.

CHAPTER 3 It is lower than you think it is: Recent TFR in Brazil

Helena Cruz Castanheira

Hans-Peter Kohler

1. Introduction

Understanding fertility trends in middle-income countries is of essential importance for understanding global population trends and patterns of global population aging. Most middle-income countries have experienced substantial declines in fertility, along with improvements in mortality, and many have attained – or are at the verge of attaining – below replacement fertility levels. For example, in South Korea during 1950–2010, life expectancy increased from 47.9 to 80 years, fertility (TFR) declined from 5.1 to 1.3 children per woman, and GDP per capita grew substantially with a growth rate of more than 5% p.a. during 1960–2010. While often seen as a sufficient condition for fertility decline, rapid economic development is not always a necessary condition: in Bangladesh during 1950–2010, for example, life expectancy increased from 45.3 to 67.8, fertility (TFR) declined from 6.4 to 2.4 children per woman, and GDP per capita grew during 1960–2010 at an average rate of only 1.5% p.a. Both India and China saw large fertility declines before the onset of rapid economic growth. Iran holds the record of the most rapid decline in fertility from 6.5 to 1.8 during the period 1980–2010 when average economic growth was relatively modest at around 1.3% p.a. (Abbasi-Shavazi et al. 2009). Perhaps even more surprisingly, countries as diverse as Argentina, Bangladesh, Mexico and South Africa are expected to *reach* net reproduction rates (NRRs) below 1– and thus below replacement fertility – within 5–10 years, and by 2015–20, more than 1 billion persons are expected to live in countries with below-replacement fertility (as measured by NRR) in sub-Saharan Africa, Southern and South-Eastern Asia, and Latin America and the Caribbean alone (57 millions in SSA, (6% of total pop), 531 millions (21% of total pop) in Southern and South-Eastern Asia, and 484 millions (75% of total pop) in Latin America and the Caribbean). Moreover, close to 1/2 billion individuals live in countries that are expected to newly *attain* below-replacement fertility (NRR < 1) in these regions within the next 10 years (United Nations 2013).

While the above UN Population Statistics indicate a remarkable spread of low fertility in middle-income countries, we argue in this paper that below-replacement fertility may have progressed even more, and that several key middle-income countries may have fertility levels that are significantly below those reflected in the UN estimates. The reason is that in many middle-income countries at the verge of below-replacement fertility, including for instance Brazil, Colombia, Ecuador, Peru, and Venezuela, fertility rates are estimated from census and related survey data, and these estimates are adjusted using the P/F-Brass method for potential underreporting of births. But we illustrate in the case of Brazil that this adjustment in the context of contemporary low fertility and fertility postponement may do more harm than good. We illustrate in the case of Brazil that in recent years, estimates of fertility based on census and registration data have become highly reliable, and that these sources suggest TFR levels for Brazil around 1.75 in 2010, which is 8% below the official TFR estimate. Moreover, our estimates suggest that 30% of Brazil's population resides in regions where SINASC-based TFR is below 1.65, and 70% of the population is in regions with TFR levels below 1.75. The reasons of why the P/F adjustment has become misleading and should no longer be used is twofold: first, underreporting in census and related surveys has been reduced to improved question framing and the fact that recall errors are less common in contexts of low-fertility low-infant-mortality contexts. Related, civil registration has improved in countries such as Brazil so that it now provides reasonable coverage of births (94%). Second, the assumptions of the P/F model no longer hold when low fertility is not only due to stopping behavior, but increasingly due to a postponement of fertility that shifts the age pattern of fertility to later ages.

While we illustrate in the case of Brazil that the P/F adjustment is misleading, and thus indicating a higher level of fertility than we believe is actually prevailing, Brazil is not the only country to which this issue applies. A wide range of countries, including Colombia, Peru, Venezuela, and Ecuador, use similar procedures in estimating their official TFR levels, and we believe that in several countries the P/F adjustment may result in an upward exaggeration of the official TFR. As a result, we believe that in many Latin American – and possibly other middle-income countries – low fertility may actually be lower than is reflected in official estimates, and that some countries such as Brazil may have attained below-replacement fertility considerably earlier than has been assumed to be the case.

Brazil is an important country for which to document the importance of accurately estimating TFR. It is the fifth largest country in the world, the largest country in Latin America. Based on UN Predictions, it contributes 1.4% to the global population growth between now and 2050 with the medium fertility variant, and almost 20% of the population growth in Latin America. Brazil has also a dominating influence to population aging in Latin America. But all of these factors and future trends are in part affected by estimates of recent and current TFR levels, and the resulting forecasts of future TFR trajectories. Our replications of the UN projections for Brazil indicate that a lowering of recent fertility estimates to levels that we perceive are accurate has important implications. For example: with the new TFR, we have a national population that is 10 million smaller in 2050 than the originally predicted by the UN estimates. This is a large difference given that we are considering only 40 years of projection.

2. The P/F Brass Method

Total fertility rates are important measures in population projections, and historical and international comparisons. It measures the average number of births a woman would have if she experiences the age-specific fertility rates observed in a specific period throughout her entire reproductive life (ages 15 to 49). Period TFR (TFR) can be very different from cohort TFR (CTFR) in places with high levels of fertility postponement (Goldstein et al. 2009) or acceleration (Parrado 2011). Most of recent research has focused on discussing the extent in which these two factors affect the period TFR, with quantum and tempo effects being extensively discussed in the literature (Bhrolchain 1992; Boongarts and Feeney 1998; Kohler and Philipov 2001; Sobotka

2004; Schoen 2004). Notwithstanding the discussion of its realistic representation of cohort fertility, the TFR remains largely used in the literature (Myrskylä et al 2011) and in world population projections (United Nations 2013), and so it is very important to estimate it accurately. Despite being a very straightforward measure, derived from the sum of age-specific fertility rates (ASFR) for the female population aged 15 to 49 years old, there is some heterogeneity across countries in the use of indirect methods for estimating the TFR.

In this paper, we ask a fundamental question related to the precision of the TFR estimation when using a demographic technique, denominated the P/F Brass ratio, to adjust the TFR in Brazil. This method was developed by Brass in the 1960s for correcting TFR levels in closed and stable populations in countries using surveys, instead of birth registries, to estimate their TFR (Brass et al. 1968; Brass 1975; United Nations 1983). However, despite the research showing the problems of this technique in countries with rapid fertility decline (Brass 1996; Moultrie and Dorrington 2008; Schmertmann et al. 2013), some countries still use this adjustment in non-stable population contexts. Brazil is one of these countries and the P/F Brass method is used to calculate its official TFR. The implications of a misleading adjustment in Brazil's TFR are beyond the estimations of the country, but affects estimations for South America and the world since it represents almost half of the population in the region and it is the fifth most populous country in the world (United Nations 2013). Consequently, it is very important to estimate its correct TFR for local, regional, and global population projections.

The P/F Brass method was primarily developed to analyze survey data from Africa, in which researchers were detecting systematic errors in its recollections (Brass et al. 1968). The first source of error is related to imprecisions in the reference period based on the question asking whether a person had a birth in the year preceding the Census. Women were reporting events that happened, on average, eight to fifteen months before the Census, leading to distorted values of the TFR. This error should be absent in recent Brazilian Census, since births in the year preceding the survey are estimated based on month and year of last birth, which are not sensitive to the reference period error. The second source of error defined by Brass et al. (1968) is the memory error, which suggests that women may forget the total number of children ever born, specially older women, because of low literacy, difficulties in counting the number of births, and high levels of mortality making it more difficult to remember and report births that have died. The memory error would affect the estimated TFR from the Brazilian Census if women aged 15 to 49 years old underreport births that occurred in the twelve months prior to the survey. In this regard, the P/F ratio is used to adjust the TFR estimated from the Census.

The adjustment used in Brazil is the P_2/F_2 ratio, which is the total parturition in the age group 20-24 in relation to the cumulated period fertility in the same age group. This ratio is multiplied by the TFR or age specific fertility rates, increasing or decreasing its level. The numerator of the ratio, P_2, is the average number of births ever had by women in the 20-24 age group. The denominator of the ratio (F_2) is calculated, first, by acquiring the fertility in the beginning of the age interval, age 20, which is given by $Ø_2 = 5*f_1$ [from the formula $Ø_i = 5*(f_1 + f_2 + ... + f_{i-1})$], f_1 is the age-specific fertility rate in the 15 to 19 age group. Next, the average number of births in the 20-24 age group is estimated, but it is not recommended to obtain it by multiplying f_2 by 2.5, because fertility is not constant in the interval, especially in the beginning of the fertility schedule. So, Brass et al. (1968) computed a series of multiplying factors to be interpolated using the observed f_1/f_2 ratio for obtaining k_2. The final value of F_2 is given by the formula $F_2 = Ø_2 + k_2*f_2$. After obtaining P_2 and F_2, the P_2/F_2 ratio is estimated. This ratio should be one in a stable population with no errors in data recollection, assuming that the fertility of women who died or emigrated is similar to the fertility of women that survived or immigrated. In this context, if the ratio is not one the data might have reference or memory errors, which is corrected by multiplying the P_2/F_2 ratio by the TFR as suggested by Brass et al. (1968). Nevertheless, the ratio can also be different from one when there are problems in satisfying the method's assumptions, specially the stable population assumption. In this case, when multiplying the Census' TFR by the P_2/F_2 ratio we might have a misleading estimate of the country TFR.

In Brazil, the estimated P_2/F_2 ratio was 1.12 in 1991, 1.10 in 2000, and 1.19 in 2010. Which means that when the adjustment is applied, the observed TFRs of the Census Surveys are increased by 12%, 10%, and 19% respectively. It is unclear whether these values are generated by the underreporting of births in the Census or by problems in not satisfying the method's assumption. The first hypothesis is not possible to measure using the Census microdata; however, it is unlikely that the underreporting of births would increase with time, as observed with the increase in the ratio from 2000 to 2010, and also that the magnitude of the underreporting of births of women aged 15-19 and 20-24 would be so high. A more plausible hypothesis is the unmet assumptions of the method. Brazil is not a stable population, its fertility decline started in the mid-1960s, but demographers defended that the P/F Brass method could be used because its fertility decline consisted of stopping behavior rather than postponement. So, arguably, the cumulative fertility of the first two age groups (ages 15 to 19 and 20 to 24) and the parturition of the age group 20 to 24 should remain stable, not affecting the P_2/F_2 ratio. Nowadays, however, Brazilian fertility is characterized by a significant postponement of fertility in the first two age groups of the fertility schedule, and so the P_2/F_2 ratio can be distorted, no longer reflecting corrections in births' underreporting. The 2000 Census already suggested that Latin America's early motherhood imperative was weakening (Rosero-Bixby et al 2009), and the 2010 Census can clearly confirm this trend for Brazil. Figure 1 shows the age specific fertility rates for first births per 1,000 women in 1991, 2000, and 2010. In 2000, we can observe a decline in first births starting at age 19, and in 2010, this decline is even more expressive and start two years earlier in the fertility schedule. In addition, we can observe an increase in first births rates after age 27, a typical configuration of postponement behavior. In this context, using the P_2/F_2 ratio can be very misleading, because the stable population assumption is clearly violated even for the first two age groups of the fertility schedule.

The stable population assumption can be relaxed with a new variant of the P/F Brass method proposed by Schmertmann et al. (2013). In this new method, P is calculated similarly to the P/F Brass method, but the denominator is estimated with a set of multipliers resultant from calibrated spline interpolation of fertility schedules using the Human Fertility database (HFD) and the US Census International Database (IDB), explained in Schmertmann (2012). The final TFR is also obtained differently from the P/F Brass method, instead of multiplying the P_2/F_2 ratio to the survey's TFR, the authors use a weighted least squares regression to calculate the final adjusted TFR. The dependent variable of this regression is the natural log of the observed TFR multiplied by the P/F Schmertmann's ratio, so each age group has a Pi/Fi value that is multiplied by the country's TFR. The independent variable of the regression is the difference between the mean age of childbearing of each age group (calculated using the M0 and M1 values resultant from the calibrated spline interpolations) and the middle age of each age group (17.5, 22.5, 27.5, 32.5, 37.5, 42.5, and 47.5). This regression is weighted by the inverse of the variance of P, F and TFR [1/(1/P + 1/F – 1/TFR)]. The exponential function of the regression's intercept, b0, provides Schmertmann's final adjusted TFR. The authors maintain that this technique performs better on countries with rapid fertility decline than the original P/F Brass method and Feeney (1996), which was tested by Moultrie and Dorrington (2008). The official estimates of the TFR from Brazilian Census surveys use the classical P_2/F_2 Brass adjustment.

3. Analyzing the Consistency of Fertility Estimates across Data Sources

We have proposed that errors in data recollection should not bias the Brazilian TFR in the magnitude of the P_2/F_2 ratio currently calculated for the country. We maintain instead that the magnitude of the ratio and its time-trend is mostly driven by the violation of the stable population assumption, which results in biased official TFRs for the country. In order to contextualize the TFR estimates obtained using the P_2/F_2 adjustment, we analyze the consistency of TFR estimates in Brazil across data sources. Figure 2 shows TFR estimates for different data sources from 1991 to 2013. The official estimate shown in the figure is the only data source that has the P_2/F_2 adjustment. It uses the information on fertility from the Census (1991, 2000, and 2010) and Pnad household surveys and apply the P_2/F_2 adjustment. The civil registry and SINASC (Live Births Information System) estimates presented in the figure are two different information sources on birth registries. If the birth occurs in the hospital, the hospital produces a form in which the first copy is sent to SINASC, which is managed by the Health Ministry, and the second copy is given to the family to registry their children in public notaries. The information from notaries is assembled and published by the National Statistics Office (IBGE) and constitutes the civil registry system. If the birth occurs in the household, the first health unit or notary visited sends the information to the SINASC system. Birth estimates from SINASC are usually greater than the civil registry estimates for a given year because of late registration. The birth certificate is free since 1997, and the civil registry has improved its coverage considerably in the last twenty years. The DHS data trend in Figure 2 is the 1996 Demographic and Health Survey (DHS) and its Brazilian equivalent in 2006, the *Pesquisa Nacional de Demografia e Saúde* (PNDS). Finally, the modified Brass data trend shown in Figure 2 is the new variant of the P/F Brass method (Schmertmann et al. 2013), explained in section 2.

It can be observed that between 1992 and 1995 the official estimates were lower than survey estimates; however, after 1996, official values, which use the P_2/F_2 Brass correction, start diverging significantly from survey estimates. For instance, the 1996 DHS presents a similar value to the official estimate, being 2.5 births per woman in 1994-1996 compared to 2.52 of the official estimate in 1995. In the DHS a decade later, official estimates contrast sharply with DHS estimates, being 2.06 in 2005 compared to 1.80 in the 2006 DHS. In this same year, the Pnad survey shows a TFR of 1.77, consistent with the DHS estimate. In 2010, the official TFR estimate is 1.90 births per woman, and the Census data shows 1.60 births per woman and a confidence interval of 1.53 to 1.66 births. The TFR of the national birth system (SINASC) is 1.71 births per woman in 2010 and the civil registry is 1.65, consistent with the boundaries of the 2010 Census estimate, but very different from official estimates. In general, the estimates across data sources are fairly consistent, but the official estimate, which uses the P_2/F_2 Brass adjustment, is considerably larger than the other estimates.

When considering the new variant of the P/F Brass method proposed by Schmertmann et al. (2013), denominated ‘modified Brass’ in Figure 2, the results in 1991 and 2000 are lower than the official estimate and greater than the Census estimates, as we would expect when correcting for birth underreporting. In 2000, for instance, the TFR using modified Brass is 2.23, which is lower than the official estimate (2.38) and closer to the upper bound of the Census' confidence interval (2.21). Nevertheless, in 2010, the modified Brass provide a greater TFR than the official estimate, being 1.94 for the former and 1.90 for the latter. If we eliminate the information of the 15-19 age group by attributing zero weight in the regression, as recommended in Schmermann et al. (2013), the TFR decreases to 1.91, but is still higher than the official estimate for 2010. The similarities between the TFRs estimated with the classical P/F Brass method and the new variant of the method in 2010, suggest that the modified Brass method may also be influenced by the rapid fertility decline observed in the country.

4. What is the best TFR estimate for Brazil?

We have observed that TFR estimates from direct methods have similar values, but the estimates using indirect methods are considerable larger. Nevertheless, it remains unclear which set of estimates are better suited to represent Brazilian fertility, as SINASC and civil registry are likely to have sub-registries. In this regard, we use a question from the Census to estimate the amount of sub-registry in the administrative records. The question asks the type of birth certificate of children aged 10 or below. The first alternative is the formal birth certificate emitted from notaries, which is the basic input of the national civil registry system. The second alternative is the form provided in health facilities in order to register the birth in the notary, this option together with the first would be the correspondent adjustment to births registered in SINASC. The third option is the indigenous birth certificate (RANI), which is not accounted in the civil registry data or SINASC. This analysis is applied only for the 2010 Census, since this question is not available for the 1991 and 2000 Census.

Table 1 shows the type of birth registry for children aged zero in August 1st 2010, and so that were born between August 2nd 2009 and August 1st 2010. We can observe that 93.94% of these children had an official birth certificate from the notary, 3.3% a health facility birth declaration (DNV), 0.2% the Indigenous birth certificate (RANI), and 2.5% didn't have a birth registry. So, a total of 2.8% births in Brazil during the Census period were not registered by the civil registry or SINASC. In order to adjust the SINASC and civil registry data for sub-registry, we assume that the Census percentages correspond to the actual sub-registry in these databases. The SINASC and civil registry data for 2010 is for the calendar year (from January 1st 2010 to December 31st 2010), and so our first assumption is that the sub-registries observed for the Census period are applied to the 2010 calendar period. Our second assumption is that the information reported by the individuals is accurate and that the SINASC and the civil registration system accurately process the information once the individuals receive the official birth certificate or the birth declaration (DNV) from health facilities.

The Brazilian TFR in 2010 using the Civil Registry data is 1.65 and its sub-registry (Table 3) is 100-93.94 = 6.06%. The correction factor for the civil registry will be 1+0.0606 = 1.0606, which is multiplied by the initial TFR resulting in a TFR of 1.754 children per women. The Brazilian TFR in 2010 using the SINASC data is 1.71 and its sub-registry is 100-97.2 = 2.76% (registries from notaries and health facilities), providing a correction factor of 1+0.0276 = 1.0276, and the final SINASC TFR is then 1.757. The two adjustments provide very similar results, which increase our confidence in the data and estimates. These results are significantly lower than the 1.90 children per women calculated with Brass P_2/F_2 ratio from the 2010 Census data, and in greater agreement with the TFR of 1.80 resultant from the 2003-2006 TFR of the PNDS, the Brazilian DHS's equivalent.

Despite being very confident of the adjusted SINASC and civil registry results at the national level, this is not the case for the state-level estimates. The classical error incurred when using different data sources in the numerator and denominator, driven mostly by interstate migration, are an issue. The problem is driven by the fact that the mid-year population assumption for person-years might not be accurate, with more or less exposure to fertility than assumed in the denominator. Places with high inter-state immigration would present the most problematic estimates. In this regard, the state ranking of TFRs presented by the Census unadjusted TFR might be the correct, but its levels incorrect. Figure 3 presents the total fertility rate across Brazilian states in 2010, ordered from the lowest to the largest TFR estimates from the 2010 Census. Overall, the adjusted SINASC-TFRs present estimates that are within the 95% confidence interval of the Census estimates, the three exceptions are Rio de Janeiro, São Paulo, and Brasília, which are states with high levels of immigration. Official estimates (Census with the P_2/F_2 adjustment), in contrast, are consistently higher than the upper end of the Census confidence interval, presenting significantly higher estimates across states. In Figure 3, we can also observe that nearly 70% of Brazil's total population is concentrated in the first eleven states in which the TFRs from SINASC are within the 1.55 – 1.75 range. Moreover, half of Brazil's population is in the states of Minas Gerais, São Paulo, Rio de Janeiro, Rio Grande do Sul, and Santa Catarina that present a SINASC TFR of, respectively, 1.55, 1.69, 1.63, 1.57, and 1.59. Studies have shown a bifurcation in fertility regimes across high-income countries (Rindfuss et al. 2015), and some countries present very low fertility levels for sustained periods of time (Kohler et al. 2002). By dropping the P/F Brass correction, we can observe that some Brazilian states might have already started in the direction of low fertility levels, being in the lower range of fertility regimes encountered in high-income countries.

In this study, we propose that the P/F Brass method should no longer be used, since it is much greater than the estimates of birth sub-registry available. Research should further investigate a solution for the state-level estimates in order to have a more accurate figure of regional TFR. Overall, we find that household surveys and the national registration systems present consistent estimates that differ sharply from the official estimate that uses the P/F Brass adjustment. The stable population assumption that is violated when using the P_2/F_2 Brass method is likely to be the main reason for the overestimation of the TFR current encountered in Brazil's official TFR estimate.

5. The Effect of the 2010 Bias on Population Projections

The UN Population Projection uses Brazil's official TFR to estimate its population projection. In this section, we estimate the UN population projection and estimate a cohort-component projection using the 2010 TFR based on Brazil's Live Birth Information System (SINASC), which provides in our perspective the most reliable data source for estimating live births in Brazil. Using the cohort component method (Preston et al. 2001), we maintain the mortality and fertility observed in 2010 for the next 40 years, and assume a closed population. With this method, we observe a 10.1% increase in the population in 2050 using a TFR of 1.71 (unadjusted SINASC-TFR) and an increase of 17.7% using the official TFR of 1.90. However, it might be unrealistic to assume that the observed fertility levels will remain low in the next forty years, and so we re-estimate the UN Median Projection, which uses Bayesian estimates, to predict the future TFR in Brazil given the current and historical observed pattern across 201 countries.

Figure 4 shows the predicted fertility trajectory for Brazil using the official TFR estimate and the SINASC unadjusted estimate in the last observed point estimate of the 2005-2010 period. We can observe that, because of the assumptions of the UN projection, the two TFR's trajectories approximate to similar levels in the period 2070-2075, however they are very different in the initial years of the projection. The lowest TFR projected when using 1.90 in 2010 is 1.68 children per women in 2030-2035, however when using the TFR of 1.71 in 2010, in 2015 we already have a value of 1.63, which is lower than any value observed when using the official estimate. The lowest TFR using the SINASC TFR for 2010 is 1.57 children per women in 2020-2025. In general, we observe very different TFR trajectories when we drop the P/F Brass adjustment from the TFR estimation.

The UN Bayesian Projection using the medium-fertility assumption shows a total projected population in 2050 of 230,188,000 with the official estimate in 2010, and 219,815,000 with the SINASC estimate, almost a 10 million people difference in only 40 years of projection, and a 5% relative difference between the two estimates. In 2100, the median projected values differ by 10%. Not only is the total population in each country different when using different TFR trajectories, but also the projected age structure of the population. The population pyramids in 2050 using the principal components projection method and the UN Bayesian method are also different, which can provide different results in terms of old age dependency ratio, and, consequently, influencing pensions and taxes prognoses.

6. Discussion

The P/F Brass adjustment overestimates Brazil's TFR and should no longer be used. This overestimation has important consequences for calculating the future trajectory of Brazil's TFR, with the official estimate predicting a lowest TFR of 1.68 children per women in the 2010-2100 period and the SINASC estimate suggesting the lowest rate of 1.55 children per women in the same period. The total population and age structure predicted from the two estimates are also different, which may result in diverging pensions and taxes prognoses. More research is needed in order to estimate the sub-reporting in the Birth Registry system and in the Live Births Information System (SINASC), but the existing estimates already suggest that it is lower than the adjustment incurred with the P/F Brass method.

Tables

Table 1: Type of Birth Registry for Children Aged Zero in August 1st 2010 in Brazil

Type of Birth Registry	Frequency	Percent	Cum.
Notary	2,563,604	93.94	93.94
Health facility	89,946	3.30	97.24
Indian birth registry	4,771	0.17	97.41
Don't have	67,427	2.47	99.88
Don't know	2,752	0.10	99.98
Ignored	419	0.02	100.00
Total	2,728,919	100	

Source: Census 2010.

Table 2: Type of Birth Registry in Brazil by Age in August 1st 2010

Age	Type of Birth Registry						Total
	Notary	Health facility	Indian birth registry	Don't have	Don't know	Ignored	
0	93.94	3.30	0.17	2.47	0.10	0.02	100
1	97.20	1.66	0.19	0.86	0.09	0.01	100
2	97.88	1.29	0.17	0.57	0.07	0.02	100
3	98.21	1.13	0.18	0.40	0.07	0.01	100
4	98.39	1.00	0.19	0.32	0.08	0.01	100
5	98.61	0.88	0.17	0.25	0.07	0.01	100
6	98.72	0.78	0.18	0.22	0.08	0.01	100
7	98.92	0.61	0.17	0.21	0.08	0.02	100
8	99.11	0.46	0.16	0.16	0.08	0.02	100
9	99.15	0.43	0.15	0.16	0.09	0.02	100
10	99.19	0.39	0.14	0.16	0.09	0.02	100
Total	98.18	1.05	0.17	0.50	0.08	0.02	100

Source: Census 2010.

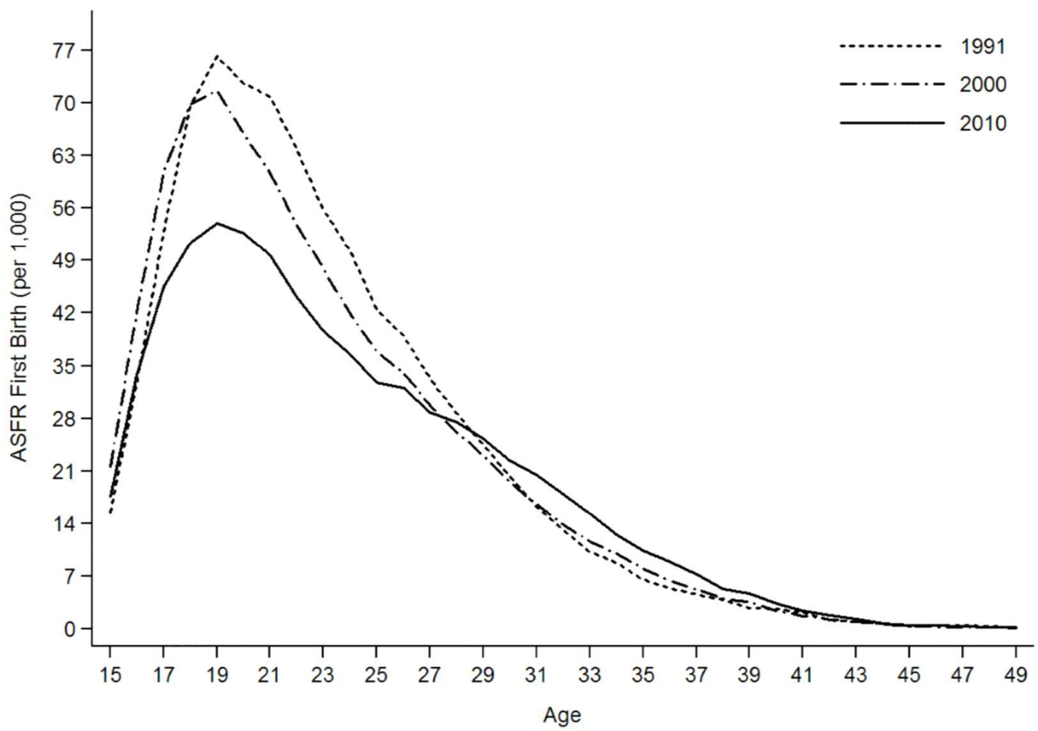
Table 3: Birth Registry and SINASC Adjustments

State	TFR Civil Registry Adjusted	TFR Civil Registry	Civil Registry Adjustment	TFR SINASC Adjusted	TFR SINASC	SINASC Adjustment
Brazil	1.75	1.65	1.06	1.76	1.71	1.03
Rondônia	1.81	1.70	1.06	1.83	1.77	1.03
Acre	2.27	1.97	1.16	2.63	2.44	1.08
Amazonas	2.31	1.91	1.21	2.58	2.29	1.13
Roraima	2.38	2.00	1.19	2.62	2.31	1.13
Pará	1.98	1.66	1.19	2.17	1.96	1.10
Amapá	2.40	2.14	1.13	2.47	2.31	1.07
Tocantins	1.98	1.81	1.09	2.01	1.93	1.05
Maranhão	1.98	1.71	1.16	2.09	1.94	1.08
Piauí	1.79	1.57	1.14	1.85	1.74	1.06
Ceará	1.75	1.62	1.08	1.76	1.69	1.04
Rio Grande do Norte	1.70	1.61	1.06	1.71	1.67	1.02
Paraíba	1.82	1.67	1.09	1.81	1.77	1.02
Pernambuco	1.76	1.66	1.06	1.77	1.72	1.03
Alagoas	1.97	1.84	1.07	1.96	1.89	1.04
Sergipe	1.80	1.69	1.07	1.83	1.78	1.03
Bahia	1.73	1.63	1.06	1.71	1.67	1.03
Minas Gerais	1.56	1.54	1.02	1.55	1.53	1.01
Espírito Santo	1.70	1.66	1.02	1.69	1.67	1.01
Rio de Janeiro	1.55	1.50	1.03	1.63	1.61	1.01
São Paulo	1.71	1.67	1.02	1.69	1.68	1.00
Paraná	1.73	1.69	1.02	1.73	1.72	1.01
Santa Catarina	1.59	1.56	1.02	1.59	1.58	1.01
Rio Grande do Sul	1.56	1.53	1.02	1.57	1.56	1.01
Mato Grosso do Sul	1.93	1.80	1.07	1.96	1.86	1.05
Mato Grosso	1.82	1.71	1.06	1.83	1.77	1.03
Goiás	1.65	1.60	1.03	1.63	1.61	1.01
Distrito Federal	1.79	1.70	1.05	1.75	1.74	1.00

Source: Census: obtained from the Census microdata. SINASC: Live Births Information System with data available online at http://tabnet.datasus.gov.br/cgi/deftohtm.exe?sinasc/cnv/nvuf.def, last accessed on May 11th 2015. The reference period is from January 1st to December 31st. Civil registry: obtained from the National Institute of Geography and Statistics (IBGE), tables 343 and 2680 available at www.sidra.ibge.gov.br. The reference period is from January 1st to December 31st.

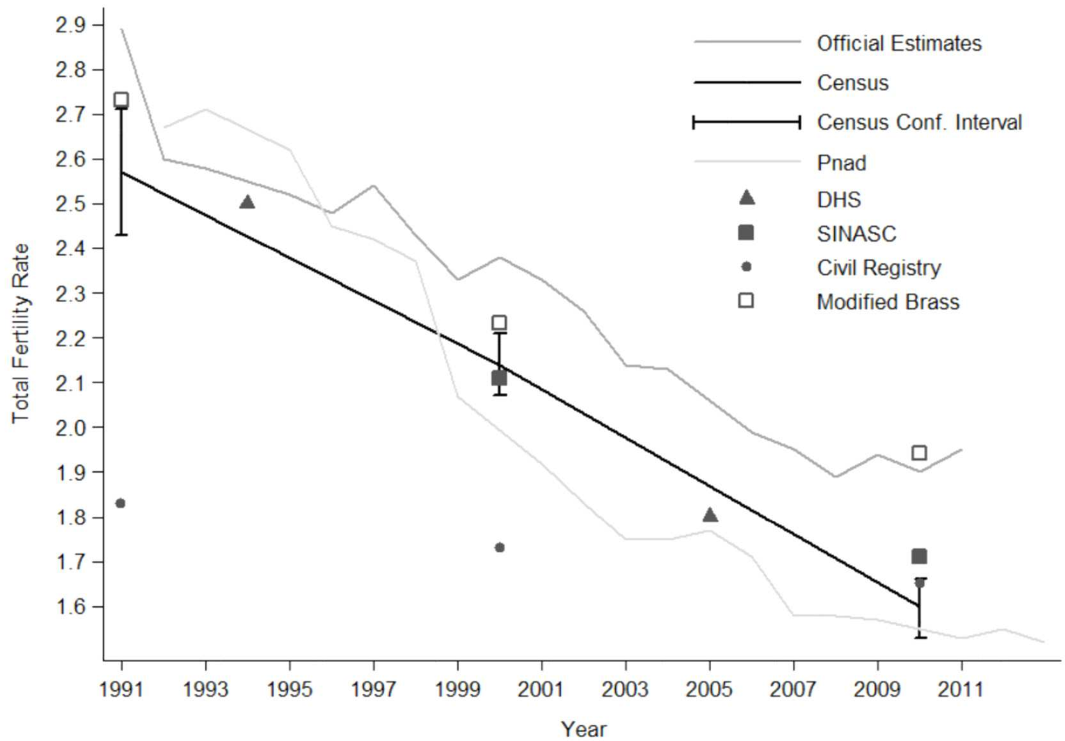
Figures

Figure 1. Age Specific Fertility Rates for First Births per 1,000 women



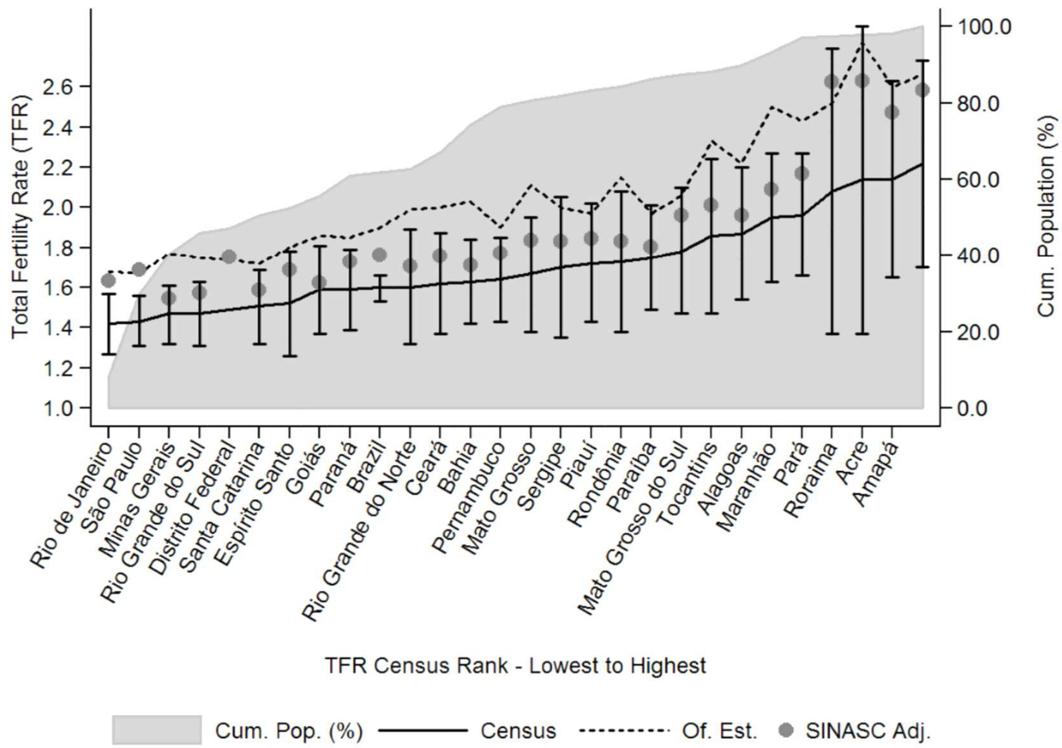
Source: IBGE. Brazilian Demographic Censuses of 1991, 2000, and 2010.

Figure 2: Estimated Total Fertility Rate in Brazil in Different Data Sources from 1991 to 2013



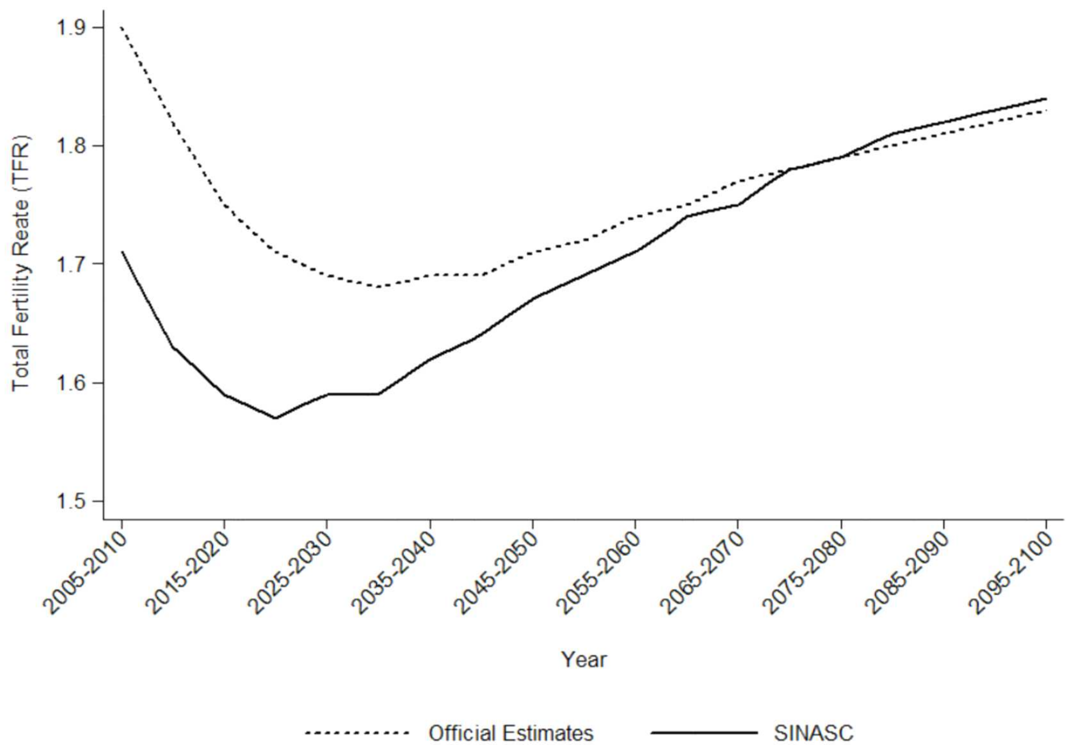
Source: Official TFR estimates: obtained from the National Institute of Geography and Statistics (IBGE), table 3727 available at www.sidra.ibge.gov.br, last accessed on May 11th, 2015. These TFRs are estimated based on the Annual Household Survey (PNAD) during non-census years and in the Demographic Census Survey for the years 1991, 2000 and 2010. The P/F Brass adjustment is applied in all years. Census and Pnad estimates: obtained from the Census microdata. Census and Pnad TFRs refer to the twelve months prior to the survey's reference date. SINASC: Live Births Information System with data available online at http://tabnet.datasus.gov.br/cgi/deftohtm.exe?sinasc/cnv/nvuf.def, last accessed on May 11th 2015. The reference period is from January 1st to December 31st. DHS: The estimated TFR of the Demographic and Health Surveys (DHS) are available at http://www.statcompiler.com/, for the 1996 DHS, and at http://bvsms.saude.gov.br/bvs/publicacoes/pnds_crianca_mulher.pdf, for its 2006 Brazilian equivalent, the PNDS. The reference period of the TFR is three years before the survey's reference date. In the graph, the data point is located in the mid-period of the three years reference period for both surveys. Civil registry: obtained from the National Institute of Geography and Statistics (IBGE), tables 343 and 2680 available at www.sidra.ibge.gov.br. The reference period is from January 1st to December 31st. Modified Brass: Uses the Demographic Census and the modified P/F Brass method elaborated in Schmertmann et al. (2013)

Figure 3: Estimated Total Fertility Rates by Brazilian States in 2010



Source: Official TFR estimates: obtained from the National Institute of Geography and Statistics (IBGE), table 3727 available at www.sidra.ibge.gov.br, last accessed on May 11th, 2015. Census estimates: obtained from the Census microdata. Census TFRs refer to the twelve months prior to the survey's reference date. SINASC: Live Births Information System with data available online at http://tabnet.datasus.gov.br/cgi/deftohtm.exe?sinasc/cnv/nvuf.def, last accessed on May 11th 2015. The reference period is from January 1st to December 31st.

Figure 4: TFR Projections Based on the UN Projection using the Medium-fertility Assumption and the Bayesian Method of Estimation, with different TFR estimates for 2010



Source: United Nations (2013) and SINASC, data available online at http://tabnet.datasus.gov.br/cgi/deftohtm.exe?sinasc/cnv/nvuf.def , last accessed on May 11th 2015.

Note: SINASC estimations using the 'bayesTFR' R program (Ševcıková et al. 2011).

Appendix 1

In the 2010 Census sample, the information on last birth is asked in the long-from questionnaire of the Census, and this from with was applied to 11% of Brazilian's household in the 2010 Census, which resulted in a sample size of more than 20 million individuals. The percentage of household randomly sampled in each municipality was 50% if less than 2,500 inhabitants (5% of total municipalities), 33% if between 2,500 and 8,000 inhabitants (34% of total), 20% if between 8,000 and 20,000 inhabitants (31% of total), 10% if between 20,000 and 500,000 inhabitants (29% of total), and 5% if 500,000 inhabitants or more (1% of total).

The total fertility rate is estimated in two steps. The first step consists of calculating age-specific fertility rates (births/person-years) for each age group of women in ages 15 to 49. For the Census survey and birth registries, the TFR single age groups were used, while for the Live Births Information System (SINASC) five-year age-groups were used because of data availability.

Appendix 2: Data from Figures

Table A1. Figure 2: Estimated Total Fertility Rate from Varied Sources

Year	Oficial Estimates	Pnad	Census	Census Conf Int	DHS	SINASC	Civil Registry	Civil Registry Census Months	Modified Brass
1991	2.89		2.57	2.43 - 2.71			1.89		2.73
1992	2.60	2.67							
1993	2.58	2.71							
1994					2.5				
1995	2.52	2.62							
1996	2.48	2.45							
1997	2.54	2.42							
1998	2.43	2.37							
1999	2.33	2.07							
2000	2.38		2.14	2.07 - 2.21		2.11	1.78		2.23
2001	2.33	1.92							
2002	2.26	1.83							
2003	2.14	1.75							
2004	2.13	1.75							
2005	2.06	1.77			1.80				
2006	1.99	1.71							
2007	1.95	1.58							
2008	1.89	1.58							
2009	1.94	1.57							
2010	1.90		1.60	1.53-1.66		1.713	1.71	1.67	1.94
2011	1.95	1.53							
2012		1.55							
2013		1.52							

Source: For a description of the data sources, see notes to Figure 2.

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