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Abstract

This study identifies the causal effect of pension generosity on women's fertility behavior. It capitalizes on Brazil's expansion of the pension system to rural workers, whose pension wealth subsequently more than tripled. Event study, difference-in-differences and instrumental variable methods show that the pension reform reduces the propensity of childbearing of women in fertile age by 10% in the short-run. Completed fertility declines by 1.3 children within 20 years after the reform, reducing the contribution base of the Pay-As-You-Go pension system in the long-run. The fertility response is strongest at higher birth parities, among older women and among mothers with sons.

JEL-Codes: J130, I380, H550, D150.

Keywords: pension wealth, fertility, old-age security hypothesis, quasi-experiment, PAYG, Brazil.

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1 Introduction

During the history of mankind, children have been natural candidates to provide old-age support for their parents. While such informal intra-family transfers have been replaced by formal pension schemes in industrialized economies, many developing and emerging economies still face the challenge of expanding coverage and generosity. Yet, how does the introduction and enhancement of public pension systems affect demography? In theoretical work, the introduction of formal old-age pensions has been associated with declining fertility rates—a link that can be rationalized by the old-age security hypothesis of fertility. Macro data seem to support the postulated trade-off between old-age pensions and the number of children per woman. Yet, there is almost no micro-economic evidence on the causal behavioral fertility responses to the introduction or expansion of pension systems.

This paper provides the first comprehensive evidence on the effect of introducing old-age pensions on the fertility behavior of women.[1] We test whether women in fertile age—who face an unanticipated, exogenous and substantial increase in public pension wealth—reduce their number of offsprings that might have otherwise served as informal old-age providers within the family. Our study is based on a quasi-experiment in Brazil, where pension eligibility rules, coverage and benefit levels were completely equalized between urban and rural workers in 1991. Rural and informal sector workers became suddenly eligible for generous and publicly provided old-age pensions—especially rural women, who had effectively been barred from the pension system.

Our first contribution is the careful identification of the causal effect of expanding old-age pensions on the fertility of women aged 15-44 using a large nationally representative data set spanning the period 1981 to 2014. Our analysis is based on an event study (ES) approach for the long-run and a difference-in-differences (DID) framework for the short-run assessment. Therefore, we exploit variation over time (pre vs. post reform) as well as variation across groups: female rural workers were strongly affected by the pension reform (treatment group), while their urban counterparts experienced almost no increase in pension wealth (control group). We complement this analysis with an instrumental variable (IV) approach to estimate the semi-elasticity of fertility with respect to pension wealth. A comprehensive assessment of alternative explanations and potential confounders, such as changes in education, child mortality, culture, etc., bolsters our identifying assumptions.

[1] Empirical studies have analyzed the effect of similar pension policies in Brazil, South Africa and Ukraine on labor supply decisions or the nutritional status of children (Duflo, 2003; De Carvalho Filho, 2008; Danzer, 2013).

As a second contribution, we study the heterogeneity of fertility responses with respect to women's age as well as the number and gender of previously born children. We expect older women, i.e., those towards the end of their fertility cycle, to respond stronger to the pension reform for five theoretical reasons: (1) they have on average older children, making their survival and, hence, their ability to provide old-age support more likely; (2) they therefore also tend to have better information about the ability of their children as future care-takers; (3) they are closer to the saturation level of fertility, prompting a more immediate fertility response; (4) they gain more in terms of pension wealth given the shorter discounting period; and (5) they are more aware and care more about reforms that will become relevant in the relatively near future. Consequently, fertility effects should be most strongly felt at higher birth parities. Regarding the sex of previously born children, we expect that women who already have a son in the family respond differently from women who have only daughters because of Brazil's inheritance tradition favoring sons.

We find that the rise in pension wealth has a disincentive effect on women's total number of births in the long-run and on the childbearing probability in the short-run. Our estimates show, first, that the pension reform reduced completed fertility at age 45 by on average 1.3 within the first 20 years after its implementation. Second, the annualized probability of childbirth falls by about 1.2 percentage points briefly after the reform, corresponding to a 10% decrease. The entire effect of the pension reform is concentrated among women in the middle and late phases of their fertility cycle (aged 30-44) whose childbearing probability declines by about 2.5 percentage points (a 25% decrease); the results for women aged 15-29 are insignificant. The fertility response is strongest at higher birth parities and among mothers with sons. A simple back-of-the-envelope calculation illustrates how these fertility responses have eroded the contribution base of the Brazilian PAYG system and, hence, introduced a structural deficit.

Historically, fertility declined at about the time when many high- and middle-income countries introduced or expanded their social security systems (Ehrlich and Kim, 2007; Schwarz, 2014). Although economic theory suggests that public old-age pensions affect individual's reproductive behavior and are a contributing cause to the fertility decline observed in modern economies, testing this link and estimating its size and significance has proven difficult, particularly at the micro-level. Most existing studies provide evidence based on theoretical OLG models; these studies use cross-country variation in aggregated time-series data to estimate the correlation between governmental public pension spending (or coverage rates) and various fertility measures (Hohm, 1975; Ehrlich and Zhong, 1998; Ehrlich and Kim, 2007; Boldrin et al., 2015). Other studies estimate the correlation between the generosity of public pension spending and fertility rates for single countries (e.g. Swidler (1983) for the US; Cigno and Rosati (1996) for Germany, Italy, the UK, and the US; Fenge and Scheubel (2017) for Germany). These papers suggest that the generosity of public old-age pensions and fertility (or population growth) are negatively correlated. Quasi-experiments are rare: Billary and Galasso (2009) use survey data on two quasi-experimental policy changes in Italy in 1992 and 1995 and compare households with worsening pension outlook and those unaffected by the reform. Their results suggest that less generous pensions increase fertility, even in industrialized economies. Surprisingly, there is little evidence on the link between old-age pensions on fertility in low- and middle-income countries where the trade-off should be especially pronounced given the reliance on informal intra-family arrangements for old-age provision. An exception is Nugent and Gillaspy (1983) who use Mexican census data and estimate a negative correlation between changes in the child-to-women ratio and the population share covered by social security at the *municipio* (county) level between 1960 and 1970.

The remainder of this paper is as follows. Section 2 discusses the theoretical background. Section 3 provides details on the Brazilian pension system and the reform of 1991 which we exploit as quasi-experiment. Section 4 describes the data of our empirical analysis, Section 5 the identification strategy. Section 6 presents the results. Section 7 discusses several robustness exercises. Section 8 concludes.

2 Theoretical Background

In caring about their economic well-being at old-age, individuals can secure sufficient and reliable provision through savings, a functioning pension system or transfers from children (Galasso et al., 2009). The so-called old-age security hypothesis of fertility (Leibenstein, 1957) implicitly posits the trade-off between children and other forms of pension savings: it postulates that parents not only procreate for the joy of parenthood (as proposed by Becker et al., 1990) but also to secure economic support from their offsprings at old-age.[2] In developing countries, widespread poverty and a lack of savings instruments imply that parents need to rely on fertility or on the public pension system (Leibenstein, 1957, 1975; Nugent, 1985). Accordingly, missing or incomplete pension institutions are associated with higher fertility rates (Boldrin and Jones, 2002; Galasso et al., 2009; Boldrin et al., 2015). The introduction of formal

[2]Empirical tests of the old-age security motive can be found in e.g. Neher (1971); Hohm (1975); Cain (1981, 1984); Jensen (1990); Hoddinott (1992); Cai et al. (2006) and Oliveira (2016). Indirect evidence shows that public transfers crowd out private transfers (Jensen, 2004; Amuedo-Dorantes and Juarez, 2015).

pension systems will, all else equal, exogenously increase the expected old-age pension wealth of individuals. This will in turn lower the need to rely on informal intra-family contracts for old-age provision and, hence, reduce fertility in the long-run. Theoretical considerations suggest, however, that the immediacy and magnitude of this fertility response differ along three dimensions: the age of the woman as well as the number and the gender of previously born children.

2.1 Age of the Woman

One crucial determinant of the response is the age of the woman, reflecting different phases of her fertility cycle.

First, there is a positive relationship between a woman's age and the precision with which she can target her desired fertility level. At the time of the exogenous pension reform older women tend to have on average more and older children than younger women. Since child mortality is most prevalent in the first five years after birth, older women face lower risks to their children's survival (*survival effect*).

Second, older parents tend to have older children and, hence, a richer set of information about their physical, mental and health constitution. They know their children's gender, ability, and attitudes. For these reasons, older women are less likely to hoard children, which lowers overall fertility (*information effect*).

Third, an immediate fertility response will be expected among women who have already achieved or are close to achieving their desired lower fertility (given the new pension wealth level), i.e., among those in the middle or late phase of their fertility cycle. Moreover, since conception cannot be perfectly controlled, older women have a shorter remaining at-risk period, which will reduce the expected number of undesired pregnancies and, hence, increase the certainty about the number of children (*saturation effect*).

Fourth, the fertility response depends on the magnitude of the pension wealth effect, i.e., how strongly the present value of future old-age pension benefits changes as a consequence of the reform. A stronger wealth increase is expected to decrease fertility more strongly. Since older women are closer to retirement age, they discount future benefit streams over shorter periods of time, leading to an overall larger pension wealth (*pension wealth effect*).

Fifth, behavioral responses to the pension reform are more likely when its consequences are more salient which can be expected for individuals closer to retirement age (*salience effect*).

Taken together, the survival, information, saturation, pension wealth and salience effects suggest that older women respond stronger to an increase in the generosity of the pension system in the short-run. Lower numbers of children, higher levels of uncertainty and a lower increase in pension wealth suggest that younger women continue having children until they have their (downwardly revised) desired number of children. In essence, a pension reform will reduce the fertility of younger women less than of older women in the short-run. In the long-run younger women can respond stronger, as they have more scope for adjusting to their new desired level of fertility.[3]

To test these theoretical predictions empirically, our analysis differentiates between long- and short-run outcomes (completed fertility at age 45 vs. annual childbearing probabilities) as well as between younger and older women (aged 15-29 and 30-44).

2.2 Number of Previously Born Children

A woman's probability of childbearing tends to be negatively correlated with the number of previously born children, i.e., with birth parity. As a direct implication of the age-specific fertility responses, women will reduce higher birth parities if the pension reform affected all women similarly. Introducing a full coverage pension scheme with a *de facto* flat pension for rural workers will cut the overall number of children rural women have. Hence, we do not expect negative effects on the first or second child but on higher parities.

2.3 Gender of Previously Born Children

Women might react differently to the pension reform depending on whether they have already given birth to a son or not. Such gender-sensitive behavior could be explained by cultural norms or inheritance traditions favoring sons. While there is no indication for gender-related abortions in Brazil (Chiavegatto Filho and Kawachi, 2013), prior evidence suggests that women who have initially given birth to girls are more likely to have additional children in the hope to conceive a boy (Reynolds, 2018); these women are also less likely to use contraception (Arnold, 1992). The presence of a son seems especially important for inheriting land: While inheritance laws are relatively gender-neutral in Brazil, land is nevertheless typically passed to a son. In consequence, only 11% of land in Brazil is owned by women (Deere and León, 2003).

[3] As a theoretical possibility, young women may postpone having more children—maybe because they want to wait and see how changes in the pension system materialize after the reform. In this case, they would reduce fertility in the short-run in the years following the reform.

3 Fertility and the Policy Reform in Brazil

3.1 Fertility Trends

Brazil has experienced sharply falling fertility since the middle of the 20th century. The decline started in the 1940s in more affluent urban areas and in the 1970s more broadly. Possible explanations are changes in the demand for fertility (owing to educational reforms or small-family norms evolving from the mass media) and in the supply of family planning services (owing to health care reforms and access to contraception), or changes in nuptial or residential patterns (Martine, 1996; Rios-Neto et al., 2018). Most of the expansion in urbanization, health care, education, marriage, and telenovelas took place between the 1960s and early 1980s (Martine, 1996; Caetano and Potter, 2004; La Ferrara et al., 2012). Starting in the 1990s, demographers observe a fertility stopping behavior among older women with sterilization; fertility becomes concentrated among women under 25 years of age with rising rates of teen pregnancies (Rios-Neto et al., 2018). In this environment of declining fertility, Brazil implemented a comprehensive pension reform.

3.2 The Pension Reform

Before the 1970s, only formally employed urban workers were entitled to old-age pensions in fragmented occupation specific programs. With the Assistance to the Rural Worker Program (FUNRURAL) in 1971, the first pension scheme for formally employed rural workers was established; however, it left the vast group of informal or self-employed rural workers uncovered.

The first free elections in 1985 marked the end of the military rule in Brazil and paved the way for the Constitution of the Federative Republic of Brazil, approved in October 1988. Unlike in most other countries, the Brazilian public social security system (including pensions) was enshrined in the Constitution, making the pension system hard to modify or even abolish. Accordingly, no significant reforms were undertaken until the early 2000s (Hunter and Sugiyama, 2009).

With the codification into law, the coverage of social security was universalized such that informal and rural workers got—for the first time in Brazilian history—access to social protection and pension benefits. As benefits were equalized for eligible individuals, the reform had a particularly strong impact on the rural population. The implementation of the pension reform started with its lawful approval in July 1991. Potential anticipation effects, e.g., that individuals foresee the reform and adjust their behavior accordingly, are weak owing to considerable information deficits about the specific design of the new pension system before its implementation in 1991.

The rules and provisions of the Brazilian pension scheme before and after the reform and the different pension formulae (old-age, lengths-of-service, disability) after the reform are illustrated in Table 1.[4]

Before the reform, rural workers had to document formal employment to receive a pension; but no more than 2-6% of rural women were legally employed in the 1980s, according to official statistics (PNAD data; Neri (2002)). Even among the formally employed, at most one household member was eligible for old-age pensions, mostly the husband. Informal rural workers were entirely barred from the pension system.

For urban workers, the reform left the eligibility rules (e.g., qualifying age thresholds) and the benefit calculation roughly unchanged. Quite differently, the pension reform expanded coverage and pension benefits dramatically for rural workers. Irrespective of being able to prove formal employment or having made any social security contributions, they became eligible for old-age retirement benefits from a qualifying age of 55 (60) years for women (men).[5] Even unpaid workers in subsistence agriculture became entitled. The pension generosity increased predominantly owing to a reduced retirement age (women minus 10 years, men minus 5 years); an increase in benefits from 50% to 100% of the minimum wage; and, the abolition of the one-beneficiary-per-household restriction. In essence, the pension system in rural areas was transformed from exclusive and meager to full-coverage and generous.

In line with unchanged rules, the indexed number of urban pension eligibles and recipients seems hardly affected by the reform in 1991, according to Figure 1. Quite differently, eligibility for rural pensions more than doubled between 1990 and 1992, as did the number of rural pension recipients with a small delay. The enrollment of newly eligible pensioners was not implemented immediately owing to bureaucratic delays (De Carvalho Filho, 2008). However, benefits were automatically adjusted and more than 2 million rural beneficiaries started receiving pensions until 1994, providing a strong signal for the credibility of the reform (Dias and Amaral, 2001).

[4] The Brazilian pension reform of 1991 also affected the rules governing maternity leave. From an individual perspective, the value of the old-age pension will be far larger than the value of a maternal leave benefit. Moreover, benefits such as paid maternity leave reduce the relative costs of children and would consequently increase fertility.

[5] Formally, an extremely soft documentation requirement (confirmation of work) was still in place. Agricultural and fishing associations, unions, government agencies, and even the church were entitled to confirm any kind of work in agriculture, under sharecropping/tenancy contracts, as co-owner of land, or as worker in small scale mining. Anecdotal evidence suggests that destitute elderly were granted proof of agricultural work even if they lacked birth certificates or other proof of age.

Table 1: The Brazilian Retirement Pension System Before and After the Reform in 1991

Occupation	Pension type		The pension system before the reform	The pension system after the reform
Rural workers	**Old age**	**Eligibility criteria:**	Minimum age of 65 years, documented formal employment for one out of past three years or three years overall.	Minimum age of 60/55 for males/females. Weak documentation requirements.
		Benefit level:	Flat and equal to 50 % of the minimum wage.	Same rules as urban workers.
		Restrictions:	Only one person per household is eligible.	No restriction in the number of beneficiaries per household.
	Length of service	**Eligibility criteria:**	Not eligible.	Same rules as urban workers.
	Disability	**Eligibility criteria:**	Available at any age.	Available at any age.
		Benefit level:	Flat and equal to 50% of the minimum wage.	Same rules as urban workers.
		Restrictions:	Needs to stop working.	Needs to stop working.
Urban workers	**Old age**	**Eligibility criteria:**	Minimum age of 65/60 for males/females.	Same age limits as before.
		Benefit level:	Minimum benefit level is 90% of the minimum wage.	Benefit is 70% of earnings-based benefit (average income of last 36 months) at the minimum eligibility age plus 6% for each additional year of service, up to 100%. Minimum benefit increased to 100% of the minimum wage.
		Restrictions:	Needs to quit the current job to apply for benefits. No earnings/retirement test after that.	Allowed to stay in the current job.
	Length of service	**Eligibility criteria:**	Eligibility after 30 years of work. Full benefits after 35 years of work. Fewer years for some types of work. No minimum age requirements.	Eligibility after 30/25 years of work for males/ females. Full benefits after 35/30 years of work for males/females. No minimum age requirements.
		Benefit level:	Benefits determined by years of documented work and recent labor earnings. Minimum benefit is 90% of the minimum wage. Bonus for continued work beyond maximum eligibility period.	Benefit is 70% of earnings-based benefit (average income of last 36 months) at the minimum eligibility age plus 6% for each additional year of service, up to 100%. Minimum benefit increased to 100% of the minimum wage.
		Restrictions:	Needs to quit the current job to apply for benefits. No earnings/retirement test after that.	
	Disability	**Eligibility criteria:**	Available at any age.	Available at any age.
		Benefit level:	Benefit is 90% of the minimum wage.	Benefit is 80% of earnings-based benefit plus 1 % for each year of payroll tax contributions, up to 100%. Minimum benefit level of 100% of the minimum wage.
		Restrictions:	Needs to stop working.	Needs to stop working.

Benefit Calculation After the Pension Reform:

After the reform, monthly benefit levels for all types of pensions (old-age, length of service and disability) were equalised for all eligible individuals and formally determined as follows:

$$Old\ Age_i = max\left\{(0.7 + 0.01 \times \tfrac{P_i}{12}) \times B_i; 1 \times minwage\right\} \qquad Length\ of\ Sevice_i = max\left\{(0.7 + 0.06 \times \tfrac{P_i}{12}) \times B_i; 1 \times minwage\right\} \qquad Disability_i = max\left\{(0.8 + 0.01 \times \tfrac{P_i}{12}) \times B_i; 1 \times minwage\right\}$$

with B_i being the benefit income determined as the average of the final 36 monthly earnings on which individual i has paid social security contributions. P_i indicates the number of months during which individual i paid social security contributions (payments of the payroll tax) in all her working career. The minimum benefit level—which is also applicable for all individuals who fail to prove formal earnings and/or social security contributions—equals 100% of the minimum wage for all types of pensions. The pension system makes 13 benefit payments per year, which are uprated along with the national CPI (the minimum pension can be uprated by more than the CPI).

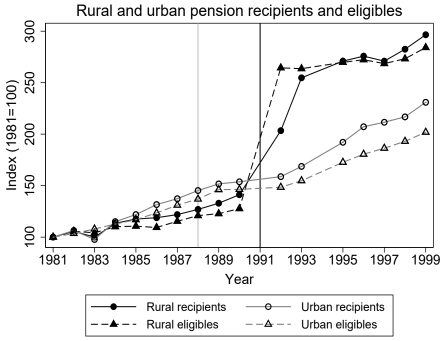

Figure 1: DEVELOPMENT OF THE INDEXED NUMBER OF PENSION ELIGIBLES AND RECIPIENTS (1981=100), BRAZIL 1981-99

Note: Graph shows the indexed number of rural and urban old-age pension eligibles and recipients (1981=100). Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Eligibility and receipt defined by age: 65 for rural elderly before the reform, 55/60 for women/men after the reform; 60/65 for urban women/men before and after the reform. Rural and urban groups are defined according to occupation and location (for details see Section 4) *Source:* PNAD 1981-90, 1992-93, 1995-99.

Figure 2 documents average pension incomes, paid out to rural and urban households in the years before and after the reform. Evidently, the reform in 1991 was comparatively more beneficial for rural pensioners. The mean pension benefits of urban pensioner households increased by 15% (118 BRL) whereas the increase was 136% (315 BRL) for rural pensioner households between 1990 and 1992.

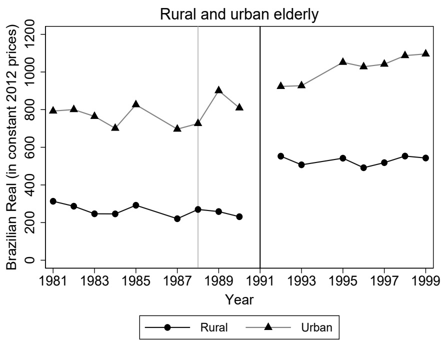

Figure 2: AVERAGE MONTHLY PENSION INCOME, BRAZIL 1981-99

Note: Graph shows average household pension income of rural and urban households with at least one old-age pension recipient, based on eligibility by age. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined according to household location. *Source:* PNAD 1981-90, 1992-93, 1995-99.

While the figure illustrates the immediate change to the cash flow of pensioners after the reform, the identifying variation relevant to our fertility analysis is the one in pension wealth of not-yet retired workers: Figure 3 shows the increase in accumulated discounted gross pension wealth of rural and urban female workers (top) and rural and urban worker couples (bottom). The pension wealth of rural worker couples rose by factor three, illustrating that rural workers—and especially women—benefitted disproportionately from the reform, while urban workers were hardly affected.

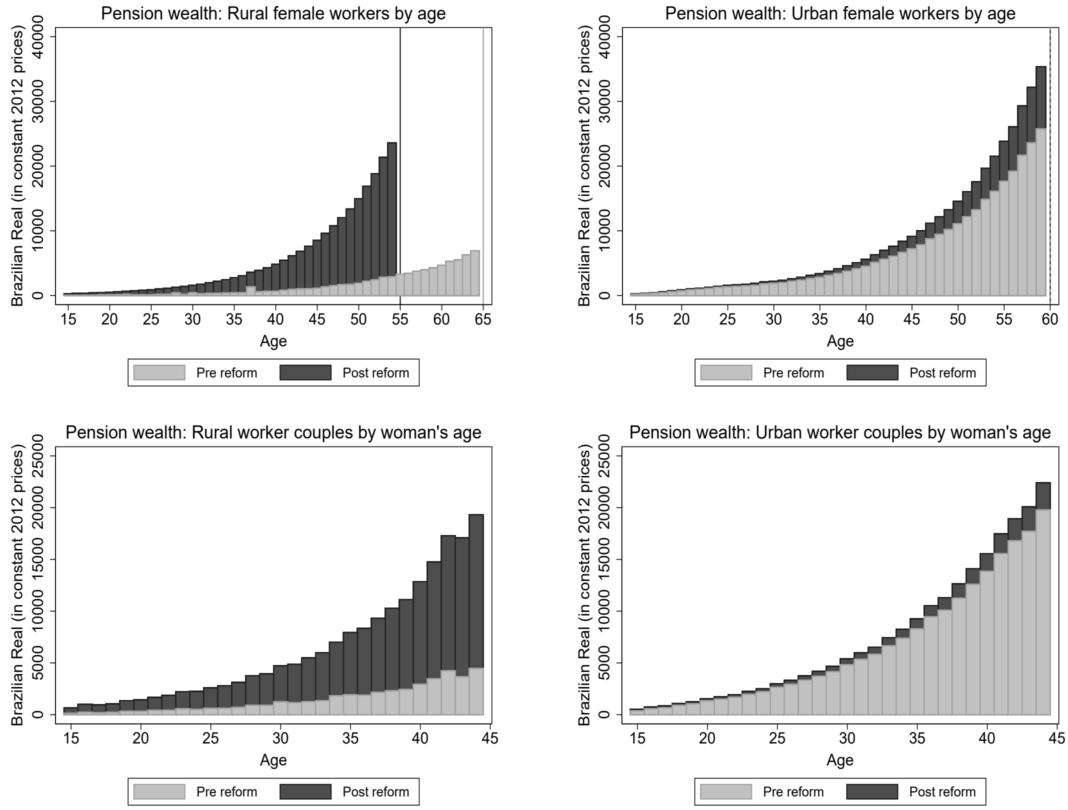

Figure 3: Gross present pension wealth of Rural and Urban Female Workers and of Rural and Urban Worker Couples with a Women aged 15-44.

Note: Pre- and post-reform pension wealth is computed as the present value of expected old age benefits before/after the reform adjusted for real interest rates and average survival probabilities (computed using IBGE mortality tables; first time available in 1998): $Pension\ Wealth = \sum_{t=0}^{T-a} s_{a,t} \times \frac{1}{(1+i)^t} \times pension_t$, with $s_{a,t}$ denoting the probability of a person of age a in a given year surviving until year t; $T-a$, indicates the remaining maximum lifespan differentiated by sex and birth cohort; i is a constant discount rate (12%); and $pension_t$ denotes the old age pension benefits in t. A non-retired person receives the pension starting in a future period $t > 0$, defined by the person's age and the regular retirement age. Rural and urban groups are defined according to occupation and location (for details see Section 4). Figure A1 shows pension wealth of rural and urban male workers. *Source:* PNAD 1981-90, 1992-93, 1995-99.

The notable increase in pension wealth of rural workers led to substantial declines in poverty among rural households (Schwarzer, 2000; Schwarzer and Querino, 2002), a spike in retirement among the rural elderly (De Carvalho Filho, 2008) and some increase in school enrollment among children co-residing with beneficiaries (De Carvalho Filho, 2012).

With the pension reform, Brazil embarked on a path from fragmented provision to inclusive universalism that relied on changes in the population's beliefs: The introduction of the new Constitution and the move towards democracy were embedded in new political organizing principles and a new and inclusive political language so that citizens could believe in the announced reforms: The success of Brazil's Constitution is exemplified in the political competition based on universal access to elections,[6] greater political stability, effective stabilization efforts, and the establishment of a relatively professional bureaucracy (Melo, 2017). Surveys regarding the effectiveness of the pension system in the 1990s reveal that 93% of rural households reported no or very little difficulties in applying for the pension and only very short waiting time periods before approval. Rural recipients reported very high levels of satisfaction with the pension system and payout of benefits (Delgado and Cardoso Jr., 2005).[7]

4 Data

Our empirical analysis exploits several cross-sections (1981-90, 1992-93, 1995-99, 2001-09, 2011-14) of the Brazilian National Sample Household Survey (PNAD), conducted annually by the Brazilian Census Bureau (IBGE) since 1971.[8] The PNAD is a nationally representative high-quality survey that collects a rich set of variables on individuals and their households;[9] its sample size is gradually increasing in line with population growth: from 1981 to 1985 (from 1986), the sample size was about 1/250 (1/500) of the Brazilian population.

Our analysis of fertility responses to the pension reform is based on two outcome variables: First, to analyze completed fertility we compute the total number of births of women aged 45: This measure allows a long-run assessment of changes in the number of children, which is important to capture the demographic significance of the reform and to rule out the possibility of timing effects in short-run fertility responses. The information to construct this variable is obtained from the PNAD survey years with a fertility section including information on the total number of births (alive or dead), i.e., the 1984-85, 1992-93, 1995-99, 2001-09 and 2011-14 waves. Our sample includes female respondents of the birth cohorts 1930 to 1969, who are between 45-69 years of age in a given survey year ($N = 767,465$).[10] We retrospectively assign the

[6]With the introduction of suffrage for illiterates the fraction of the total population (not eligible voters) voting in presidential elections rose from below 20% in the 1960s to more than 50% in 1989.

[7]While interviews in 2012 reveal a relatively high level of distrust in legal and political institutions in Brazil in general, respondents articulated trust in the persistence of the pension system (Lui Gallassi, 2012).

[8]The PNAD was not conducted in the years 1980, 1991, 1994, 2000 and 2010. Since 1981, the entire country is covered in the survey, apart from extremely remote rural areas of the following States of the Brazilian Federation: Acre, Amapa, Amazonas, Para, Rondonia and Roraima (accounting for about 1.7% of the Brazilian population). The PNAD relies on the definition of rural areas as of 2004.

[9]For variable descriptions and descriptive statistics, see Table 2 and Tables A1-A5.

[10]In 1984, the fertility section is only answered by women up to 54 years old. We exclude women above age 69 to account for the selective mortality of women with respect to their total number of births, see e.g. Barclay and Kolk (2019). Further, we exclude observations if the total number of reported births exceeds 20.

reported total number of births of women older than 45 years to the year in which they were 45 years old. We exclude individuals living in urban areas of those states that do not cover rural areas in the PNAD ($N = 50,386$). The sample size is $N = 716,215$.

Second, to assess the short-run effect of the reform on the probability of childbearing, we construct a dummy variable taking on the value of one if a woman has given birth to a child in the past 12 months, zero otherwise (based on the household roster and birth data).[11] Our sample includes female respondents in fertile age between 15 and 44 in the years 1981-99 ($N = 1,548,011$).[12] We exclude the 2000s from our analysis owing to later reforms of the pension and health care systems. We exclude individuals living in urban areas of those states that do not cover rural areas in the PNAD ($N = 118,851$) and female rural workers from mixed occupation households ($N = 102,341$).[13] This leaves us with a working sample of $N = 1,326,819$.

Our research strategy exploits the fact that the reform was disproportionately beneficial for rural workers (treatment group) compared to urban workers (control group), in absolute and relative terms. In line with eligibility rules of the old-age pension scheme we categorize rural and urban workers as follows: First, we rely on the classification of the individual's current occupation performed during the reference week (45% of the long-run and 46% of the short-run sample). Second, for individuals not employed during the reference week we use retrospective occupational information available for up to 4 years prior to the reference year (5% and 6% of the samples). Third, for individuals with insufficient personal occupational information we assign the rural status based on the occupation of the family head (21% and 41% of the samples). Finally, for individuals with insufficient personal and household head information we assign the status using information on the household location, i.e., rural or urban residence (29% and 7% of the samples).[14]

[11] Multiples or children born in the same year are coded like single births. This coding scheme also implicitly accounts for the fact that multiple births are more common among older mothers.

[12] We keep women who are younger than 15 (N =251; 0.03% of the sample) or older than 44 (N =1,171; 0.21% of the sample) and who give birth to a child in the sample and recode their ages to 15 or 44, respectively. In a robustness check we re-run our main regressions excluding women outside the 15-44 age range and find very similar results (Table A6).

[13] In a robustness check we re-run our main regressions including mixed occupation households and find very similar results (Table A7).

[14] As we allocate the urban or rural group predominantly based on occupation, selective residential migration is not a threat to our identification strategy. Note, that internal migration in Brazil was low in the relevant time period: Below 8% (5%) of urban (rural) inhabitants had changed their residence across federal states in the past eight years, according to the censuses 1991 and 2000. Further, we investigate the robustness of our classification with alternative approaches, i.e., classifying the status only with personal occupation and household location (excluding step 3) or only with household location (excluding steps 1-3); alternative approaches provide similar results (Table A8). We acknowledge that some forward-looking workers may have self-selected into specific occupations based on differences in the pension system. For women in the middle and late phase of their fertility cycle this selection is, however, a predetermined characteristic.

Table 2: Descriptive Statistics

Sample A: Long-run analysis	Urban occupation (N=596,928)				Rural occupation (N=119,287)			
	mean	sd	min	max	mean	sd	min	max
Total number of births	3.56	2.96	0	20	5.83	4.03	0	20
Childlessness (number of births=0)	0.11	0.31	0	1	0.08	0.26	0	1
Years of education	6.16	4.88	0	17	2.27	2.93	0	17
Married	0.58	0.49	0	1	0.77	0.42	0	1
Sample B: Short-run analysis	**Urban occupation (N=1,103,371)**				**Rural occupation (N=223,448)**			
	mean	sd	min	max	mean	sd	min	max
Newborn child under 1 year old	0.08	0.27	0	1	0.13	0.34	0	1
Age	27.79	8.39	15	44	27.24	8.86	15	44
Age partner	32.20	9.22	15	99	32.70	10.39	15	98
Years of education	6.79	4.08	0	17	2.93	2.78	0	17
Married	0.52	0.50	0	1	0.62	0.49	0	1
Household Income	2,806	4,658	0	820,542	1,106	2,530	0	183,294
Woman's income share	0.19	0.30	0	1	0.07	0.20	0	1
Wealth index	0.15	0.93	-8.69	12.13	-0.22	1.01	-7.93	9.98
Not worked in reference week	0.51	0.50	0	1	0.70	0.46	0	1
Worked 1-10 hours ref. week	0.01	0.11	0	1	0.00	0.04	0	1
Worked 11-20 hours ref. week	0.05	0.22	0	1	0.04	0.20	0	1
Worked 21-30 hours ref. week	0.07	0.26	0	1	0.07	0.26	0	1
Worked 31-40 hours ref. week	0.14	0.35	0	1	0.08	0.27	0	1
Worked 41-50 hours ref. week	0.15	0.36	0	1	0.07	0.25	0	1
Worked 51-60 hours ref. week	0.04	0.20	0	1	0.03	0.16	0	1
Worked >60 hours ref. week	0.02	0.15	0	1	0.01	0.09	0	1
No. of adults in household	2.90	1.43	0	22	2.84	1.35	0	16
Caretaker in household	0.10	0.30	0	1	0.08	0.28	0	1

Note: Sample A consists of Brazilian women of the birth cohorts 1930 to 1969, aged 45-69. Sample B consists of Brazilian women aged 15-44. Rural and urban groups are defined by: (1) occupation during the reference week, (2) occupation up to 4 years prior to the reference year, (3) occupation of the family head, (4) household location. Mixed urban-rural couples excluded. For more information see the variable descriptions in Tables A1-A5. *Source:* Sample A: PNAD 1984-85, 1992-93, 1995-99, 2001-09, 2011-14); Sample B: PNAD 1981-90, 1992-93, 1995-99.

5 Identification Strategy

The Brazilian social security reform of 1991 provides plausibly exogenous variation for the identification of the effect of the pension reform on reproductive behavior: As rural workers received at most sketchy and meager pensions before the reform, the old-age security motive was probably a key factor in their fertility decisions. Accordingly, the substantial pension reform is expected to induce a disproportionate negative fertility response among rural women.

To capture these effects, we employ three strategies: First, we implement an event study (ES) approach for completed fertility of women aged 45. Second, we estimate policy effects by exploiting the difference in annual birth probabilities over time (pre vs. post reform) between treatment and control group (rural vs. urban women). This difference-in-differences (DID) estimator captures the deviation of actual rural fertility trends from expected fertility trends that would have been observable without the pension reform. Third, we estimate the semi-elasticity of fertility with respect to pension wealth in an instrumental variable (IV) setup. Here, we regress the incidence of childbearing on the natural log of accumulated gross present pension wealth, which we instrument with the (treatment group × post reform) interaction.

5.1 Event Study Estimation

We explore the long-run effect of the pension reform in 1991 on completed fertility by analyzing the total number of births of women aged 45 using an ES approach. Our sample covers the period 1975-2014. We include interactions of the time dummies and the treatment indicator, $RURAL_g \times YEAR_t$, for all periods in our model.

In practice we estimate the following equation:

$$\begin{aligned} y_{igtr} = \alpha + & \sum_{t=-16}^{23} \beta_t (RURAL_g \times YEAR_t) \\ & + \gamma RURAL_g + \lambda_t + \phi_r + \psi' X_{igtr} + \epsilon_{igtr}, \end{aligned} \tag{1}$$

with the dependent variable y_{igtr} representing the number of births of woman i at age 45 of group g in region r in year t; $RURAL_g$ is the indicator for the treatment group. $YEAR_t$ are indicator variables for years $t \in [-16, 23]$. The specification allows for sixteen $(\beta_{-1}, \beta_{-2}, ..., \beta_{-16})$ pre-treatment (lead) effects and twenty-three $(\beta_{+1}, \beta_{+2}, ..., \beta_{+23})$ post-treatment (lag) effects; hence, the β_t coefficients capture the differential fertility trend between treatment and control groups for each year between 1975 and 2014. If the pension reform led to a decline in the total number of births, the post-reform β_t's carry a negative sign. The model includes birth cohort fixed effects λ_t and region (States of the Brazilian Federation) fixed effects ϕ_r to capture time-invariant regional specific factors.[15] Finally, our regression incorporates the covariate vector X_{igtr}, including years of schooling and a dummy for married. The idiosyncratic disturbance term is denoted by ϵ_{igtr}. Standard errors are clustered at the region level (States of the Brazilian Federation: 20 clusters).

An attractive feature of an event study is that the interactions of post-treatment time dummies with the treatment indicator reflect dynamics of the completed fertility response after the reform. For instance, the lag coefficients indicate whether the treatment effect fades out, stays constant, or increases over time. The effect is expected to increase after the reform since women close to 45, who have revised their fertility target downwards, have only limited capacity to restrict their fertility given

[15] We implicitly assume that women have not changed their state of residence during their fertile age. In fact, migration in Brazil was low in the relevant time period.

their proximity to age 45 and their previously born children; younger women have a longer remaining adaption period. Further, fertility norms adjust slowly. The lead coefficients shed light on the common trend in completed fertility between the rural (treatment) and urban (control) group before the reform; insignificant pre-treatment coefficients are indicative of a common trend before the reform.

5.2 Difference-in-Differences Estimation

We investigate the short-run effect of the pension reform on the reproductive behavior of women in fertile age using the DID method. The research design is not sensitive to time-invariant differences between rural (treatment) and urban (control) workers, like persistent family roles or tastes for fertility. At the same time, DID accounts for macro trends that affect rural and urban areas similarly. This is relevant as Brazil underwent a period of macroeconomic stabilization in the early 1990s. The universal and exogenous change in pension generosity permits the estimation of causal fertility responses by comparing childbearing trends between the treatment and control groups before and after the pension reform (the treatment).

Formally expressed, our OLS reduced-form model with repeated cross-sections can be written as:

$$\begin{aligned} y_{igtr} &= \alpha + \beta(RURAL_g \times POST_t) \\ &\quad + \gamma RURAL_g + \lambda_t + \phi_r + \psi' X_{igtr} + \epsilon_{igtr}, \end{aligned} \tag{2}$$

with the dependent variable y_{igtr} representing a dummy equal to one if woman i of group g in region r in year t has given birth to a child within the last 12 months before the interview, zero otherwise. $RURAL_g$ is an indicator variable for rural female workers and $POST_t$ a dummy variable that indicates observations in the years after the pension reform in July 1991. The interaction of interest $RURAL_g \times POST_t$ indicates rural female workers eligible for a future pension after the reform. The coefficient of interest β captures the differential fertility trend between treatment and control groups. If the pension reform led to a decline in fertility, β carries a negative sign. The model includes a full set of time fixed effects λ_t; hence, our approach relies on the assumption that no shock affects the two groups differently. Further, we add region (States of the Brazilian Federation) fixed effects ϕ_r to capture time-invariant regional specific factors. Finally, our regression set-up incorporates time-varying covariates represented by the vector X_{igtr}. The set of controls includes individual characteristics such as years of schooling[16], a dummy for married, the age of the woman, the age of

[16] Our results are robust when using other definitions of education, like literacy or the highest degree obtained (see Table A9).

the woman squared and dummies for birth parity, i.e., 1 to 5+ previously born children (zero being the omitted category); job related characteristics are dummies for 1-10, 11-20, 21-30, 31-40, 41-50, 51-60 and 60+ hours of professional work in the reference week (zero hours, i.e., not working, being the omitted category) and the share of household income (without pension income) earned by the woman; household characteristics are the natural logarithm of monthly household income (without pension income)[17] and wealth (using the first principal component from a set of 45 indicators for housing characteristics and durables), dummies for the number of adults in the household and a dummy for the presence of a potential caretaker in the household (i.e., a non-working pensioner aged 60 or above) as well as age of the partner and age of the partner squared. Further we include time-varying control variables measured at the group and regional level: At the regional level we control for the prevalence of TV reception and for regional industry-specific trade shocks, as telenovelas (see La Ferrara et al., 2012) and trade liberalization (see Braga, 2018) have been identified as influencing factors of fertility in Brazil. At the regional level and differentiated by urban and rural location, we control for the shares of protestants, of individuals of other faiths, or of no religion (catholic being excluded as omitted category) and for shares of individuals of different races, i.e., black, mixed-race and other (white being excluded as omitted category); at a macro-regional level, we also control for child mortality rates (under 60 months of age), based on Demographic and Health Survey (DHS) data.

The idiosyncratic disturbance term is denoted by ϵ_{igtr}. Standard errors are clustered at the region level (States of the Brazilian Federation: 20 clusters). Since the Brazilian pension reform can be considered unanticipated and exogenous, the inclusion of covariates should not change our estimates of β, while improving precision. To illustrate this, we include covariates in the results tables in a step-wise fashion.

Taking advantage of the large sample size, we perform subgroup analyses to test whether the effect of the pension reform differs across women at different phases of their fertility cycle. First, we split the estimation sample into younger (aged 15-29) and older (aged 30-44) women and estimate model (2) for both sub-samples. In the robustness section, we also address even more granular age groups. Second, we investigate the intensive margin of fertility by splitting the sample into groups of women at different birth parities: This analysis sheds light on the reform effect on childbearing for women with zero, one, two, three, four, and at least five previously born children. Finally, we also separate the sample into groups of women who already have at least one son, at least two sons, only daughters and those who have no prior

[17] Including pension income does not change the results, see Table A9, col. 2.

children at all. Comparing these estimates sheds light on the sensitivity of fertility responses to the presence of a boy—and, hence, a potential heir—in the family.

The validity of DID rests on the assumption that rural and urban women would not have had any systematic different fertility trends in the absence of the reform. We show the plausibility of the common trend assumption using a test method in the spirit of Granger (1969). The test sample contains ten pre-reform (1981-90) and seven post-reform (1992-93, 1995-99) years. The interaction terms, $RURAL_g \times YEAR_t$, test for differences between treatment and control groups, conditional on a full set of group and time effects and including all covariates.

Each estimate of Table A10 reports the year specific difference between treatment and control for the full sample (col. 1), the sample of younger women (col. 2) and the sample of older women (col. 3). The fertility trends are not significantly different between rural and urban women in the pre-treatment period, as evidenced by the insignificant pre-reform interaction coefficients, almost all of which are very close to zero. Only one estimate (in 1982 for older women) out of 27 (3.7%) is statistically significant at the 10% level, which is a fully acceptable incidence rate for Type-I errors. This suggests that rural and urban women were on the same fertility trajectory before the pension reform.

5.3 Instrumental Variable Estimation

Beside the policy effect of the pension reform, we aim at estimating the semi-elasticity of fertility with respect to pension wealth. Since pension wealth is potentially endogenous, we exploit the timing of the policy and the differential impact across population groups in Brazil, i.e., the $RURAL_g \times POST_t$ interaction, as an instrumental variable. Therefore, we apply a 2SLS estimation strategy that allows us to crystallize the exogenous variation in pension wealth.

In the first stage, we regress the age-specific accumulated discounted gross pension wealth ($pensionwealth_{igtr}$) on the instrument as well as the full set of covariates:

$$\begin{aligned} ln(pensionwealth)_{igtr} &= \sigma + \theta(RURAL_g \times POST_t) \\ &\quad + \eta RURAL_g + \delta_t + \tau_r + \kappa' X_{igtr} + \mu_{igtr}. \end{aligned} \tag{3}$$

In the second stage, we regress birth probabilities on the predicted values of pension wealth:

$$\begin{aligned} y_{igtr} &= \alpha + \beta ln(\widehat{pensionwealth})_{igrt} \\ &\quad + \gamma RURAL_g + \lambda_t + \phi_r + \psi' X_{igtr} + \epsilon_{igtr}. \end{aligned} \tag{4}$$

This estimation procedure requires a sufficiently strong and exogenous instrument. We assess the strength of the instrument with the F-statistic of the first stage, which exceeds the critical value multiple times across all samples. Also, t-statistics of the excluded instrument in the first stage are close to or over 60, suggesting a very strong instrument. The exogeneity of the instrument is plausible since the universal pension reform was implemented exogenously and without anticipation. Also, the pension scheme did not create incentives to work. Essentially, the reform affects fertility levels only through its effect via pension wealth.

5.4 Addressing Potential Confounders

Our identification strategy rests on the assumption that no other policy changes or alternative determinants of fertility explain the disproportionate decline in rural fertility behavior. Therefore we analyze trends in marriage rates, education, health care, incomes, wealth and employment between rural and urban women and show that the treatment and control groups were on very similar trajectories regarding these factors over time. We also discuss the role of cultural change in Brazil with respect to religion and telenovelas, the role of trade liberalization and patterns regarding the quantity-quality trade-off between the number of children and the investment made in their quality, i.e., education. In our regression analysis we control for these factors.

Marriage. Marriage and fertility are positively correlated. But neither the Constitution nor the pension reform contain any changes affecting the incentives to marry, not least since pension eligibility is independent of marital status. In Figure 4 (top left) we compare marriage rates of rural and urban female workers. Average marriage rates are higher for rural women but the rates tend to be very stable over the 19 year observation period.

Education. As more educated women face greater opportunity costs of having children, and as education is in general negatively correlated with fertility, an unbalanced educational expansion between rural and urban areas may directly influence the observed fertility rates. The 1988 Constitution mentions education as a basic right. To reduce the illiteracy rate elementary education was made compulsory regardless of age implying compulsory education for adults in so-called popular schools. However, a lack of schools delayed the desired expansion by more than ten years.[18] In Figure 4 (top right) we compare years of schooling of female rural and urban workers. Average years of schooling increase over time; however, rural and urban women are on similar

[18] Only in 1998, Brazil set up the Fund for the Maintenance and Development of Primary Education and Valorisation of Teachers (FUNDEF) to improve enrolment rates, with some effect (De Mello and Hoppe, 2005; Borges, 2008).

paths. Another threat to identification might stem from the educational expansion for children. Since higher education increases the economic value of a child, parents could afford their targeted level of old-age support with fewer children, as predicted by the quantity-quality trade-off theory. However, since the educational expansion was severely delayed and limited to primary education (which had been free and—in theory—compulsory before) we expect no differential growth in education among children of rural women. When analyzing years of education of children aged 15 or 7-24, we find positive secular trends, but no differences between urban and rural households or between boys and girls (Figure A4).

Health. Changes in health care might explain the fertility decline if improvements in child survival were stronger in rural areas. However, the infant and child mortality decline in the critical period between 1986 and 1996 was stronger in urban (38%) than rural (32%) areas, according to data from the DHS (Figure 4, middle left). Consequently, we would underestimate the true fertility decline owing to the pension reform. Another explanation might be a disproportionate expansion in access to health care (e.g., contraception) in rural areas; this might improve rural women's ability to control fertility. The major reform for the provision of family planning services, however, took place only after the pension reform in 1997 (Caetano and Potter, 2004).

Economic empowerment. Fertility may depend on women's relative economic potential and power within the household. Comparing the income of women as a share of household income and, therefore, the relevance of women's paid work for the household, we do not find any disproportionate gain among rural women (Figure 4, middle right).

Income and wealth. The tightening of minimum wage rules in the Constitution of 1988 might be problematic if the rural population experienced disproportionate income gains. Figure A5 (top left) compares average monthly household incomes (excluding pension income) for households with rural or urban women in childbearing age. Household income fluctuates over time, but there is no differential trend in household income after the Constitution was approved; further, the level difference between rural and urban households is stable for absolute (Figure A5, top left) and relative (Figure 4, middle right) income as well as for wealth (Figure A5, top right).

Labor force participation. Greater female labor force participation implies greater opportunity costs of children and may, hence, lower fertility. Brazil underwent comprehensive unilateral trade liberalization during the late 1980s (for agricultural products and mining) and 1990s (for manufacturing), possibly affecting the labor market. However, since the sectoral distribution of women remained almost constant during the 1990s and 2000s and since female workers were mostly employed in the non-tradable sector (with no discernible spill-over effects), women were at most mildly affected by trade liberalization (Gaddis and Pieters, 2017). Given that urban and rural workers were exposed to similar trade shocks in terms of tariff reductions (Figure A6), trade liberalization should not confound the fertility effect of the pension reform. In the regressions we control for the region-industry specific trade shocks by incorporating annual tariff rates taken from De Paiva Abreu (2004).[19] The trajectory of labor force participation between rural and urban women is quite comparable, suggesting that the fertility pattern cannot be driven by differential employment growth (Figure 4, bottom left). At the same time, average working hours of women exhibit a declining trend already before the reform. After the reform we observe a disproportionate fall in the working hours of female rural workers (Figure 4, bottom right), which is attributable to the emergence of part-time jobs (11-20 hours) in rural areas (Figure A7); if anything, this pattern works against declining fertility of female rural workers.

Religion. Brazil experienced a phase of cultural modernization during the 1980s and 1990s. The influence of the catholic church, with its promotion of traditional family roles and the rejection of modern contraception, diminished slowly. Since the PNAD does not include consistent information on religious affiliations of individuals over time, we use Brazilian Census data from 1980, 1991 and 2000 to generate the regional composition of religious affiliations, for which we control in our regressions. In graphical analysis the most marked observation is a rising share of protestants, of people of other faiths and of people without religion, at the expense of catholicism (Figure A8). This trend is more pronounced in urban than in rural areas and cannot explain the stronger decline of fertility in rural areas. However, membership in the church may be an imprecise determinant of fertility. Based on World Value Survey (WVS) data from 1991 and 1997 we analyze the self-reported importance of religion in an individual's life and confidence in the church (see Figure A9). For about 90% of rural and 80% of urban individuals, religion is very or rather important in life. These shares slightly increase from 1991 to 1997 for both groups. With about 80% of rural and about 70% of urban individuals there is also high trust in the churches. There are no disproportionate shifts in these variables between rural and urban individuals that could plausibly explain the stronger decline of fertility in rural areas.

Culture. The general modernization of the Brazilian society is also reflected in the expansion of the coverage of Brazilian territory with Rede Globo television. This led to almost nationwide broadcasting of telenovelas which tend to portray modern, small families. While this expansion mainly took place in the early 1980s, we control

[19]Methodologically based on Dix-Carneiro and Kovak (2017). See Appendix A5 for details.

for Rede Globo-coverage in our regressions to account for earlier findings that soap operas have contributed to a decline in fertility (La Ferrara et al., 2012).

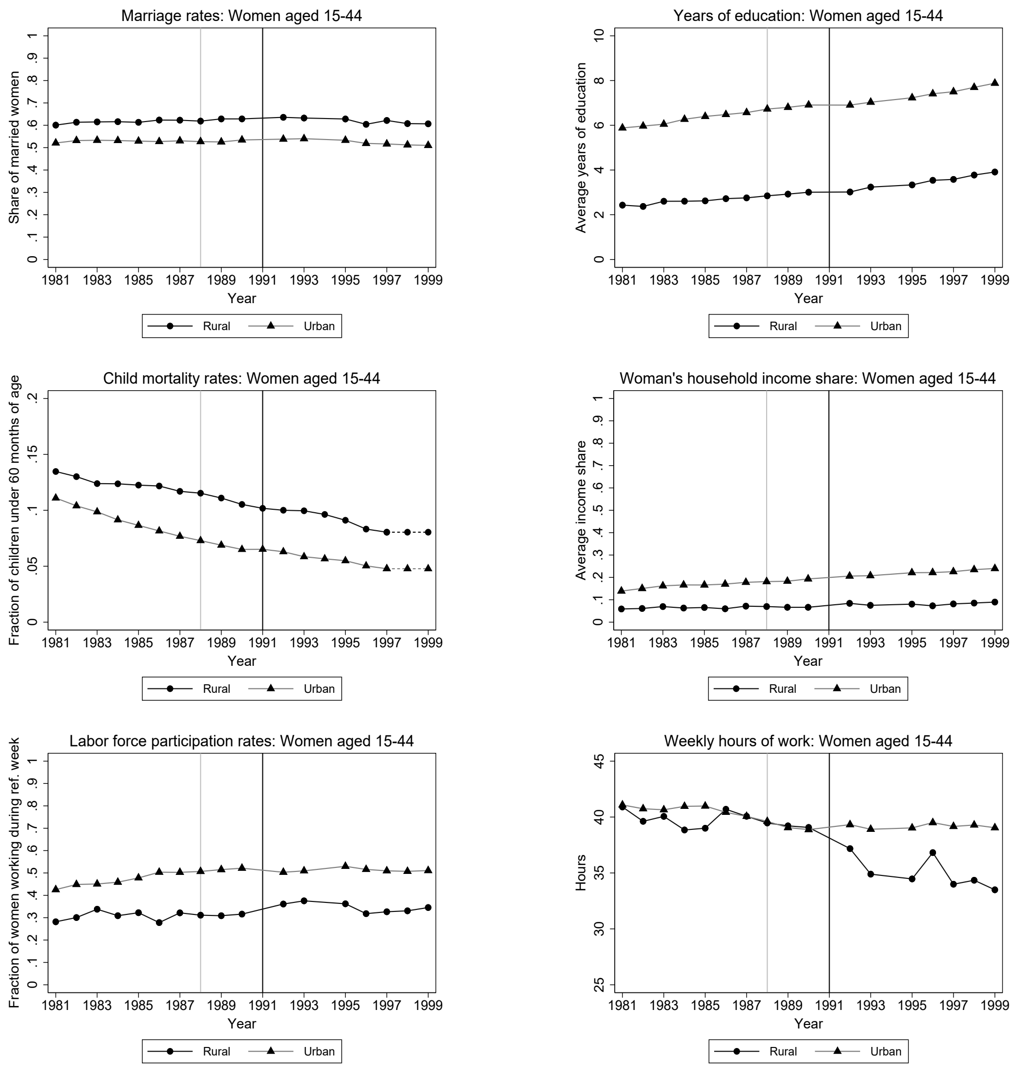

Figure 4: Confounding factor trends: Women aged 15-44

Note: Graphs show average marriage rates, years of education, child mortality rates, income of women as a share of household income, labor force participation rates, and weekly hours of work for rural and urban female workers. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. Figure A2 shows the trends for women aged 15-29 and Figure A3 for women aged 30-44. *Source:* PNAD 1981-90, 1992-93, 1995-99.

6 Fertility Response to the Pension Increase

The plausible behavioral response of women to the pension reform is a reduction in fertility. Indeed, a comparison of DHS data on the number of desired children indicates that urban women lowered their planned fertility by 0.51 children between 1986 and

1996 while rural women reduced it by 0.59 children (Table A11). The differential is especially large among women aged 30 to 44 who have fewer than the number of children they desire (-0.43 vs. -1.02).[20] Yet, how does this change in the desired number of children translate into actual fertility? Older women may immediately stop having children. In essence, we expect a negative short- and long-run response of these women. Younger women may continue having children until they have born their (revised) desired number of children. This implies no short-run, but a negative long-run response. In the following we assess the realized fertility outcomes.

6.1 Long-Run Effects: ES Estimation

The introduction of the Brazilian rural old-age pension led to a significant decline in women's completed fertility rate (total number of births at age 45), as depicted in the rural vs. urban time series (Figure 5, left panel): While rural women had on average about 7.1 children in the 1970s and early 1980s, the comparable number was 4.8 children for urban women. This sizeable gap of 2.3 children between both groups remained constant up to the year of the pension reform (1991) albeit fertility levels declined by one child per woman. In 2010, rural (urban) women now have on average about 3.6 (2.3) children (for sample size reasons we trim the graph in 2010).

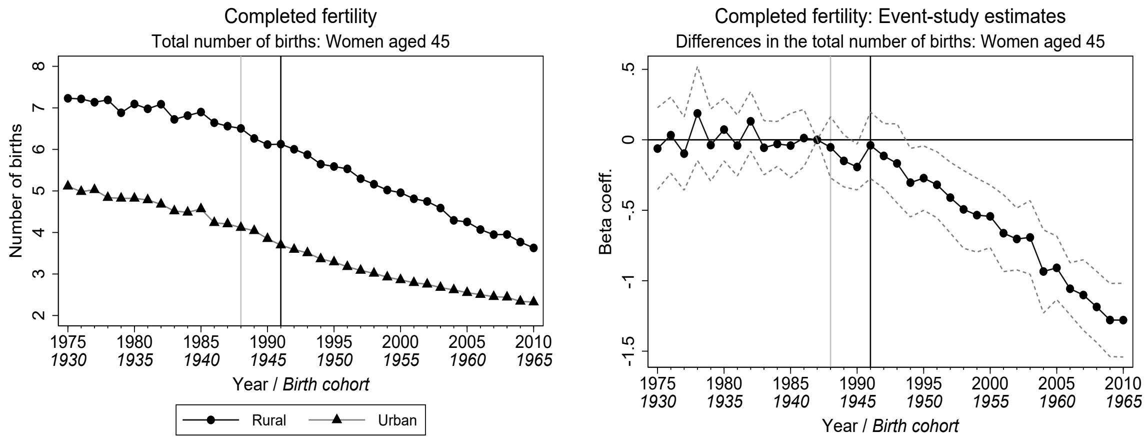

Figure 5: Completed fertility: Total births and Differences in total births at the age of 45

Note: Left graph shows the average total number of births (alive and dead) of women at age 45. Right graph shows ES estimates of the pension reform for the total number of births (alive and dead) of women at the age of 45 before and after the reform. 90% confidence interval based on standard errors clustered at the regional (federal state) level. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. *Source:* PNAD 1984-85, 1992-93, 1995-99, 2001-09, 2011-14.

[20] Accordingly, we find a stronger increase in contraceptive use among rural compared to urban women from 1986 to 1996 (see Table A12). Note, that this is not driven by differences in access to contraception (Caetano and Potter, 2004).

The right panel of Figure 5 depicts estimates of our ES approach, i.e., the completed fertility gap between urban and rural women compared to the base year 1987: The estimates fluctuate closely around zero until the year when the rural pension was introduced. After 1991, completed fertility rates declined at much faster pace in rural than in urban areas (with a modest and plausible lag of one to two years). Until 2010, the rural-urban gap has shrunk by about 1.3 births per women. The first year in which the estimator for completed fertility turns significant is 1994. This implies a quick fertility reduction among 45 year old women after the pension reform: The introduction of rural pensions has an economically significant and lasting effect on fertility in Brazil.

6.2 Short-Run Effects: DID Estimation

Next, we turn to short-run behavioral responses in childbearing by analyzing the probability that a woman in childbearing age (15-44 years old) delivers a baby in a given year. Across all samples the childbearing probabilities of rural women exceed those of urban women and childbearing rates decline over the entire observation period (Figure 6): The top panel of Figure 6 compares fertility rates for all rural and urban women, with the dashed line representing the urban counterfactual trend fixed at the year 1987. In other words, the dashed line depicts the fertility trend of urban women adjusted to rural women's fertility level as of 1987. Before the reform in 1991 fertility rates between rural and urban women develop similarly. After 1991 and in line with the disproportionate change in pension wealth we observe a relatively stronger (i.e., steeper) decline of childbearing probabilities of female rural workers in the full sample. There is an immediate drop in the annualized childbearing probability of rural women by 1 percentage point, which increases to 2 percentage points by the end of the 1990s (equivalent to -17%). This decline is not driven by younger women in the 15-29 year age range (bottom left panel), but by women aged 30-44 (bottom right panel). The overall unconditional childbearing decline among older rural women exceeds 30%. These differences by age are consistent with theoretical predictions: Older women are expected to respond more immediately due to survival, information, saturation, pension wealth and salience effects. Younger women seem to continue having babies until they reach their (downwardly revised) desired fertility level.

This finding is reflected in DID estimates regarding the probability that a woman has given birth in the past 12 months, for three different samples (Table 3): all women aged 15-44 (top), young women aged 15-29 (middle) and older women aged 30-44 (bottom). The six columns represent different specifications: without controls (col. 1), with year and region fixed effects (col. 2), with individual (col. 3), job related (col.

4), household (col. 5), and with group/region controls (col. 6). The highly significant reform effect in the full sample of women in childbearing age is around one percentage point irrespective of the chosen specification. This is equivalent to a decline in the probability of childbearing by 7-10%.

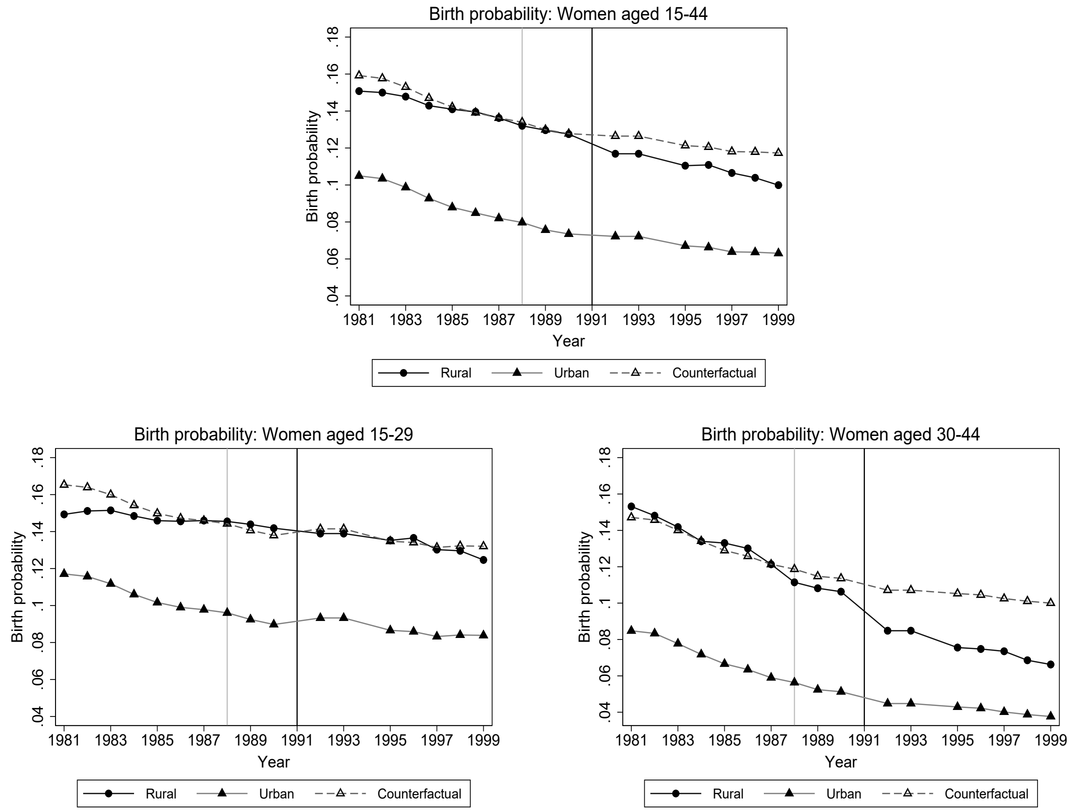

Figure 6: Birth Probabilities: Pooled (Women aged 15-44) and Subgroups (Women aged 15-29 vs. 30-44), Brazil 1981-99

Note: Graphs show 3-year moving averages (2-years at the edges: 1981-82, 1989-90, 1992-93 and 1998-99) of birth probabilities, i.e., average childbearing (0/1) rates within the last 12 months. Counterfactual trend as of 1987: gray dashed line. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

In the young cohort the estimated effects are all very close to zero and none is statistically different from zero. In the older cohort the effects are large: all coefficients are highly significant at around -0.03, equivalent to a decline of 25% and 28% in the propensity of childbearing in the past year. These estimates are very similar to the aggregate, unconditional drop in birth rates depicted in Figure 6; their stability across specifications suggests that the pension effect is orthogonal to the large number of individual, job related, household, regional and group controls. This supports the identifying assumption of the reform being a valid source of exogenous variation.

Table 3: DiD Regression Results: Pooled (Women Aged 15-44) and Subgroups (Women aged 15-29 vs. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
DID women aged 15-44	-0.008**	-0.008**	-0.012***	-0.011***	-0.011***	-0.012***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
in % to baseline	-7.0	-6.6	-9.9	-9.6	-8.8	-9.8
DID women aged 15-29	0.004	0.005	0.001	0.001	0.002	0.000
	(0.004)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
in % to baseline	3.0	3.5	0.4	0.1	0.9	0.1
DID women aged 30-44	-0.030***	-0.029***	-0.029***	-0.027***	-0.025***	-0.025***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
in % to baseline	-28.4	-28.0	-27.5	-26.4	-25.2	-24.6
Year and region FE	No	Yes	Yes	Yes	Yes	Yes
Covariates (see note):						
Individual	No	No	Yes	Yes	Yes	Yes
Job	No	No	No	Yes	Yes	Yes
Household	No	No	No	No	Yes	Yes
Regional/Group	No	No	No	No	No	Yes
N:	15-44 (1,326,819); 15-29 (780,682); 30-44 (546,137)					
R^2 15-44:	0.006	0.009	0.103	0.109	0.114	0.115
R^2 15-29:	0.004	0.006	0.148	0.156	0.161	0.161
R^2 30-44:	0.011	0.016	0.057	0.061	0.067	0.068

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Individual covariates: years of schooling, dummy for married, age of the woman, age of the woman squared and dummies for 1 to 5+ prior children (0 omitted). Job related covariates: dummies for 1-10, 11-20, 21-30, 31-40, 41-50, 51-60 and 60+ hours of professional work in reference week (0 hours, i.e., not working omitted) and share of household income (excl. pensions) earned by the woman. Household covariates: log of monthly household income (excl. pensions), household wealth, dummies for the number of adults in the household, dummy for the presence of a potential caretaker in the household, i.e., a non-working pensioner aged 60 or older, age of the partner and age of the partner squared. Regional/group covariates: dummy for Rede Globo coverage in the past year (based on La Ferrara et al. (2012) data), regional shares of religious affiliations, i.e., protestant, other faith and no religion (catholic omitted, based on Census data), regional race shares, i.e., Black, Mixed-race and Others including Asian, Indigenous and Other (white omitted, based on Census data), macro-regional child mortality (mortality rates of children under 60 months of age by macro-region, i.e., north-east, mid-west, south, and south-east, based on DHS data) and regional industry-specific trade shocks (based on methodology of Dix-Carneiro and Kovak (2017) and annual tariff rates taken from De Paiva Abreu (2004)). Rural and urban groups are defined as in Table 2. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

A number of covariates are worthy of attention (see Table A13): Among the individual level covariates marriage is positively associated with fertility, as expected. Controlling for marriage increases the R-squared substantially. Age and age of the woman's partner are positively correlated with fertility (with negative square terms). Dummies for the number of previously born children as well as dummies for working hours brackets (if working) are negatively related to fertility. Years of schooling are negatively correlated with fertility in most of our regressions, as are the woman's household income share, household income (excluding old-age pension income of co-residing pensioners) and household wealth. The number of adults in the household and the presence of a potential caretaker are positively associated with fertility. Rede Globo (telenovelas) coverage is negative especially for older women, as found in earlier studies. Child mortality rates are positively related to fertility, as expected.

6.3 Pension Wealth: IV Estimation

While the effect of the Brazilian pension reform of 1991 on fertility is policy relevant, economists also care about the generally more comparable effect of pension wealth on the childbearing propensity. This effect can be uncovered as the semi-elasticity of pension wealth using a 2SLS IV regression in which the log pension wealth is instrumented with the $RURAL_g \times POST_t$ interaction in the first stage. The instrument is sufficiently strong across all samples (F-statistic $>$ 3475, T-statistic of the excluded instrument $>$ 58). The estimates confirm the earlier results in a more generalized set-up (Table 4): A 100% increase in pension wealth reduces the probability that a child was born in the past year by 0.6 percentage points. Given that a couple's pension wealth increases by almost 300%, the overall effect is larger than the estimated policy effect in Table 3 ($3 \times -0.006 > -0.012$), although the two estimates do not statistically differ from each other. Below the main results we provide estimates for alternative discount rate specifications. If assuming—in line with the literature—that less wealthy households (Becker and Mulligan, 1997) or rural households (Anderson et al., 2004) have higher discount rates, the estimated semi-elasticities get closer to the policy effect.

The semi-elasticity for the younger and older sub-samples estimates the response for a comparable pension wealth *increase* (in %) between younger and older women. This approach deviates from estimating the policy effect for the two age groups which exhibit different pension wealth *levels* owing to the discounting horizon; it implicitly switches off the pension wealth effect described in the theoretical section. As columns (2) and (3) illustrate IV estimates for both age groups are negative; however, only the estimate for the older groups is significantly different from zero. This result suggests that differences in pension wealth alone cannot explain the differential response of younger and older women.

Table 4: IV Regression Results: Semi-elasticity of birth probabilities to Pension Wealth

Dependent Variable	Newborn child under 1 year old (0/1)		
	(1)	(2)	(3)
	Full Sample	Women aged 15-29	Women aged 30-44
Log of pension wealth	-0.006***	-0.000	-0.012***
	(0.002)	(0.002)	(0.002)
First-stage statistics:			
Dependent Variable	Pension wealth of woman and her partner		
Instrument	Rural work x After reform interaction		
Coefficient	1.951***	1.873***	2.084***
	(0.030)	(0.032)	(0.028)
F-test	4119.95	3478.38	5673.48
T-statistic of excl. instrument	64.19	58.98	75.32
Partial R^2 of excl. instrument	0.101	0.116	0.091
Estimates for alternative discount rates:			
Log of pension wealth	-0.005***	0.000	-0.012***
30% pre vs. 12% post	(0.002)	(0.001)	(0.002)
Log of pension wealth	-0.004***	0.000	-0.008***
30% rural vs. 12% urban	(0.001)	(0.001)	(0.001)
Year and region FE	Yes	Yes	Yes
Covariates (see note)	Yes	Yes	Yes
N:	1,326,819	780,682	546,137
R^2:	0.115	0.161	0.065

Note: IV estimates of the pension reform. 'First-stage' regression of log pension wealth on the instrument ($RURAL_g \times POST_t$) and all other covariates. Power of the first-stage regression reported with F-statistics. Dependent variable in the 'second-stage': Dummy, whether a child was born in the last 12 months. Full set of covariates, as in Table 3, column 6. Pension wealth computed as in Figure 3 with a constant discount rate of 12% for the main specification (c.f. Azzoni and Isai (1994)). Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

6.4 Number of Previously Born Children: DID Estimation

If women do not change their fertility behavior as long as their family is small, but stop having children once they reach their downwardly adjusted desired number of children, we observe stronger fertility effects at higher birth parities. This is expected to be more prevalent among women in the middle or late phase of their fertility cycle who have—on average—already born more children than younger women. Indeed, we find that most of the reduction in childbearing appears at higher birth parities (Figure 7 and Table 5).

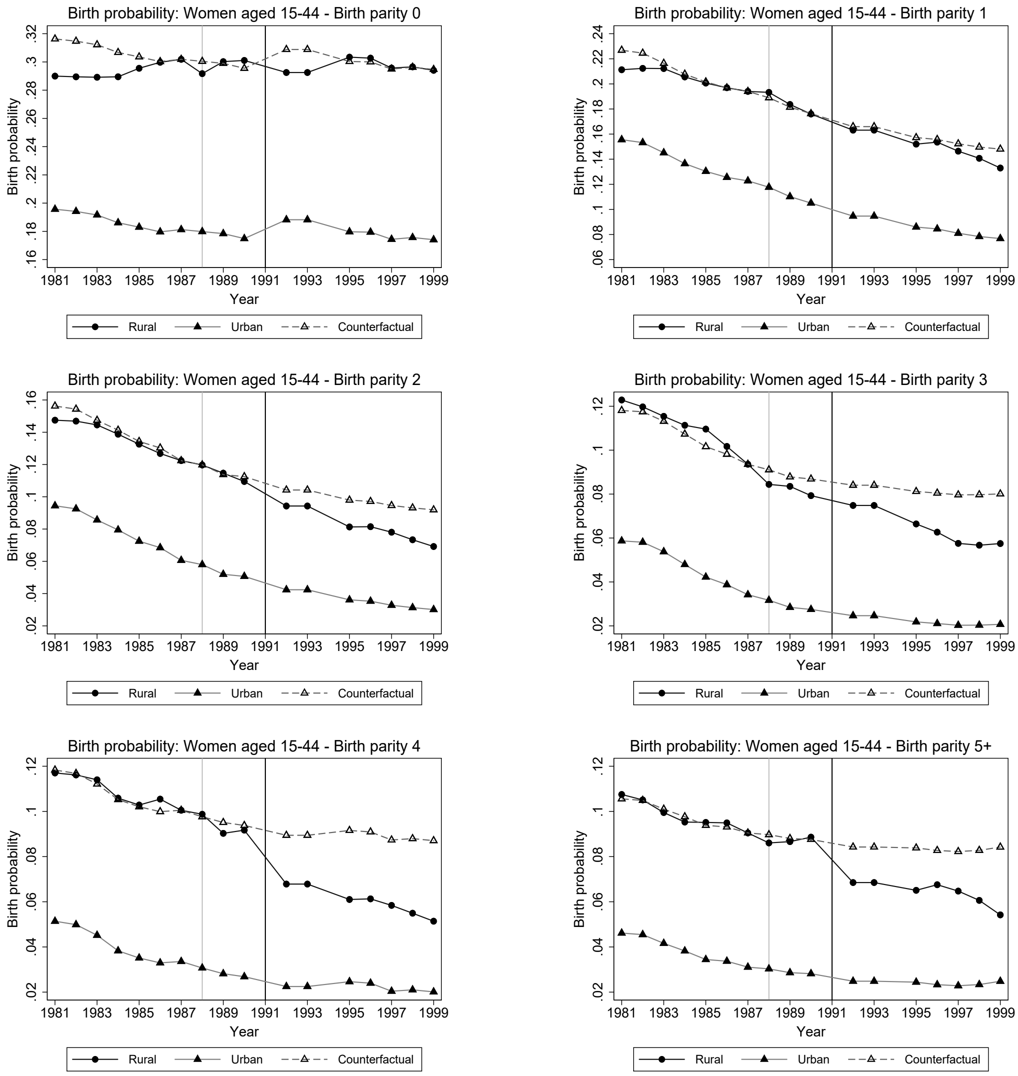

Figure 7: Birth Probabilities by Birth Parity

Note: Graphs show 3-year moving averages (2-years at the edges: 1981-82, 1989-90, 1992-93 and 1998-99) of birth probabilities, i.e., average childbearing (0/1) rates within the last 12 months, by birth parity (the number of previous children) up to 5+. Counterfactual trend as of 1987: gray dashed line. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

While the probability of having a first or second child has not significantly declined after the pension reform (top panel of Figure 7 and columns 1 and 2 of Table 5), the fertility effects turns negative at the third child and increasingly so at higher birth parities (middle and bottom panel of Figure 7 and columns 3 to 6 of Table 5). Our estimates (Table 5) show that the childbearing probability declines by 12% (2nd birth parity), 21% (3rd birth parity), 35% (4th parity), and 28% (5th+ parity), respectively. These findings support the stopping hypothesis and can at least partly explain the stronger adjustment effect among older women (who tend to have already more children).

Table 5: DiD Regression Results: By Birth Parity (Women aged 15-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
	Parity 0	Parity 1	Parity 2	Parity 3	Parity 4	Parity 5+
DID women aged 15-44	0.025***	0.002	-0.012**	-0.018***	-0.032***	-0.025***
	(0.007)	(0.006)	(0.005)	(0.005)	(0.006)	(0.004)
	7.8	1.0	-12.4	-21.3	-34.5	-27.7
Year and region FE	Yes	Yes	Yes	Yes	Yes	Yes
Covariates (see note)	Yes	Yes	Yes	Yes	Yes	Yes
N:	208,539	273,684	307,041	222,715	125,239	189,601
R^2:	0.135	0.096	0.071	0.070	0.085	0.109

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Full set of covariates, as in Table 3, column 6 (dummies for 0 to 5+ prior children omitted). Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

While the results suggest that the pension reform made it more likely for women to have fewer children, does it also make women stop having children altogether? If children were predominantly reared for old-age provision, the need to have any children would be dramatically reduced after the reform. However, we find no evidence for an increase in childlessness after the pension expansion: Figure A10 shows that the fraction of women aged 45 who have never given birth to a child remained quite stable at 11% for urban women and 8% for rural women, although the modest sample sizes lead to bumpy time series. Nevertheless, there is no clear indication of a higher prevalence of childlessness in the ES set-up.

6.5 Gender of Previously Born Children: DID Estimation

We find significant differences in the probability of childbearing between women who have at least one son, at least two sons, only daughters or no children at all. For these four samples, Figure 8 shows the probabilities of childbearing in the past 12 months (Figure A11 for women aged 15-29 and women aged 30-44 separately). Figure 8 shows consistent declines in childbearing only for women with at least one or two previously born sons. Once we turn to the econometric evaluation in a DID framework (Table 6), we find strong and significant effects only for the overall sample of women with at least one or two boys, a result that is entirely driven by older women aged 30-44. This group reduces the probability of childbearing in the past 12 months by 30%. Women with daughters or without children do not reduce fertility in any of the three age samples.

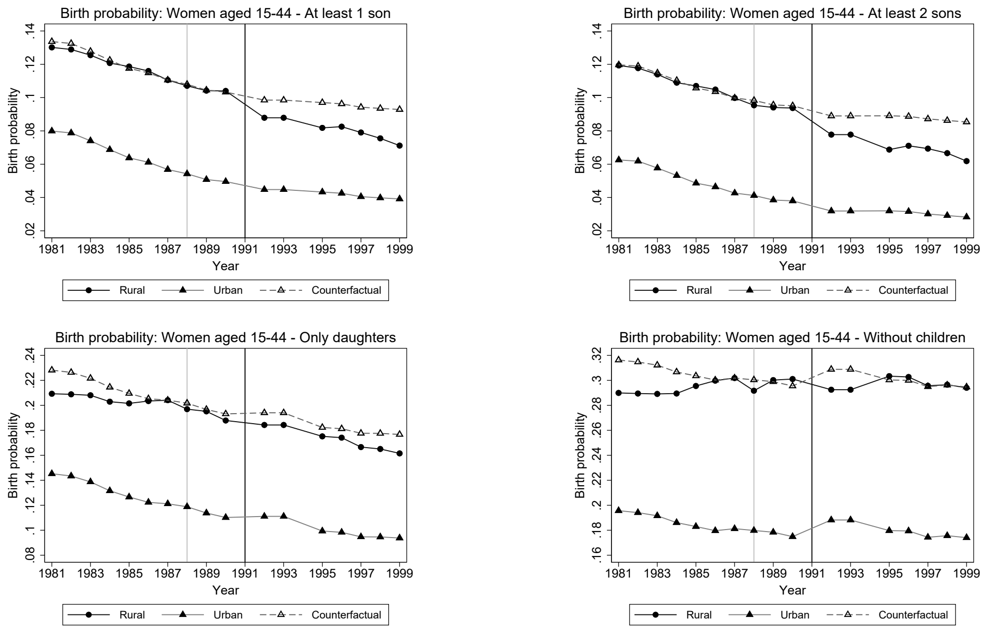

Figure 8: Birth Probabilities by Older Children's Gender: Women aged 15-44

Note: Graphs show 3-year moving averages (2-years at the edges: 1981-82, 1989-90, 1992-93 and 1998-99) of birth probabilities, i.e., average childbearing (0/1) rates within the last 12 months, conditional on whether women had already given birth to at least one son, at least two sons, only daughters or no children at all. Counterfactual trend as of 1987: gray dashed line. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table 6: DiD Regression Results: by Older Children's Gender (Women aged 15-44)

Dependent Variable	Newborn child under 1 year old (0/1)				
	(1)	(2)	(3)	(4)	(5)
	Full sample	1+ sons	2+ sons	Only daughters	No children
DID women aged 15-44	-0.012***	-0.016***	-0.019***	0.003	0.025***
	(0.003)	(0.004)	(0.004)	(0.003)	(0.007)
in % to baseline	-9.9	-16.3	-21.2	1.5	7.8
DID women aged 15-29	-0.000	-0.004	-0.004	0.011**	0.041***
	(0.003)	(0.003)	(0.003)	(0.004)	(0.010)
in % to baseline	-0.2	-3.9	-6.6	4.7	9.8
DID women aged 30-44	-0.025***	-0.030***	-0.036***	-0.002	0.008
	(0.004)	(0.005)	(0.006)	(0.004)	(0.006)
in % to baseline	-24.6	-28.1	-29.9	-2.8	7.8
Year and region FE	Yes	Yes	Yes	Yes	Yes
Covariates (see note)	Yes	Yes	Yes	Yes	Yes
N 15-44:	1,326,819	820,066	396,070	506,753	208,539
N 15-29:	780,682	442,466	201,306	338,216	151,566
N 30-44:	546,137	377,600	194,764	168,537	56,973
R^2 15-44:	0.115	0.086	0.077	0.137	0.135
R^2 15-29:	0.161	0.140	0.141	0.158	0.127
R^2 30-44:	0.068	0.069	0.078	0.080	0.102

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Full set of covariates, as in Table 3, column 6. Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

6.6 Financial Gap in the PAYG Pension System

Given its favorable demographic composition in the decades prior to 1988, Brazil opted (like most developing countries) for a Pay-As-You-Go (PAYG) pension scheme. In PAYG systems large cohorts of young workers provide for the smaller cohorts of concurrent retirees. The Brazilian pension reform, however, does not only increase the benefit (i.e., expenditure) side of the social security system by expanding generosity; it also reduces the contribution (i.e., income) base by lowering the number of future contributors through today's reduced fertility. To illustrate the total financial gap in the Brazilian PAYG system we conduct a simple static back-of-the-envelope calculation, after fertility levels have adjusted to the new equilibrium. We decompose the total reform effect into a 'generosity effect' and a 'contribution base effect'.

Assuming that the Brazilian PAYG system is actuarially fair the benefit side with average pension payment B being paid out to N recipients for 13 monthly installments will equal the contribution side with L contributors paying twelve monthly contributions C. The pension reform now makes pensions more generous and, hence, costly. It introduces a rise in the benefit level B and a growth in the number of pension recipients N. At the same time the reform reduces the fiscal base of contributions by reducing the future number of contributors L through the fertility effect (-1.3 children per rural woman).

The total long-run financial gap resulting from the reform in the Brazilian PAYG system is:

$$-(\Delta B \times 13 \times \Delta N) \times (1 + AC) + (C \times 12) \times \Delta L \tag{5}$$

with the administrative cost AC of the pension system being assumed to equal either 7% (Mitchell, 1998), 10.75% (Zockun, 1983) or 14.5%.[21]

We now estimate the relative contributions of the 'generosity effect' vs. the 'contribution base effect' using PNAD data, based on the assumption that environmental factors such as earnings or immigration and policy variables such as tax rates or social security contributions remain constant. Expressed as a fraction of GDP, the total funding gap with $AC = 10.75\%$ is around 0.7% in 1995, 1.0% in 2001, 1.2% in 2005 and 1.8% in 2011. These results are consistent with the World Bank (2003), which reports that the Brazilian Pension System for the private sector reached a deficit of 0.9% of GDP in 1999. Importantly, the 'generosity effect' accounts for only a modest part of the total financial gap (about 20% in 2011), while the 'contribution base effect' of the fertility decline weighs heavily from an actuarial perspective.

[21] The World Bank (2005) reports administrative costs of the Brazilian pension system of 10-15%.

Table 7: Funding Gap of the Brazilian PAYG Pension System as a Result of the Reform (in % of GDP)

Year	1995	2001	2005	2011
AC =7%				
Generosity effect	0.07	0.12	0.18	0.35
Contribution base effect	0.61	0.86	0.96	1.42
Total financial gap	0.68	0.97	1.14	1.77
AC =10.75%				
Generosity effect	0.07	0.12	0.19	0.36
Contribution base effect	0.61	0.87	0.98	1.46
Total financial gap	0.68	0.99	1.16	1.81
AC =14.5%				
Generosity effect	0.08	0.13	0.19	0.37
Contribution base effect	0.61	0.88	0.99	1.49
Total financial gap	0.69	1.00	1.19	1.86

Note: The table shows calculations of the long term financial gap of the Brazilian PAYG pension system as a result of the reform (in % of GDP as of 2012), decomposed into a 'generosity effect' and a 'contribution base effect'. The long term loss in rural payroll contributors is computed as the difference between the observed number of children and the hypothetical number of children if completed fertility was 1.3 children higher per rural women. Average pension payments and pension recipients are taken from the PNAD. Social security contributions: Monthly average gross income $\times$ 9.3% (employees average contribution rate) + monthly average gross income $\times$ 36.3% (employers contribution rate). Average gross income: Average net income + Average net income $\times$ 17.5% (the average marginal tax rate). *Source:* PNAD 1981-90, 1992-93, 1995-99, 2001-09, 2011-14.

This computation illustrates the economic significance of the estimated fertility effect. It highlights that fertility consequences of public pension policies are economically costly and can affect the sustainability and political acceptance of a PAYG pension system in the long-run.

7 Discussion and Robustness

In this section we discuss the robustness of our results. First, we analyze more granular age groups. Second, we investigate alternative specifications regarding control variables and time trends. Third, we test the sensitivity of our results with respect to regional, occupational and age-specific sample choices. Fourth, we address alternative error correlation structures.

7.1 Analysis of Granular Age Groups

Our main findings suggest that younger and older women respond differently to the pension reform. To investigate the relationship between age and effect size further we assess DID estimators in six five-year age cohorts (15-19; 20-24; 25-29; 30-34;

35-39; 40-44). Separate regressions suggest (see Table A14) no significant negative fertility response among the youngest two groups. The negative effect for the age group 25-29 is marginally (in)significant, depending on the specification. The three oldest age groups exhibit highly significant negative fertility responses to the pension reform. The relative reduction in fertility (conditional on the cohort-specific baseline fertility rate) is increasing with age. While women aged 30-34 reduce their probability of childbearing by 23%, the corresponding values are 26% and 35% for women aged 35-39 and 40-44.

In a further refinement we estimate (2) with the full set of control variables for fifteen 2-year age cohorts. The results are graphically depicted in Figure A12. As before the pension reform has no systematic significant impact on the fertility of rural women up to age 30, with the 90% confidence interval mostly containing the zero. The pension reform, however, has a consistent significant negative effect on fertility of women above age 30. The point estimates between age 30 and 44 are all remarkably similar and range between -0.2 and -0.4: Only women in the middle or late phase of their fertility cycle exhibit a significantly negative fertility response to the pension reform.

7.2 Alternative Specifications

Alternative model specifications tend to provide very similar results: In Table A9, col. 1 we control for alternative educational variables. While average years of schooling (our control variable in the main analysis) progressed on similar trajectories between urban and rural women, we observe growth in primary school degrees by 15 percentage points in rural and 10 percentage points in urban areas between 1981 and 1991 (see Figure A14); the main contributor of this rise stems from the introduction of compulsory education for adults after 1988. This was accompanied by rising shares of secondary and tertiary attainments among urban women. Hence, we amend our previous regression (2) by adding dummy variables for the highest degree obtained (primary, secondary or tertiary) as well as a dummy for literacy. In the following columns we control for the share of retirement income in total household income (col. 2), for family income variables such as the log of monthly family income without pensions and the share of family income earned by the woman of childbearing age (col. 3), for two dummies indicating whether the woman worked unpaid or in the public sector (col. 4), for additional square terms of years of schooling, log of household income without pensions, share of family income earned by the woman of childbearing age without pensions and household wealth (col. 5), for age and age of the woman's partner (col. 6), and for dummies of age and age of the woman's partner (col. 7). We also assess the sensitivity of our results with respect to different time trend specifications. In Table A15 we re-estimate model (2) with linear time trends or linear region-specific time trends. None of the alternative specifications alters our main results.

7.3 Sample Selection

We conduct a series of robustness checks regarding sample choices: First, we re-run our analysis in all three types of residential areas separately: So-called not self-representative municipalities (Table A16, upper panel), which are included in the PNAD sample only with a probability proportional to the inhabitant population; self-representative areas (Table A16, middle panel), which are completely sampled; and Metropolitan areas (Table A16, lower panel). Despite substantially smaller sample sizes, our results are preserved in all three subgroups: Birth probabilities are negatively affected in the full sample (statistical significance is not always maintained); the effect is not different from zero among younger women and significantly negative among older women.

Second, we re-define treatment vs. control groups: (1) using only household residence and (2) using only personal occupational information and household residence (Table A8). Our results are robust across different treatment specifications.

Third, we include couples in which wife and husband work in different occupational areas, i.e., in a rural and an urban occupation (Table A7). This increases the sample size by 7.7%, but does not change any of the results no matter whether we control with a dummy indicator for these added women (top panel) or not (bottom panel).

Fourth, we exclude 1422 women (0.1% of the sample) who gave birth at ages younger than 15 years or older than 44 years and who were previously recoded as 15 or 44 years old as is standard in much of the demographic literature (Table A6). Again, the omission of these women has no effect on the results.

Fifth, we can also rule out that women postpone childbearing beyond the cut-off age of 45 years (for instance, if the oldest mothers were 38 years old in 1985 but 47 years old in 1995). Figure A13 shows age density-plots for childbearing women in 1985 and 1995. Over time, female rural and urban women tend not to give birth later, but earlier as described in the literature (Rios-Neto et al., 2018).

7.4 Standard Errors

The error terms in (2) are likely to be correlated between individuals. To account for this intra-group correlation in our main analysis we cluster standard errors at the regional level (States of the Brazilian Federation). As a robustness check we cluster standard errors at the level of regions and settlement types: Increasing the number of clusters from 20 to 47 does not alter the results (Table A17).

As we exploit data covering a relatively long observation period, fertility rates may suffer from autocorrelation. As illustrated by Bertrand et al. (2003), this can inflate the incidence of Type-I error and lead to significant estimates even in the absence of a true effect. To account for potential autocorrelation in fertility rates we reduce the annual observations into one pre- and one post-treatment mean observation per group and region. We then repeat our analysis with this 'collapsed' data set containing only 80 observations. Despite the drastically reduced estimation power, we still find highly significant negative policy effects for older women and a weaker average effect for all women (see Table A18, col. 6). The final specification (col. 6) remains marginally insignificant at conventional levels. These findings based on only two time periods suggests that autocorrelation is not a major threat to our analysis.

8 Conclusions

This study estimates the causal effect of expanding public old-age pensions on individual fertility behavior. Exploiting the Brazilian pension reform of 1991 we find that the associated increase in pension wealth has a significant negative effect on fertility: The reform reduced the probability of childbirth by 10% in the short-run and completed fertility by about 1.3 within the first 20 years after the reform. The decrease of the childbearing probability is fully concentrated among women in the middle and late phases of their fertility cycle (aged 30-44), among women with more children and among women who already have a son in the family; in fact, the effect is insignificantly different from zero for women below age 30. A number of robustness checks and falsification exercises support the causal interpretation of the policy effects of the Brazilian pension reform.

Our results lend micro evidence to the long established view that the introduction of pension systems in emerging and developing countries is associated with declining fertility rates at the macro level. The findings bolster the old-age security hypothesis and confirm the credibility of negative correlations found in previous empirical studies using aggregate data. Of course, a note of caution is warranted with respect to the external validity: Our quasi-experiment in Brazil relies on an unusual equalization of pension benefits between urban and rural areas. Still, many pension schemes in emerging and developing countries disadvantage (or fully exclude) rural populations similar to the system that underwent reform in Brazil in the early 1990s.

Our results are politically relevant in the light of population aging in low- and middle-income countries (LMICs). According to the United Nations (2017), the most rapid demographic aging is projected to take place in LMICs: by 2050, 80% of the world's elderly (age 60+) will be living in those countries. Since most LMICs are insufficiently prepared for the upcoming demographic change, they may experience dramatic economic and social repercussions. While longer lives represent one of the greatest benefits from economic development, governments and societies around the globe have to consider how to provide for the growing number of old citizens. For most countries, public old-age pension systems are a major pillar to grant citizens economic security during retirement (Barr and Diamond, 2009; Holzmann and Hinz, 2009; Willmore, 2007). However, in LMICs these programs are often not generous, exclusive, and complicated with respect to legal eligibility criteria. In particular, only few of these programs cover the rural population and low-income or informal-sector workers. Improving coverage and benefit levels of pension systems in LMICs is one of the key challenges of the next decades given widespread old-age poverty (World Bank, 2017). From our perspective it is essential that policymakers consider the fertility effects of reforming public pensions, especially when declining fertility can erode the basis of PAYG systems. Avoiding unintended demographic consequences of reforms will make pension systems more sustainable in the long-run.

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Appendix

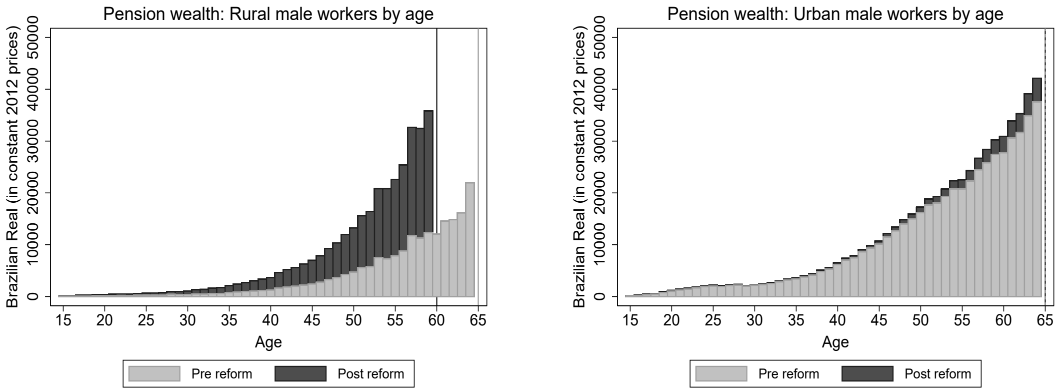

Figure A1: Gross present pension wealth of Rural and Urban Male Workers.

Note: Pre- and post-reform pension wealth is computed as the present value of expected old age benefits before/after the reform adjusted for real interest rates and average survival probabilities (computed using IBGE mortality tables; first time available in 1998): $Pension\ Wealth = \sum_{t=0}^{T-a} s_{a,t} \times \frac{1}{(1+i)^t} \times pension_t$, with $s_{a,t}$ denoting the probability of a person of age a in a given year surviving until year t; $T-a$, indicates the remaining maximum lifespan differentiated by sex and birth cohort; i is a constant discount rate (12%); and $pension_t$ denotes the old age pension benefits in t. A non-retired person receives the pension starting in a future period $t > 0$, defined by the person's age and the regular retirement age. Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

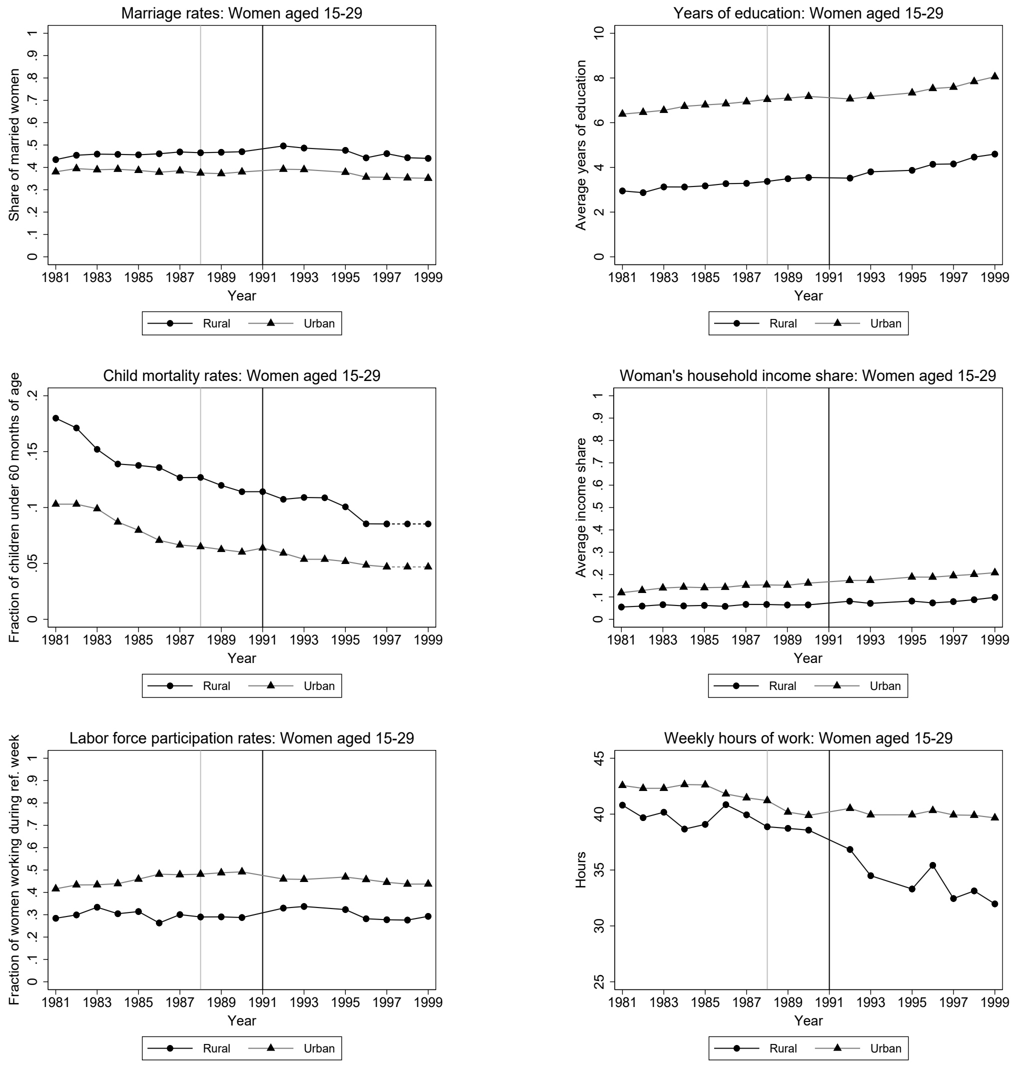

Figure A2: Confounding factor trends: Women aged 15-29

Note: Graphs show average marriage rates, years of education, child mortality rates, income of women as a share of household income, labor force participation rates, and weekly hours of work for rural and urban female workers. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

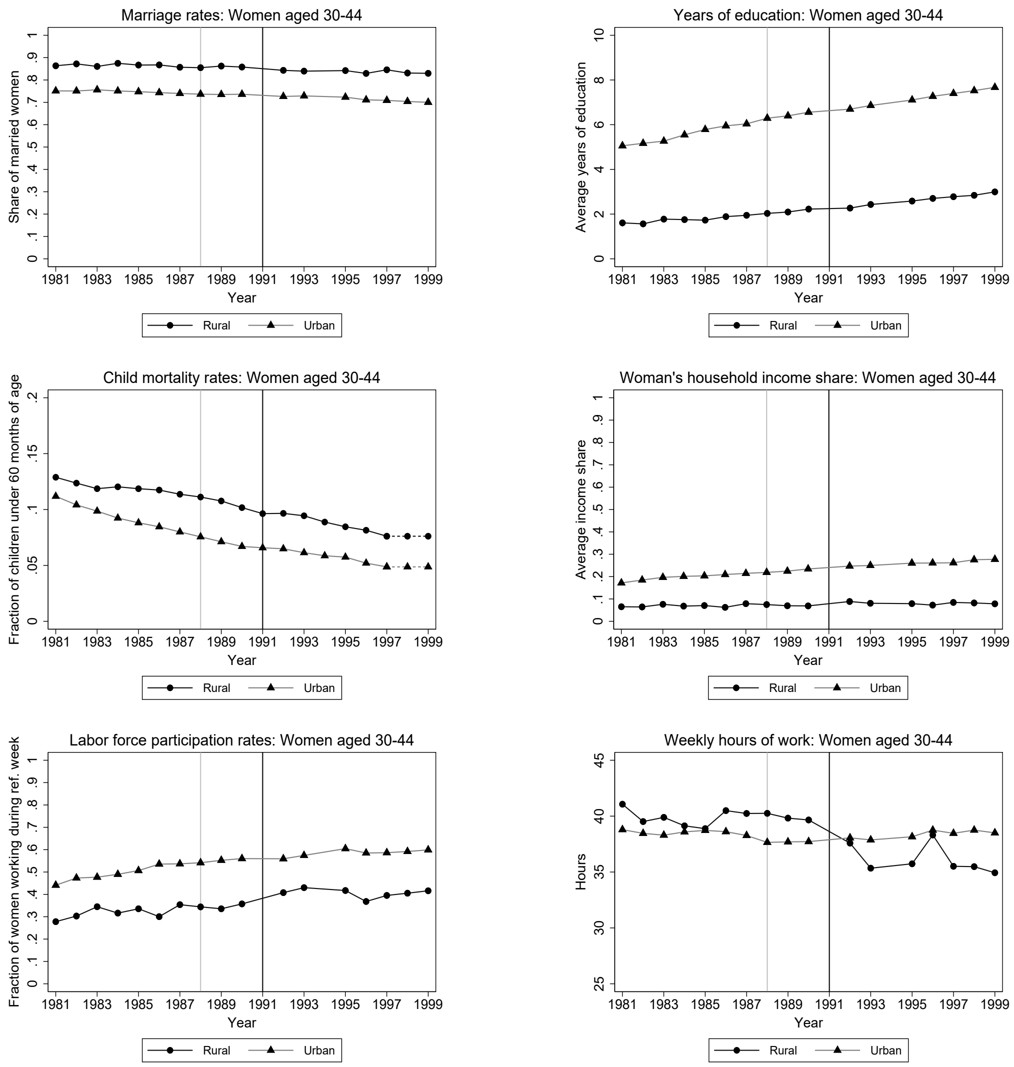

Figure A3: Confounding factor trends: Women aged 30-44

Note: Graphs show average marriage rates, years of education, child mortality rates, income of women as a share of household income, labor force participation rates, and weekly hours of work for rural and urban female workers. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

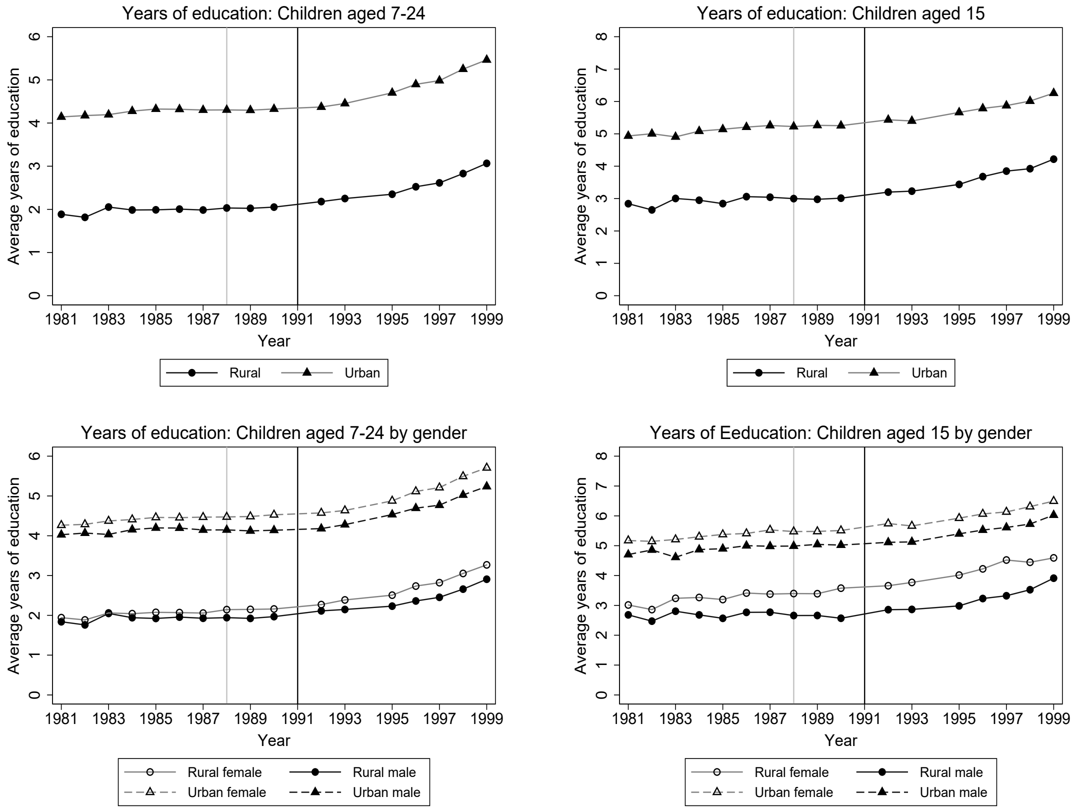

Figure A4: Educational Outcomes of Children - Years of Schooling, Brazil 1981-99

Note: Graphs show average years of education of children living in households with female rural and urban workers aged 15-44. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined by the parents occupation, defined as in Table 2. *Source:* PNAD 1981-90, 1992-93, 1995-99.

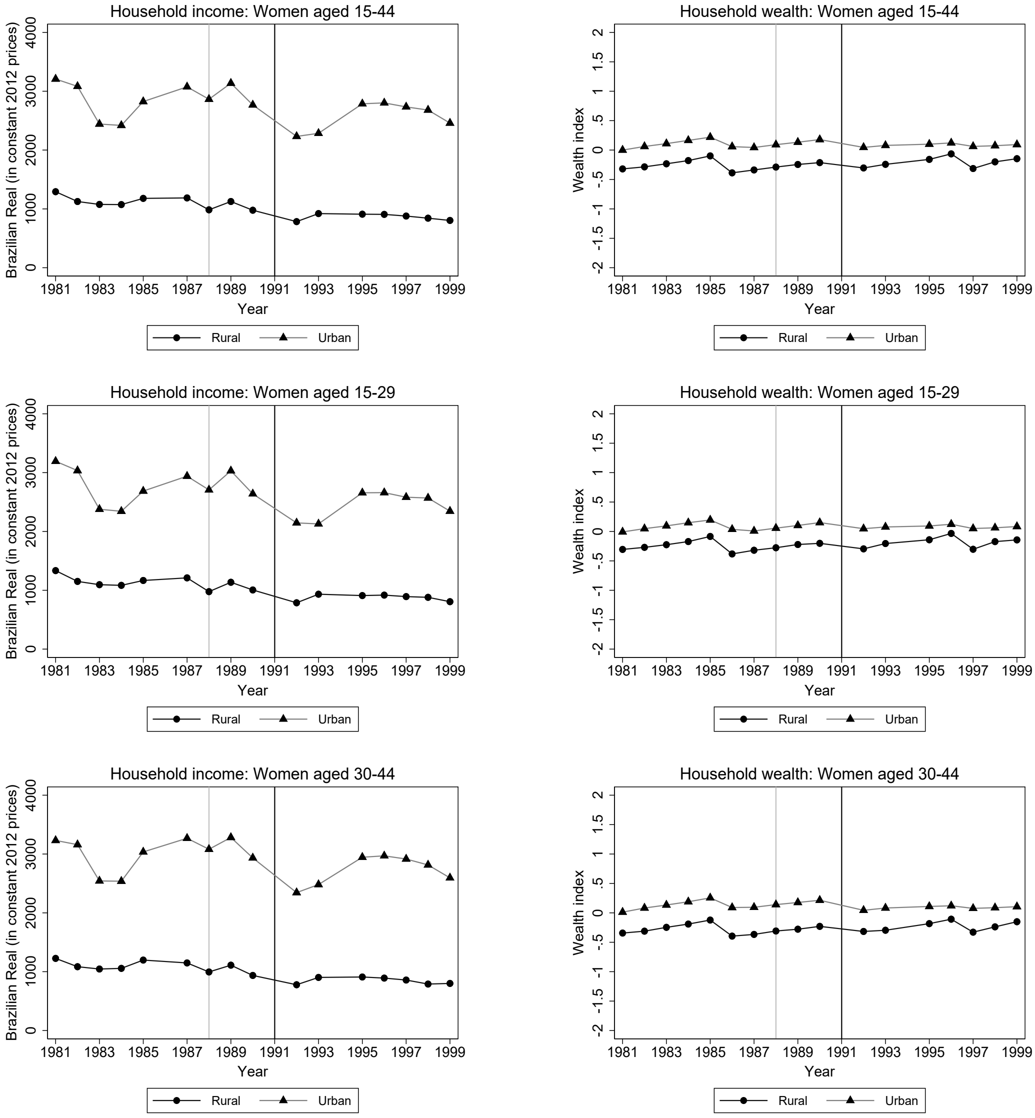

Figure A5: Household Income (without Pensions) and household Wealth index, Brazil 1981-99

Note: Left graphs show average household income (excl. retirement pensions) for rural and urban worker households. Right graphs show average wealth index levels of rural and urban worker households. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

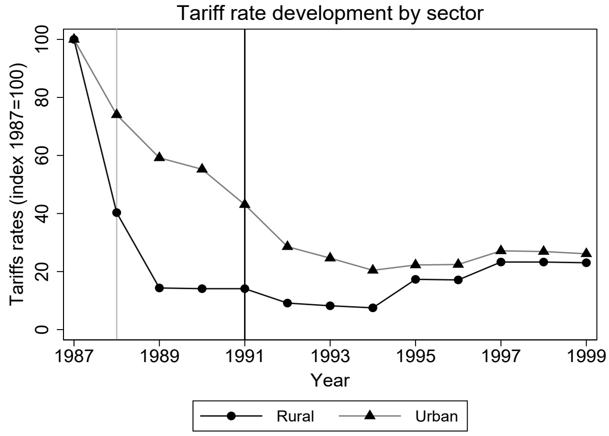

Figure A6: Development of the indexed Tariff rates (1987=100): by urban and rural industrial sectors, Brazil 1987-1999

Note: Graph shows the indexed annual indexed tariff rates applicable for urban and rural industries (1987=100). Tariff rates for the agricultural industry are assigned to rural, tariff rates of all other manufacturing industries are assigned to urban. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). *Source:* De Paiva Abreu (2004).

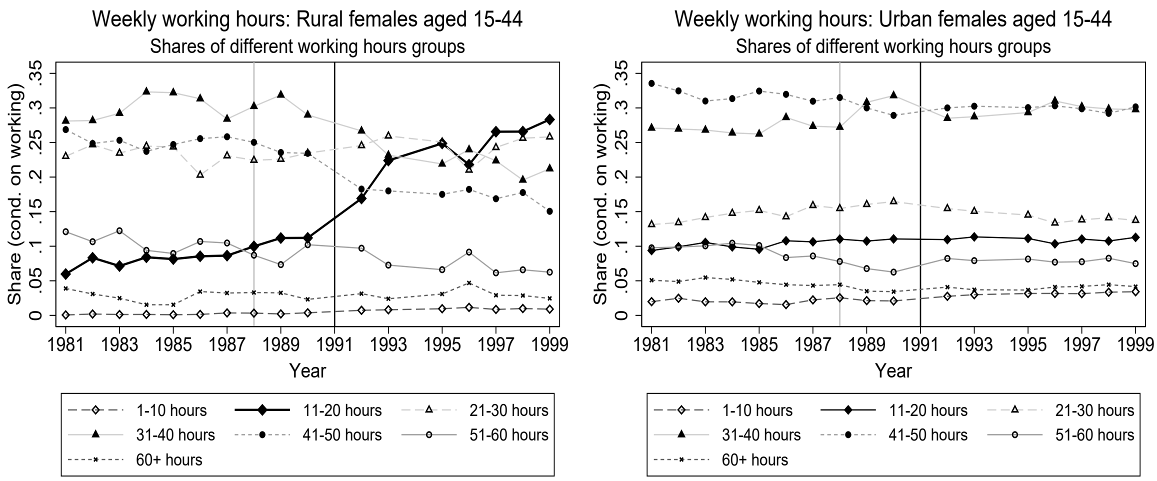

Figure A7: Full- and Part-Time Employment Shares, Brazil 1981-99

Note: Graphs show the share of women in 10-hour brackets of working hours for female rural and urban workers. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. *Source:* PNAD 1981-90, 1992-93, 1995-99.

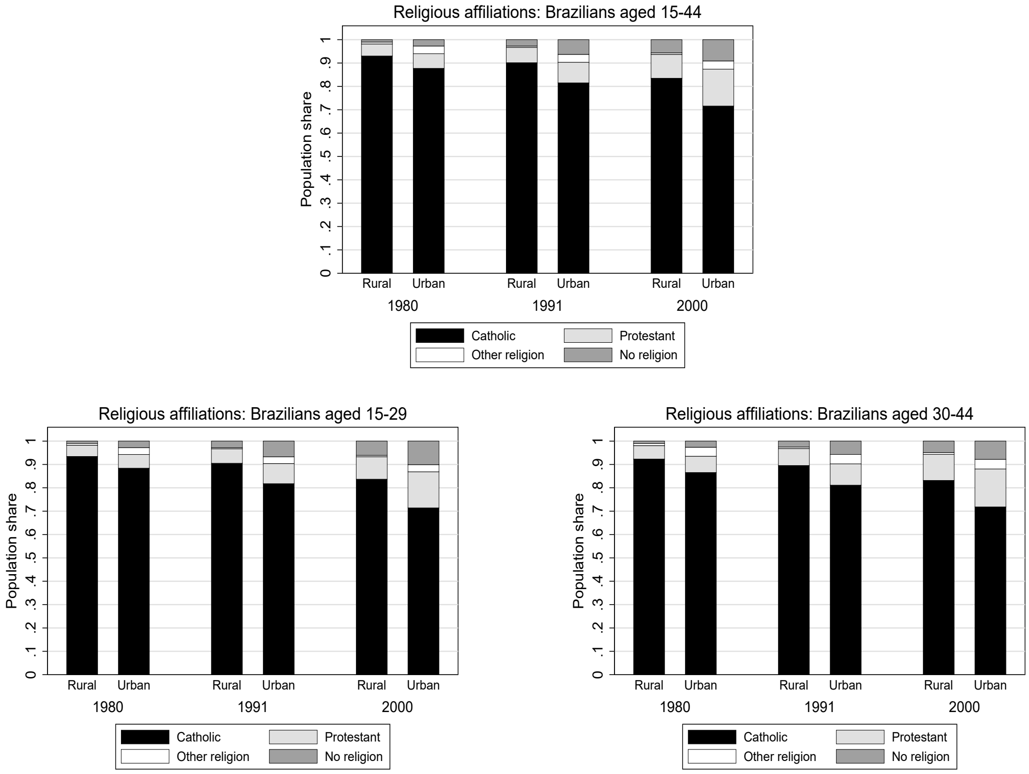

Figure A8: Religious Affiliations, Brazil 1980, 1991, 2000

Note: Sample consists of the Brazilian population aged 15-44 participating in the Census. Rural and urban groups defined by household location. *Source:* Brazilian Census 1980, 1991, 2000.

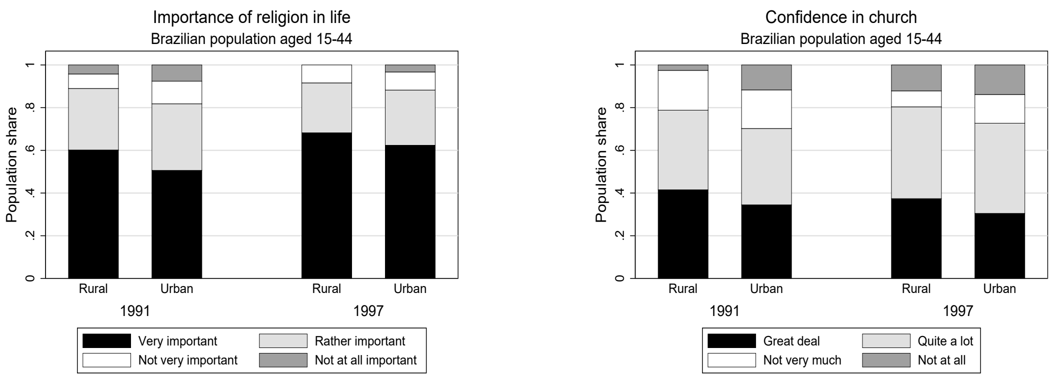

Figure A9: Importance of religion in life and confidence in church, Brazil 1991, 1997

Note: Sample consists of the Brazilian population aged 15-44 participating in the World Values Survey. Rural and urban groups defined by household location (for the WVS we approximate this by the size of the town: smaller or larger than 10.000 inhabitants). *Source:* WVS 1991, 1997.

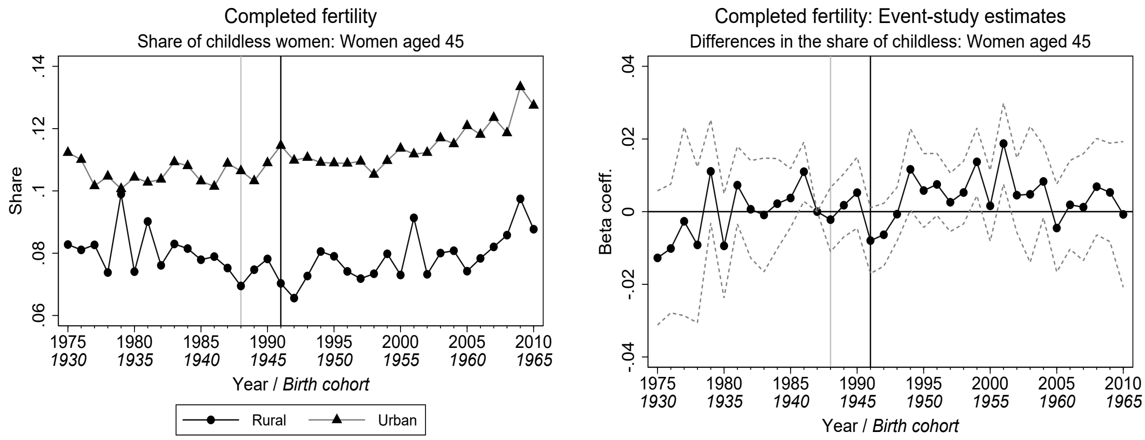

Figure A10: Childlessness: Differences in the share of childless women at the age of 45

Note: Left graph shows the share of childless women at age 45. Right graph shows ES estimates of the pension reform for the share of childless women at age 45 before and after the reform. Dependent variable: Dummy equal to one if a woman never gave birth to a child (alive or dead). 90% confidence interval based on standard errors clustered at the regional (federal state) level. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1984-85, 1992-93, 1995-99, 2001-09, 2011-14.

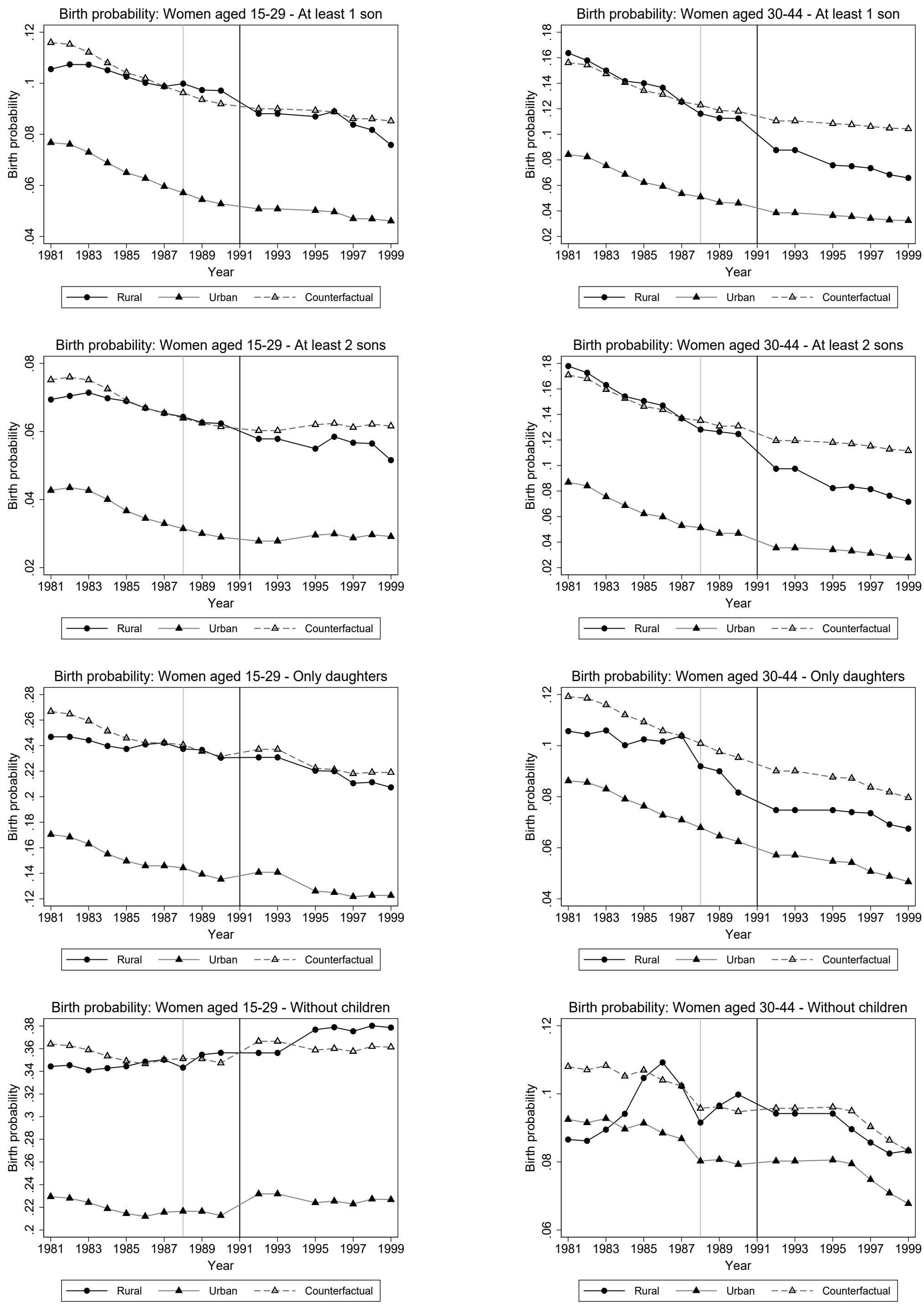

Figure A11: Birth Probabilities by Older Children's Gender: Women aged 15-29 vs. Women aged 30-44

Note: Graphs show 3-year moving averages (2-years at the edges: 1981-82, 1989-90, 1992-93 and 1998-99) of birth probabilities, i.e., average childbearing (0/1) rates within the last 12 months, conditional on whether women already gave birth to at least one boy, at least two boys, had only girls born before or had no children before. Left graphs for women aged 15-29, right graphs for women aged 30-44. Counterfactual trend as of 1987: gray dashed line. Vertical lines: new Constitution approved in 1988 (light-gray); Pension reform implemented in 1991 (black). Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

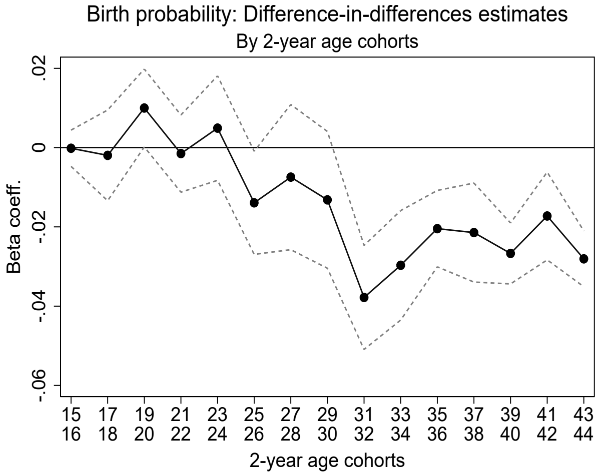

Figure A12: DiD Regression Results: Women by Age

Note: DID estimates of the pension reform for 2-year age groups of women aged 15-44. Dependent variable: Dummy, whether a child was born in the last 12 months. Full set of covariates, as in Table 3, column 6. Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. 90% confidence interval based on standard errors clustered at the regional (federal state) level. *Source:* PNAD 1981-90, 1992-93, 1995-99.

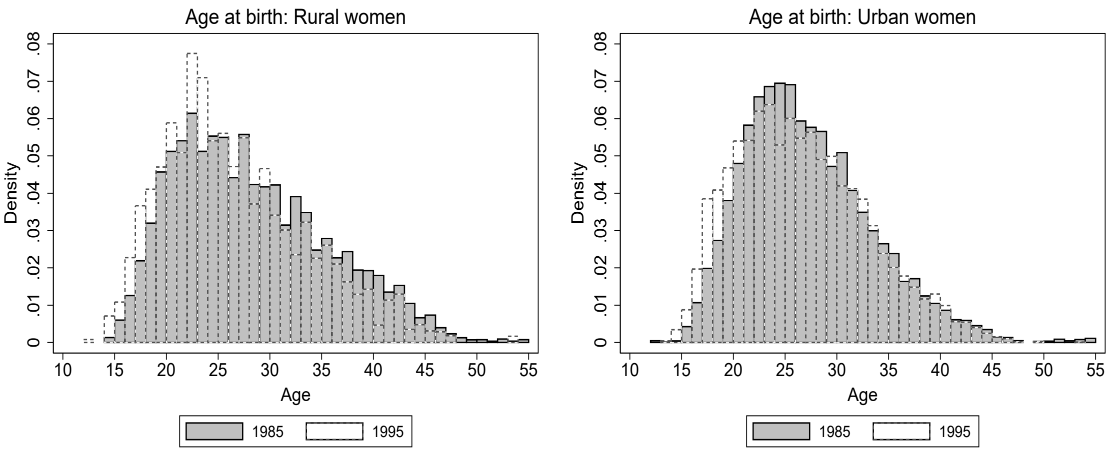

Figure A13: Age of mothers at birth, density plots

Note: Graphs show age of mothers at birth, rural and urban, for 1985 and 1995. Sample consists of Brazilian females aged 10-55. Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

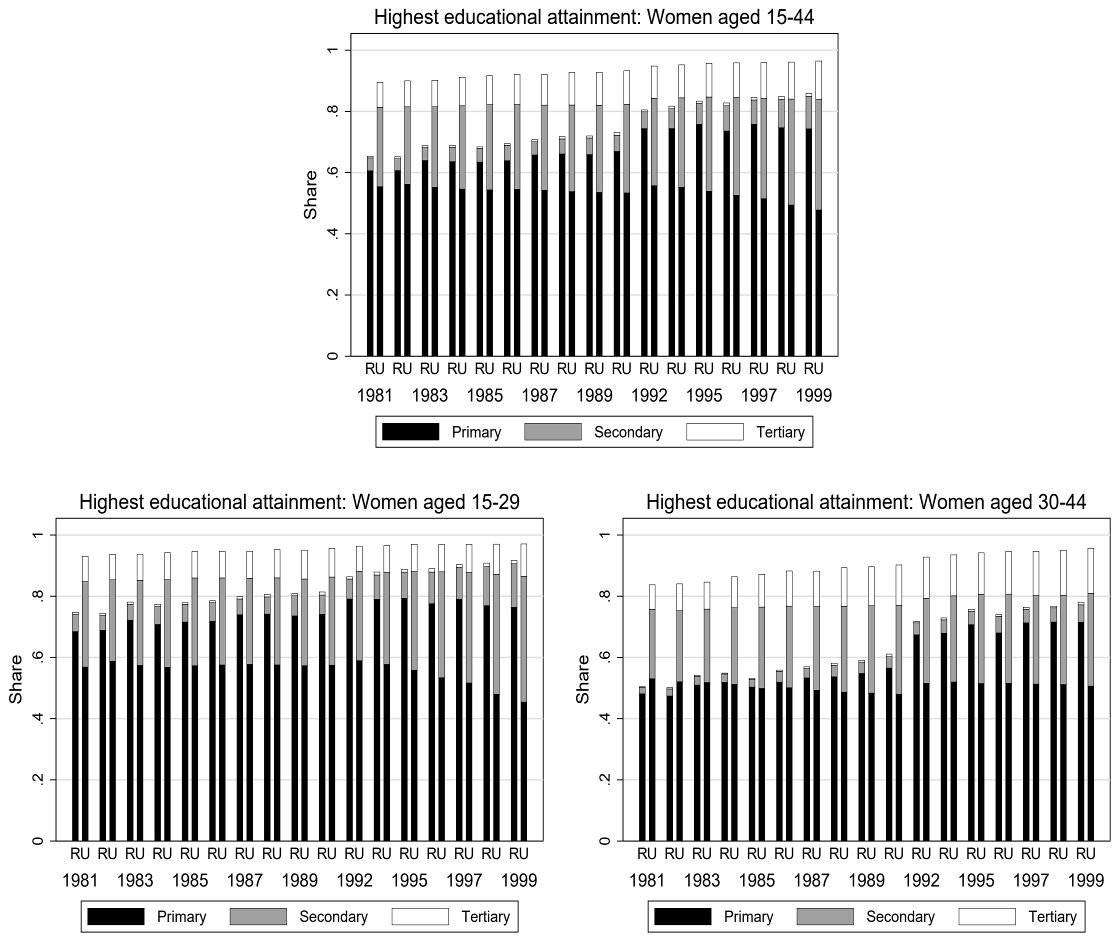

Figure A14: Highest Educational Attainment, Brazil 1981-99

Note: Graphs show shares of female rural (R) and urban (U) workers primary, secondary or tertiary educational attainments. Rural and urban groups are defined as in Table 2. Mixed urban-rural couples excluded. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A1: OUTCOME VARIABLES (FERTILITY MEASURES)

Variables	Description
Total number of children born (alive or dead) after completed fertility phase (45 years or above)	This variable is constructed from the PNAD survey years with a fertility section including information on the total number of births (alive or dead), i.e., the 1984-85, 1992-93, 1995-99, 2001-09 and 2011-14 waves. Our sample includes female respondents of the birth cohorts between 1930 to 1969, who are between 45-69 years of age in a given survey year (In 1984, the fertility section is only answered by women up to 54 years old. We exclude women above 69 years to account for selective mortality). We assign the reported total number of births of women older than 45 years of age in a given year retrospectively to the year in that they were 45 years old.
New-born Child under 1 year old (0/1)	Binary indicator variable that equals 1 if a woman has a new-born child under 12 months of age and 0 otherwise. Constructed using household composition and birth data. Multiples and two birth events per year are coded single births.
Childlessness ()	Binary indicator that equals 1 if a woman aged 45 years or above never gave birth to a child. Based on the variable 'total number of children born (alive or dead) after completed fertility phase'.

Table A2: TREATMENT VARIABLES

Variables	Description
RURAL	Binary indicators that equals 1 if an individual is a rural worker, 0 otherwise. Assignment of rural worker status according to an (legal) occupation-based definition, as follows: 1.) by the legal classification of the individual's current occupation performed during the reference week. 2.) for individuals not employed during the reference week we use retrospective occupational information available for up to 4 years prior to the reference year. 3.) for individuals with insufficient personal occupational information we assign the treatment status based on the occupation of the family head. 4.) for individuals with insufficient personal and household head information we assign the status using information on the household location, i.e., rural or urban residence.
POST	Binary indicators that equals 1 for observations in the years after the pension reform in July 1991, 0 otherwise.
Gross present pension wealth	Pre- and post-reform pension wealth is computed as the present value of predicted old age benefits before and after the reform adjusted for real interest rates and average survival probabilities (computed using IBGE mortality tables; first time available in 1998): $Pension\ Wealth = \sum_{t=0}^{T-a} s_{a,t} \times \frac{1}{(1+i)^t} \times pension_t$, with $s_{a,t}$ denoting the probability of a person of age a in a given year surviving until year t; $T-a$, indicates the remaining maximum lifespan differentiated by sex and birth cohort; i a constant discount rate (a plausible rate of 12% in the main analysis, see Azzoni and Isai (1994)); and $pension_t$ denotes the old age minimum pension benefit. A retired person receives the pension from period $t = 0$ until death. A non-retired person receives the pension starting in a future period $t > 0$, defined by the person's age and the regular retirement age. We compute pre- and post reform pension wealth for each individual and every year on the basis of individuals income from its main job. We assume that this income is the relevant base income for individuals regarding their retirement pension benefits. Further we assume that the future benefit level is 100% of this income, i.e., that individuals work until they are eligible for full benefits. If an individual is not working (i.e., has no current main job income) or its main job income falls below the minimum pension level (minimum wage), we impute the average pension benefit level of current pensioners by year, men and women and by rural and urban (in more detail, we take the conditional mean of pension benefits that lie between the 1st and 99th percentile of the distribution; for rural workers this is about the applicable minimum pension benefit level since almost all rural pensioner receive minimum pension benefits). To compute couples pension wealth, the pension wealth of rural women with a partner eligible for the regular old age pension before the reform is set to 0.

Table A3: Individual level covariates

Variables	Description
Age of the woman	Age of the female respondent (in years).
Age of the woman's partner	Age of the female respondent's partner (in years). For female respondents who currently do not have a partner, we impute the value by adding the samples (by year, federal state and rural residence) average difference in ages between women and their partners to the age of the woman.
Dummies for 10-hour brackets of reported number of hours of work per week	Binary indicators (8) that equals 1 if the individual is working 0 hours (not working), 1-10 hours, 11-20 hours, 21-30 hours, 31-40 hours, 41-50 hours, 51-60 hours or more than 60 hours on average per week, 0 otherwise. Average number of hours worked per week are conditional on employment during reference-week. Employed during reference-week refers to a binary indicator (see variable description for details).
Can read and write (literacy)	Binary indicator that equals 1 if the individual knows how to read and write (self-assessed), 0 otherwise.
Employed during reference week	Binary indicator variable that equals 1 if the respondent worked during the reference-week, 0 otherwise. Following Helfand and Brunstein (2001), our variable is harmonized by using the IBGE definition until 1992. The PNAD captures various types of employment, including employees, domestic workers, self-employed or own account workers, employers and unpaid family workers. To ensure consistency over time, our variable considers only economic activities enumerated in cash or in kind as well as unremunerated activities of at least 15 hours per week. Activities related only to the production for own consumption purposes and unpaid family work for the purpose of producing a non-marketed services (e.g. care for elderly/children) were not considered as employment. Individuals having a job but were temporarily away from work during the reference week (e.g. due to vacation or strike, etc.) are considered employed as long as they report an economic activity.
Highest educational attainment	Four binary indicators (No, Primary, Secondary, Tertiary education) for the highest educational attainment.
Married	Binary indicator that equals 1 if female is married, 0 otherwise. Based on the PNAD household and family roster.
Woman's income share	The female respondents income as a share of total household income (excluding retirement pension income).
Years of education	Years of education, i.e., all years of school attendance (Primary, Secondary, Tertiary) reported by the respondent. For female respondents who do not report their years of education, we impute the mean value by year, region and rural residence.

Table A4: HOUSEHOLD LEVEL COVARIATES

Variables	**Description**
Caretaker in the household	Binary indicator that equals 1 if a non-working (indicated by employment during reference-week variable) pension-eligible person or actual pensioner is co-residing in the household.
Household/Family pension income share	Retirement pension income (old-age, length of service and disability) of all family members/household members as a share of household/family income. Households are defined as potentially extended multigenerational family units, which may include, in addition to first-degree relatives, more distant relatives or other co-residing persons, who live in the same household and under the same roof and make their living through contributions from the household members (i.e., a potentially extended multigenerational family). Since we do not know anything about caring arrangements and intra-family bargaining within those households, we additionally use a narrower concept than households, i.e., self assessed families within those households.
Log household/family income	Natural logarithm of the income of all household/family members excluding retirement pension income. Income sources: labor income from main and other jobs, merchandise income, wage bonuses, income from rent, interest, donations, transfers and other sources.)
Number of adults in the household	Number of adults in the household, i.e., individuals aged 18 years or older.
Dummies for prior children	Six dummies indicating the number of previously born children (birth parity), i.e., born at least one year before the survey: 0, 1, 2, 3, 4, 5+.
Wealth Index	The wealth index is constructed as the first principal component from the following variables: Six dummies indicating the form of water supply used by the household: main network with internal/external plumbing, well or spring with internal/external plumbing, other sources with internal/external plumbing. Four dummies indicating the sewage treatment: sanitation network, septic tank, rudimentary tank, other forms of sewage treatment. Two dummies indicating access to sanitary instalments: use of sanitary instalments is exclusive to household, sanitary instalments are shared by more than one household. One dummy for whether the household has electricity. Five dummies for household ownership of durable goods: refrigerator, stove, water filter, radio, television. Three dummies indicating the type of garbage disposal: collected, burned or buried, other (dumped into empty land or into river/lake/sea). Fifteen interaction terms of three dummies for household type (house, apartment, room) with five dummies for ownership status (paid for, still paying, rented, ceded, other). Three dummies indicating the main material used on walls of the household type: masonry, wood, other materials. Four dummies indicating main material used on roof of the household type: shingle roof, concrete roof, zinc roof, other materials. Two variables for the number of rooms per capita and the number of bedrooms per capita. We construct the wealth index separately by rural and urban household location to measure relative living conditions in rural and urban areas - this is relevant regarding structural differences in housing conditions (cf. McKenzie (2005) and Vyas and Kumaranayake (2006) for the problem of clumping and truncation in PCA). In the case of item non-response we impute averages by federal state and rural/urban residence in the time spans from 1981-85, 1986-90, 1992-96 and 1997-99 (if for example a value is missing for a rural household in a federal state in one year it is replaced by the time spans mean value of the variable of rural households in this region.).

Table A5: Region/Group level covariates

Variables	Description
Child mortality	Constructed from Demographic and Health Surveys (DHS) data (1986 and 1996) for each year between 1981-97 (constant 1997 level for 1998-99) for four different macro-regions, i.e., north-east, mid-west, south and south-east (region south is represented by Rio de Janeiro and Sao Paulo in the DHS, Minas Gerais and Espirito Santo are not covered). Generated as the moving average of the number of children having dying before 60 months old in the past 10 years divided by the total number of children being born in the past 10 years (The same procedure is applied by DHS to generate child mortality statistics on regional levels.).
Race shares	Four variables constructed from Brazilian Census data (1980, 1990 and 2000). Shares are matched to the PNAD using the nearest neighbour census year. Share of black, mixed-race, white or other skin colour/race within a federal state and by rural/urban residence.
Catholic/Protestant/ Other/No religion shares	Four variables constructed from Brazilian Census data (1980, 1990 and 2000). Shares are matched to the PNAD using the nearest neighbour census year. Definition of religious affiliation includes several religious groups reported by the individual. Catholic's include: Catholic roman, Brazilian Catholic Apostolic, other Catholic, Anglican. Protestant's include: Protestant, undetermined Protestant, Adventist, Baptist, Lutheran, Methodist, Presbyterian, other traditional Protestant, Mennonite, Congregational, Pentecostal, Assembly of God, Christian Congregation of Brazil, God is Love, Foursquare Gospel, Universal of the Kingdom of God, New Life Evangelical Protestant Pentecostal, Evangelical Protestant Biblical Revival Pentecostal, Chain Of Prayer Pentecostal, House of the Blessing, Brazil for Christ, Church of the Nazarene, House of Prayer, Maranata, Salvation Army, Religion of God, Undetermined Evangelical Protestant, Renewed Evangelical Protestant without institutional ties, Pentecostal Evangelical without institutional ties. Other include: Buddhist, Hindu, Jewish, Muslim, Orthodox, Neo-Christian, other Neo-Christian, undetermined Christian, Christian without institutional ties, Jehovah's Witnesses, Latter Day Saints (Mormons), Spiritist, other afro Brazilian, oriental Brazil, new oriental, other oriental, esoteric, indigenous, other minorities, other not classified.
Rede Globo Coverage	Constructed using the La Ferrara et al. (2012) data set on Rede Globo coverage. Share of individuals in a federal state receiving the Rede Globo broadcasting signal (television reception) in the year preceding the survey.
Share of women using modern/other/no contraceptive methods	Constructed from Demographic and Health Surveys (DHS) data (1986 and 1996). Share of female respondents self-reported use of modern contraceptive methods (Pill, IUD, Injections, Diaphragm/Foam/Jelly, Condom, Female or Male Sterilization), other contraceptive methods (Periodic abstinence, withdrawal, other methods, country-specific traditional/folkloric methods) or no contraceptive method. Constructed for four different macro-regions, i.e., north-east, mid-west, south and south-east (region south is represented by Rio de Janeiro and Sao Paulo in the DHS, Minas Gerais and Espirito Santo are not covered).
Regional industry specific trade shocks	Constructed as in Dix-Carneiro and Kovak (2017) but not aggregated on regional level (therefore, regional industry specific) using annual tariff reductions taken from De Paiva Abreu (2004). Region-industry specific trade shocks for 20 industry groups: 1) agriculture; 2) mineral mining; 3) petroleum and gas extraction and coal mining; 4) non-metallic mineral goods manufacturing; 5) iron and steel nonferrous and other metal production and processing; 6) machinery, equipment, commercial installation manufacturing and tractor manufacturing; 7) electric, electronic and communication equipment and component manufacturing; 8) automobile, transportation and vehicle parts manufacturing; 9) wood products, furniture manufacturing and peat production; 10) paper manufacturing, publishing and printing; 11) rubber product manufacturing; 12) chemical product manufacturing; 13) petroleum refining and petrochemical manufacturing; 14) pharmaceutical products, perfumes and detergents manufacturing; 15) plastic products manufacturing; 16) textiles manufacturing; 17) apparel and apparel accessories manufacturing; 18) footwear and leather and hide products manufacturing; 19) food processing (coffee, plant products, meat, dairy, sugar, oils, beverages and other); 20) miscellaneous other products manufacturing.

Table A6: Sensitivity Analysis - Women Outside Age Range Excluded: Pooled (Women aged 15-44) and Subgroups (15-29 vs. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
DID women aged 15-44	-0.008**	-0.007**	-0.012***	-0.012***	-0.010***	-0.011***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
DID women aged 15-29	0.004	0.005	0.001	0.000	0.001	-0.000
	(0.004)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
DID women aged 30-44	-0.028***	-0.028***	-0.027***	-0.025***	-0.024***	-0.023***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Year and region FE	No	Yes	Yes	Yes	Yes	Yes
Covariates (see note):						
Individual	No	No	Yes	Yes	Yes	Yes
Job	No	No	No	Yes	Yes	Yes
Household	No	No	No	No	Yes	Yes
Regional/Group	No	No	No	No	No	Yes
N:	15-44 (1,325,397); 15-29 (780,431); 30-44 (544,966)					
R^2 15-44:	0.006	0.008	0.108	0.113	0.119	0.119
R^2 15-29:	0.004	0.006	0.148	0.156	0.160	0.161
R^2 30-44:	0.011	0.015	0.047	0.051	0.057	0.058

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Covariates as in Table 3. Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A7: SENSITIVITY ANALYSIS - MIXED OCCUPATION COUPLES: POOLED (WOMEN AGED 15-44) AND SUBGROUPS (15-29 VS. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Women from mixed occupation couples included, dummy for mixed couple						
DID women aged 15-44	-0.008**	-0.007**	-0.013***	-0.014***	-0.013***	-0.012***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
DID women aged 15-29	0.007*	0.008**	-0.000	-0.002	-0.001	-0.001
	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
DID women aged 30-44	-0.031***	-0.030***	-0.028***	-0.028***	-0.027***	-0.024***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
N:	15-44 (1,429,160); 15-29 (844,420); 30-44 (584,740)					
R^2: 15-44	0.006	0.008	0.104	0.109	0.115	0.116
R^2: 15-29	0.004	0.005	0.150	0.158	0.162	0.163
R^2: 30-44	0.011	0.016	0.058	0.061	0.067	0.069
Panel B: Women from mixed occupation couples included no dummy for mixed couple						
DID women aged 15-44	-0.007**	-0.006**	-0.013***	-0.013***	-0.012***	-0.012***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
DID women aged 15-29	0.007**	0.008**	-0.001	-0.002	-0.001	-0.001
	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
DID women aged 30-44	-0.028***	-0.028***	-0.028***	-0.027***	-0.026***	-0.024***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
N:	15-44 (1,429,160); 15-29 (844,420); 30-44 (584,740)					
R^2 15-44:	0.006	0.008	0.104	0.109	0.115	0.116
R^2 15-29:	0.004	0.005	0.150	0.158	0.162	0.163
R^2 30-44:	0.011	0.016	0.057	0.061	0.067	0.069
Controls in Panel A and Panel B:						
Year and region FE	No	Yes	Yes	Yes	Yes	Yes
Covariates (see note):						
Individual	No	No	Yes	Yes	Yes	Yes
Job	No	No	No	Yes	Yes	Yes
Household	No	No	No	No	Yes	Yes
Regional/Group	No	No	No	No	No	Yes

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Covariates as in Table 3. Rural and urban group definition as in Table 2. Panel A: Dummy for mixed couple=1 if occupation of the woman is not equal the occupation of her partner, 0 otherwise; Panel B: Mixed couples included according to the woman's occupation. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A8: SENSITIVITY ANALYSIS - RURAL AND URBAN GROUP DEFINITION: POOLED (WOMEN AGED 15-44) AND SUBGROUPS (WOMEN AGED 15-29 VS. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Defined by household location						
DID women aged 15-44	-0.010***	-0.001***	-0.015***	-0.014***	-0.014***	-0.011***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)	(0.003)
DID women aged 15-29	0.002	0.002	-0.004	-0.003	-0.004	-0.001
	(0.003)	(0.003)	(0.003)	(0.002)	(0.003)	(0.003)
DID women aged 30-44	-0.031***	-0.031***	-0.030***	-0.028***	-0.027***	-0.021***
	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)
N:	15-44 (1,429,160); 15-29 (844,420); 30-44 (584,740)					
R^2 15-44:	0.006	0.008	0.104	0.110	0.116	0.116
R^2 15-29:	0.004	0.005	0.150	0.158	0.163	0.163
R^2 30-44:	0.011	0.016	0.057	0.062	0.068	0.069
Panel B: Defined by personal occupational information and household location						
DID women aged 15-44	-0.009**	-0.008**	-0.013***	-0.013***	-0.012***	-0.013***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.003)	(0.004)
DID women aged 15-29	0.006	0.007*	0.001	0.000	0.001	-0.001
	(0.004)	(0.004)	(0.003)	(0.002)	(0.003)	(0.004)
DID women aged 30-44	-0.032***	-0.032***	-0.031***	-0.029***	-0.028***	-0.026***
	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)	(0.005)
N:	15-44 (1,289,689); 15-29 (757,592); 30-44 (532,097)					
R^2 15-44:	0.006	0.008	0.102	0.108	0.114	0.114
R^2 15-29:	0.004	0.006	0.146	0.155	0.160	0.160
R^2 30-44:	0.011	0.015	0.056	0.061	0.067	0.067
Controls in Panel A and Panel B:						
Year and region FE	No	Yes	Yes	Yes	Yes	Yes
Covariates (see note):						
Individual	No	No	Yes	Yes	Yes	Yes
Job	No	No	No	Yes	Yes	Yes
Household	No	No	No	No	Yes	Yes
Regional/Group	No	No	No	No	No	Yes

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Covariates as in Table 3. Rural and urban group definition as in Table 2. Panel A: Rural and urban group definition by household location; Panel B: Rural and urban group definition by personal occupational information and household location (not using occupational information of the household head). Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A9: Results of the DiD Regression Analysis - Alternative Specifications: Pooled (Women Aged 15-44) and Subgroups (Women aged 15-29 vs. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
DID women aged 15-44	-0.011***	-0.011***	-0.013***	-0.010***	-0.011***	-0.012***	-0.012***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
DID women aged 15-29	0.0001	0.000	-0.002	0.001	-0.000	-0.001	-0.000
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
DID women aged 30-44	-0.023***	-0.024***	-0.024***	-0.023***	-0.023***	-0.024***	-0.025***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Year and region FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Covariates (see note):							
Individual	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Job	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Household	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Regional/group	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Other indicators for educational achievements	Yes	No	Yes	No	No	No	No
Household retirement income share	No	Yes	No	Yes	No	No	No
Income variables on family basis	No	No	Yes	No	No	No	No
Dummies working unpaid or in public sector	No	No	No	Yes	No	No	No
Squared terms	No	No	No	No	Yes	No	No
Linear Age and age partner	No	No	No	Yes	No	Yes	No
Age and age partner dummies	No	No	No	Yes	No	No	Yes
N:	15-44 (1,326,819); 15-29 (780,682); 30-44 (546,137)						
R^2 15-44:	0.116	0.115	0.118	0.115	0.117	0.114	0.121
R^2 15-29:	0.162	0.162	0.165	0.161	0.163	0.159	0.162
R^2 30-44:	0.069	0.068	0.068	0.068	0.069	0.055	0.059

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Individual, job, household and region/group level covariates as in Table 3. Other indicators for educational achievements are a dummy indicating the ability to read and write and three dummies indicating highest educational stage (primary, secondary, college/university; has not attended any school dummy omitted) attended/attending. Income variables on family basis are log of family income and share of family income earned by the woman (both excl. pensions) for the nuclear family, i.e., only the woman, her partner and own children. Additional squared terms are years of schooling, log of household income (excl. pensions), share of household income (excl. pensions) earned by the woman and household wealth. Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A10: Test of the Common Trend: Pooled (Women aged 15-44) and Subgroups (Women aged 15-29 vs. 30-44)

Dep. Variable	Newborn child under 1 year old (0/1)		
	(1)	(2)	(3)
Period	Women aged 15-44	Women aged 15-29	Women aged 30-44
1981	0.001	-0.003	0.009
	(0.007)	(0.007)	(0.009)
1982	0.006	0.001	0.017*
	(0.008)	(0.008)	(0.009)
1983	0.007	0.008	0.006
	(0.006)	(0.006)	(0.007)
1984	0.004	0.004	0.005
	(0.006)	(0.006)	(0.007)
1985	0.009	0.008	0.011
	(0.006)	(0.006)	(0.008)
1986	0.003	-0.000	0.009
	(0.006)	(0.006)	(0.008)
1988	-0.001	0.002	-0.005
	(0.005)	(0.006)	(0.006)
1989	0.000	0.007	-0.008
	(0.005)	(0.006)	(0.007)
1990	0.005	0.006	0.003
	(0.004)	(0.004)	(0.007)
1992	-0.003	-0.000	-0.004
	(0.004)	(0.004)	(0.007)
1993	-0.004	0.007	-0.016**
	(0.006)	(0.007)	(0.007)
1995	-0.001	0.012*	-0.012**
	(0.005)	(0.006)	(0.007)
1996	-0.012*	-0.000	-0.021**
	(0.006)	(0.007)	(0.008)
1997	-0.010	0.005	-0.024***
	(0.008)	(0.010)	(0.008)
1998	-0.010	0.002	-0.019**
	(0.008)	(0.010)	(0.008)
1999	-0.020***	-0.005	-0.033***
	(0.006)	(0.006)	(0.009)
Year and region FE	Yes	Yes	Yes
Covariates (see note)	Yes	Yes	Yes
N:	1,326,819	780,682	546,137
R^2:	0.115	0.161	0.068

Note: Year-specific treatment effects for the full sample and by age groups 15-29 and 30-44. Dependent variable: Dummy, whether a child was born in the last 12 months. Full set of covariates, as in Table 3, column 6. Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A11: Desired Number of Children, by urban/rural and age groups

All women						
	1986		1996		Change	
Age groups	Urban	Rural	Urban	Rural	Urban	Rural
15-44	2.70	3.10	2.19	2.50	-0.51	-0.60
15-29	2.50	2.83	2.02	2.25	-0.47	-0.58
30-44	3.00	3.50	2.40	2.81	-0.60	-0.69
Women with at least one child						
	1986		1996		Change	
Age groups	Urban	Rural	Urban	Rural	Urban	Rural
15-44	2.86	3.29	2.31	2.66	-0.55	-0.63
15-29	2.53	2.95	2.02	2.32	-0.51	-0.63
30-44	3.09	3.57	2.48	2.89	-0.61	-0.68
Women with number of children<children desired						
	1986		1996		Change	
Age groups	Urban	Rural	Urban	Rural	Urban	Rural
15-44	2.93	3.42	2.51	2.76	-0.42	-0.66
15-29	2.68	3.10	2.25	2.49	-0.43	-0.61
30-44	3.70	4.73	3.27	3.70	-0.43	-1.03

Note: Sample consists of Brazilian females aged 15-44. Rural and urban groups are defined by household location. The corresponding DHS questionnaire from which we derive the numbers ask for the ideal number of children is the number of children the respondent would have liked to have in her whole life, irrespective of the number she already has. *Source:* DHS 1986, 1996.

Table A12: Contraceptive Methods, by urban/rural and age groups

Panel A: Modern Methods								
All women								
	1986		1996		Change		Change in %	
Age groups	Urban	Rural	Urban	Rural	Urban	Rural	Urban	Rural
15-44	0.395	0.308	0.512	0.441	0.117	0.134	29.7	43.4
15-29	0.395	0.307	0.508	0.438	0.114	0.131	28.8	42.8
30-44	0.396	0.309	0.517	0.445	0.122	0.136	30.8	44.1
Panel B: Traditional Methods								
All women								
	1986		1996		Change		Change in %	
Age groups	Urban	Rural	Urban	Rural	Urban	Rural	Urban	Rural
15-44	0.060	0.077	0.039	0.057	-0.021	-0.020	-34.8	-26.7
15-29	0.060	0.076	0.039	0.056	-0.021	-0.020	-35.1	-26.6
30-44	0.061	0.079	0.040	0.057	-0.020	-0.022	-34.4	-27.1

Note: Numbers show the shares of current contraceptive methods used by women. Pregnant women are coded as not currently using. Classification of contraceptive methods according to DHS definition: Modern methods are Pill, IUD, Injections, Diaphragm/Foam/Jelly, Condom, Female Sterilization, Male Sterilization and Norplant. Traditional methods are Periodic Abstinence (Rhythm), Withdrawal, Abstinence, and any other country specific methods. Rural and urban groups are defined by household location. *Source:* DHS 1986, 1996.

Table A13: DiD Regression Results - Covariates: Pooled (Women Aged 15-44) and Subgroups (Women aged 15-29 vs. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)		(2)		(3)	
	Women aged 15-44		Women aged 15-29		Women aged 30-44	
DID estimator	-0.012***	(0.003)	0.000	(0.003)	-0.024***	(0.004)
Years of education	-0.001**	(0.000)	-0.002***	(0.000)	0.001*	(0.000)
Dummy married	0.140***	(0.006)	0.164***	(0.007)	0.048***	(0.003)
Age of the woman	0.002	(0.001)	0.043***	(0.003)	-0.048***	(0.003)
Age of the woman squared	-0.000***	(0.000)	-0.001***	(0.000)	0.001***	(0.000)
Dummy birth-parity 1	-0.075***	(0.004)	-0.096***	(0.005)	-0.014***	(0.002)
Dummy birth-parity 2	-0.128***	(0.004)	-0.144***	(0.005)	-0.046***	(0.003)
Dummy birth-parity 3	-0.143***	(0.004)	-0.151***	(0.005)	-0.052***	(0.003)
Dummy birth-parity 4	-0.140***	(0.003)	-0.150***	(0.005)	-0.037***	(0.003)
Dummy birth-parity 5+	-0.139***	(0.003)	-0.163***	(0.004)	0.004	(0.005)
Dummy 1-10 hours of work	-0.033***	(0.003)	-0.039***	(0.005)	-0.027***	(0.002)
Dummy 11-20 hours of work	-0.034***	(0.002)	-0.040***	(0.003)	-0.027***	(0.002)
Dummy 21-30 hours of work	-0.037***	(0.002)	-0.044***	(0.002)	-0.031***	(0.002)
Dummy 31-40 hours of work	-0.042***	(0.002)	-0.056***	(0.002)	-0.032***	(0.002)
Dummy 41-50 hours of work	-0.050***	(0.002)	-0.065***	(0.002)	-0.039***	(0.002)
Dummy 51-60 hours of work	-0.059***	(0.002)	-0.076***	(0.003)	-0.042***	(0.002)
Dummy 60+ hours of work	-0.062***	(0.003)	-0.082***	(0.004)	-0.039***	(0.002)
Household income share of the woman	0.005*	(0.003)	-0.014**	(0.006)	0.014***	(0.001)
Dummy caretaker	-0.000	(0.002)	-0.015***	(0.001)	-0.001	(0.001)
Log of household income excl. pensions	-0.000	(0.000)	0.000**	(0.000)	0.000*	(0.000)
Wealth index	-0.020***	(0.003)	-0.021***	(0.003)	-0.013***	(0.002)
Age of woman's partner	0.004***	(0.001)	-0.003***	(0.001)	-0.010***	(0.000)
Age of woman's partner squared	-0.000***	(0.000)	0.000***	(0.000)	0.000***	(0.000)
Dummy Globo coverage	0.004	(0.009)	0.009	(0.007)	-0.011	(0.011)
Protestants share	-0.183***	(0.059)	-0.105	(0.061)	-0.245***	(0.061)
Other religions share	-0.632***	(0.167)	-0.609***	(0.168)	-0.475**	(0.186)
No religion share	0.037	(0.082)	-0.052	(0.080)	0.169*	(0.089)
Race share black	0.091	(0.159)	0.115	(0.146)	-0.083	(0.167)
Race share mixed	0.130**	(0.056)	0.123**	(0.050)	0.098	(0.061)
Race share other	-0.222	(0.154)	-0.235	(0.155)	-0.270	(0.175)
Under 5 years mortality rate	0.188*	(0.107)	0.120	(0.104)	0.320***	(0.107)
Year and region FE	Yes		Yes		Yes	
Other Covariates (see note)	Yes		Yes		Yes	
N:	1,326,819		780,682		546,137	
R^2:	0.115		0.161		0.068	

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Covariates as in Table 3, explained in detail in Appendix A1-A5. Other household covariates: dummies for the number of adults in the household. Other regional/group covariates: regional industry specific trade shocks (variable based on methodology of Dix-Carneiro and Kovak (2017) and annual tariff reduction data taken from De Paiva Abreu (2004)). Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A14: DiD Regression Results: Subgroups (5-Year Age Cohorts)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
DID women aged 15-19	0.003	0.004	0.001	0.001	0.001	-0.001
	(0.004)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)
in % to baseline	4.0	4.9	0.9	1.5	1.6	-0.6
DID women aged 20-24	0.007	0.008	0.005	0.004	0.006	0.005
	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)
in % to baseline	3.3	3.9	2.7	2.2	2.9	2.6
DID women aged 25-29	-0.011	-0.011	-0.012*	-0.011*	-0.001	-0.010
	(0.007)	(0.007)	(0.006)	(0.006)	(0.006)	(0.008)
in % to baseline	-6.7	-6.5	-6.9	-6.5	-5.6	-6.1
DID women aged 30-34	-0.037***	-0.036***	-0.033***	-0.032***	-0.031***	-0.030***
	(0.006)	(0.006)	(0.006)	(0.005)	(0.006)	(0.006)
in % to baseline	-26.2	-26.0	-24.1	-23.3	-22.7	-22.6
DID women aged 35-39	-0.033***	-0.031***	-0.028***	-0.026***	-0.024***	-0.023***
	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)	(0.005)
in % to baseline	-31.1	-30.2	-27.6	-26.3	-25.0	-24.2
DID women aged 40-44	-0.027***	-0.027***	-0.026***	-0.025***	-0.024***	-0.023***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
in % to baseline	-37.3	-37.5	-36.6	-35.9	-34.6	-33.2
Year and region FE	No	Yes	Yes	Yes	Yes	Yes
Covariates (see note):						
Individual	No	No	Yes	Yes	Yes	Yes
Job	No	No	No	Yes	Yes	Yes
Household	No	No	No	No	Yes	Yes
Regional/Group	No	No	No	No	No	Yes
N:	15-19 (289,695), 20-24 (257,998), 25-29 (232,989), 30-34 (209,944), 35-39 (181,595), 40-44 (154,598)					
R^2 15-19:	0.008	0.004	0.246	0.252	0.256	0.256
R^2 20-24:	0.009	0.013	0.148	0.158	0.164	0.164
R^2 25-29:	0.008	0.012	0.072	0.081	0.086	0.087
R^2 30-34:	0.011	0.015	0.047	0.053	0.060	0.061
R^2 35-39:	0.014	0.019	0.035	0.038	0.045	0.047
R^2 40-44:	0.015	0.020	0.115	0.116	0.121	0.122

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Covariates as in Table 3. Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A15: Other Specifications - Time trends and region-time trends: Pooled (Women Aged 15-44) and Subgroups (Women aged 15-29 vs. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Time-trend						
DID women aged 15-44	-0.009**	-0.008**	-0.010***	-0.010***	-0.011***	-0.010***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)	(0.003)
DID women aged 15-29	0.004	0.005	0.001	0.001	-0.000	0.001
	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
DID women aged 30-44	-0.030***	-0.030***	-0.030***	-0.029***	-0.028***	-0.027***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
N:	15-44 (1,326,819); 15-29 (780,682); 30-44 (546,137)					
R^2: 15-44	0.007	0.009	0.078	0.082	0.083	0.088
R^2: 15-29	0.004	0.006	0.123	0.125	0.125	0.133
R^2: 30-44	0.013	0.016	0.043	0.045	0.047	0.063
Panel B: Region-Time-Trend						
DID women aged 15-44	-0.008**	-0.007*	-0.006***	-0.006**	-0.009**	-0.008**
	(0.003)	(0.003)	(0.002)	(0.003)	(0.004)	(0.003)
DID women aged 15-29	0.004	0.004	0.002	0.003	0.000	0.002
	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)	(0.003)
DID women aged 30-44	-0.030***	-0.023***	-0.023***	-0.022***	-0.024***	-0.023***
	(0.004)	(0.003)	(0.003)	(0.003)	(0.004)	(0.003)
N:	15-44 (1,326,819); 15-29 (780,682); 30-44 (546,137)					
R^2 15-44:	0.008	0.009	0.079	0.082	0.083	0.088
R^2 15-29:	0.006	0.006	0.123	0.125	0.125	0.133
R^2 30-44:	0.015	0.017	0.044	0.047	0.047	0.063
Panel A: Time trend						
Panel B: Region-Time trend						
Controls in Panel A and Panel B:						
Year and region FE	No	Yes	Yes	Yes	Yes	Yes
Covariates (see note):						
Individual	No	No	Yes	Yes	Yes	Yes
Job	No	No	No	Yes	Yes	Yes
Household	No	No	No	No	Yes	Yes
Regional/Group	No	No	No	No	No	Yes

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Covariates as in Table 3. Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A16: SENSITIVITY ANALYSIS - CENSIT AREAS: POOLED (WOMEN AGED 15-44) AND SUBGROUPS (15-29 VS. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Only Not Self-Representative Censit Areas						
DID women aged 15-44	-0.007**	-0.007**	-0.009***	-0.009***	-0.009***	-0.007***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)
DID women aged 15-29	0.002	0.001	0.001	-0.001	0.000	0.002
	(0.003)	(0.003)	(0.003)	(0.002)	(0.003)	(0.002)
DID women aged 30-44	-0.023***	-0.023***	-0.022***	-0.020***	-0.019***	-0.015***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.003)
N:	15-44 (469,375); 15-29 (280,560); 30-44 (188,815)					
R^2: 15-44	0.007	0.010	0.106	0.110	0.116	0.117
R^2: 15-29	0.003	0.006	0.156	0.161	0.165	0.165
R^2: 30-44	0.016	0.025	0.066	0.069	0.076	0.077
Panel B: Only Self-Representative Censit Areas						
DID women aged 15-44	-0.016*	-0.014*	-0.015*	-0.013	-0.010	-0.012
	(0.008)	(0.008)	(0.007)	(0.008)	(0.008)	(0.008)
DID women aged 15-29	-0.007	-0.006	-0.006	-0.003	-0.000	-0.002
	(0.011)	(0.011)	(0.010)	(0.011)	(0.011)	(0.011)
DID women aged 30-44	-0.032***	-0.030***	-0.027***	-0.025***	-0.022***	-0.025***
	(0.008)	(0.008)	(0.007)	(0.007)	(0.008)	(0.007)
N:	15-44 (233,898); 15-29 (138,093); 30-44 (95,805)					
R^2: 15-44	0.004	0.005	0.104	0.110	0.117	0.116
R^2: 15-29	0.002	0.003	0.142	0.151	0.155	0.156
R^2: 30-44	0.006	0.009	0.053	0.057	0.062	0.063
Panel C: Only Metropolitan Censit Areas						
DID women aged 15-44	-0.013**	-0.012*	-0.009	-0.009	-0.003	-0.008
	(0.005)	(0.006)	(0.006)	(0.005)	(0.009)	(0.010)
DID women aged 15-29	0.003	0.004	0.008	0.009	0.015	0.010
	(0.013)	(0.013)	(0.011)	(0.011)	(0.014)	(0.015)
DID women aged 30-44	-0.038**	-0.037**	-0.034**	-0.034**	-0.030**	-0.033**
	(0.013)	(0.013)	(0.013)	(0.012)	(0.013)	(0.014)
N:	15-44 (15-44 (623,546); 15-29 (362,029); 30-44 (261,517)					
R^2 15-44:	0.002	0.005	0.098	0.104	0.111	0.111
R^2 15-29:	0.001	0.003	0.142	0.152	0.158	0.159
R^2 30-44:	0.003	0.007	0.049	0.053	0.059	0.060
Controls in Panel A, Panel B and Panel C:						
Year and region FE	No	Yes	Yes	Yes	Yes	Yes
Covariates (see note):						
Individual	No	No	Yes	Yes	Yes	Yes
Job	No	No	No	Yes	Yes	Yes
Household	No	No	No	No	Yes	Yes
Regional/Group	No	No	No	No	No	Yes

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Covariates as in Table 3. Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01,^{**} p < 0.05,^{*} p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A17: Sensitivity Analysis - Region-Censit Area Cluster Variable: Pooled (Women aged 15-44) and Subgroups (15-29 vs. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
DID women aged 15-44	-0.008**	-0.008**	-0.012***	-0.012***	-0.011***	-0.012***
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
DID women aged 15-29	0.004	0.005	0.001	0.000	0.001	-0.000
	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)	(0.003)
DID women aged 30-44	-0.030***	-0.029***	-0.029***	-0.027***	-0.025***	-0.025***
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)
Year and region FE	No	Yes	Yes	Yes	Yes	Yes
Covariates (see note):						
Individual	No	No	Yes	Yes	Yes	Yes
Job	No	No	No	Yes	Yes	Yes
Household	No	No	No	No	Yes	Yes
Regional/Group	No	No	No	No	No	Yes
N:	15-44 (1,326,819); 15-29 (780,682); 30-44 (546,137)					
R^2 15-44:	0.006	0.009	0.103	0.108	0.114	0.115
R^2 15-29:	0.004	0.006	0.148	0.156	0.161	0.161
R^2 30-44:	0.011	0.016	0.057	0.061	0.067	0.068

Note: DID estimates of the pension reform. Dependent variable: Dummy, whether a child was born in the last 12 months. Covariates as in Table 3. Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the level of censit areas by regions in parentheses, $^{***}p < 0.01,^{**} p < 0.05,^{*} p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.

Table A18: Sensitivity Analysis - Two-period model: Pooled (Women aged 15-44) and Subgroups (15-29 vs. 30-44)

Dependent Variable	Newborn child under 1 year old (0/1)					
	(1)	(2)	(3)	(4)	(5)	(6)
DID women aged 15-44	-0.007**	-0.007*	-0.010*	-0.003	-0.010	-0.053
	(0.003)	(0.004)	(0.005)	(0.008)	(0.007)	(0.043)
DID women aged 15-29	0.005	0.005	0.003	0.004	0.007	0.017
	(0.004)	(0.005)	(0.004)	(0.006)	(0.008)	(0.049)
DID women aged 30-44	-0.027***	-0.027***	-0.047***	-0.032*	-0.042***	-0.087***
	(0.004)	(0.005)	(0.014)	(0.016)	(0.011)	(0.027)
Year and region FE	No	Yes	Yes	Yes	Yes	Yes
Covariates (see note):						
Individual	No	No	Yes	Yes	Yes	Yes
Job	No	No	No	Yes	Yes	Yes
Household	No	No	No	No	Yes	Yes
Regional/Group	No	No	No	No	No	Yes
N:	15-44 (80), 15-29 (80), 30-44 (80)					
R^2 15-44:	0.722	0.907	0.945	0.963	0.975	0.992
R^2 15-29:	0.641	0.871	0.921	0.941	0.951	0.973
R^2 30-44:	0.649	0.861	0.914	0.938	0.968	0.997

Note: DID estimates of the pension reform in a two-period model: all variables are pre- and post reform averages by region and rural and urban occupation. Dependent variable: Dummy, whether a child was born in the last 12 months. Covariates as in Table 3. Rural and urban group definition as in Table 2. Mixed urban-rural couples excluded. Standard errors clustered at the regional level in parentheses, $^{***}p < 0.01,^{**} p < 0.05,^{*} p < 0.1$. *Source:* PNAD 1981-90, 1992-93, 1995-99.