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Social Pensions and Intimate Partner Violence Against Older Women

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Social Pensions and Intimate Partner Violence against Older Women*

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Abstract

The prevalence and determinants of intimate partner violence (IPV) among older women are understudied. This paper documents that the incidence of IPV remains high at old ages and provides the first evidence of the impact of access to income on IPV for older women. We leverage a Mexican reform that lowered the eligibility age for a non-contributory pension and a difference-in-differences approach. Women's eligibility for the pension increases their probability of being subjected to economic, psychological, and physical/sexual IPV. In contrast, we show that IPV does not increase when men become eligible. Looking at potential mechanisms, we find suggestive evidence that men use violence as a tool to control women's resources. Additionally, women reduce paid employment after becoming eligible for the pension, which may indicate that they spend more time at home, leading to greater exposure to potentially violent partners.

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> *Globally, evidence is lacking on the prevalence, patterns and types of violence against women aged 50 years and older, particularly in low- and middle-income countries.*
>
> WHO, *Violence against women 60 years and older*, 2024

1 Introduction

Globally, one in six people aged 60 and above has experienced some form of abuse in the past year (Yon et al., 2017). The incidence, prevalence, and complexity of elder abuse are projected to increase as many countries are facing rapid population aging (Sethi et al., 2011). Nonetheless, violence against older people remains severely understudied, partly because of a lack of data. In particular, very little is known about the prevalence or determinants of intimate partner violence (IPV) for women of post-reproductive age.[1] The implications of existing studies of IPV against prime-age women might not necessarily help combat violence against older women, who may face unique challenges and barriers to recognizing, reporting, and seeking help for abuse.[2]

This paper addresses these gaps by documenting the prevalence of IPV beyond reproductive age and by providing the first evidence of the impact of women's access to income in old age on IPV. We show that in 2016 in Mexico, the country studied in this paper, 7.06% of women aged 50-79 with a current partner have reported experiencing physical or sexual IPV within the past 12 months, which is comparable to the 8.35% of women aged 15-49 who reported similar experiences. The incidence of other forms of IPV is even higher: 22.55% of women aged 50-79 are subjected to psychological IPV, and 10.25% of them are victims of economic abuse inflicted by their partner. To understand the effect of exogenous changes in women's income on IPV, we investigate whether social pensions that aim to improve the standards of living of the elderly may also inadvertently affect IPV against older women.

A growing number of Latin American countries have recently introduced or expanded non-contributory (social) pensions, which are an important social protection tool in settings with high labor informality. In particular, due to the large gender coverage gaps in old-age pensions, social pensions are an effective tool to reach women in old-age poverty (UN, 2015). While a growing

[1] In this paper, IPV is defined as violence perpetrated by the intimate partner or spouse of the victim. It can happen in many ways, including physical, verbal, emotional, economic, and sexual abuse.

[2] IPV is associated with various adverse consequences for the victims and their children. Several studies have shown associations between IPV and various outcomes (Lloyd, 1997; Farmer and Tiefenthaler, 2004; Pollak, 2004; Tolman and Wang, 2005). More recent studies show causal estimates of adverse effects on victims and their children (Aizer, 2011; Bindler and Ketel, 2022; Currie et al., 2022; Bhuller et al., 2024) and negative spillovers within the school environment (Carrell and Hoekstra, 2010; Carrell et al., 2018).

body of literature studies the impact of social pensions on consumption, retirement, poverty, and health, the extant literature does not speak to the impact of social pensions on IPV. On the one hand, additional household income could reduce IPV by reducing stress and triggering events (Heath et al., 2020). On the other hand, social pensions could lead to an unintended increase in IPV, as additional income received by women may bolster their economic autonomy and strengthen their agency within households, prompting male partners to use IPV as an instrument to extract resources or reduce women's bargaining power (e.g. Angelucci, 2008; Bobonis et al., 2013; Erten and Keskin, 2018). Therefore, the effect of pensions on IPV predicted by theory is ambiguous, and an empirical investigation is needed to shed light on this issue.

To estimate the causal impact of women's eligibility for social pensions on IPV, we leverage the expansion of a non-contributory pension program in Mexico. Several features of this program make it an ideal natural experiment to study this question. First, the reform lowered the eligibility age for the non-contributory pension from 70 to 65 in 2013, which allows us to estimate the plausibly exogenous variation in pension eligibility across age groups and over time using a Difference-in-Differences (DID) approach. Second, the amount of transfer is non-trivial. Beneficiaries received a permanent, stable cash transfer of 1,160 Mexican pesos bimonthly (580 per month on average), approximately 50% of their monthly per capita income. This newly found and steady income stream could persistently shift women's economic status within the household. Third, the research setting offers a rare opportunity to investigate IPV among older women, thanks to data availability. We take advantage of a large Mexican household survey that asks questions about the experiences of IPV to women aged 15 and older, unlike most reproductive health surveys that only include women up to 49 years old. Fourth, Mexico is a relevant country for studying IPV against older women because the incidence of IPV is high in Mexico. Based on 2011 Mexican data, which were collected before the reform, 21.57% of women over 60 had been victims of any form of IPV in the past year, including physical, psychological, economic, and sexual abuse.

We employ three waves (2006, 2011, and 2016) of the National Survey on the Dynamics of Household Relationships (ENDIREH), a cross-sectional national and state-level representative survey that contains detailed information on IPV for women above 15 throughout the life cycle. The richness of the data allows us to examine the impact of social pensions on different types of IPV, including physical, sexual, and psychological violence, as well as measures of economic abuse, a form of IPV that is also pervasive but understudied. The data also permit us to analyze heterogeneity by subgroups of the population of interest and look at additional outcomes, including labor market outcomes, co-residence patterns, and decision-making within the household, to shed light on the mechanisms driving these treatment effects. Our empirical analysis focuses on women who do not receive a contributory pension, are married or in a union, and whose partners live at home at the time of the interview.

To isolate the causal effect of social pensions on IPV, we use a DID method that leverages the quasi-random assignment of social pensions across age groups and over time. We compare the outcomes of women aged 66 to 69 with those aged 61 to 64 before and after the program's implementation.[3] We demonstrate that after the 2013 reform, women aged 66 to 69 have a substantially higher likelihood of receiving a monetary transfer from the government than before the reform: the DID estimate shows that the reform increased the probability by 45.6 percentage points. Regarding IPV, we investigate the impact of the reform on the likelihood of having experienced abuse from an intimate partner in the past 12 months. We find that becoming eligible for a non-contributory pension significantly impacts physical or sexual violence, economic violence, and psychological violence. More specifically, our estimates show that the incidence of IPV among women aged between 66 and 69 increases relative to women aged between 61 and 64.

Within the group of women who are eligible for the pension, we find that the estimated effect of women's eligibility on IPV varies with women's age. For all three types of IPV, there is a larger impact for women aged 66 to 67 than for women aged 68 to 69, however, the coefficients are not statistically different, except for economic IPV. For psychological IPV, we estimate a statistically significant increase also for women aged 68 to 69. Importantly, we find no evidence that the take-up of the program varies with women's age. We propose two potential explanations for why the effect of the reform on IPV is larger for women aged 66 and 67. First, from the institutional setting perspective, in 2016, women aged 66 to 67 have recently become eligible for at most two years, while women aged 68 to 69 have been eligible for the pension for three years since 2013 (see Appendix Figure A1). Secondly, household dynamics may change and adjust rapidly once the new financial resources become accessible.

Several robustness checks confirm the validity of our estimates. The results of a DID using a placebo reform with only pre-reform waves provide support for the common-trends assumption. An event study by age further corroborates the causal interpretation of our estimates. We demonstrate the lack of anticipation effects by showing no impact of the reform on younger women (comparing the outcomes of women aged 61-64 with those aged 56-59). Furthermore, we show that the reform does not affect marital status or change the predetermined characteristics of women in our sample, suggesting that our results are not biased by endogenous changes in the composition of the sample.

[3] We drop women aged 65 because we can only observe respondents' age at integral in the survey; therefore, the treatment status of respondents aged 65 cannot be determined with certainty. Previous research also uses this approach to study the impact of eligibility for the program on extreme poverty (Ávila-Parra et al., 2024). Additionally, because the survey asks about the incidence of IPV in the 12 months before the interview, women aged 65 may report episodes that happened before they gained eligibility for the program.

Next, we examine whether the reform's effect on IPV varies with other predetermined characteristics of the women and their partners. The increase in the incidence of economic and psychological IPV after women gain eligibility is primarily driven by women who experienced family violence during their childhood. The increase in the incidence of physical/sexual IPV is larger for women in poorer households and women married to a man with less than primary education.

As the 2013 reform also expanded pension eligibility for men, we next study the effect of men's eligibility on the probability that their partners experience IPV. Interestingly, we find that women's risk of being a victim of IPV is not affected by their husbands' eligibility for the pension.

To shed light on the potential mechanisms behind the estimated effects of women's eligibility on IPV, we take advantage of the granular information on the types of IPV experienced by women and find that the estimated increase in IPV is driven by increases in economic abuse combined with other types of violence (physical/sexual IPV and/or psychological IPV). In contrast, the probability of experiencing physical/sexual IPV and/or psychological IPV alone without economic abuse is not affected. These findings corroborate the interpretation that partners of women who have recently become eligible for the pension use violence as a tool to extract economic rents (Eswaran and Malhotra, 2011; Bobonis et al., 2013; Erten and Keskin, 2018). Turning to labor market outcomes, we find that women reduce paid employment and are more likely to report not working because they are dedicated to household work after becoming eligible for the social pension, suggesting the possibility that an increase in time spent at home and greater exposure to violent partners may also play a role in explaining the increases in IPV. This finding echoes the results of previous research that found an increase in IPV following job loss in Brazil (Bhalotra et al., 2021). Moreover, we show that the increase in psychological IPV is driven by women married to men whose eligibility status is unchanged between 2011 and 2016. This finding is consistent with the male backlash/status inconsistency story. In addition, we explore changes in women's bargaining power as a possible channel. We first show that becoming eligible for the pension does not impact women's decision-making. The null result is consistent with the household bargaining channel being weaker among older women than younger women, which may be due to older women having fewer outside options. To further test the importance of outside options, we show that the effect of women's eligibility for the pension on IPV does not vary between states that allow unilateral divorce and states that do not, further corroborating the conjecture that the threat of leaving the marriage may be less relevant for older women. Finally, becoming eligible for the pension does not impact women's household composition.

Our paper makes three contributions. First, we shed light on the severely understudied topic of violence against older adults by providing some of the first quantitative evidence on the prevalence of IPV among women of post-reproductive age. As the aging population presents challenges worldwide, it is imperative to study family violence against older women as an urgent and relevant topic. Older women may be more vulnerable to abusive relationships, for example, because they face different challenges in reporting abuse or because they experience life transitions such as retirement. Although older victims of IPV have been systematically overlooked in the literature, in recent years, development professionals at the World Bank and the United Nations have recognized that more research is needed to address this knowledge gap.[4] Our data indicate that, despite the risk of IPV being lower in older ages, the incidence for women aged 50-79 is nonetheless high and comparable to the younger age groups in Mexico (see Figure 1).

Second, this paper closely relates and contributes to the literature on the economic determinants of violence against women by specifically examining risk factors for older populations. By analyzing the effect of an exogenous increase in women's income, provided by the non-contributory pension program, on the prevalence of IPV, we significantly add to the existing research that has focused on labor market opportunities (Aizer, 2010; Erten and Keskin, 2021; Kotsadam and Villanger, 2022; Sanin, 2023) and cash transfers (Buller et al., 2018; Baranov et al., 2021).[5] Our paper addresses a critical gap in understanding how financial changes affect IPV dynamics among an often overlooked but still vulnerable demographic group. It also broadens the scope of research by providing novel evidence of the effect of exogenous income changes on economic abuse, an overlooked form of IPV. In doing so, it adds to recent research showing the impact of cohabiting with an abusive partner on women's employment and earnings (Adams et al., 2024).

Third, we add to the literature on the effects of non-contributory pensions. Between 2000 and 2013, at least 18 countries in Latin America introduced reforms of non-contributory pension programs (Villagómez et al., 2014), and government expenditure on non-contributory pensions is at least 1% of GDP in several countries in Latin America (Bando et al., 2022). A growing body of literature has examined the impact of non-contributory pensions on various measures of well-being, including consumption, physical and mental health, labor supply, and the outcomes of other family members (e.g. Case and Deaton, 1998; Pfutze and Rodríguez-Castelán, 2019; Águila et al., 2018; Huang and Zhang, 2021; Miglino et al., 2023).[6] Only a handful of papers study the effect of pensions on the relationship between eligible women and their partners: Ambler (2016) exam-

[4]For instance, in 2021, the World Health Organization wrote that "While existing evidence indicates that younger women and women of reproductive age are at the highest risk of IPV and sexual violence, the magnitude, patterns, and forms of violence experienced by older women need to be better understood and researched...". In particular, development professionals point to the lack of data on women over 50 as a key challenge in documenting and understanding IPV among this vulnerable population (Meyer et al., 2020; Sardinha et al., 2022).

[5]A growing literature has also examined how IPV varies with policy interventions, shocks, and historical factors. For example, previous studies have examined how IPV responds to compulsory education reforms (Erten and Keskin, 2018), arrest and prosecution policies (Aizer and Dal Bo, 2009; Chin and Cunningham, 2019; Iyengar, 2009), conflict (LaMattina, 2017), and historical traditions (Alesina et al., 2021; Tur-Prats, 2019).

[6]For a summary of this literature, see Bando et al. (2022).

ines the impact of the South African old-age pension on household decision-making, and Berniell et al. (2020) look at the effect of a non-contributory pension in Argentina on divorce and participation in household chores. We expand this knowledge by investigating the effect of eligibility for a non-contributory pension on IPV, which is a harmful and pervasive form of violence against women.

The remainder of the paper proceeds as follows: Section 2 discusses the institutional setting in Mexico and the 2013 reform of the non-contributory pension system. Section 3 describes the data and Section 4 the empirical strategy. Section 5 presents the main results on IPV. Section 6 combines a conceptual framework together with empirical analyses to better understand the mechanisms. Section 7 concludes.

2 Background

Despite recent progress toward gender equality, substantial gender gaps remain in Mexico, especially regarding the labor market and women's agency. At only 45%, Mexican women's labor force participation rate is 6 percentage points lower than the average in all Latina American and Caribbean countries. Conservative gender norms about the division of household chores and who controls household resources are still prevalent. On average, women spend six hours doing unpaid housework daily compared to two hours for men. Furthermore, half of Mexican women agree that "women earning more than their husbands is problematic." Violence against women is high in general, with 66% of women aged 15 years and older having experienced at least one violent incident. The share of divorced and separated women has increased in the past two decades, but remains low among low-educated and older women (Inchauste Comboni et al., 2019).

Mexico is one of the countries in Latin America with the lowest coverage rates for contributory pensions. This is due to several factors, including a requirement to contribute for a minimum of 1,250 weeks (around 24 years), a high incidence of informality in the labor market, and a significant mobility between the formal and informal sectors. Consequently, a considerable proportion of the Mexican population aged 65 and older has no income. Women are more likely to work part-time and in the informal sector compared to men,[7] and these two factors contribute to a wide gender gap in access to a pension fund or savings for retirement, leading to a higher risk of poverty for women in old age (Inchauste Comboni et al., 2019). Specifically, in 2010, 26.8% of women and 9.8% of men above 65 had no income. In addition, for those with some income, the coverage rate for contributory pensions was low, with only 24.3% of women and 40.1% of men having a

[7] The gap in informal work between women and men is larger in the northern states (Inchauste Comboni et al., 2019).

contributory pension (Villagómez et al., 2014).

Due to the inadequate coverage of contributory pension schemes, in 2007 the Mexican government introduced a non-contributory pension program called "El Programa de Adultos Mayores" (PAM) to provide financial support to the elderly population. Initially, only individuals aged 70 and above who were not eligible for a contributory pension and lived in small villages were eligible for PAM. In subsequent years, the program was gradually expanded to all individuals aged 70 and above.[8]

We study the 2013 PAM expansion, which broadened eligibility to all adults aged 65 and above who were not receiving a contributory pension. Retiring at age 65 is not mandatory in Mexico, and eligibility for PAM does not require beneficiaries to be out of the labor force. The expansion was announced in December 2012 and quickly implemented in February 2013. Beneficiaries received a bimonthly cash transfer of 1,160 Mexican pesos, about 50% of their monthly per capita income,[9] and the money could be delivered in person or electronically.[10] There was a strong program take-up. By the end of 2014, a third of the population aged 65-69 were beneficiaries of PAM. In 2017, PAM was one of the main social protection instruments in the country, with a budget of 17.6 billion pesos, about 0.16% of Mexico's GDP, and 5.1 million beneficiaries.

3 Data and Descriptive Statistics

3.1 Data

To estimate the impact of becoming eligible for the social pension on women's experience of IPV, we use three rounds (2006, 2011, and 2016) of the "*Encuesta Nacional sobre la Dinámica de la Relaciones en los Hogares*" (ENDIREH).[11] ENDIREH is a cross-sectional national and state-level representative survey that collects information on the experiences of IPV for women aged 15 and older. One woman per household is interviewed to ensure privacy and confidentiality. A total of 133,398, 152,636, and 111,256 women were interviewed in 2006, 2011, and 2016, respectively.

[8] In 2008, PAM covered individuals aged 70 and above living in villages with up to 20,000 inhabitants. From 2009 to 2012, the program was expanded to eligible individuals living in villages of up to 30,000 inhabitants. Finally, in 2012, the program was available to all individuals aged 70 and above who did not qualify for a contributory pension. Those individuals eligible for the program received a cash transfer of 500 Mexican pesos per month (40 US dollars at the time of the policy). For more details on the introduction of PAM in 2007, see previous studies by (e.g. Amuedo-Dorantes and Juarez, 2015; Galiani et al., 2016; Pfutze and Rodríguez-Castelán, 2019).

[9] Calculated by the authors using MCS-ENIGH.

[10] The social pension program also provided assistance in opening a bank account and promoted social engagement by encouraging beneficiaries to participate in personal development groups and information sessions on topics such as healthcare, human rights, etc.

[11] The English name of the survey is "*National Survey on the Dynamics of Household Relationships*".

The ENDIREH questionnaire contains detailed questions on various forms of abuse, including physical, sexual, psychological, and economic abuse. Table A1 shows the survey questions used to define the IPV variables, while Appendix A defines all variables used in the analysis. We group physical and sexual IPV because this is the definition typically used by the World Health Organization to describe global prevalence. Moreover, sexual violence is infrequent in our sample. We also report the disaggregated results when we investigate potential mechanisms in Section 6.

In our analysis, we measure IPV using women's self-reported experiences of IPV in the 12 months before the interview. This allows us to isolate the episodes of IPV that occurred after the 2013 reform in the 2016 survey wave. We focus on incidents of violence inflicted by the current partner on women who were married or in an informal union and whose partners were living in the house at the time of the interviews. We exclude 203 women (1.95% of our sample) whose partners are not living in the house. We also restrict our sample to women without contributory pension income, as these women are more likely to be eligible for the program.[12] Finally, we exclude four states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011 from our main analysis. Our results are robust to including the states in the sample (Table A8).[13]

We cannot observe whether ENDIREH respondents or their partners are PAM beneficiaries because the survey does not collect detailed information on the various types of government transfers received by household members. Therefore, when using ENDIREH, we rely on the information on whether the respondents receive any monetary transfer from the government other than Progresa as a proxy for PAM take-up.[14] This information is not available for the respondents' partners.

To supplement the analysis of the effect of eligibility on PAM takeup and household income, we use the 2008, 2010, 2012, 2014, and 2016 waves of the Socioeconomic Conditions Module of the National Survey of Household Income and Expenditure (MCS-ENIGH), following Ávila-Parra et al. (2024).[15] MCS-ENIGH respondents are asked a specific question on whether they receive PAM, allowing us to measure take-up precisely for both men and women.

[12] In 2013, women without a contributory pension were eligible. In 2014, eligibility was further expanded to those who receive a contributory pension below 1092 Mexican pesos per month (SEDESOL, 2014). In ENDIREH, we define women without contributory pension income if they report having no income from retirement and pension, but we do not observe the amount of the contributory pension received. We exclude 1,715 women (14.4% of our sample) who receive a contributory pension. Our estimates are qualitatively very similar when we include women who receive a contributory pension in the sample (See Column 10 of Table A8).

[13] We exclude regions that implemented monetary, noncontributory pension programs for individuals aged 61 to 69 between 2006 and 2011, following Aguila et al. (2012), which details the various regional programs and their eligibility criteria in Mexico. Accordingly, Chiapas, Tabasco, Tlaxcala, and Zacatecas are excluded from the analysis.

[14] This survey question explicitly excludes the conditional cash transfer program known as *Progresa/Oportunidades* from the government transfer. There is a specific survey question that asks about whether the respondents are beneficiaries of Progresa/Oportunidades.

[15] Ávila-Parra et al. (2024) use the 2008, 2010, 2012, and 2014 waves.

As we explain more in detail in the next section, our identification strategy relies on comparing changes in outcomes over time for women aged 66-69, who become eligible for the pension after the 2013 reform, and women ages 61-64, who have not yet been eligible for the pension in the data (but will be eligible once they turn 65). We define the treatment and control groups based on the respondents' age at the time of the interviews in each survey round. We follow the same procedure to define eligibility for the women's husbands. The available data allows us to examine women's eligibility and their husbands' eligibility separately, but it does not permit us to separate the case in which only one spouse is eligible from the case in which both spouses are eligible. The main limitation is that because women marry older men on average, the sample size becomes very small when we look at women who are already eligible for the pension and whose husbands are not yet eligible.

We exclude women aged 65 from the sample for two reasons. First, can only observe respondents' age at integral in the survey; therefore, as highlighted by Ávila-Parra et al. (2024), women aged 65 may be "partially treated" at the time of the interviews; second, the IPV questions refer to episodes of IPV that happened in the 12 months before the interview, meaning that some women who are 65 may report incidents of IPV that occurred before they became eligible for the pension. We also exclude women older than 69 because some of them were already eligible for the non-contributory pension before the reform.[16]

Figure 1 displays how the incidence of IPV evolves with age in the two survey waves that preceded the 2013 reform (2006 and 2011). Among all age groups, psychological abuse is the most frequent type of violence, followed by economic abuse. The risk of psychological, financial, and physical or sexual abuse is highest for women younger than 20 (aged 15 to 19). While the incidence of abuse decreases with age, the decline is not steep and appears to be slower for physical or sexual violence.

Appendix Figure A2 further unpacks the time trend in IPV incidence by plotting the means for the control and treatment groups in waves 2006, 2011, and 2016 separately. We find that IPV decreased between 2006 and 2011 for both the treatment and control groups, but afterward, it remained constant or slightly increased for the control group between 2011 and 2016. In contrast, women in the treatment group experienced a large increase in IPV between 2011 and 2016.

[16] In the 2011 wave of ENDIREH (before the 2013 reform), women older than 69 in villages of up to 30,000 residents were already eligible for the "70 y más" program.

4 Empirical Strategy

Our identification strategy exploits the natural experiment created by Mexico's expansion of its non-contributory pension program for women aged 65 and older. We perform a DID approach, comparing women aged 66 to 69 with women aged 61 to 64 before and after the 2013 program was established. Women aged 66 to 69 were not eligible before 2013 (survey waves 2006 and 2011) but became eligible after 2013 (survey wave 2016). Women aged 61 to 64 constitute our control group, as they were not yet eligible for the non-contributory pension at the time of the interview. We estimate the following equation:

$$Y_{iat} = \beta_1 Age\,66\text{-}69_{ia} X\, Wave\,2016_{it} + \alpha Age_a + \lambda\, Wave_t + X'_{iat}\gamma + \varepsilon_{iat} \tag{1}$$

where Y_{iat} represents our main outcome variables for woman i, who is aged a and is interviewed in survey year t. First, we measure program receipt using a binary variable that is equal to 1 if the woman i at age a received any government program at the time of the interview t. To measure the effect of the 2013 pension reform on IPV, we use binary indicators that are equal to 1 if the woman i has experienced at least once any physical/sexual, psychological, or economic abuse from her partner at the age of a during the 12 months before the interview t, and 0 otherwise.

$Age\,66\text{-}69_{ia}$ is a dummy variable equal to 1 if woman i is aged 66 to 69 and 0 if woman i is aged 61 to 64. $Wave\,2016_{it}$ is also a dummy equal to 1 if woman i is interviewed in the year 2016 and 0 otherwise. Age_a are age fixed effects that control for time-invariant factors that affect women of the same age. $Wave_t$ are survey wave fixed effects, which account for factors that change over time and may affect IPV similarly for women in all age groups. X_{iat} is a vector of individual characteristics, including state fixed effects, rural residency, education, partner's education, number of children, a binary indicator for speaking an indigenous language (for both the woman and her partner), partner's age,[17] age at marriage, age when they started the current relationship, and experience of violence in childhood (for both the woman and her partner). We cluster the standard errors at the age level (eight groups) and report the p-values for wild cluster bootstrap with Webb weights in square brackets in all tables (Webb, 2023).[18]

Next, we look separately at ages 66-67 and 68-69 to distinguish between women who have been eligible for the pension for at least 3 years and women who have been eligible for the pension for at most two years. As illustrated in Appendix Figure A1, women who are aged 68-69 in 2016 were

[17]We control for partner's age as a continuous variable in our baseline specification. As a robustness check, we control for the partner's age more flexibly using age fixed effects and find that the results are unchanged (Table A8).

[18]The Webb weights are more reliable for inference with a few clusters (Webb, 2023). We utilize the Stata procedure "boottest" developed by Roodman et al. (2019).

aged 65 and 66 when the reform was implemented in 2013 and, in 2016, have been eligible for the pension for 3 years (since 2013). Women who are aged 66-67 in 2016 were younger than 65 in 2013 and have been eligible for at most 2 years. Thus, distinguishing between these two groups helps us differentiate between those who immediately became eligible and those who became eligible in recent years. We estimate the following regression:

$$\begin{gathered} Y_{iat} = \rho_1 Age66\text{-}67_{ia} \times Wave2016_{it} + \rho_2 Age68\text{-}69_{ia} \times Wave2016_{it} + \alpha Age_a \\ + \lambda Wave_t + X'_{iat}\gamma + \varepsilon_{iat} \end{gathered} \tag{2}$$

The coefficient estimate ρ_1 identifies short-term effects, while ρ_2 identifies medium-term effects.

5 Results

5.1 Impact of Eligibility on Takeup

Columns (1) and (2) of Table 1 show the impact of non-contributory pension eligibility on the probability of receiving a monetary transfer from the government other than Progresa/Oportunidades, which is our measure of program take-up in the ENDIREH survey. In Column (1), women aged 66 to 69 are 45.6 percentage points more likely to receive government aid after the reform. When we examine the impact by age group, Column (2) indicates that women aged 66 to 67 experience a 41.6 percentage points increase in the probability of receiving government aid, while women aged 68 to 69 experience a larger increase of 50.4 percentage points. However, the coefficients for the two age groups are not statistically different from each other, as indicated by the p-value of 0.206.

The MCS-ENIGH data, which allow precise measurement of both PAM participation (take-up) and the amount received, confirm that PAM take-up was relatively high after the 2013 reform. Table A2 shows that married women aged 66 to 69 are 57.6 percentage points more likely to receive PAM following the reform. Consistent with findings from the ENDIREH survey, take-up was slightly higher among women aged 68 to 69 (63.2 percentage points) compared to women aged 66 to 67 (52.8 percentage points), although this difference is not statistically significant. Regarding non-contributory pension amounts, Columns (3) and (4) of Table A2 reveal that married women aged 66 to 69 received, on average, an additional 297.84 pesos per month after the reform. Given the 57.6 percentage point take-up rate observed in Column (1), this implies that women participating in the non-contributory pension program were generally receiving the full amount provided by the program, 580 pesos per month.

5.2 Impact of Women's Eligibility on IPV

5.2.1 Main Results

Columns (3)-(8) of Table 1 show the reduced-form (*intent-to-treat*) estimates of non-contributory pension eligibility on the likelihood of experiencing different types of IPV in the past 12 months. The odd-numbered columns, which report the effects on women aged 66 to 69, show that non-contributory pension eligibility has a significant impact on all types of violence. Specifically, eligibility raises the probability of experiencing any violence, physical or sexual violence, economic violence, and psychological violence by 4.1, 1.8, 2.9, and 4.9 percentage points, respectively. Compared to the average probability of experiencing violence in the last 12 months for women aged 66-69 before the reform, these effects correspond to increases of 17.7%, 24%, 29%, and 28.8%, respectively. The coefficient estimates are statistically significant at the 10 percent level except for the estimate of the effect on psychological violence, which is statistically significant at the 5 percent level. These findings remain statistically significant after adjusting for multiple hypothesis testing using sharpened q-values, following Benjamini et al. (2006).[19]

The even-numbered columns, which report the effects on women aged 66 to 67 and women aged 68 to 69 separately, reveal an interesting pattern. For women aged 66 to 67, pension eligibility significantly increases all forms of IPV examined. Specifically, the probability of experiencing any violence, physical or sexual violence, economic violence, and psychological violence increases by 5.4, 2.7, 4.6, and 6.3 percentage points, respectively. Compared to baseline probabilities for women aged 66-69 before the reform, these effects represent increases of 23.4%, 36%, 46%, and 37.1%, respectively. The estimated effects on economic and psychological violence remain statistically significant at the 5 percent level even after adjusting for multiple hypothesis testing, while the effects on any violence and physical or sexual violence are statistically significant at the 10 percent level. In contrast, for women aged 68 to 69, the impact of eligibility on IPV is minimal and not statistically significant for most violence types, except for psychological violence. The increase in psychological violence is roughly half the magnitude observed for the 66 to 67 age group and is statistically significant at the 10 percent level. However, the difference in the estimated effects on IPV between the two age groups (66-67 and 68-69) is generally statistically insignificant, as indicated by the p-values testing the equality of coefficients, which are reported in the last row. Only the difference in the effects on economic violence between these age groups is statistically significant at the 10 percent level.

These findings align with the hypothesis that women who recently became eligible for non-

[19] We include Benjamini et al. (2006) sharpened q-values to control the false discovery rate (FDR) for the regressions that use the three types of violence (physical or sexual, economic, and psychological) as dependent variables. The FDR is defined as *"the expected proportion of rejections that are type I errors"* (Anderson, 2008).

contributory pensions (aged 66 to 67) might initially face a higher risk of IPV, potentially because the newly introduced income disrupts established household power dynamics. Over time, however, households may reach a new equilibrium, reflected by the smaller effects observed among women aged 68 to 69.

We also examine the reform's impact on the frequency of IPV by analyzing the probability of experiencing at least two episodes of violence in the past 12 months as an outcome variable.[20] In Table A3, the dependent variable equals one if the woman experienced IPV more than once and zero if she experienced it only once or did not experience it at all. The estimates are very similar to the effect on the incidence of experiencing IPV at least once in Table 1, suggesting that the observed increase in IPV incidence following pension eligibility is primarily driven by a rise in repeated episodes rather than isolated occurrences.

5.2.2 Threats to Validity and Robustness

Parallel Trend Assumption The DID coefficient (β_1) identifies the impact of pension eligibility under the assumption that, had the 2013 reform not occurred, trends in outcomes for women aged 66-69 would have been similar to trends in outcomes for women aged 61-64. In this Section, we perform several empirical exercises to support the interpretation of our estimates as causal and ensure the reliability of our findings.

First, we conduct a placebo test, removing the 2016 wave from the sample and using 2011 as the placebo treatment year. Table A4 presents the estimates. We find a statistically significant increase in the probability of receiving government aid for women aged 68 to 69, although, at 2.1 percentage points, the magnitude of the effect is much smaller than our main estimate. Looking at IPV, the estimated impact of the placebo reform on any violence, physical or sexual violence, economic violence, and psychological violence is small and statistically insignificant, supporting the parallel-trends assumption.

Secondly, we estimate an age-specific event-study in Figure A3 and Appendix Table A5.[21] This specification compares the change in IPV for each age group (61, 62, 63, 64, 66, 67, 68, 69) between 2016 and the two pre-reform waves, using 64-year-olds as the reference group. If treated ages (66–69) and control ages (61–64) would have moved in parallel absent the reform, the 2016 cross-section should display a discrete jump in IPV at the eligibility cutoff but no systematic pattern below it; any pre-existing age-related drift would instead appear as a smooth gradient that

[20] The survey asks respondents whether they have experienced IPV once or more than once in the past 12 months. We do not have information on the exact number of episodes in the past year.

[21] Figure A3 shows 95 percent confidence intervals based on standard errors that are robust to heteroskedasticity, while Table A5 reports standard errors clustered at the age level and p-values from wild cluster bootstrap procedures.

begins before age 66. Consistent with this logic, lead coefficients for ages 61–63 are statistically indistinguishable from zero, whereas coefficients turn positive exactly at age 66. The absence of differential pre-trends, together with the sharp post-eligibility divergence, supports the validity of the parallel-trends assumption.

Lastly, we rule out that the reform causes endogenous sample selection in the group of women in our sample. Table A6 shows that the reform does not affect women's probability of entering the analytical sample, which includes women who are currently married or in a union, living with their partner, and not receiving a contributory pension.

Compositional Changes in the Sample We also verify that the reform does not lead to compositional changes in our sample, for example, due to endogenous sample selection caused by an effect of the reform on mortality. Specifically, Appendix Figure A4 shows the effect of the reform on various predetermined characteristics of female respondents and their partners, such as educational attainment, the number of children, language, having experienced violence as a child, women's age at marriage, and their partner's age. The graph reports the coefficient for the short-term effect, i.e., the impact on women aged 66 to 67. Notably, all coefficients are small and statistically insignificant.[22] This result suggests that our results cannot be driven by compositional changes affecting the treatment and control group differently after the reform.

Anticipation Effects A potential concern is that anticipation of future eligibility for the non-contributory pension may affect younger women in the control group, influencing their risk of experiencing IPV and biasing our estimates. If anticipation increases IPV among control women, our estimates would understate the true effect. More problematically, if anticipation reduces IPV in the control group, it could make IPV appear to rise among treated women, overstating the effect even if their actual IPV rates remain unchanged.

Two analyses help us rule out the possibility that anticipatory changes in IPV occur for women in the control group. First, we use a placebo test to examine the anticipation effects of the reform. To this aim, we use age 60 as the placebo eligibility age and compare respondents in our control group (women aged 61 to 64) with younger respondents who are further away from age 65 and are less likely to change their behavior in anticipation of the reform: those aged 56-59. Table A7 reports the results. The estimates of the effect on IPV are very small and statistically insignificant, suggesting the lack of anticipation effects in IPV among the control group.

[22] Moreover, when we regress the main explanatory variable of interest (the interaction between the "Age 66–69" and "2016 Survey Wave" indicators) on all pre-determined characteristics and compute the F-statistic for their joint significance, we obtain a value of 0.464. This indicates that the pre-determined characteristics are not jointly significantly different between women in the treatment and control groups, before or after the reform.

Second, the event study by age (Figure A3 and Appendix Table A5) also supports the lack of anticipation effects. It examines whether IPV changes differently for each age group in the treatment and control groups after the reform, relative to the pre-reform period. The event study shows that changes in IPV for women aged 61, 62, and 63 in the 2016 wave, relative to the 2006 and 2011 waves, are similar to those observed for women aged 64 (the reference age group). If we expect the anticipation effect to be larger as women get closer to the eligibility age, these results suggest a lack of anticipation effects.

Alternative Specifications and Sample Restrictions Finally, several exercises further establish the robustness of the estimates. Table A8 shows that our main findings are robust to using numerous alternative specifications: choice of fixed effects (Columns 2 and 3), controlling for fewer household characteristics (Column 4), controlling for partner's age fixed effects (Column 5), adding state-specific age fixed effects (Column 6), adding state-specific survey-year fixed effects (Column 7), weighting the regressions using survey weights (Column 8), including women from the four excluded states into the sample (Column 9), including women that are receiving a contributory pension (Column 10), dropping survey wave 2006 (Column 11), and using the doubly-robust estimator by Sant'Anna and Zhao (2020) (Column 12). We describe these robustness exercises and their results in detail in Appendix B.

5.2.3 Heterogeneous Effects

Next, we investigate whether the reform has differential effects on specific groups of women based on several characteristics. Figure 2 reports the results of the heterogeneity analysis for women aged 66 and 67, and Appendix Figure A5 displays the results for women aged 68 to 69.

The purpose of the subgroup analysis is three-fold. First, we aim to identify the groups of women who are more likely to experience IPV after the reform. To this end, we examine two factors associated with a higher risk of IPV *ex-ante*: having been exposed to violence in childhood and low household wealth. The importance of intergenerational transmission of IPV is well-documented in various contexts, including Mexico (Kalmuss, 1984; Pollak, 2004; Jeyaseelan et al., 2007; Sánchez Argüelles, 2018). In our sample, 48.34 percent of women reported experiencing physical or verbal abuse from people they lived with during childhood or witnessing such abusive behavior among family members, and before the reform, this group of women was 18.3 percentage points more likely to have experienced IPV in the last 12 months compared to women who were not exposed to violence in childhood. Women in poorer households also have a higher baseline IPV risk (Aizer, 2010). We measure household wealth using an asset index that captures the living standards by considering home infrastructure, access to utilities, and asset ownership.[23] In our sample, older women in households with a low asset index were 3 percentage points more likely to have experienced physical or sexual IPV in the last 12 months than women in households with a high asset index before 2013. In contrast, household assets are not a significant factor for other types of IPV in the context under study. Figure 2 reveals that the rise in the incidence of physical or sexual IPV for women aged 66 and 67 is primarily driven by those in poorer households. In contrast, the increase in economic and psychological violence for women aged 66 to 67 is larger for women who experienced family violence during childhood. In figure A5, we find no evidence that the effect of the reform on IPV for women aged 68 to 69 varies with exposure to family violence in childhood or household wealth.[24]

Second, we attempt to uncover mitigating or aggravating factors that may protect women or enable their abusers: the presence of other family members in the home. Other family members may lower the risk of IPV by protecting from abuse, or may increase the risk of IPV by enabling the abuser. In our sample, 66.2 percent of women aged 66 to 69 before 2013 lived with other family members, including children (56%), grandchildren (27.9%), siblings (0.9%), and parents (1.33%).[25] Figures 2 and A5 show that co-residence with other family members does not impact the reform's effects on IPV.

Lastly, we explore heterogeneity by the woman's and her partner's educational attainment to understand the importance of outside options and bargaining power, as women's education is correlated with economic opportunities (Quisumbing and Maluccio, 2003; Doss, 2013; Hidrobo and Fernald, 2013; Heath, 2014) and men's education has been shown to affect IPV.[26] Additionally, previous research has shown that the effect of women's eligibility for cash transfers on IPV varies with their educational attainment (Hidrobo and Fernald, 2013) or the educational attainment of their husbands (Angelucci, 2008). We find no evidence that the impact of the reform on IPV varies with the woman's education. However, their husbands' educational attainment does seem to matter. We find that the impact on physical or sexual IPV for women aged 66 to 67 is larger for those married to a man who has less than a primary education. This result is consistent with the predictions of the male backlash theory, which predicts a larger increase in IPV in couples where men have a lower status relative to their wives (Heath, 2014) or more traditional attitudes toward gender

[23] See Appendix A for details on how we construct the asset index and the binary indicators for high and low wealth index.

[24] The results of the heterogeneity analysis by childhood violence and poverty for the age group 66 to 69 are qualitatively similar to the estimates for the age group 66 to 67 but less precisely estimated.

[25] While there could be a concern that the reform may affect co-residence patterns, in Section 6, we find limited evidence that this is the case.

[26] For example, using a natural experiment to establish causality, Özer et al. (2023) find that the husband's education decreases physical, economic, and psychological violence against violence in Turkey.

norms (Angelucci, 2008). It is also consistent with the findings of a previous study on Progresa, which found that drunken IPV increased among beneficiaries entitled to large transfers and whose husbands had no education, possibly because of male backlash (Angelucci, 2008).[27] However, it is important to point out that older cohorts have low variability in education level. In our sample, men and women aged 66 to 69 have a low educational attainment overall: 52% of women and 48% of their partners did not complete primary education, with the rest having mostly primary education.

There may be a concern that the heterogeneous results are driven by different program take-up rates among different sub-groups in the population. Appendix Figure A6 presents estimates of the effect of the 2013 pension reform on the probability of receiving a monetary transfer from the government for the different subgroups analyzed in Figures 2 and A5. We observe a relatively consistent program take-up among women aged 66-67 across different subgroups, except for higher take-up rates among women with less than primary education compared to women with at least primary education. However, we did not find that the impact of the reform on IPV differs along this dimension. Therefore, the diversity in the reform's impact on IPV is not attributable to differing adoption rates among these sub-groups.

5.3 Impact of Men's Eligibility on IPV

So far, our analysis has focused on the impact of women's eligibility for the non-contributory pension on IPV. However, men aged 65 and older are also eligible for the program after the 2013 reform. The ENDIREH survey lacks information on whether the partners of female respondents receive a pension, either contributory or non-contributory. As a result, we cannot estimate directly the take-up rate for men using the ENDIREH data. However, we can obtain precise estimates from the MCS-ENIGH dataset. The MCS-ENIGH data show that married men aged 66-69 who do not receive a contributory pension have a take-up rate of 54.5 percentage points, which is very similar to the estimated take-up rate of 57.6 percentage points among married women of the same age (Appendix Table A2). In this section, we examine whether women's experiences of IPV change when their husbands become eligible for the non-contributory pension program.

To study the impact of men's eligibility, we analyze a sample of women currently married or in a union with men aged 61-64 and 66-69. Husbands aged 61-64 are not eligible for the pension at the time of the interview, while husbands aged 66-69 become eligible after 2013 (2016 survey wave). We exclude women who are married to men aged 65 to mirror the analysis of women's eligibility. We obtain the *intent-to-treat* estimates of the effect of the husband's eligibility for the

[27] As shown in A5, there are no heterogeneous effects by partner's education for women aged 68 to 69. For the pooled age group 66 to 69, the estimates of the heterogeneity analysis by husband's education are qualitatively similar to the estimates for the age group 66-67, but are not precisely estimated.

pension on IPV by estimating a regression similar to Equation 1, where we replace $Age66-69$ with $HusbandAge66-69$ and the woman's age-fixed effects with the husband's age-fixed effects. Additionally, we replace the partner's age with the woman's age in the control variables X'_{iat}.

To isolate the impact of men's eligibility from the effect of women's eligibility, we restrict the sample to women who are not yet eligible for the pension based on their age at the time of the survey (64 or younger). Table 2 reports the results. The odd-numbered columns report the effects on women married to men aged 66 to 69, and the even-numbered columns report the effects on women married to men aged 66 to 67 and women married to men aged 68 to 69 separately.

Looking at women married to men aged 66 to 69, we find no evidence that the reform affected their probability of being subjected to IPV. The coefficient estimates are small and statistically insignificant for all types of IPV. When we look at women aged 66 to 67 and 68 to 69 separately, the coefficient estimates of the effect of the reform on "any IPV" and "psychological IPV" are bigger in magnitude for the age group 68 to 69 but remain statistically insignificant.

In sum, we have shown that the probability that women are victims of IPV is unchanged after their husbands become eligible for the pension. One potential limitation of this analysis is that, differently from the analysis of women's eligibility, we cannot exclude women married to men who receive contributory pensions from our empirical analysis due to the lack of information on husbands' pensions in the ENDIREH survey. Men are considerably more likely than women to receive contributory pensions; data from MCS-ENIGH indicate that 40.84% of men aged 66 to 69 are enrolled in a contributory pension program, compared to just 9.2% of women. Consequently, the portion of men receiving non-contributory pensions in the empirical analysis based on ENDIREH is likely much lower than that of women. The absence of a discernible effect may be due to lower take-up rates among men.

6 Mechanisms and Discussions

6.1 Potential Mechanisms

This section discusses the channels through which a woman's eligibility for a social pension may affect IPV among older couples and tests the various mechanisms empirically. First, we discuss three mechanisms that predict an increase in IPV after women gain eligibility for the pension: instrumental violence, increased exposure and male backlash; second, we discuss two mechanisms that predict a decrease in IPV after women become eligible: an increase in women's bargaining power and stress reduction; third, we analyze changes in co-residence patterns. Finally, we compare our results with the findings of previous research on cash transfer on IPV,[28] with a particular attention to how the various mechanisms may operate differently for older couples relative to younger couples.

Instrumental Violence In economic models where violence is viewed as *instrumental* or *extractive*, men use violence to “extract rents”, control women’s behavior or appropriate their resources (Bloch and Rao, 2002; Eswaran and Malhotra, 2011; Bobonis et al., 2013; Heath, 2014; Haushofer et al., 2019; Calvi and Keskar, 2023). In this framework, becoming eligible for a pension puts women at increased risk of IPV. Men may exercise violence to influence how their partners use their pension income, take their money, or control their behavior, which may have changed after the increase in financial autonomy provided by the pension.

We analyze this mechanism in two ways. First, since the instrumental theory of violence postulates that one of the perpetrator’s motives to utilize IPV is to extract rents from their partners, we exploit the granular information on the various forms of IPV in the ENDIREH survey to test whether the reform only increases psychological and physical/sexual IPV when economic IPV is also present. To this aim, we first examine the impact of becoming eligible for the non-contributory pension on the probability of experiencing economic IPV alone or in combination with other types of violence (Columns (1) to (8) of Table 3). We find that women do not experience an increase in economic IPV alone after gaining eligibility. In contrast, the probability of experiencing economic IPV and psychological IPV jointly increases by 37%, while the impact of eligibility on the likelihood of experiencing economic IPV jointly with physical or sexual IPV is mostly imprecisely estimated. We interpret these findings as suggestive evidence that psychological IPV, but not physical/sexual IPV, may be used as a tool to extract women’s new economic resource. Having established that women’s eligibility increases the probability of psychological IPV in conjunction with economic IPV, we next examine whether the probability of experiencing psychological or physical/sexual IPV without being subjected to economic IPV also increases after the reform. Columns (9) and (10) of Table 3 show that it is not the case. When economic IPV is not present, we find that the impact of women’s eligibility on the probability of experiencing any other type of IPV is statistically insignificant and relatively small (about 10% of the sample mean). Altogether, these findings suggest that the extraction and control of newly found women’s economic resources are drivers of the estimated increase in IPV.

Second, we investigate what specific types of violent acts drive the estimated increase in IPV by looking at the disaggregated subcategories defined by the National Statistical Institute (Insti-

[28] We refer to Buller et al. (2018) and Baranov et al. (2021) for a comprehensive overview of the pathways through which cash transfer programs may affect IPV.

tuto Nacional de Estadística, Geografía e Informática—INEGI).[29] Table A9 reports the results. For physical or sexual IPV, we look at moderate physical violence, severe physical violence, and sexual violence separately and find no increases in any of these types of IPV (Panel A). We disaggregate economic IPV into three sub-categories: theft or economic coercion, economic control or blackmail, and failure to fulfill economic responsibility (Panel B). We find a statistically significant increase in the likelihood that a woman's partner fails to provide economic support and uses economic control against her, for example, by complaining about how she spends money. Lastly, we divide psychological IPV into five types (Panel C): indifference, degradation, intimidation, isolation, and threat. We find a statistically significant increase in the probability that women experience indifference, degradation, and intimidation from their partners. These behaviors include, for example, shaming, humiliating, ignoring, throwing away her belongings, and spying on her. Overall, these findings suggestively corroborate the interpretation that when a woman becomes eligible for the pension, her partner uses moderate physical violence and forms of psychological abuse to control her newly found economic resources and decrease his economic contributions to the household budget. However, we interpret these results cautiously because we are testing many hypotheses in Table A9 and none of the coefficients is statistically significant after adjusting for multiple hypothesis testing.[30]

Impact on Labor Supply and Exposure The *exposure* theory posits that a woman's risk of IPV increases as she spends more time with her partner (Dugan et al., 2003). Previous studies have demonstrated the relevance of this mechanism in the context of unemployment in Brazil, revealing that job loss among both men and women contributes significantly to the increase in IPV (Bhalotra et al., 2021). In the context of our study, a woman's eligibility for a social pension could potentially lead to an increase in IPV if she reduces her labor supply and consequently spends more time in the presence of her potential abuser. Qualitative studies also highlight retirement as a key factor for triggering IPV among older couples (Pathak et al., 2019).

We analyze this mechanism by estimating the impact of women's eligibility on labor market outcomes in Table 4. Columns (1) and (2) show estimates of the effect of women's eligibility for the pension on women's paid employment; in Columns (3) and (4), the outcome variable is a binary indicator that equals one if the woman asserts that she does not work because she is dedicated to household work; in Columns (5) and (6) the dependent variable is a binary indicator that equals one if the woman's husband works for a pay.

The results indicate that eligible women are 21% less likely to work for pay after the reform.

[29] We follow the definitions in Ramírez (2007). See Appendix Table A1 for definitions of the variables.

[30] We computed the sharpened q-values for the coefficient estimates in Table A9 and found that they are always above 0.1. These q-values are not reported in the table.

Additionally, the probability that they don't work because they are committed to household work increases by 6% after they gain eligibility for the non-contributory pension. These findings suggest that women's exposure to IPV risk may increase as a function of more time spent at home and being more exposed to violent partners.

These results should be interpreted with two caveats. First, we estimate an increase in IPV jointly with a reduction in paid work after the reform, and we cannot determine from the data which variable changes first. Given previous research findings that IPV reduces victims' employment and earnings (Bindler and Ketel, 2022; Bhuller et al., 2024), there may be a concern that women's eligibility for the pension first leads to an increase in IPV, which in turn reduces women's labor supply. Our data do not allow us to dismiss this possibility entirely. Second, because information on the time spent with partners is not available in the ENDIREH data set, we infer an increase in exposure from an observed decrease in labor supply. However, it is important to underscore that the estimated reduction in paid employment and the increase in the probability of being dedicated to household work are merely suggestive indicators that exposure to the perpetrator may increase.

Male Backlash/Status Inconsistency The observed increase in IPV after women gain eligibility for the non-contributory pension is also consistent with the sociological theory of *male backlash/status inconsistency*. In contexts where the traditional gender norm of the male breadwinner is prevalent, men may feel that an increase in women's income threatens their leadership role within the family and may use IPV to restore their status (Macmillan and Gartner, 1999; Erten and Keskin, 2018). The backlash response to a program that increases women's economic opportunities is potentially stronger in couples where men have a lower status relative to their wives (Heath, 2014; Angelucci, 2008) or more traditional attitudes toward gender norms (Angelucci, 2008). In this regard, the finding of the heterogeneity analysis that the reform increases physical or sexual IPV more for women married to a man who has less than primary education is potentially suggestive of male backlash for this subgroup of women.

To further examine this mechanism empirically, we differentiate between newly eligible women whose husbands' eligibility changes at the same time as theirs (i.e. couples without relative status inconsistency) and newly eligible women whose husbands' eligibility stays unchanged (i.e. couples with relative status inconsistency). With this aim, we estimate the impact of the 2013 reform separately for women married to men aged 66 to 69 years, who are eligible in 2016, and women married to men older than 69, who were already eligible before the reform, or younger than 65, who are not yet eligible in 2016. Figure 3 displays the results obtained by estimating the model separately for women aged 66 to 67 (Panel a) and women aged 68 to 69 (Panel b). For women aged 66 to 67 years, we find that the impact of the reform is greater for women married to men whose eligibility status is unchanged between 2011 and 2016. Among these couples, the reform increases the economic resources of the women relative to their husbands. The estimated coefficient for these women is positive and statistically significant, while the coefficient estimate for women married to men who also gain eligibility (aged 66 to 69) is zero and imprecisely estimated. The differential impacts of the reform on psychological abuse are statistically different at the 1 percent level. For older women aged 68 and 69 at the time of the survey, we find no evidence of heterogeneity by husband's age.

In sum, we find that psychological IPV increases only in couples where the wife gains eligibility for the pension while her husband does not, either because they were already eligible or because they are not yet eligible. This result is consistent with a male backlash response to the change in husbands' status relative to their wives, where men resort to psychological violence to restore their status.[31]

Women's Bargaining Power and Threat Point In the household bargaining model, an increase in the woman's potential income relative to her husband's improves her outside options, increasing her bargaining power and reducing IPV (Farmer and Tiefenthaler, 1997; Aizer, 2010). We hypothesize that this channel is less relevant in our context for a few reasons. First, women's bargaining power has been shown to decline steadily at post-reproductive ages (Calvi, 2020), and women's eligibility for the pension may not be sufficient to make the threat of leaving the marriage credible after age 65. Second, the threat of divorce or a non-cooperative equilibrium must be credible for this mechanism to be relevant, and recent data from Mexico indicate that it may not be the case for older women. Hoehn-Velasco et al. (2023) report the average percentage of divorces by age calculated using INEGI marriage and divorce microdata from March until December 2019. They show that the percentage of wives who divorce declines steeply after age 44, and for wives aged 65 and over, it is only 1.88%.

Finally, we test the hypothesis that the reform improves the outside options of eligible women empirically using two analyses. First, we study the impact of eligibility on women's decision-making power within the household, which we measure using a summary index (see Appendix A for details). We find very small and statistically insignificant changes in women's decision-making power within the household after women become eligible for the pension (Columns 13 and 14 of Appendix Table A10). Second, we study whether the impact of the reform on IPV differs in states where unilateral divorce law is allowed at the time of the survey, following the classification of states used by García-Ramos (2021). According to the household bargaining model, credible

[31] We also do a similar exercise for the impact of men gaining eligibility in Figure A7. We do not find any differences in the impact of the reform on IPV between men married to women who also gain eligibility and men married to women whose eligibility status remains unchanged.

threats of divorce might reduce IPV. If this channel is important, we would expect the effect of the reform on IPV to be muted in states that passed a unilateral divorce law by 2016. Figure A8 compares the estimates. We do not find evidence that a unilateral divorce law matters for the effect of the reform on IPV. If anything, states without unilateral divorce laws have a stronger increase in IPV, which is consistent with the household bargaining channel, but the difference is statistically insignificant. As the divorce rate among older women is very low, it is not surprising that this outside option is not credible. Overall, the evidence supports the hypothesis that the bargaining channel is weak in the context of our study.

Impact on Household Income and Stress Reduction According to the *absolute resource theory* and the *stress theory*, an increase in total household income increases the resources available to the household and reduces stress, which in turn decreases the risk of IPV (Fox et al., 2002; Buller et al., 2018; Heath et al., 2020). By relaxing the household budget, a monetary transfer may also decrease IPV by reducing conflicts arising from scarce household resources (Buller et al., 2018). We explore these mechanisms by estimating the impact of eligibility for the pension on total household monetary income, net of transfers from the government and family members, in Appendix Table A11 using the MCS-ENIGH sample. In Columns (1)-(2), we consider the impact of women's eligibility for the pension on household monetary income using a logarithmic specification, while in Columns (3)-(4), we consider the impact of men's eligibility. We find that household monetary income increases by 7.5% (12.5%) after women (men) become eligible for the non-contributory pension, and the coefficient is statistically significant at the 5% level.[32] However, the absolute resources theory and the stress reduction theory both predict that IPV *decreases* due to increased household income, while we find that women's eligibility for the pension *increases* IPV. This result suggests that other mechanisms may be at play to countervail the potential reduction in IPV coming from the increase in household income.

Impact on Co-residence Patterns Lastly, we investigate the reform's effect on living arrangements. We create six binary indicators to measure the various co-residence patterns observed in the data: Respondents who live in a household with more than two household members (Columns 1 and 2); respondents who live with their parents (Columns 3 and 4); respondents who live with

[32]Behavioral responses from the respondents and other family members may attenuate the effect of women's eligibility for the pension on total household income. We have already shown that women reduce paid employment after gaining eligibility (Table 4). Previous research on the "70 y más" program has found that the non-contributory pension crowds out intra-family transfers to the elderly (Amuedo-Dorantes and Juarez, 2015). Using ENDIREH, in Columns (7) and (8) of Table 4, we show that a woman's eligibility for the pension reduces the probability of receiving transfers from relatives by 1.6 percentage points, or 7% relative to the sample mean. Consistent with previous research findings, this result suggests that the relatives of the women who become eligible for the pension reduce monetary transfers to them after the reform. In Columns (9) and (10) of Table 4, we find no change in the probability of being a Progresa beneficiary.

their children (Columns 5 and 6); respondents who live with their grandchildren (Columns 7 and 8); respondents who live with their siblings (Columns 9 and 10); and respondents who live with other individuals (Columns 11 and 12). Appendix Table A10 displays the results. We find no evidence that the reform changes co-residence patterns. The only exception is a 36% decrease in the probability that siblings live in the house, which is statistically significant at the 10% level, for women aged 66 to 67.

To summarize, our empirical investigation into how a woman's pension eligibility may influence IPV identifies four key mechanisms: the use of violence as an instrument to control women's resources, greater exposure to the risk of abuse due to increased time spent at home, male backlash, and increased household income. Given the estimated increase in IPV due to the reform, we infer that the rise in IPV predicted by instrumental, exposure, and backlash theories offsets the potential decline in IPV suggested by the absolute resource theory.

6.2 Discussion and Comparison with Previous Studies

Two meta-analyses concluded that the evidence from available studies points toward a decrease in IPV on average after women receive cash transfers (Baranov et al., 2021; Buller et al., 2018). In contrast, we find that women's eligibility for a social pension increases IPV. In this section, we discuss potential reasons why our results differ from the findings of most existing studies on cash transfers and IPV.

First, the disparate findings may be explained by our focus on older women, whose risk and experiences of IPV may differ from those of younger women in several ways. For example, older victims are more likely to have disabilities and depend financially on their partners or other family members, making it harder to recognize abuse, seek help, and threaten to leave an abusive relationship. They may also face more stigma due to gender roles being more conservative among older generations.

Second, our findings align with previous research showing that cash transfers increase emotional IPV in key subgroups of women. Hidrobo and Fernald (2013) find that an unconditional cash transfer reduces psychological IPV for women with more than primary school education in Ecuador but increases it for women with primary education or less who do not have less education than their husbands, possibly because the outside options are not a credible threat for this group of women. This finding is consistent with our results, as older women have lower outside options, as suggested by their low divorce and separation rates (Hoehn-Velasco et al., 2023).

Third, a few studies that found a decrease in IPV after women receive cash transfers show that the increased decision-making power of women beneficiaries is one of the mechanisms driving the decline in IPV (Buller et al., 2016; Hidrobo et al., 2016; Ritter Burga, 2014). In Section 6.1, we argued that the household bargaining framework does not apply to our context because the affected older women in our setting have a low propensity to divorce. This assumption is supported by the null findings on the reform's effect on decision-making. Despite increasing the economic resources available to women, the reform fails to significantly improve their decision-making power as predicted by the household bargaining model (Appendix Table A10).

Fourth, regarding the other theoretical channels discussed in Section 6, the predictions of the instrumental theory, the male backlash theory, and the exposure theory are consistent with larger increases in IPV among older women than younger women. Some of the instrumental models of violence predict that a program that provides women with income or work opportunities may increase IPV for women with low *ex-ante* outside options (Eswaran and Malhotra, 2011; Heath, 2014), as is the case for older women. Regarding male backlash, if gender norms are more conservative among older couples than younger couples, older women may be at higher risk of male backlash after becoming eligible for cash transfers than younger women. Looking at the exposure channel, a non-contributory pension may generate a larger income effect on labor supply than a cash transfer for younger women, as it targets older women close to the common retirement age.

Next, we discuss how our results relate to the results of past studies on the determinants of IPV in Mexico. The evidence on the conditional cash transfer Progresa/Oportunidades and IPV is mixed. Angelucci (2008) finds that the impact of Progresa on a specific type of IPV — aggressive behavior after drinking — varies depending on the size of the transfer and the educational attainment of the husband. She estimates a decrease in drunken IPV among women entitled to the minimum transfer and women whose husbands completed primary school. In contrast, she finds an increase in drunken IPV among women entitled to large transfers if their husband has no education. Bobonis et al. (2013) show that physical IPV decreases, but psychological IPV increases, including threats of physical IPV without associated physical IPV, among Progresa beneficiaries in the short run. Bobonis et al. (2013) interpret their results as suggestive that partners use psychological IPV as an instrument to extract rent from their female partners. However, Bobonis et al. (2025) find no effect of Progresa on IPV in the long run.

Turning to other determinants of IPV, García-Ramos (2021) studies the role of unilateral divorce using the enactment of the policy by some Mexican states as a natural experiment. She finds that easier access to divorce increases physical, psychological, and economic IPV in the long run, and interprets this result as suggestive that partners use IPV as an instrument to prevent women from leaving the marriage.

In sum, the findings of previous research on conditional cash transfers and unilateral divorce suggest that IPV tends to be instrumental in Mexico, which may be due to the high prevalence of IPV throughout the country. Our finding that women's pension eligibility increases IPV also suggests that men use violence as a tool to control women's resources.

7 Conclusion

Although gender-based violence remains a prevalent global issue, there is limited understanding of this issue among women of post-reproductive age. This study sheds light on the prevalence of IPV among older women and provides the first estimates of the effect of old-age pensions on IPV. Using detailed survey data from Mexico, we first document that although IPV decreases with age, older women still face significant risks. Second, we investigate the effect of a Mexican pension reform that expanded the coverage of a non-contributory program to individuals aged 65 and older using a DID approach. We show that the reform substantially increases the program take-up among men and women aged 66 to 69, but the effects on IPV vary with the gender and age of the recipient. We find that women's eligibility for the pension increases the probability of experiencing physical or sexual IPV, economic IPV, and psychological IPV, compared to younger women who are not yet eligible. For all three types of violence, the estimated effects are bigger for women aged 66 to 67 than for women aged 68 to 69. In contrast, men's eligibility for the pension does not affect the probability that their wives are subjected to IPV. Further analysis suggests that the use of violence as a tool to extract or control women's newly found economic resources may play a role in explaining the increase in IPV after women become eligible for the pension. Additionally, we find that the effect of women's eligibility on IPV is greater for women whose husbands do not become eligible, which is consistent with male backlash. We also show that women reduce their paid employment after becoming eligible, suggesting a potential increase in time spent at home and greater exposure to abusive partners.

The results of this paper show the unintended consequences of a social protection program that has been shown to reduce extreme poverty among the elderly in Mexico (Ávila-Parra et al., 2024). Our findings help identify the groups of women who are at higher risk of experiencing IPV after becoming eligible for a social pension: women who experienced family violence in childhood, women with low socioeconomic status, and women whose husbands have low levels of education. To mitigate the risk of increased IPV, the government may target interventions for this high-risk population. Numerous studies conducted in Asia and Sub-Saharan Africa suggest that community-level and group-based programs designed to combat IPV, such as gender training, couples' dialogue, and family coaching, are effective in reducing IPV (Chang et al., 2020; Leight et al., 2023). However, more research is needed to examine whether these anti-IPV interventions are effective in reducing IPV in women of post-reproductive age, for a couple of reasons. First, most of these studies include only men and women of reproductive age. Second, in terms of pathways, some of these interventions reduced physical and/or sexual IPV while also changing attitudes toward IPV, suggesting that a shift in attitudes may be a channel. However, it may be more difficult to influence attitudes in elderly couples as attitudes may be less malleable in old age (Glenn, 1980; Krosnick and Alwin, 1989).

This study contributes to a better understanding of the complex relationship between income, social programs, and IPV for older women. It highlights the importance of considering age-specific factors when implementing policy responses to address IPV and emphasizes the need for more research on this vulnerable population. In particular, it is imperative to build more evidence on the forms of abuse that are specific to old age, as well as documenting abuse by family members other than the intimate partner, including children (WHO, 2024).

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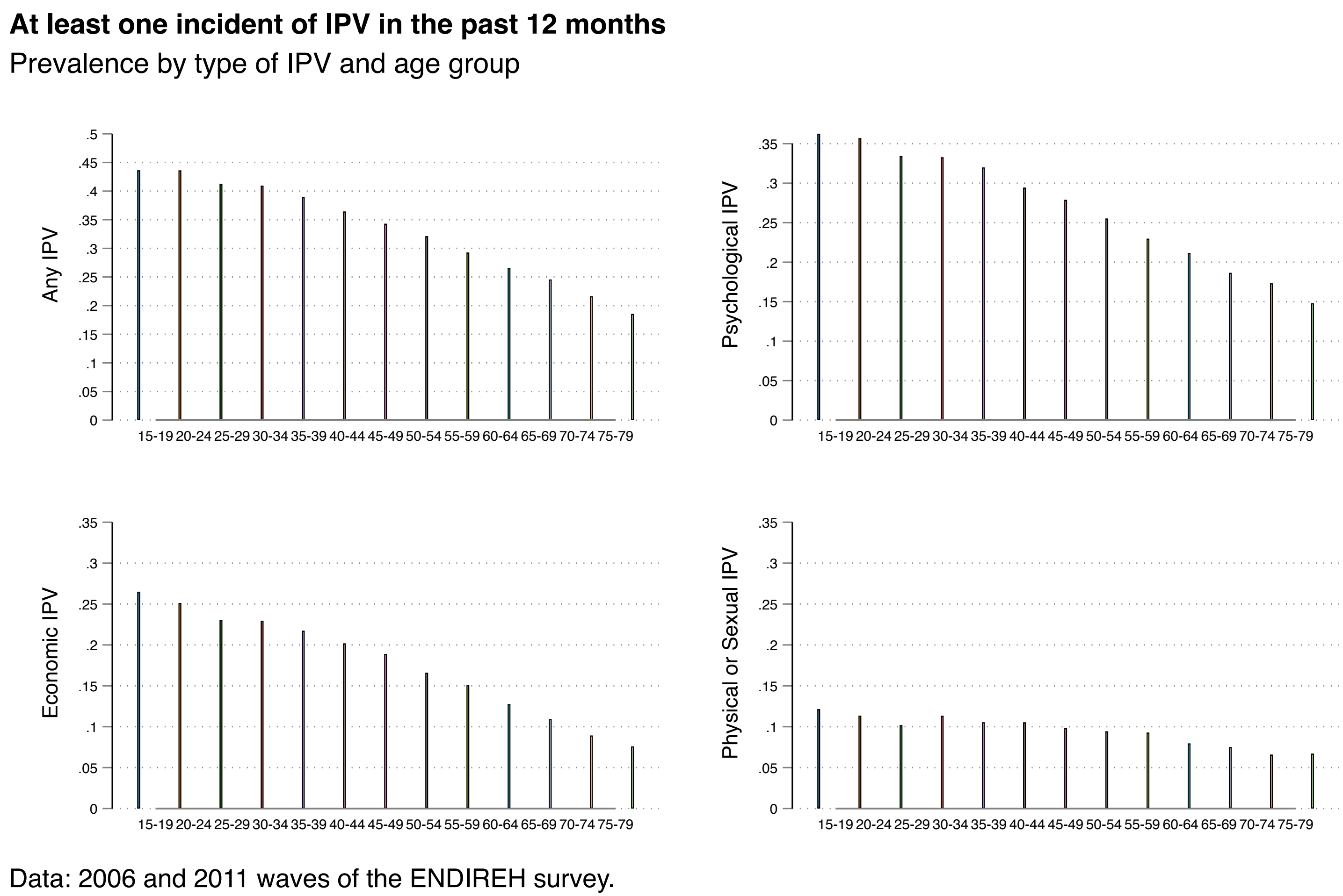
Figures and Tables

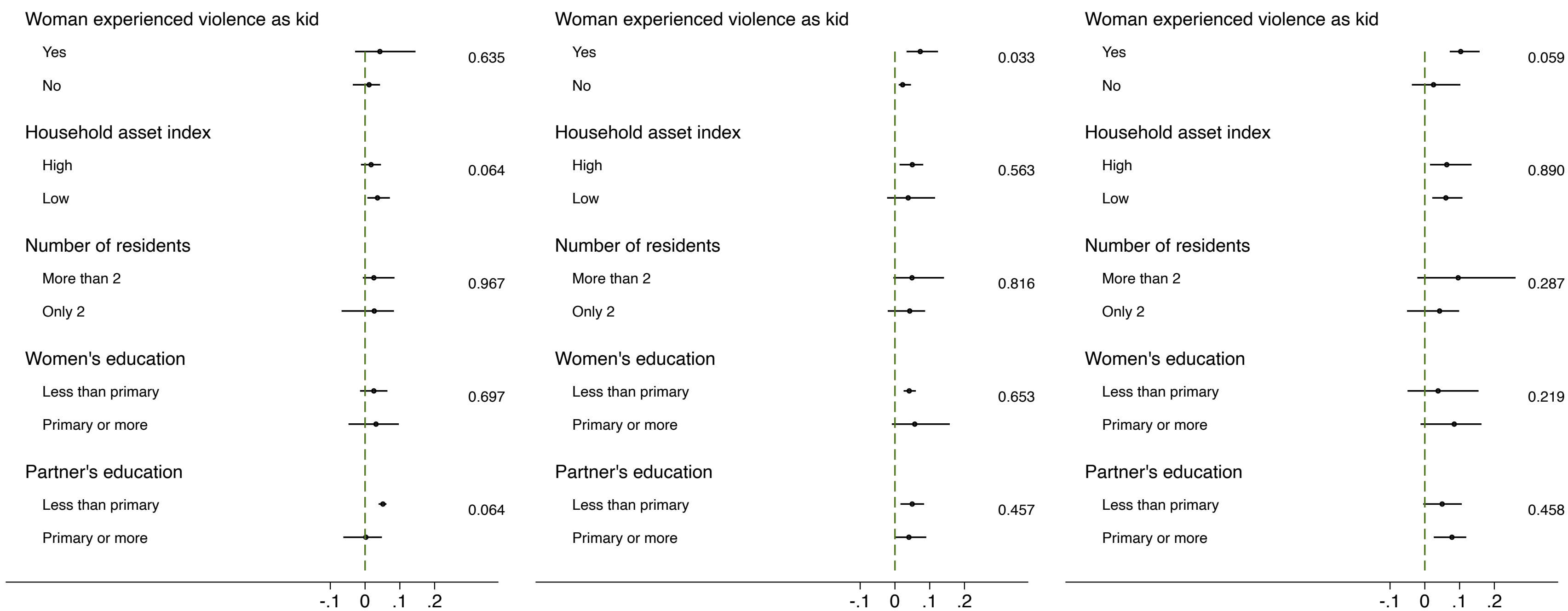
Figure 1: Descriptive Statistics: The Age Gradient in Intimate-Partner-Violence Before the Reform



Notes: These figures show the incidence of intimate partner violence for women aged 15 to 79 who are currently married or in a union, live with their partners and do not receive a contributory pension in the 2006 and 2011 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that introduced similar reforms between 2006 and 2011) are excluded. The respondents are classified as victims of violence if they experienced an act of psychological, physical or sexual, or economic violence by their partner at least once in the last 12 months prior to the interview.

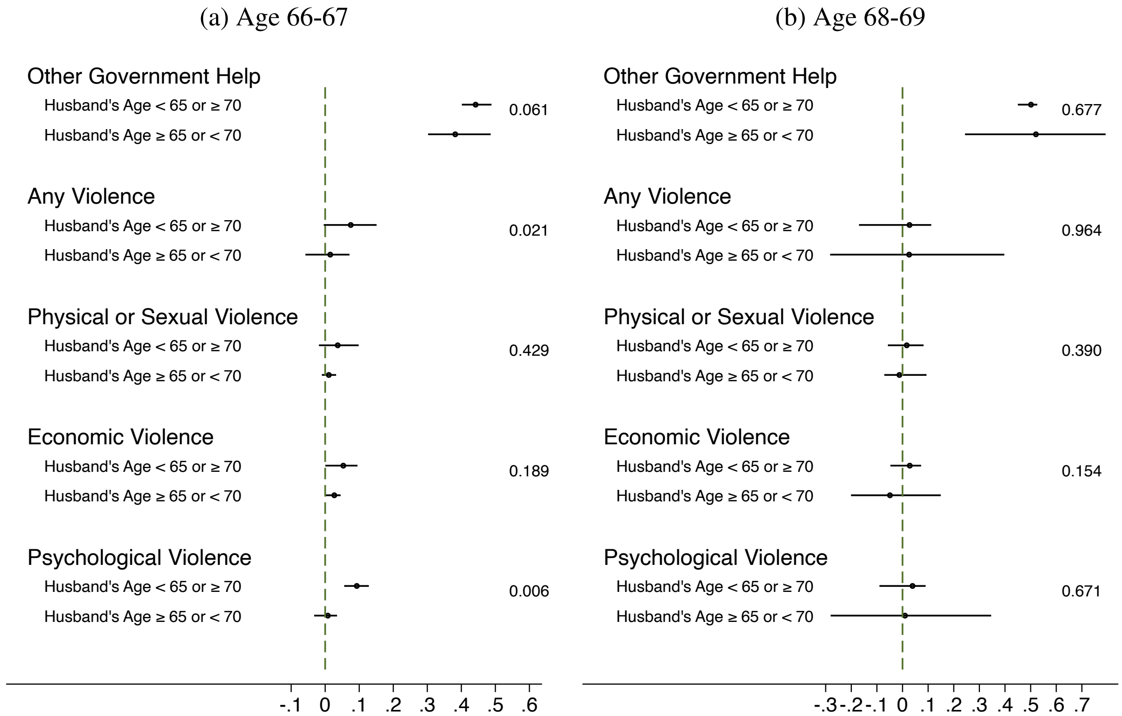
Figure 2: Heterogeneity: Impact of Eligibility on IPV for Women aged 66-67

(a) Physical or Sexual Violence (b) Economic Violence (c) Psychological Violence



Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. For each heterogeneity variable, the figure reports the coefficient estimate of the difference-in-differences coefficient, i.e., the interaction between "Age 66-67" and the "2016 Survey wave" indicators, for stratified regressions on the split samples. The figure shows the 95 percent confidence intervals estimated using Wild Cluster Bootstrap and the p-value for the test of equality of coefficients in a fully interacted regression.

Figure 3: Mechanisms: Impact of Eligibility on IPV by Husband's Age



Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Figure (a) reports the coefficient estimate of the interaction between "Age 66-67" and the "2016 Survey wave" indicators, and Figure (b) reports the coefficient estimate of the interaction between "Age 68-69" and the "2016 Survey wave" indicators. The figures show the 95 percent confidence intervals estimated using Wild Cluster Bootstrap and the p-value for the test of equality of coefficients in a fully interacted regression.

Table 1: Impact of Eligibility on the Probability of Receiving Government Aid and IPV in the Past 12 Months

	Receives Government Aid		Any Violence		Physical or Sexual Violence		Economic Violence		Psychological Violence	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Age 66-69 X Wave 2016	0.456***		0.041*		0.018*		0.029*		0.049**	
	(0.032)		(0.018)		(0.009)		(0.011)		(0.012)	
	[0.000]		[0.055]		[0.095]		[0.059]		[0.019]	
					{0.068}		{0.063}		{0.061}	
Age 66-67 X Wave 2016		0.416**		0.054*		0.027*		0.046**		0.063**
		(0.035)		(0.015)		(0.005)		(0.005)		(0.006)
		[0.012]		[0.069]		[0.052]		[0.013]		[0.014]
						{0.061}		{0.044}		{0.044}
Age 68-69 X Wave 2016		0.504***		0.025		0.008		0.009		0.032*
		(0.001)		(0.026)		(0.010)		(0.009)		(0.015)
		[0.000]		[0.535]		[0.550]		[0.514]		[0.057]
						{0.225}		{0.225}		{0.061}
Observations	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193
R^2	0.350	0.351	0.077	0.077	0.038	0.039	0.047	0.047	0.065	0.065
Mean Dep. Variable	0.029	0.029	0.231	0.231	0.075	0.075	0.100	0.100	0.170	0.170
P-value 66-67 vs. 68-69		0.206		0.512		0.175		0.053		0.151

Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. Benjamini et al. (2006) sharpened q-values are calculated following Anderson (2008) and are shown in curly brackets. For regressions with only one treated age group (66-69), the q-values are calculated for all 3 hypotheses (3 outcomes: physical or sexual violence, economic violence, and psychological violence). For regressions with two treated age groups (66-67 and 68-69), the q-values are calculated for all 6 hypotheses (3 outcomes and 2 treatments). *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 2: Impact of Husband's Eligibility on IPV

	Any Violence		Physical or Sexual Violence		Economic Violence		Psychological Violence	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age Partner 66-69 X Wave 2016	0.012		-0.001		-0.005		0.010	
	(0.014)		(0.007)		(0.010)		(0.013)	
	[0.529]		[0.861]		[0.628]		[0.466]	
Age Partner 66-67 X Wave 2016		-0.002		-0.001		-0.004		0.001
		(0.007)		(0.004)		(0.009)		(0.012)
		[0.857]		[0.828]		[0.726]		[0.959]
Age Partner 68-69 X Wave 2016		0.034		-0.001		-0.007		0.026
		(0.025)		(0.017)		(0.018)		(0.021)
		[0.116]		[0.868]		[0.748]		[0.294]
Observations	12,240	12,240	12,240	12,240	12,240	12,240	12,240	12,240
R^2	0.089	0.089	0.047	0.047	0.053	0.053	0.071	0.071
Mean Dep. Variable	0.257	0.257	0.075	0.075	0.131	0.131	0.203	0.203
P-value 66-67 vs. 68-69		0.315		0.993		0.878		0.490

Notes: The sample includes women younger than 65 who are married or in a union and live with men ages 61-64 and 66-69 in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for the husband's age-fixed effects, survey year-fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the husband's age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** p<0.01, ** p<0.05, * p<0.1.

Table 3: Mechanisms: Combinations with and without Economic Violence

	Experiencing IPV with Economic Violence								Experiencing IPV without Economic Violence	
	Only Economic		With Only Psychological		With Only Physical or Sexual		With Psychological and Physical or Sexual		Any Combination	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Age 66-69 X Wave 2016	-0.002		0.014*		0.002		0.014		0.012	
	(0.007)		(0.006)		(0.002)		(0.008)		(0.013)	
	[0.819]		[0.055]		[0.364]		[0.222]		[0.428]	
Age 66-67 X Wave 2016		-0.003		0.019**		0.004		0.026**		0.008
		(0.009)		(0.007)		(0.002)		(0.005)		(0.016)
		[0.748]		[0.029]		[0.146]		[0.033]		[0.748]
Age 68-69 X Wave 2016		-0.000		0.009		0.000		-0.000		0.017
		(0.009)		(0.007)		(0.002)		(0.006)		(0.017)
		[0.964]		[0.566]		[0.870]		[0.940]		[0.543]
Observations	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193
R^2	0.008	0.008	0.026	0.027	0.009	0.009	0.026	0.026	0.032	0.032
Mean Dep. Variable	0.030	0.030	0.038	0.038	0.003	0.003	0.031	0.031	0.130	0.130
P-value 66-67 vs. 68-69		0.859		0.516		0.403		0.050		0.742

Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** p<0.01, ** p<0.05, * p<0.1.

Table 4: Mechanisms: Impact of Eligibility on Labor Market Outcomes and Transfers

	Women's Labor Market				Husband's Labor Market		Transfers			
	Receives Money from Employment		Doesn't Work Because Dedicates to Housework		Partner Receives Money from Employment		Receives Monetary Help from Family		Receives Money from Progresa	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Age 66-69 X Wave 2016	-0.033*		0.045***		-0.001		-0.016*		0.020	
	(0.014)		(0.013)		(0.017)		(0.007)		(0.011)	
	[0.051]		[0.002]		[0.964]		[0.062]		[0.117]	
Age 66-67 X Wave 2016		-0.028*		0.033**		-0.019		-0.017*		0.019
		(0.009)		(0.004)		(0.021)		(0.009)		(0.014)
		[0.095]		[0.021]		[0.577]		[0.067]		[0.292]
Age 68-69 X Wave 2016		-0.040		0.059**		0.020		-0.014		0.022
		(0.029)		(0.027)		(0.011)		(0.007)		(0.011)
		[0.311]		[0.040]		[0.169]		[0.175]		[0.221]
Observations	10,185	10,185	10,193	10,193	10,193	10,193	10,193	10,193	10,192	10,192
R^2	0.043	0.043	0.047	0.047	0.086	0.086	0.083	0.083	0.265	0.265
Mean Dep. Variable	0.156	0.156	0.751	0.751	0.463	0.463	0.230	0.230	0.151	0.151
P-value 66-67 vs. 68-69		0.666		0.558		0.174		0.806		0.809

Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** p<0.01, ** p<0.05, * p<0.1.

Online Appendix

Social Pensions and Intimate Partner Violence against Older Women

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Appendix Content

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A Definition of Key Variables

A.1 Main Outcome Variables

A.1.1 Intimate Partner Violence

See Appendix Table A2 for the definition of the following key variables: *Any IPV, Physical or Sexual IPV, Economic IPV, Psychological IPV, Theft and coercion, Economic control, Failure to provide, Moderate physical violence, Severe physical violence, Indifference, Intimidation, Isolation, Threats, and Sexual violence*.

A.1.2 Other Outcome Variables

- *Receives government aid*: It is a binary indicator that equals one if the respondent reports receiving a government transfer other than Oportunidades/Progresa, at the time of the interview.
- *Receives money from employment*: It is a binary indicator that equals one if the respondent reports receiving money from employment, at the time of the interview.
- *Doesn't work because dedicates to housework*: Binary indicator that equals one if the respondent reports that she does not work because she is dedicated to doing housework, at the time of the interview.
- *Partner receives money from employment*: It is a binary indicator that equals one if the respondent reports that her partner is receiving money from employment, at the time of the interview.
- *Receives monetary help from family*: It is a binary indicator that equals one if the respondent reports receiving money from relatives or acquaintances who live abroad or within the country, at the time of the interview.
- *Receives money from Progresa*: It is a binary indicator that equals one if the respondent reports receiving money from the government program Oportunidades/Progresa, at the time of the interview.
- *Bargaining power index*: Index created by the authors to measure the female respondent's decision-making power within her household. ENDIREH collects information on household decision-making in seven different situations in women's lives: when to work, when to leave the house, expenditures, when to move, decisions about money, what to buy, and social life. We summarize these variables using an index. First, we create binary indicators that equal one if the respondent alone decides in each of these situations. Second, we calculate the standardized z-score of each binary indicator using the mean and standard deviations of the control group (women aged 61 to 64). Finally, we compute the summary index by taking the arithmetic average of the seven standardized z-scores.

A.2 Variables Used to Define the Sub-groups for the Heterogeneity Analysis

- *Woman experienced violence as a kid*: It is a binary indicator that equals one if the respondent reports that when she was a child, she was hit, insulted, or offended by family members, or witnessed family members hitting or insulting each other.
- *Household asset index*: Continuous variable created by the authors following García-Ramos (2021). We use the first component of principal component analysis of the following variables: binary indicators for having an earth floor, a cement floor, a wood floor, access to public water in the house, access to public water out of the house, other types of water access, a public drain system, a septic tank, other types of drain; binary indicators for owning a radio, a computer, a landline, a mobile phone, a washing machine, a car; a crowding index, defined as the number of residents divided by number of rooms.
  - *Low*: It is a binary indicator that equals one if the respondent's household asset index belongs to the first and second quintile of the index distribution in our sample.
  - *High*: It is a binary indicator that equals one if the respondent's household asset index belongs to the third, fourth, or fifth quintile of the index distribution in our sample.
- *Number of residents*: Total number of residents living in the same house as the respondent, at the time of the interview.
- *Women's education*:
  - *Less than primary*: It is a binary indicator that equals one if the respondent reports that she has no education or has some years of education but did not complete primary education.
  - *Primary or more*: It is a binary indicator that equals one if the respondent reports that she completed primary education, secondary education, or tertiary education.
- *Partner's education*:
  - *Less than primary*: It is a binary indicator that equals one if the respondent reports that her partner has no education or has some years of education but did not complete primary education.
  - *Primary or more*: It is a binary indicator that equals one if the respondent reports that her partner completed primary education, secondary education, or tertiary education.

A.3 Variables Used to Examine Sample Selection

- *Married/union*: It is a binary indicator that equals one if the respondent is married or cohabitates with a partner at the time of the interview.
- *Married/union with partner*: It is a binary indicator that equals one if the respondent is married or cohabitates with a partner and the partner lives with the respondent, in the same house, at the time of the interview.
- *Receives pension*: It is a binary indicator that equals one if the respondent reports receiving money from a retirement plan or a pension at the time of the interview.

A.4 Variables Used to Define Co-residence Patterns

- *More than 2 members in the house*: It is a binary indicator that equals one if the respondent reports that more than two household members are living with her, in the same house, at the time of the interview.
- *Parents live in the house*: It is a binary indicator that equals one if the respondent reports that her parents or her partner's parents are living with her, in the same house, at the time of the interview.
- *Children live in the house*: It is a binary indicator that equals one if the respondent reports that her children or her partner's children are living with her, in the same house, at the time of the interview.
- *Grandchildren live in the house*: It is a binary indicator that equals one if the respondent reports that her grandchildren are living with her, in the same house, at the time of the interview.
- *Siblings live in the house*: It is a binary indicator that equals one if the respondent reports that her siblings or her partner's siblings are living with her, in the same house, at the time of the interview.
- *Others live in the house*: It is a binary indicator that equals one if the respondent reports that other family members are living with her, in the same house, at the time of the interview.

B Robustness to Alternative Specifications

In this Section, we conduct several robustness tests to examine the sensitivity of our results to the choice of specification. Table A8 shows how the coefficient estimates change when we vary the independent variables in the regression. Panel A reports the estimates for the effect of eligibility on program take-up: the probability of receiving government aid. Panel B reports the estimates for any IPV, Panel C for physical or sexual IPV, Panel D for economic IPV, Panel E for psychological IPV. Each column reports estimates obtained using a different specification. In Column 1, we control only for the non-interacted main effects $Age66-67$ and $Wave2016$ (without including interview year fixed effects or age fixed effects). In Column 2, we control only for interview year fixed effects and age fixed effects. In Column 3, we add state fixed effects to the specification in Column 2. In Column 4, we add the following controls to the specification in Column 3: educational attainment (respondent's and partner's), respondent's number of children, a binary indicator for speaking an indigenous language (for the respondent and her partner), partner's age, respondent's age at marriage, respondent's age at the start of the current relationship, experience of violence as a kid (for the respondent and her partner), and rural residency. The specification in Column 4 is our baseline specification. In Column 5, we replace the partner's age, which is entered as a continuous variable, with fixed effects for the partner's age. Next, we add to our baseline specification by including state-specific age fixed effects (Column 6), which control for any state policy that only applies to individuals of a certain age; and state-specific survey-year fixed effects (Column 7), which control for state policies that were implemented during our study period (between 2006 and 2016) for all individuals in our sample. In Column 8, we re-estimate the model with weighted least squares using survey weights. In Column 9, we expand the sample to include the states of Chiapas, Tabasco, Tlaxcala, and Zacatecas, which implemented reforms similar to the 2013 PAM expansion between 2006 and 2011, and are excluded from the sample used in the main results. In Column 10, we also include in the sample women who are receiving a contributory pension. In Column 11, we restrict the sample to the survey waves closest to the 2013 reform (2011 and 2016), excluding the 2006 wave. Finally, in Column 12 we use the *doubly-robust* difference-in-differences (DID) estimator proposed by Sant'Anna and Zhao (2020), which is consistent when either a propensity score model or an outcome regression model is correctly specified, in cases where the parallel-trends assumption is satisfied after conditioning on covariates.[1]

Our results are not altered by using any of these different specifications, suggesting that they are not sensitive to the choice of control variables and are robust to holding constant time-varying or age-varying characteristics at the state level.

[1] We apply the *doubly-robust* estimator to the restricted sample of Column 11 that includes only the 2011 and 2016 survey waves because the estimator requires a two-by-two research design (Sant'Anna and Zhao, 2020).

C Appendix Figures and Tables

Figure A1: Short-run vs. Long-run Effects

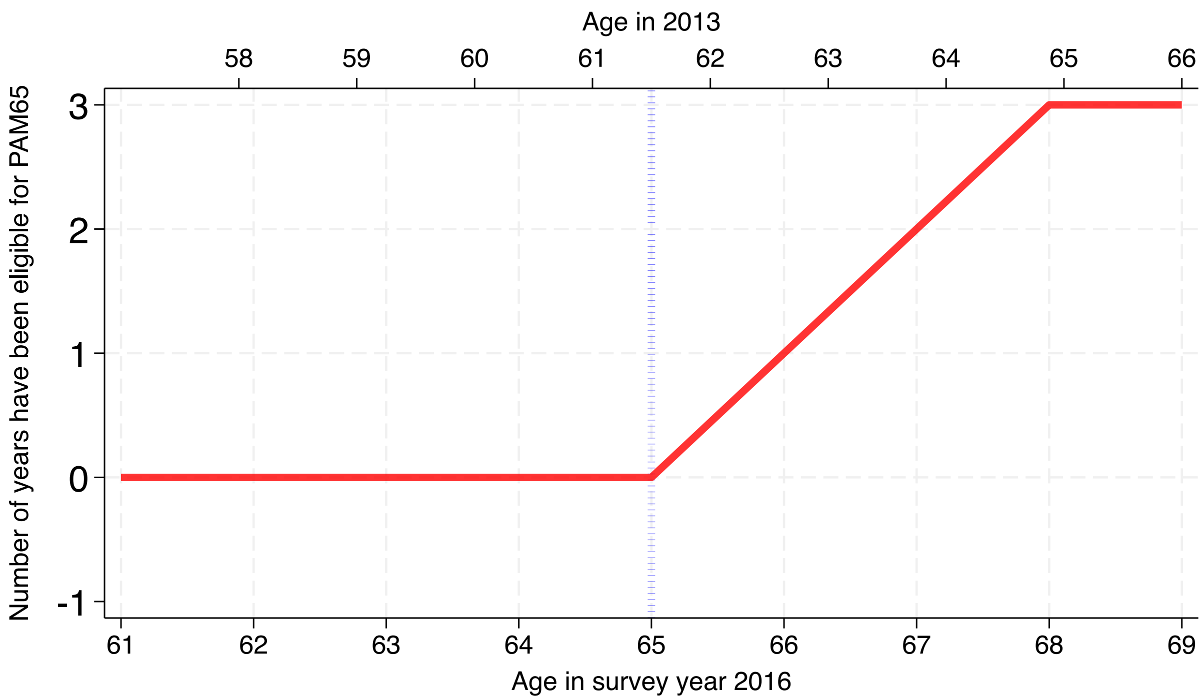
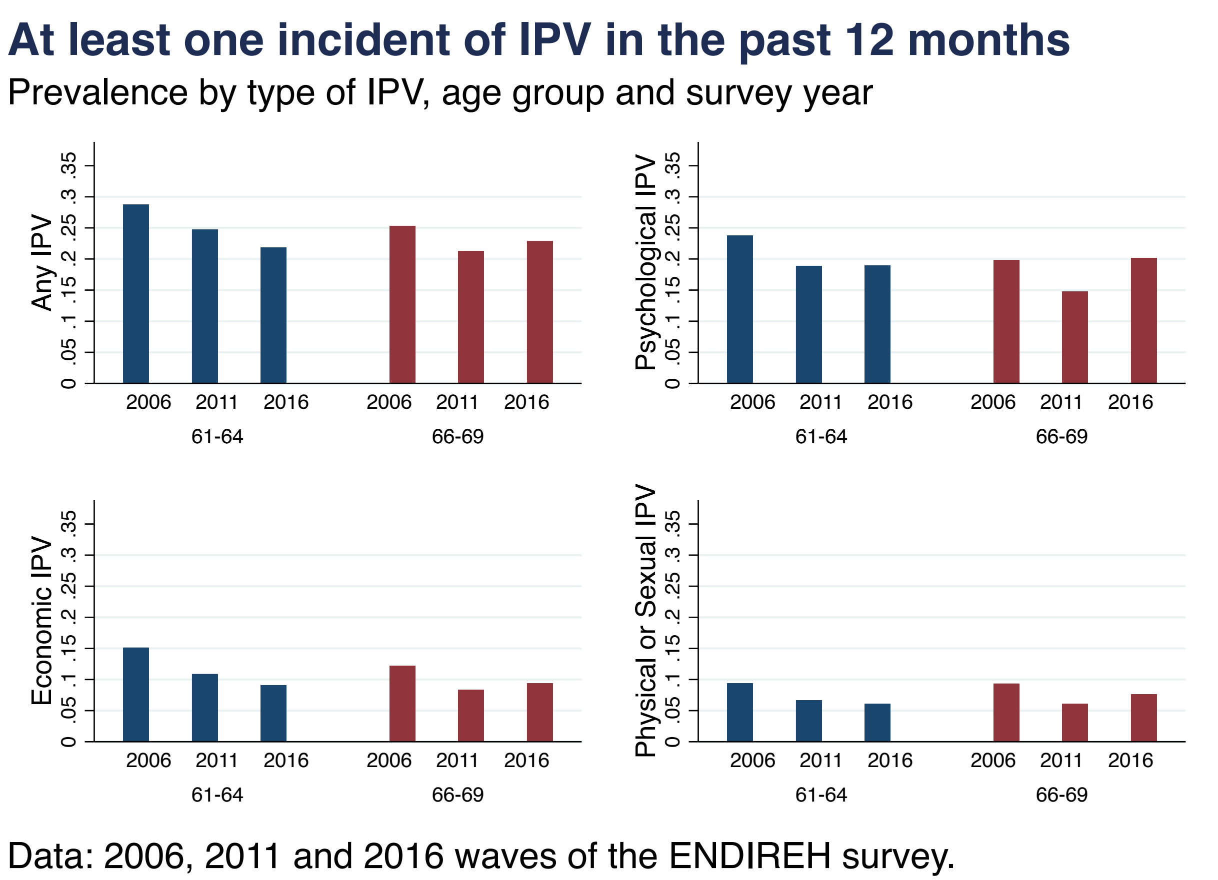
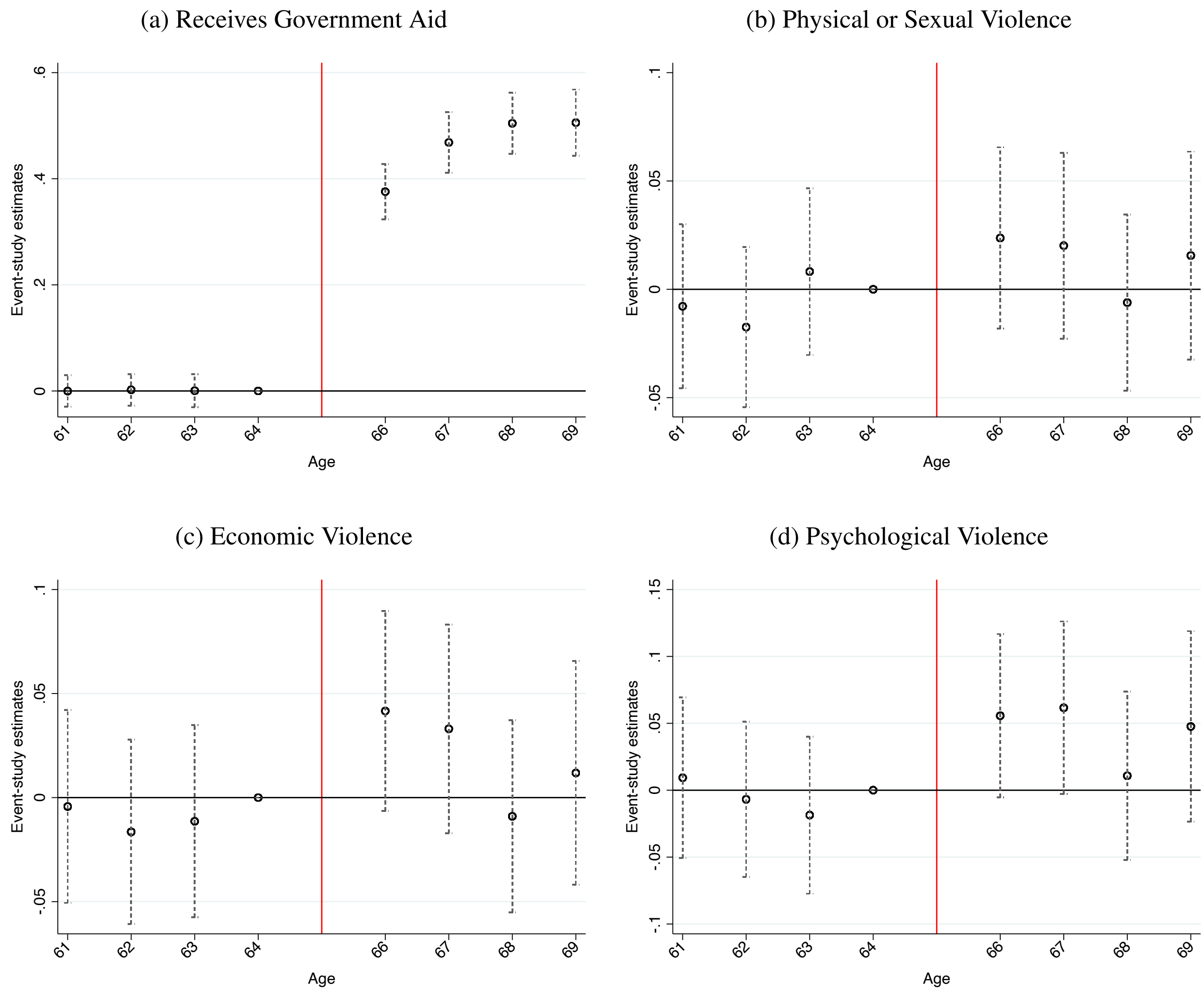


Figure A2: Descriptive Statistics: Prevalence of IPV by Type, Age Groups and Survey Waves



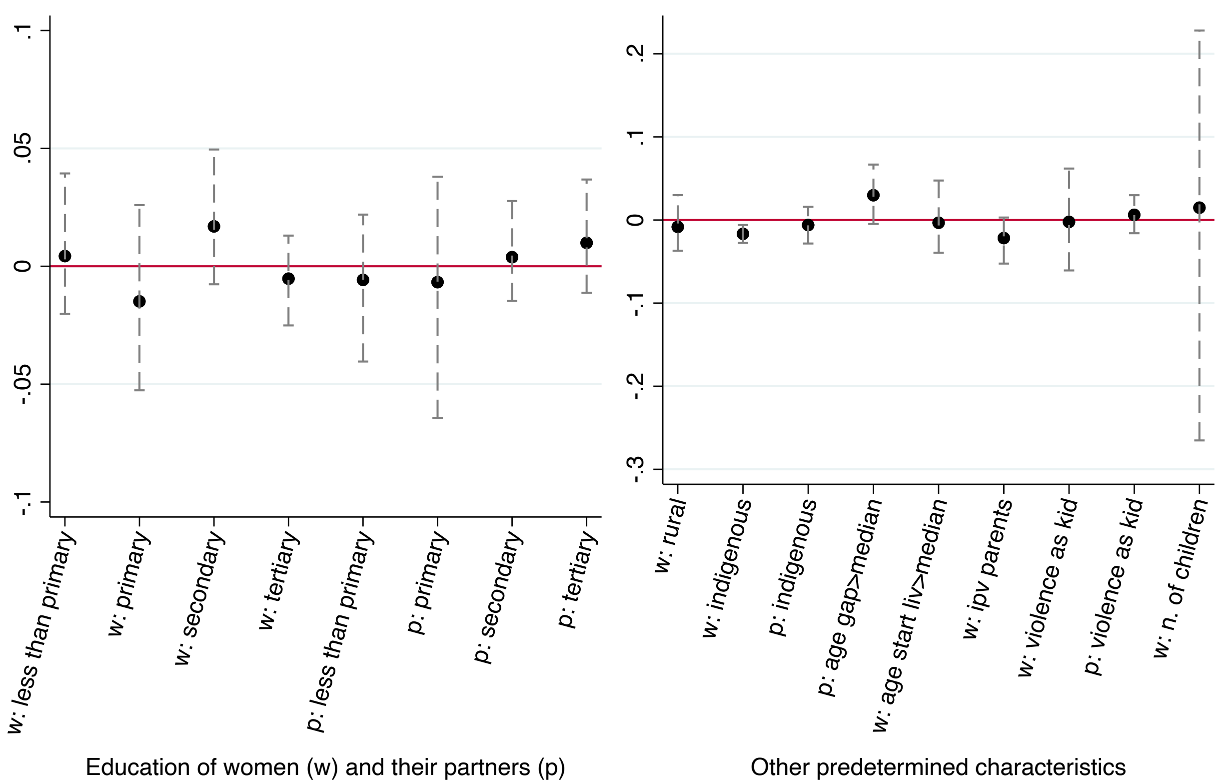
Notes: These figures show the incidence of intimate partner violence for women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. The respondents are victims of violence if they have experienced at least once an act of psychological, physical or sexual, or economic violence by their partner in the last 12 months before the interview.

Figure A3: Event Study by Age



Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Each figure reports the coefficient estimate of the interaction between each age dummy and the "2016 Survey wave" indicator. The figure shows the 95 percent confidence intervals estimated using robust standard errors. The coefficient estimates are also reported in Table A7 along with standard errors clustered at the age level and P-values for wild cluster bootstrap.

Figure A4: Impact of Eligibility on Predetermined Characteristics



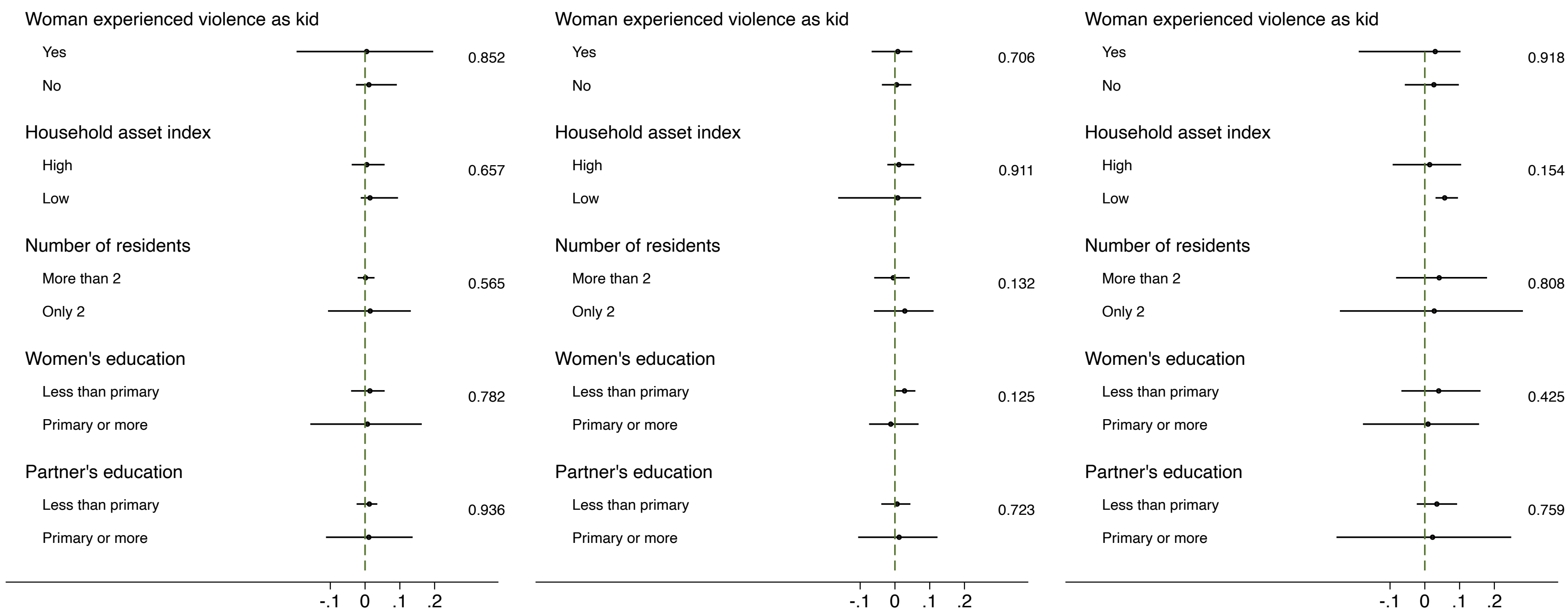
Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, and state fixed effects. The figures report the coefficient estimate of the interaction between "Age 66-69" and the "2016 Survey wave" indicators of the different outcomes described in the x-axis. w: refers to the respondent's characteristics, p: refers to her partner's characteristics. The figures show the 95 percent confidence intervals estimated using Wild Cluster Bootstrap.

Figure A5: Heterogeneity: Impact of Eligibility on IPV for Women aged 68-69

(a) Physical or Sexual Violence

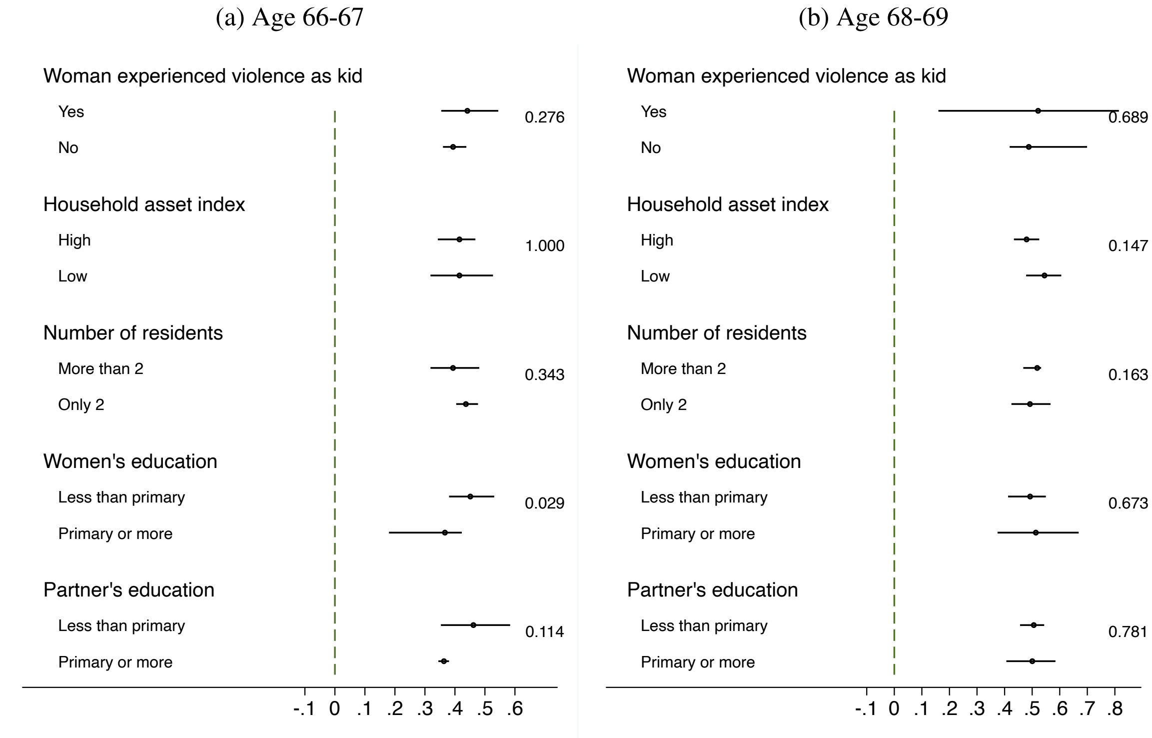
(b) Economic Violence

(c) Psychological Violence



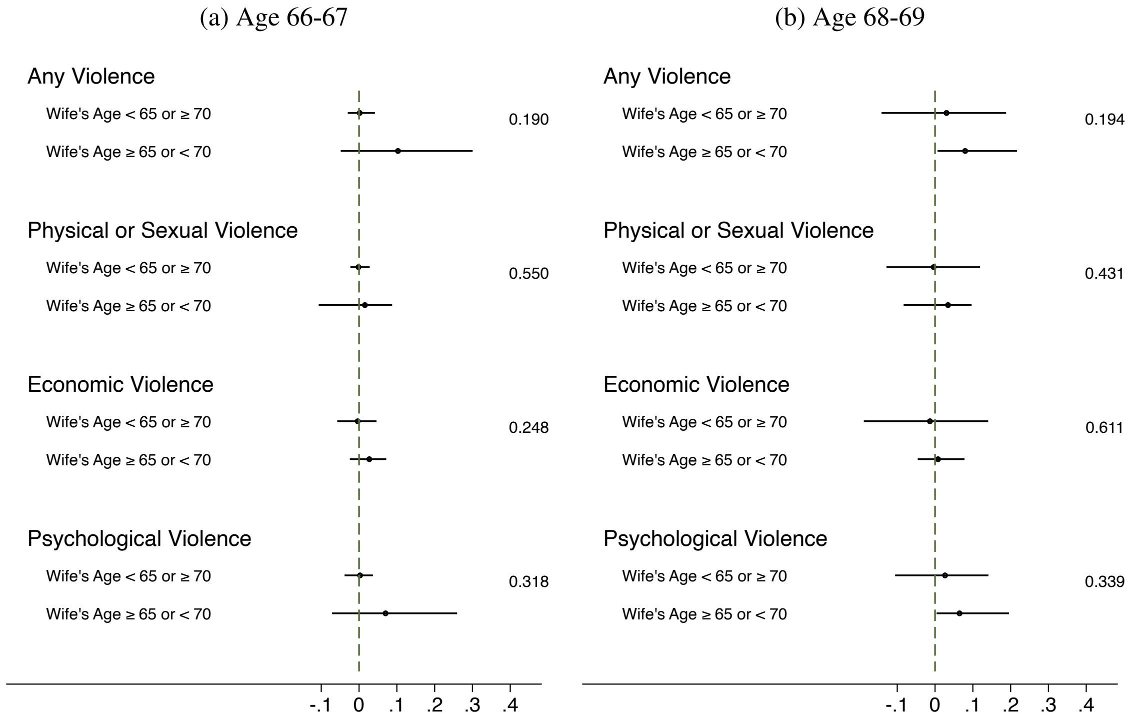
Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. For each heterogeneity variable, the figure reports the coefficient estimate of the interaction between "Age 68-69" and the "2016 Survey wave" indicators, for stratified regressions on the split samples. The figure shows the 95 percent confidence intervals estimated using Wild Cluster Bootstrap and the p-value for the test of equality of coefficients in a fully interacted regression.

Figure A6: Heterogeneity: Impact of Eligibility on the Probability of Receiving Government Aid



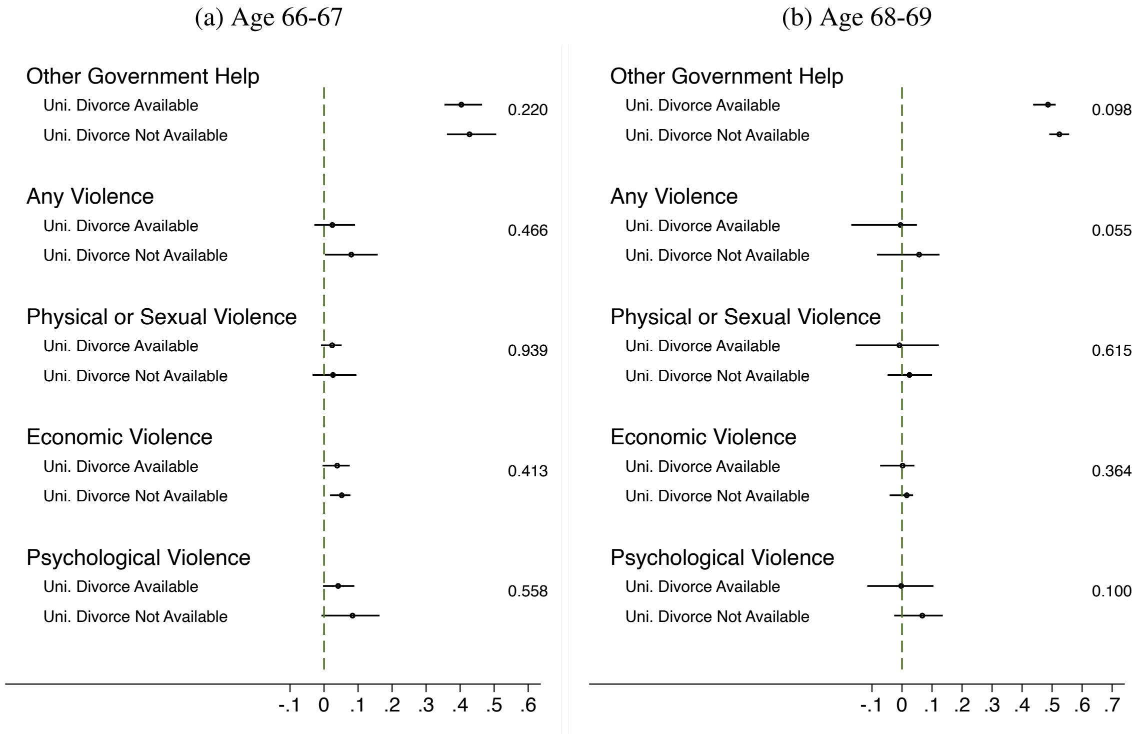
Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Figure (a) reports the coefficient estimate of the interaction between "Age 66-67" and the "2016 Survey wave" indicators, and Figure (b) reports the coefficient estimate of the interaction between "Age 68-69" and the "2016 Survey wave" indicators. The figures show the 95 percent confidence intervals estimated using Wild Cluster Bootstrap and the p-value for the test of equality of coefficients.

Figure A7: Mechanisms: Impact of Husband's Eligibility on IPV by Wife's Age



Notes: The sample includes women who are currently married or in a union and live with their partners ages 61-64 and 66-69 in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for the husband's age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Figure (a) reports the coefficient estimate of the interaction between "Husband's Age 66-67" and the "2016 Survey wave" indicators, and Figure (b) reports the coefficient estimate of the interaction between "Husband's Age 68-69" and the "2016 Survey wave" indicators. The figures show the 95 percent confidence intervals estimated using Wild Cluster Bootstrap and the p-value for the test of equality of coefficients in a fully interacted regression.

Figure A8: Mechanisms: Outside Options and Divorce Laws



Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Figure (a) reports the coefficient estimate of the interaction between "Age 66-67" and the "2016 Survey wave" indicators, and Figure (b) reports the coefficient estimate of the interaction between "Age 68-69" and the "2016 Survey wave" indicators. The figures show the 95 percent confidence intervals estimated using Wild Cluster Bootstrap and the p-value for the test of equality of coefficients.

Table A1: Definition of IPV Variables

Category	**Sub-category**	**Question**: In the past 12 months, the respondent's partner ...
Economic IPV	Theft or coercion	Appropriated or took money or possessions from her
Economic IPV	Economic control	Complained about how she spent money
Economic IPV	Economic control	Has been stingy with the household expenses even if he has money
Economic IPV	Economic control	Has forbidden the respondent from working or studying
Economic IPV	Failure to provide	Did not provide the respondent with economic support or threatened the respondent to not support her financially
Economic IPV	Failure to provide	Spent money needed for the household
Psychological IPV	Degradation	Shamed her, underestimated or humiliated her
Psychological IPV	Degradation	Said she cheated on him
Psychological IPV	Degradation	Became angry because household chores were not done like he wanted
Psychological IPV	Indifference	Ignored her, did not show his affection
Psychological IPV	Indifference	Stopped talking to her
Psychological IPV	Intimidation	Made her feel fear
Psychological IPV	Intimidation	Destroyed, threw away, or hid things belonging to her or the household
Psychological IPV	Intimidation	Watched over, spied on, followed
Psychological IPV	Isolation	Locked her in, forbid her from going out or being visited
Psychological IPV	Isolation	Turned her children or relatives against her
Psychological IPV	Threat	Threatened her with a weapon
Psychological IPV	Threat	Threatened to kill her, himself or the children
Psychological IPV	Threat	Threatened to leave her, hurt her, take her children away or kick her out
Physical or Sexual IPV	Moderate physical IPV	Pushed her or pulled her hair
Physical or Sexual IPV	Moderate physical IPV	Kicked her
Physical or Sexual IPV	Moderate physical IPV	Kicked her
Physical or Sexual IPV	Moderate physical IPV	Threw objects at her
Physical or Sexual IPV	Moderate physical IPV	Beat her with his hands or an object
Physical or Sexual IPV	Severe physical IPV	Tied her up
Physical or Sexual IPV	Severe physical IPV	Tried to choke her or hang her
Physical or Sexual IPV	Severe physical IPV	Assaulted her with a knife or a blade
Physical or Sexual IPV	Severe physical IPV	Fired a weapon at her
Physical or Sexual IPV	Sexual IPV	Demanded that you have sexual intercourse
Physical or Sexual IPV	Sexual IPV	Forced you to do sexual things that you do not like
Physical or Sexual IPV	Sexual IPV	Used physical strength to force you to have sexual intercourse

Notes: The definitions are based on the questionnaires in the ENDIREH surveys.

Table A2: Impact of Eligibility on PAM Take-up and Amount (ENIGH)

	Married Women				Married Men			
	Take Up		Monthly Amount		Take Up		Monthly Amount	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age 66-69 X Post	0.576***		297.839***		0.545***		285.163***	
	(0.040)		(21.393)		(0.049)		(25.891)	
	[0.001]		[0.001]		[0.002]		[0.002]	
Age 66-67 X Post		0.528**		275.764**		0.490**		258.129**
		(0.052)		(29.856)		(0.065)		(35.251)
		[0.014]		[0.013]		[0.021]		[0.021]
Age 68-69 X Post		0.632***		323.673***		0.614***		319.264***
		(0.004)		(6.255)		(0.011)		(3.044)
		[0.000]		[0.000]		[0.000]		[0.000]
Observations	12,002	12,002	12,002	12,002	10,520	10,520	10,520	10,520
R^2	0.535	0.538	0.462	0.464	0.511	0.515	0.378	0.381
Mean Dep. Variable	0.129	0.129	81.400	81.400	0.124	0.124	79.378	79.378
P-value 66-67 vs. 68-69		0.186		0.343		0.185		0.229

Notes: The sample includes women and men who are currently married or in a union, live with their partners and do not receive a contributory pension in 2008, 2010, 2012, 2014 and 2016 waves of Mexico's Socioeconomic Conditions Module of the National Survey of Household Income and Expenditure (ENIGH). The survey asks about income in the past 6 months as a diary. We use the average monthly income from the past 6 months. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A3: Impact of Eligibility on Experiencing IPV More than Once in the Past 12 Months

	Any Violence		Physical or Sexual Violence		Economic Violence		Psychological Violence	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age 66-69 X Wave 2016	0.042*		0.019*		0.030**		0.045**	
	(0.017)		(0.009)		(0.011)		(0.012)	
	[0.070]		[0.098]		[0.050]		[0.034]	
			{0.082}		{0.082}		{0.082}	
Age 66-67 X Wave 2016		0.060**		0.029**		0.048**		0.056**
		(0.012)		(0.005)		(0.004)		(0.008)
		[0.048]		[0.039]		[0.011]		[0.035]
				{0.071}		{0.071}		{0.071}
Age 68-69 X Wave 2016		0.019		0.007		0.008		0.031*
		(0.021)		(0.010)		(0.005)		(0.017)
		[0.588]		[0.594]		[0.251]		[0.096]
				{0.247}		{0.169}		{0.085}
Observations	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193
R^2	0.069	0.070	0.031	0.031	0.040	0.040	0.060	0.060
Mean Dep. Variable	0.199	0.199	0.054	0.054	0.086	0.086	0.141	0.141
P-value 66-67 vs. 68-69		0.166		0.153		0.061		0.341

Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011 and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. Benjamini et al. (2006) sharpened q-values are calculated following Anderson (2008) and are shown in curly brackets. For regressions with only one treated age group (66-69), the q-values are calculated for all 3 hypotheses (3 outcomes: physical or sexual violence, economic violence, and psychological violence). For regressions with two treated age groups (66-67 and 68-69), the q-values are calculated for all 6 hypotheses (3 outcomes and 2 treatments). *** p<0.01, ** p<0.05, * p<0.1.

Table A4: Placebo Test: Using Only the 2006 and 2011 Waves and the 2011 Wave as Post-reform Wave

	Receives Government Aid		Any Violence		Physical or Sexual Violence		Economic Violence		Psychological Violence	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Age 66-69 X Wave 2011	0.011		-0.004		-0.007		0.004		-0.005	
	(0.008)		(0.015)		(0.012)		(0.015)		(0.014)	
	[0.260]		[0.830]		[0.593]		[0.830]		[0.731]	
Age 66-67 X Wave 2011		0.002		-0.001		-0.010		0.002		0.004
		(0.008)		(0.018)		(0.014)		(0.017)		(0.014)
		[0.832]		[0.932]		[0.578]		[0.920]		[0.787]
Age 68-69 X Wave 2011		0.021*		-0.007		-0.004		0.006		-0.017
		(0.004)		(0.016)		(0.012)		(0.014)		(0.015)
		[0.097]		[0.725]		[0.794]		[0.668]		[0.348]
Observations	6,854	6,854	6,854	6,854	6,854	6,854	6,854	6,854	6,854	6,854
R^2	0.026	0.027	0.081	0.081	0.046	0.046	0.052	0.052	0.070	0.070
Mean Dep. Variable	0.025	0.025	0.253	0.253	0.094	0.094	0.122	0.122	0.198	0.198
P-value 66-67 vs. 68-69		0.204		0.806		0.689		0.710		0.356

Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006 and 2011 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** p<0.01, ** p<0.05, * p<0.1.

Table A5: Impact of Eligibility on IPV: Age Event Study

	Receives Government Aid (1)	Any Violence (2)	Physical or Sexual Violence (3)	Economic Violence (4)	Psychological Violence (5)
Wave 2016* Age 61	-0.000	-0.019	-0.008	-0.004	0.009
	(0.002)	(0.001)	(0.001)	(0.001)	(0.002)
	[0.947]	[0.254]	[0.300]	[0.275]	[0.300]
Wave 2016* Age 62	0.002	-0.044	-0.017	-0.016	-0.007
	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
	[0.253]	[0.296]	[0.292]	[0.266]	[0.316]
Wave 2016* Age 63	0.000	-0.046	0.008	-0.011	-0.019
	(0.001)	(0.002)	(0.001)	(0.002)	(0.002)
	[0.729]	[0.299]	[0.241]	[0.395]	[0.259]
Wave 2016* Age 66	0.376	0.012	0.024	0.042	0.056
	(0.001)	(0.002)	(0.001)	(0.001)	(0.002)
	[0.136]	[0.351]	[0.247]	[0.275]	[0.313]
Wave 2016* Age 67	0.468	0.043	0.020	0.033	0.062
	(0.002)	(0.002)	(0.001)	(0.001)	(0.002)
	[0.185]	[0.290]	[0.310]	[0.249]	[0.301]
Wave 2016* Age 68	0.504	-0.031	-0.006	-0.009	0.011
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
	[0.149]	[0.233]	[0.256]	[0.232]	[0.429]
Wave 2016* Age 69	0.506	0.033	0.016	0.012	0.048
	(0.001)	(0.002)	(0.002)	(0.002)	(0.003)
	[0.150]	[0.347]	[0.325]	[0.261]	[0.349]
Observations	10,193	10,193	10,193	10,193	10,193
R^2	0.353	0.077	0.039	0.047	0.066
Mean Dep. Variable	0.028	0.250	0.073	0.116	0.204

Notes: The table reports the results of an age event study where each age dummy is interacted with the "2016 Survey wave" indicator. The sample includes women who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2011 and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A6: Impact of Eligibility on Sample Selection

	Married/Union		Married/Union with Partner		Receives Pension		Selected in our Sample	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Age 66-69 X Wave 2016	0.005		0.004		0.010		-0.006	
	(0.016)		(0.017)		(0.009)		(0.014)	
	[0.734]		[0.822]		[0.343]		[0.700]	
Age 66-67 X Wave 2016		-0.005		-0.005		0.013		-0.021
		(0.019)		(0.022)		(0.005)		(0.013)
		[0.773]		[0.739]		[0.138]		[0.295]
Age 68-69 X Wave 2016		0.017		0.015		0.007		0.012
		(0.022)		(0.023)		(0.017)		(0.013)
		[0.611]		[0.618]		[0.726]		[0.465]
Observations	20,806	20,806	20,806	20,806	20,806	20,806	20,806	20,806
R^2	0.022	0.022	0.022	0.022	0.120	0.120	0.049	0.049
Mean Dep. Variable	0.547	0.547	0.538	0.538	0.227	0.227	0.461	0.461
P-value 66-67 vs. 68-69		0.562		0.634		0.621		0.147

Notes: The sample includes women aged 61-64 and 66-69 in the 2006, 2011, and 2016 ENDIREH surveys that have been selected for the domestic violence modules. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A7: Anticipation Effect: Comparing Women Aged 56-59 and 61-64

	Receives Government Aid		Any Violence		Physical or Sexual Violence		Economic Violence		Psychological Violence	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Age 61-64 X Wave 2016	0.024***		0.006		0.004		0.003		-0.004	
	(0.002)		(0.015)		(0.008)		(0.007)		(0.011)	
	[0.000]		[0.739]		[0.592]		[0.701]		[0.719]	
Age 61-62 X Wave 2016		0.025***		0.002		-0.004		0.001		0.001
		(0.002)		(0.014)		(0.007)		(0.007)		(0.011)
		[0.004]		[0.900]		[0.671]		[0.921]		[0.904]
Age 63-64 X Wave 2016		0.023**		0.010		0.013		0.005		-0.010
		(0.002)		(0.021)		(0.006)		(0.008)		(0.012)
		[0.015]		[0.719]		[0.161]		[0.589]		[0.514]
Observations	15,481	15,481	15,481	15,481	15,481	15,481	15,481	15,481	15,481	15,481
R^2	0.026	0.026	0.088	0.088	0.050	0.050	0.054	0.054	0.073	0.073
Mean Dep. Variable	0.017	0.017	0.265	0.265	0.079	0.079	0.128	0.128	0.210	0.210
P-value 61-62 vs. 63-64		0.201		0.761		0.059		0.607		0.452

Notes: The sample includes women aged 56-59 and 61-64 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table A8: Robustness Tests: Alternative Specifications and Sample Selections

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Panel A: Receives ***government aid*** (mean=0.048)												
Age 66-67 X Wave 2016	0.419**	0.418**	0.417**	0.416**	0.416**	0.413**	0.416**	0.378**	0.419**	0.359**	0.413**	0.406***
	(0.036)	(0.035)	(0.035)	(0.035)	(0.035)	(0.036)	(0.034)	(0.046)	(0.041)	(0.023)	(0.032)	(0.010)
	[0.015]	[0.016]	[0.013]	[0.012]	[0.014]	[0.015]	[0.012]	[0.020]	[0.021]	[0.011]	[0.011]	
Age 68-69 X Wave 2016	0.504***	0.504***	0.503***	0.504***	0.506***	0.504***	0.504***	0.472***	0.505***	0.448***	0.496***	0.491***
	(0.002)	(0.002)	(0.001)	(0.001)	(0.002)	(0.003)	(0.002)	(0.017)	(0.006)	(0.002)	(0.003)	(0.005)
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.006]	[0.000]	[0.000]	[0.000]	
P-value 66-67 vs. 68-69	0.187	0.183	0.199	0.206	0.206	0.195	0.212	0.283	0.198	0.080	0.212	
Panel B: Experienced ***any type*** of violence in the last 12 months (mean=0.231)												
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Age 66-67 X Wave 2016	0.056*	0.056*	0.055*	0.054*	0.055*	0.051*	0.055*	0.076**	0.037	0.055**	0.056**	0.053
	(0.020)	(0.020)	(0.019)	(0.015)	(0.014)	(0.015)	(0.016)	(0.013)	(0.013)	(0.013)	(0.019)	(0.034)
	[0.069]	[0.071]	[0.072]	[0.069]	[0.075]	[0.077]	[0.071]	[0.031]	[0.106]	[0.050]	[0.038]	
Age 68-69 X Wave 2016	0.031	0.030	0.026	0.025	0.027	0.023	0.028	0.007	0.021	0.024	0.030	0.050
	(0.023)	(0.024)	(0.021)	(0.026)	(0.027)	(0.029)	(0.027)	(0.010)	(0.024)	(0.026)	(0.021)	(0.042)
	[0.275]	[0.310]	[0.348]	[0.535]	[0.506]	[0.624]	[0.473]	[0.529]	[0.555]	[0.632]	[0.253]	
P-value 66-67 vs. 68-69	0.540	0.547	0.464	0.512	0.537	0.583	0.577	0.031	0.551	0.476	0.532	
Panel C: Experienced ***physical or sexual*** violence in the last 12 months (mean=0.075)												
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Age 66-67 X Wave 2016	0.027*	0.027**	0.027*	0.027*	0.027**	0.029**	0.027*	0.049**	0.011	0.034**	0.031**	0.033**
	(0.006)	(0.006)	(0.006)	(0.005)	(0.006)	(0.005)	(0.006)	(0.006)	(0.006)	(0.008)	(0.006)	(0.014)
	[0.051]	[0.050]	[0.057]	[0.052]	[0.032]	[0.047]	[0.061]	[0.022]	[0.200]	[0.048]	[0.027]	
Age 68-69 X Wave 2016	0.009	0.008	0.007	0.008	0.009	0.008	0.009	0.013	0.008	0.009	0.010	0.002
	(0.008)	(0.009)	(0.009)	(0.010)	(0.010)	(0.011)	(0.011)	(0.009)	(0.011)	(0.010)	(0.011)	(0.021)
	[0.402]	[0.461]	[0.508]	[0.550]	[0.530]	[0.610]	[0.545]	[0.206]	[0.595]	[0.449]	[0.622]	
P-value 66-67 vs. 68-69	0.078	0.095	0.072	0.175	0.276	0.175	0.238	0.053	0.659	0.059	0.197	
Panel D: Experienced ***economic*** violence in the last 12 months (mean=0.100)												
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Age 66-67 X Wave 2016	0.047***	0.047***	0.047***	0.046**	0.048**	0.047**	0.046***	0.075*	0.038**	0.049**	0.046**	0.050***
	(0.004)	(0.004)	(0.004)	(0.005)	(0.005)	(0.007)	(0.004)	(0.017)	(0.008)	(0.006)	(0.003)	(0.012)
	[0.005]	[0.004]	[0.004]	[0.013]	[0.017]	[0.026]	[0.008]	[0.063]	[0.040]	[0.019]	[0.007]	
Age 68-69 X Wave 2016	0.009	0.008	0.007	0.009	0.011	0.010	0.010	0.003	0.013	0.004	0.007	0.020
	(0.007)	(0.007)	(0.007)	(0.009)	(0.009)	(0.011)	(0.010)	(0.018)	(0.009)	(0.007)	(0.009)	(0.013)
	[0.319]	[0.376]	[0.446]	[0.514]	[0.241]	[0.588]	[0.519]	[0.830]	[0.202]	[0.673]	[0.587]	
P-value 66-67 vs. 68-69	0.046	0.056	0.060	0.053	0.052	0.054	0.044	0.014	0.221	0.045	0.046	
Panel E: Experienced ***psychological*** violence in the last 12 months (mean=0.170)												
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Age 66-67 X Wave 2016	0.065**	0.065**	0.063**	0.063**	0.063**	0.062**	0.064**	0.086**	0.038*	0.058**	0.062**	0.051*
	(0.010)	(0.009)	(0.008)	(0.006)	(0.007)	(0.006)	(0.008)	(0.010)	(0.009)	(0.007)	(0.008)	(0.028)
	[0.027]	[0.027]	[0.022]	[0.014]	[0.012]	[0.012]	[0.020]	[0.017]	[0.068]	[0.023]	[0.031]	
Age 68-69 X Wave 2016	0.037*	0.036*	0.032*	0.032*	0.032*	0.034*	0.033*	0.039**	0.018	0.024	0.040*	0.042
	(0.013)	(0.014)	(0.011)	(0.015)	(0.016)	(0.018)	(0.016)	(0.014)	(0.014)	(0.015)	(0.012)	(0.030)
	[0.072]	[0.072]	[0.051]	[0.057]	[0.061]	[0.061]	[0.072]	[0.038]	[0.326]	[0.146]	[0.053]	
P-value 66-67 vs. 68-69	0.143	0.169	0.046	0.151	0.201	0.327	0.162	0.022	0.351	0.160	0.272	
Survey Year FE & Age FE		✓	✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Reg FE			✓	✓	✓	✓	✓	✓	✓	✓	✓	✓
Controls				✓	✓	✓	✓	✓	✓	✓	✓	✓
Partner Age FE					✓							
State X Age FE						✓						
State X Survey Year FE							✓					
Weighted								✓				
Including Chiapas, Tabasco Tlaxcala and Zacatecas									✓			
Including Individuals With Pension										✓		
Dropping 2006 Wave											✓	
Doubly Robust DID												✓
Observations	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	11,635	11,924	7,182	5,857

Notes: See notes to Table 2.

Table A9: Impact of Eligibility on Disaggregated IPV Sub-Categories

Panel A: Physical and Sexual Violence

	Moderate Physical		Severe Physical		Sexual Violence	
	(1)	(2)	(3)	(4)	(5)	(6)
Age 66-69 X Wave 2016	0.012		0.003		0.006	
	(0.007)		(0.003)		(0.008)	
	[0.139]		[0.352]		[0.601]	
Age 66-67 X Wave 2016		0.018		0.004		0.012
		(0.005)		(0.002)		(0.010)
		[0.105]		[0.145]		[0.603]
Age 68-69 X Wave 2016		0.002		0.006		0.003
		(0.004)		(0.004)		(0.006)
		[0.643]		[0.836]		[0.691]
Mean Dep. Variable	0.058	0.058	0.006	0.006	0.028	0.028
P-value 66-67 vs. 68-69		0.287		0.423		0.484

Panel B: Economic Violence

	Theft or Coercion		Economic Control		Fail to Provide	
	(7)	(8)	(9)	(10)	(11)	(12)
Age 66-69 X Wave 2016	0.002		0.023*		0.011	
	(0.003)		(0.010)		(0.009)	
	[0.518]		[0.089]		[0.307]	
Age 66-67 X Wave 2016		0.005		0.037**		0.025**
		(0.003)		(0.008)		(0.005)
		[0.171]		[0.047]		[0.032]
Age 68-69 X Wave 2016		0.004		0.023		0.005
		(0.004)		(0.015)		(0.009)
		[0.849]		[0.328]		[0.260]
Mean Dep. Variable	0.003	0.003	0.090	0.090	0.051	0.051
P-value 66-67 vs. 68-69		0.304		0.065		0.049

Panel C: Psychological Violence

	Indifference		Degradation		Intimidation		Isolation		Threats	
	(13)	(14)	(15)	(16)	(17)	(18)	(19)	(20)	(21)	(22)
Age 66-69 X Wave 2016	0.028		0.038**		0.010		0.007		0.010	
	(0.014)		(0.013)		(0.006)		(0.005)		(0.009)	
	[0.135]		[0.040]		[0.177]		[0.174]		[0.307]	
Age 66-67 X Wave 2016		0.050**		0.050**		0.018*		0.010		0.011
		(0.006)		(0.012)		(0.005)		(0.005)		(0.006)
		[0.027]		[0.049]		[0.076]		[0.109]		[0.193]
Age 68-69 X Wave 2016		0.004		0.023		0.008		0.006		0.010
		(0.008)		(0.015)		(0.016)		(0.016)		(0.016)
		[0.843]		[0.185]		[0.845]		[0.631]		[0.688]
Mean Dep. Variable	0.135	0.135	0.086	0.086	0.039	0.039	0.021	0.021	0.0036	0.036
P-value 66-67 vs. 68-69		0.044		0.444		0.065		0.595		0.794

Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** p<0.01, ** p<0.05, * p<0.1.

Table A10: Impact of Eligibility on Bargaining Power and Co-residence Patterns

	Household Composition												Bargaining Power	
	More than 2 members in the house		Parents live in the house		Children live in the house		Grandchildren live in the house		Siblings live in the house		Others live in the house		Bargaining Power Index	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
Age 66-69 X Wave 2016	0.004		0.005		0.000		0.006		-0.003		-0.010		0.009	
	(0.023)		(0.003)		(0.019)		(0.014)		(0.002)		(0.018)		(0.014)	
	[0.849]		[0.180]		[0.983]		[0.731]		[0.164]		[0.594]		[0.579]	
Age 66-67 X Wave 2016		0.011		0.003		-0.004		0.002		-0.004*		0.010		0.014
		(0.021)		(0.004)		(0.012)		(0.014)		(0.002)		(0.012)		(0.014)
		[0.636]		[0.661]		[0.811]		[0.928]		[0.053]		[0.496]		[0.507]
Age 68-69 X Wave 2016		-0.004		0.007		0.005		0.012		-0.001		-0.035		0.002
		(0.029)		(0.002)		(0.038)		(0.013)		(0.002)		(0.018)		(0.022)
		[0.892]		[0.131]		[0.822]		[0.469]		[0.576]		[0.214]		[0.897]
Observations	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,193	10,186	10,186
R^2	0.062	0.062	0.011	0.011	0.065	0.065	0.042	0.042	0.009	0.009	0.045	0.046	0.074	0.074
Mean Dep. Variable	0.597	0.597	0.009	0.009	0.487	0.487	0.271	0.271	0.011	0.011	0.112	0.112	-0.096	-0.096
P-value 66-67 vs. 68-69		0.603		0.505		0.764		0.059		0.490		0.066		0.734

Notes: The sample includes women aged 61-64 and 66-69 who are currently married or in a union, live with their partners, and do not receive a contributory pension in the 2006, 2011, and 2016 ENDIREH surveys. Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for age fixed effects, survey year fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** p<0.01, ** p<0.05, * p<0.1.

Table A11: Impact of Eligibility on Household Income (ENIGH)

	Log HH Monetary Income			
	Women's Eligibility for PAM		Men's Eligibility for PAM	
	(1)	(2)	(1)	(2)
Age 66-69 X Post	0.075**		0.125**	
	(0.028)		(0.043)	
	[0.028]		[0.030]	
Age 66-67 X Post		0.082*		0.159*
		(0.032)		(0.045)
		[0.097]		[0.064]
Age 68-69 X Post		0.067		0.084
		(0.025)		(0.051)
		[0.117]		[0.142]
Observations	12,002	12,002	10,520	10,520
R^2	0.260	0.260	0.235	0.235
Mean Dep. Variable	7.314	7.314	7.037	7.037
P-value 66-67 vs. 68-69		0.617		0.438

Notes: The sample includes women and men who are currently married or in a union, live with their partners and do not receive a contributory pension in 2008, 2010, 2012, 2014 and 2016 waves of Mexico's Socioeconomic Conditions Module of the National Survey of Household Income and Expenditure (ENIGH). Observations from Chiapas, Tabasco, Tlaxcala, and Zacatecas (states that implemented monetary welfare reforms targeting individuals aged 65 and above between 2006 and 2011) are excluded. All regressions control for survey year fixed effects, age fixed effects, state fixed effects, and individual and household characteristics. Standard errors clustered at the age level are reported in parentheses. P-values for wild cluster bootstrap with Webb weights are reported in square brackets. *** p<0.01, ** p<0.05, * p<0.1.