UMEÅ UNIVERSITY

Essays on Health, Labor Market Behavior, and Economic Incentives.

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Umeå 2024

Doctoral thesis
ISBN print: 978-91-8070-225-6
ISBN PDF: 978-91-8070-226-3
ISSN: 0348-1018
Umeå Economic Studies No: 1018
Cover by Alejandro Vega
Electronic version available at: http://umu.diva-portal.org/
Printed by: CityPrint i Norr AB
Umeå, Sweden 2024

To my family

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Abstract

Paper [1] analyzes how labor force participation changes in response to major health shocks, such as new cancer diagnoses, heart attacks, and strokes, in middle-aged to elderly Mexican couples, and how the spouses interact in their responses to these shocks. The data originates from the Mexican Health and Aging Study and provides information on how couples coordinate their labor market activities in response to major health shocks. The results show that women's labor force participation is negatively affected by a major health shock to their husbands. In contrast, men's labor force participation does not change significantly in response to a major health shock to their wives.

Paper [2] focuses on the correlation between negative health shocks and the households' share of wealth held in risky assets. By using U.S. data from the Health and Retirement Study, we try to establish a link between negative health shocks and financial outcomes such as the household's probability of owning risky assets and the share of risky assets held. We define a recent negative health shock to include cancer or malignant tumor diagnoses, stroke or transient ischemic attack, heart attack, coronary heart disease, angina, congestive heart failure, or other heart problems. We find that both the probability of owning risky assets, and the share of risky assets, are significantly lower among households where the women has experienced a negative health shock. In contrast, neither the probability of owning risky assets nor the share of risky assets held by the household are significantly associated with a negative health shock to the man.

Paper [3] investigates whether job loss can cause symptoms of depression in later life. We focus on couples aged 50 or older. We use data from the Health and Retirement Study, which provides longitudinal information about changes in labor market status and mental health outcomes among respondents and their spouses in the United States. To deal with potential reverse causality problems, we utilize data on job loss resulting from business closures. We find that job loss can lead to depressive symptoms for the affected individual's partner. The effects are gendered, as women are negatively affected by job losses experienced by their husbands, but we do not observe such harmful effects among men whose wives lose their jobs. We also show how the effects of job loss vary across couples with differing levels of economic resources and health care needs, as well as differential access to health care.

Paper [4] estimates the labor supply response to an increase in the marginal wage rate among middle-aged to elderly Mexican women. Using data from the National Survey of Occupation and Employment, I find that an increase in the marginal wage rate is associated with an increase in worked hours. The results suggest that the marginal wage rate elasticities are larger for older women than for their younger counterparts.

Keywords

Labor force participation, labor supply, health status, financial risk-taking, gender differences, job loss.

Acknowledgements

I would like to express my deepest gratitude to my three esteemed supervisors, Thomas Aronsson, David Granlund, and Gauthier Lanot, whose guidance, support, and expertise were invaluable throughout the journey of completing this doctoral thesis. Your unwavering dedication to my academic growth, your insightful feedback, and your encouragement were instrumental in shaping this research. Your mentorship has left an indelible mark on my scholarly development.

I am also profoundly grateful to my family, who have been a constant source of love, encouragement, and support. My parents have been my pillars of strength, always believing in me and providing the unwavering encouragement that enabled me to embark on this academic pursuit. To my brother, your faith in me has been a driving force in my perseverance. Your sacrifices and understanding during this demanding period are deeply appreciated.

I would also like to extend my gratitude to my friends and colleagues, Alejandro Egüez, Aemiro Melkamu, Mattias Vesterberg, the Johans, Hanna Lindström, Linn Karlsson, Sef Meens-Eriksson and Golnaz Amjadi, who offered their support, comments to my papers and camaraderie during this challenging journey. Your moral support and shared experiences made the research process more meaningful and memorable.

This thesis represents the culmination of years of dedication and hard work, and it could not have been achieved without the collective support of these wonderful individuals. Thank you for being there for me, for believing in me, and for sharing in the joy of this accomplishment.

Umeå, 2024.

Alejandro

This thesis contains a summary and four self-contained papers.

Paper [1]

Vega, A. (2019). *Major health shocks and decisions about labour force participation amongst Mexican couples.* (Umeå Economic Studies 967). Umeå School of Business, Economics and Statistics, Umeå University.

Paper [2]

Vega, A. & Velli, E. (2020). *Health and the share of wealth held in risky assets.* (Umeå Economic Studies 972), Umeå School of Business, Economics and Statistics, Umeå University.

Paper [3]

Vega, A., & Baranowska-Rataj, A. (2021). *The effects of involuntary job losses on depressive symptoms among couples*. (CEDAR Working Papers, 2021:5). Umeå School of Business, Economics and Statistics, and the Department of Sociology, Umeå University.

Paper [4] Vega, A. (2023). *The Labor Supply of Elderly Women in Mexico*. (Umeå Economic Studies 1017). Umeå School of Business, Economics and Statistics, Umeå University.

1 Introduction and Background

In an era of demographic shifts and evolving economic landscapes, the interplay between health outcomes, labor supply, and financial risk-taking has become an increasingly important research area. Aging populations, changing family dynamics, and economic challenges have underscored the need to understand how individuals and households respond to critical life events. Among these, health shocks and their implications for labor market behavior and financial behavior have gained prominence. The decisions individuals make regarding their labor force participation and financial behavior are not isolated events but are intertwined with their health status and the well-being of their families.

This collection of research papers explores the multifaceted relationships between health, labor market choices and outcomes, and financial behavior. It offers insights into how households allocate resources when facing health shocks, how job loss affects the mental well-being of the affected individual's spouse/partner, and how elderly women's labor supply reacts to economic incentives. The insights, which will be described in more detail below, are crucial as they can be used to tailor public policies aimed at enhancing the well-being of individuals and society.

Paper [1] contributes to this understanding by examining how major health shocks, such as new cancer diagnoses, heart attacks, or strokes, impact the labor force participation decisions of middle-aged to elderly Mexican couples. Building upon and extending earlier studies (Berger, 1983; McClellan, 1998; Charles, 1999; Coile, 2004; Van Houtven & Coe, 2010), this research not only measures the extent to which health shocks affect labor force participation, but also presents a theoretical model for understanding how unexpected health events may impact the underlying time-allocation choices in the household. Earlier studies have yielded varying results. Charles (1999) finds that the likelihood of labor force participation among men decreases slightly when their spouse experiences a negative health shock, whereas Coile (2004) observes a substantial increase in the likelihood of labor force participation among men in similar situations. It is worth noting that these studies are based on U.S. and European populations and that none of them place their results in a theoretical context. In contrast, I present a theoretical model that explains possible mechanisms behind the results, and, to my knowledge, this is the first paper that focuses on middle-aged to elderly Mexican couples. Understanding how couples adapt to health shocks is of paramount importance for policymakers striving to enhance labor force participation among elderly women and improve access to formal caregiving, particularly in the context of Mexico's changing family structures and demographic transitions.

Paper [2] analyzes the relationship between negative health shocks and the ownership and allocation of risky financial assets in households where both spouses are 65 years or older. Utilizing data from the Health and Retirement Study (HRS), this study explores how health shocks affect households' financial choices. By introducing a theoretical framework, it also contributes to our understanding of the incentives driving these decisions. This research is particularly pertinent given the substantial share of wealth held by aging couples and the elevated likelihood of experiencing health shocks in later life. Earlier studies (Wu, 2003; Rosen and Wu, 2004; and Berkowirtz & Qiu, 2006) focus their work only on empirical modelling and estimation and do not examine the households' choice of consumption and investment in risky assets. We go further and present a theoretical model, which serves as a background to the empirical analysis. Furthermore, our study employs an objective measure of health, in contrast to previous literature that is mostly based on self-rated health measures. By using conditions diagnosed by a medical doctor, we are able to sidestep the subjectivity that is inevitable in any type of self-rated variable. Moreover, compared with Wu (2003), who examined the same measure of negative health shock as we do, we use more waves of the HRS, i.e., from 1994 to 2014. We also focus on an older age group since people aged 65 and above hold a substantial share of the net wealth in the U.S.

Paper [3] shifts the focus to the impact of involuntary job losses on depressive symptoms among couples aged 50 or older. Leveraging longitudinal data from the Health and Retirement Study (HRS), this paper investigates the ripple effects of job losses on mental health outcomes, emphasizing the interconnectedness of family members' well-being in the face of economic adversity. By scrutinizing the role of economic resources, health, and access to insurance, this research extends our understanding of how job losses reverberate through households, particularly impacting the well-being of the affected individual's spouse. The paper adds to the literature by bringing together insights from different disciplines to explain the mechanisms of the impact of job loss on the mental health of couples, as well as by discussing potential buffering mechanisms. A substantial body of research—including studies by Bubonya et al. (2017), Marcus (2013), Mendolia (2014), and Jolly (2022) focusing on the prime-age population—has shown that men's job losses have negative effects on their wives' mental health, while job loss among women has no significant effect on the mental health of their husbands. However, more knowledge is needed about the consequences of job losses on the mental health among older couples. The only study that addresses this topic specifically was carried out by Siegel et al. (2003) and found no evidence of spillover effects, possibly due to a small sample size. In the context of population aging, it is clearly important to understand the interrelatedness of the working lives and health outcomes of people approaching retirement. Older workers with fewer assets can be viewed as particularly vulnerable because job losses may have more significant financial consequences later in life (Couch et al., 2009). Moreover, re-employment becomes less likely for older workers (Chan & Huff Stevens, 2001), meaning that employability, which moderates the negative consequences of missed earnings in the working-age population (Green, 2011; Silla et al., 2009), is less protective.

Paper [4] directs attention to the labor supply decisions of elderly Mexican women and their responsiveness to changes in the marginal wage rate. As Mexico grapples with demographic shifts and fiscal challenges related to an aging population, understanding the labor market behavior of elderly women is critical for assessing the potential impact on tax revenues and social security. Drawing upon data from the Encuesta Nacional de Ocupación y Empleo (ENOE), this study employs robust methodologies to estimate the elasticity of labor supply with respect to the marginal wage rate, providing insights into the labor market behavior of older women. While many studies have examined the labor supply behavior of men (Pencavel, 2002; Devereux, 2004) and women (Eissa & Hoynes, 2004; Bargain, 2005; Blau & Kahn, 2007), these studies mostly focus on the prime-aged to pre-retirement age population in the U.S. and Europe. There are only a small number of studies addressing the labor supply behavior of women in Mexico. Gómez and Campos-Vázquez (2010) investigate the labor supply elasticity of married women who are between 25 and 55 years old. To my knowledge, there are only two earlier studies focusing on the labor supply of the Mexican elderly, although without reporting on or addressing the relationship between labor supply and wages. Juárez (2010) estimates the effect of a so-called nutrition transfer for individuals aged 70 or older and finds that the transfer reduces the labor force participation of single men aged 70 and older by 46 percent. For women, the point estimate is negative, but not significantly different from zero. Avila-Parra and Escamilla-Guerrero (2017) find that expanding the pension program did not reduce the labor force participation of Mexican adults between 66 and 69 years of age.

The summary of the papers in the following section sheds light on their individual contributions, while the main results and insights are summarized and discussed in a concluding section.

2. Summary of the Papers

Paper [1]: Major health shocks and decisions about labour force participation amongst Mexican couples

This study employs data from the Mexican Health and Aging Study (MHAS) to investigate how middle-aged to elderly Mexican couples respond to major health shocks, including new cancer diagnoses, heart attacks, or strokes. According to MHAS, approximately 2.9% of Mexicans aged 50 to 75, who are married or in partnerships, experience major health shocks. These events impact not only the directly affected individuals but also their spouses. The spouses of the affected individuals may respond by altering their labor force participation in primarily two ways. Firstly, if the affected individual withdraws from the labor force, the household faces reduced income. To offset this loss, the healthy spouse might enter the labor force if previously inactive or increase their work efforts if already participating. Secondly, health shocks can redefine the value of non-working time, prompting the healthy spouse to exit the labor force, or to decrease the number of hours they work, in order to care for their sick partner or to take over household responsibilities previously managed by the partner.

The paper introduces a theoretical model that provides a framework for interpreting the relationships between health shocks and household labor supply behavior. The model's comparative statics show that the healthy spouse's labor supply may either increase or decrease following a negative health shock to their partner. The theoretical model also serves as a starting point for the empirical analysis, which is based on four waves of the MHAS data to estimate the probability that the non-affected spouse participates in the labor force. The findings indicate a reduction in the probability of labor force participation among women when their husband faces a major health shock. In contrast, there is no significant relationship between men's labor force participation and negative health shocks to their wives. The analysis is also extended to examine the effects of health shocks on part-time and full-time labor force participation probabilities separately. The results highlight a decrease in the probability of full-time participation among women when their husbands experience a major health shock, while the effect on part-time participation remains statistically insignificant. In line with the results described above, a major health shock to the wife affects neither full-time participation nor part-time participation among men.

Further investigation includes an examination of additional health issues such as diabetes, hypertension, and arthritis. Interestingly, these health problems do not significantly affect the labor force participation of the non-affected partner. One possible explanation is that these conditions may not substantially hinder work when adequately treated. The study's main results suggest that Mexican women take their husband's health shocks into account in their labor force participation decisions, while the labor force participation among men largely remains unaffected by similar health shocks to their wives. Thus, women may prioritize caregiving and time spent with their ill spouse over work, highlighting their potential role as caregivers during such circumstances. These findings hold practical implications for policymakers aiming to formulate strategies to enhance the labor supply among elderly women, possibly by facilitating better access to formal caregiving within Mexican households.

Paper [2]: Health and the share of wealth held in risky assets

This study examines the relationships between negative health shocks and the households' ownership of risky financial assets alongside their allocation of these assets within their portfolios. To comprehensively explore this connection, the analysis uses a sample of couples aged 65 or above in the U.S., extracted from 11 waves (1994–2014) of the Health and Retirement Study (HRS). This dataset encompasses crucial information such as recent negative health shocks diagnosed by medical professionals, households' financial wealth, and a range of demographic and economic variables.

The paper's focus on the financial choices of couples aged 65 years and older is particularly pertinent due to their substantial share of the total U.S. household net worth. This demographic group holds a significant portion of the nation's wealth, as evidenced by median net worth figures from 2004, where households represented by individuals aged 65 and older exhibited a median net worth of $190,100, markedly surpassing the median net worth of $93,100 for all U.S. households. Furthermore, as the propensity for aging couples to face negative health shocks is elevated, an examination of whether these shocks influence their financial decisions is warranted. This becomes particularly interesting in the context of an aging population, where a growing number of older individuals will hold substantial wealth in the future.

The analysis aligns with previous research by Rosen and Wu (2004) and Edwards (2008), which anticipates a diminished allocation of wealth to risky financial assets among households experiencing negative health shocks. This phenomenon could be attributed to the sale of risky assets to cover medical expenses. Distinguishing itself from prior studies, the current paper employs an objective health measure—conditions diagnosed by a medical practitioner—thus sidestepping the subjectivity inherent in self-rated variables.

The research also differs from earlier work that predominantly focused on empirical modelling and estimation by introducing a theoretical framework to examine households' investments in risky assets, and how these investments are affected by a health shock to one of the household members. The model shows that such responses can go in either direction, suggesting that further examination requires empirical analysis. Through meticulous numerical analysis, considering the model's comparative statics, it is discerned that the change in risky asset allocation hinges on the functional form of households' utility functions, and risk aversion does not inherently lead to reduced investment in risky assets.

The study establishes that households in which the woman experienced a negative health shock subsequently have a lower probability of owning risky assets and a lower share of wealth allocated to such assets. Conversely, no statistically significant correlation is detected between negative health shocks to men and the observed outcomes.

Paper [3]: The effects of involuntary job losses on depressive symptoms among couples

This study investigates whether job loss may lead to depressive symptoms experienced by the affected individual's spouse/partner. While numerous studies have shed light on the repercussions of unemployment on mental health, often focusing on individuals directly affected by job loss, this study seeks to illuminate the broader ripple effects on family members, particularly spouses, who share an emotional bond and are intertwined economically.

Employing longitudinal data sourced from the Health and Retirement Study (HRS), the paper estimates the influence of job losses on mental health outcomes in households where both spouses are 50 to 65 years old. It assesses these effects across various dimensions, including age, employment status, and wealth, all while scrutinizing the potential mitigating role played by economic resources. Moreover, the analysis extends its scope to evaluate whether these effects diverge among couples characterized by varying levels of health needs and access to health insurance. To avoid the potential problem of reverse causality, the study relies on data regarding job loss due to business closures, thereby sidestepping potential biases.

The outcomes of the investigation reinforce earlier findings regarding the gendered patterns of the spillover effects of job losses, while extending these insights to the context of older couples. The findings underscore a conspicuous rise in depressive symptoms among female spouses of respondents who have undergone job losses, in alignment with the tenets of the family stress literature. Furthermore, among women, the study identifies that the detrimental impacts on mental health due to a partner's job loss are particularly pronounced in cases where the woman is unemployed, coupled with lower accumulated wealth and a lack of access to health insurance—highlighting the protective potential of economic resources. The research also reveals that women's job losses do not manifest in a corresponding negative impact on the mental well-being of their male spouses, irrespective of these men's participation in paid work, socioeconomic circumstances, or access to health insurance.

Overall, the study advances our understanding of the interplay between job loss, spousal dynamics, and mental health outcomes. It unravels a nuanced portrait in which gendered and economic considerations intertwine to shape the repercussions of job losses, fostering a deeper understanding of the complex web of implications arising from significant life-course events within families.

Paper [4]: The labor supply of elderly Mexican women

This study examines how middle-aged to elderly women change their labor supply in response to an increase in the marginal (net-of-tax) wage rate. The investigation stems from the crucial need to understand how women's engagement in the labor market shifts as they become older, in response to economic incentives that could potentially foster increased the labor force participation of women in post-retirement age, especially given that women participate in the labor force to a much lesser extent than men as they get older. If economic stimuli incentivize women to join or persist in the labor force, and if those already employed augment their work hours later in life, tax revenue will increase, thereby potentially easing the constraints on social security.

The relevance of this inquiry stems from the backdrop of an aging population, declining fertility rates, and an extended life expectancy. In Mexico, the proportion of women aged 50 to 70 witnessed an ascent from 13 percent of all women in 2010 to approximately 17 percent in 2018. This upward trajectory is projected to persist, with the share of women in this age bracket projected to approach 24 percent by 2050. Concurrently, the population of both men and women aged 65 and above constituted 7.2% of the total population in 2018, with estimates predicting a surge to 10.3% by 2030 and a remarkable 16.8% by 2050. These demographics are likely to exert pressures on social security expenditures, as an increased number of elderly individuals may necessitate augmented pension benefits or government transfers.

Using data from the Encuesta Nacional de Ocupación y Empleo (ENOE) surveys spanning 2005 to 2017, this study adheres to the methodological framework outlined by Blundell and Macurdy (1999) to counteract selection biases. Beginning with the estimation of the participation equation and computation of the inverse Mills ratio, the study then utilizes this ratio in the log-wage estimation to control for selection among labor force participants. This culminates in the derivation of imputed wages for non-participants. Subsequently, employing an instrumental variable tobit model, the study links these imputed wages to the labor supply equation, thereby estimating the labor supply of both participants and non-participants. The findings unveil marginal wage rate elasticities for participating women ranging from 0.02 to 0.10, and from 0.29 to 0.66 within a pooled sample of all women.

Delving deeper, the study develops a second model tailored to participants, introducing controls for random effects. Although this adjustment does not alter the qualitative findings, it does lead to lower elasticities when compared to the pooled model. Ultimately, the results lend support to the premise that older cohorts of women exhibit more pronounced responses to shifts in the marginal wage rate compared to their younger counterparts, with the decision to participate in the labor force proving more sensitive to economic incentives than the intensive margin of labor supply.

3. Conclusions

In this collection of research papers, I have investigated various facets of the relationships between health shocks, labor supply behavior, and financial behavior among aging populations. I have also examined how job loss influences the mental well-being of the affected individual's spouse/partner, and how changes in the marginal wage affect the labor supply of elderly women. Across the first three papers, a pattern of gender differences in response to various life events emerges, highlighting the influence of gender roles on individual choices in response to health shocks and job losses.

Both Paper [1] and Paper [2] find gender difference in the responses to health shocks. Women—but not men—respond to health shocks to their spouse/partner by reducing their labor supply, and the household's share of risky assets is reduced when the woman—but not when the man—faces a negative health shock. These findings align with earlier studies conducted in different contexts. Notably, Coile (2004) found that females in the U.S. decreased their labor force participation when their spouse faced major health shocks accompanied by a substantial decrease in functioning. Similarly, Siegel (2006), using U.S. data, found that wives of men with heart problems reduced their labor force participation. Rosen and Wu (2004) and Berkowitz and Qiu (2006) found that a negative health shock affecting the female partner is associated with a lower likelihood of owning risky assets.

The findings of Paper [1] suggest that women act as caregivers for their spouses, which, in turn, affects household labor income. The loss in labor income, along with potential medical expenses, contribute to reduce the household's well-being. Policymakers could, therefore, develop strategies to make formal caregiving more accessible to Mexican households to maintain women's labor supply and counteract the negative effects on households' well-being. In the context of Paper [1], future research may address the access to medical attention and formal caregiving, examine the potential role of household characteristics, and extend the theoretical framework to include medical expenses. Regarding Paper [2], an interesting topic for future research would be to consider the households' intertemporal decisions regarding consumption and investment in risky assets. Another would be to extend the household model to take into account and examine potential differences in preferences between spouses. Finally, it is also important to allow for a broader set of determinants of investments in risky assets, such as variation in life expectancy and bequest motives, in the theoretical and empirical analyses.

Expanding on the theme of gender differences, Paper [3] shifts the focus to the impact of job losses, where gender differences are found in how depressive symptoms manifest. Consistent with the family stress literature (Lavee, 2013), the study reveals an increased level of depressive symptoms in female spouses of respondents who lost their jobs. Furthermore, the negative effects of job loss on mental health are stronger among women who are not employed themselves and among couples with less accumulated wealth and no access to health insurance. The results highlight the buffering role of economic resources (Kalleberg et al., 2000). In contrast, male spouses of respondents who experienced job losses exhibit a decrease in depressive symptoms, aligning with predictions from the economic health capital model of Bolin et al. (2001). More specifically, when women lose their jobs, this may trigger a reallocation of resources within the household towards improving male spouses' well-being. Thus, the woman uses the time gained, after ceasing paid work, to care for her spouse, and the beneficial effects of this reallocation of resources may offset the negative effects of job loss. While the theoretical models presented in Bolin et al. (2001) include the possibility of divorce, our current analyses are limited to the impact of job losses on spousal health within couples that remain together. We do not track couples after divorce. Therefore, an interesting topic for future research would be to analyze partnership dynamics in response to job loss.

Collectively, these three papers find gender differences in economic behavior and health outcomes when analyzing household responses to health shocks and job losses, respectively. This emphasizes the influence of traditional gender roles, which shape not only behavior but also the allocation of resources within households during challenging life events. These findings contribute to the discussion in the literature about the impact of gender on decision-making and well-being in various contexts.

Finally, Paper [4] focuses on the labor supply behavior of elderly women in Mexico. The findings align with research in developed countries, such as Mroz (1987) and French (2005), which suggests that older individuals tend to have larger wage elasticities of labor supply than their younger counterparts. It is worth noting that the participation decision appears to be more responsive to economic incentives than the intensive margin of labor supply among this group of older women. These results emphasize the significance of considering economic incentives, particularly changes in the marginal wage rate, when assessing the labor supply decisions of older women. The latter is interesting in the sense that policymakers could tailor policies to increase the marginal wage rate of older women to incentivize non-participants to join the labor market. Eventually, this may lead to increased tax revenue, which would relieve the increasing pressure on the Mexican social welfare system. Future research should preferably use richer data. The data used in Paper [4] does not report non-labor income, and the severity of this omitted variable problem is difficult to assess. Moreover, the data lacks information on work experience, which is a potentially important determinant of wages.

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Major health shocks and decisions about labour force participation amongst Mexican couples

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Abstract

The purpose of this study is to analyze the labour force participation response to major health shocks, such as new cancer diagnoses, heart attacks and strokes, in middle-aged to elderly Mexican couples. The data originate from the Mexican Health and Aging Study (MHAS), and provides information on how couples coordinate their labour market activities in response to major health shocks. I find that female labour force participation is negatively affected by a major health shock to her spouse. In contrast, there is not a significant effect of a female negative health shock on her spouse's labour force participation. I find the same result when the labour force participation is split between part-time work and full-time work.

JEL code: I1 J01

Keywords: health shock, labour force participation, household model, labour supply, elderly couples.

*I would like to thank Thomas Aronsson, Gauthier Lanot, David Granlund, Katharina Jenderny and Magnus Wikström for valuable comments.

1 Introduction

The purpose of this study is to analyze the labour force participation response of Mexican middle-aged to elderly couples to major health shocks, such as new cancer diagnoses, heart attacks or strokes. Furthermore, the analysis provides a rationale for understanding how couples coordinate their labour market activities in response to such health shocks. Hence, the results show that females reduce their labour force participation when their spouses face major health shocks, which may suggest that females act as caregivers when males fall ill. These results are of practical relevance to policy makers when designing strategies to maintain, and even increase, elderly women's labour supply, i.e. by making formal caregiving more accessible to Mexican households.[1]

According to the Mexican Health and Aging Study (MHAS) 2.9% of Mexicans between 50 and 75 years of age who were married or living in a partnership had a new cancer diagnosis, a heart attack, or a stroke.[2] These negative health shocks will likely affect the labour force participation of the affected individual and of his or her spouse. The spouse or partner of the affected individual can respond in at least two ways. First, if the ill spouse stops participating in the labour force then the household would experience a reduction in its earned income. To overcome the loss in income, the healthy spouse may join the labour force if he/she was originally not participating, or increase his/her work effort if participating. Second, negative health shocks may change the value of non-working time, driving the healthy spouse to drop out of the labour force to take care of the ill spouse or to complete household activities that were previously the responsibility of the ill spouse.

The evidence in the literature regarding spousal labour supply responses to negative health shocks is not always consistent. The diversity in results could be driven by many

[1] Mexico is one of three OECD countries without universal health coverage. In 2016, out of pocket medical expenditure represented nearly 50% of total health spending, the highest for an OECD country (OECD, 2016).

[2] The percentage takes into account the periods 1999-2003 and 2010-2015.

factors such as the measurement of the health shock,[3] the timespan of the data, the age cohorts of the selected individuals, etc.

Several studies use the U.S. Health and Retirement Study (HRS), yet the results vary across these studies (Berger, 1983; McClellan, 1998; and Van Houtven and Coe, 2010). For example, Charles (1999) finds that the likelihood of labour force participation among men decreases by 1 percentage-point when their spouse experiences a negative health shock, while the likelihood that women participate increases by 1 percentage-point if their spouse experiences a negative health shock. In contrast, Coile (2004) finds that the likelihood of labour participation among men increases by approximately 16 percentage-points when their spouses face an acute health shock,[4] while the likelihood of labour force participation among females decreases by approximately 6 percentage-points.

It is important to note that the latter two studies have different settings. Charles (1999) uses the first two waves of the HRS, 1992-1994, a self-rated health measure, selects household heads of between 45 and 65 years old, and uses an IV method where the health measure is instrumented to estimate the probability of participation in the labour force. Coile (2004) uses the first six waves of the survey, 1992-2002, an objective health measure, selects the household heads of between 50 and 69 years old, and estimates a probit model for the change in labour participation.

Using Mexican data, Van Gameren (2008) uses the first wave of the MHAS and finds that better health causes the Mexican elderly who are married to have a greater attachment to the labour market. Although he considers married couples, he does not take into account spousal health shocks. To my knowledge, there are no earlier studies on the consequences of spousal health shocks for decisions about labour force participation among Mexican

[3] Siegel (2006) finds that the estimated response of the spouses depends on the health measure used in the estimation. The most common measures used in the literature are self-reported measures, objective measures, and self-reported measures instrumented with objective measures.

[4] Coile (2004) defines an acute health shock as a new cancer diagnosis, a heart attack, or a stroke experienced by the individual in the last two years, prior to each wave.

middle-aged to elderly couples.

The present paper adds to the existing literature on the effects of health shocks on spousal labour force participation in several ways. First, I present a theoretical model on which to base the interpretation of the results. The model focuses on how spousal labour force participation responds to negative health shocks. Second, to my knowledge, the present study is the first one that uses all waves of the MHAS, since all waves were not available before 2017. Finally, the study also extends the analysis by considering part-time and full-time employment.

In the theoretical framework I present a model that analyzes the effect of a negative health shock to the individual on the labour supply of his/her partner. The theoretical model is important in the analysis in two ways. First, it provides a framework for interpreting the relationship between health shocks and labour supply behaviour among the household members. The comparative statics of the model show that the healthy spouse may increase (decrease) the labour supply if his or her marginal utility of leisure decreases (increases) with respect to a negative health shock to her or his spouse. Second, it constitutes the starting point for the empirical model used in the estimation. In the empirical analysis I use all four waves of the MHAS and find that the female's probability of participating in the labour force decreases by 37.8% when her spouse faces a major health shock. In contrast, there is no significant effect on the male's probability of participating in the labour force when his spouse faces a major health shock.

I also extend the analysis by examining the effects of a major health shock for the spouse on the probabilities of participating part-time and full-time in the labour force. I use an empirical specification derived from the theoretical model and find that the female's probability of participating full-time decreases with a major health shock to her spouse (the decrease in the full-time probability is 34.7%, while the effect on the part-time probability is not statistically significant). A negative major health shock for the spouse does not have a significant effect on the male's labour force participation behaviour even when I distinguish between part-time and full-time.[5] This suggests that Mexican females take into account negative shocks to their spouse's health in their decisions about labour force participation and that the labour force participation of Mexican males is not significantly affected when their spouses face a major health shock.

I also estimate a specification that includes other health problems such as diabetes, hypertension and arthritis. The results of the latter specification suggest that there is not a significant effect of a spouse facing other health problems on the labour force participation of his or her partner. A possible explanation is that these health problems may not represent a significant impediment to work when treated and controlled (Testa and Simonson, 1998).

The rest of this paper is structured as follows. The theoretical model is developed in section 2. Section 3 discusses the data, and section 4 specifies the empirical model. The results are presented in section 5. Finally, conclusions are drawn in section 6.

2 Theoretical framework

2.1 Theoretical model

The objective of this section is to develop a simple model to characterize the impact of major health shocks on spousal labour supply. I present a household model and adopt an additive utility function represented by the sum of the utility of household consumption and the utilities of each household member's leisure. For simplicity, I assume that major health shocks only affect the marginal utility of leisure, and not the marginal utility of consumption.

Following Ransom (1987), I assume that the household members behave as if they maximize a household utility function subject to a family budget constraint. They maxi-

[5]The point elasticities are small and not significant, -0.001 and -0.01 respectively for part-time and full-time participation.

mize the utility function with respect to three components: household consumption, male's leisure and female's leisure. Let l_m and l_f denote the male and female leisure, respectively, c denote the household consumption, while h_m and h_f denote work hours. Let s_m indicate the male's poor health, i.e. the larger the s_m the poorer the health of the male, and similarly for the female (s_f).

The exogenous wage rate is defined by w_m for the male, and w_m for the female . Household behaviour is thus the solution to the following programme:

$$\max_{\{c,l_m,l_f,h_m,h_f\}} U = r(c) + p(l_m, s_m, s_f) + q(l_f, s_m, s_f)$$

$$s.t. \quad w_m h_m + w_f h_f \geq c,$$

$$T = l_i + h_i \quad i \in \{m, f\},$$

$$h_m \geq 0, h_f \geq 0.$$

The utility function is increasing, twice differentiable in c, l_m, and l_f, as well as strictly quasi-concave. Furthermore, p is decreasing in s_m, and q is decreasing in s_f. Using the budget and time constraints, the household chooses h_m and h_f to maximize

$$r(w_m h_m + w_f h_f) + p(T - h_m, s_m, s_f) + q(T - h_f, s_m, s_f).$$

The first order conditions for an interior maximum are

$$w_m r_c(w_m h_m + w_f h_f) - p_{lm}(T - h_m, s_m, s_f) = 0, \tag{1}$$

$$w_f r_c(w_m h_m + w_f h_f) - q_{lf}(T - h_f, s_m, s_f) = 0, \tag{2}$$

where r_c is the marginal utility of consumption, p_{lm} is the marginal utility of the male's leisure and q_{lf} is the marginal utility of the female's leisure. The issue is how an increase in s_m or s_f, respectively, affects the labour supply of the members in the household.

The expressions for the change in labour supply with respect to a major health shock are obtained by totally differentiating the first order conditions (1) and (2) with respect to

h_m, h_f, s_m, s_f and solving the resulting system of equations

$$dh_m(w_m^2 r_{cc} + p_{l_m l_m}) + dh_f(w_m w_f r_{cc}) - ds_m(p_{l_m s_m}) - ds_f(p_{l_m s_f}) = 0,$$
$$dh_m(w_f w_m r_{cc}) + dh_f(w_f^2 r_{cc} + q_{l_f l_f}) - ds_m(q_{l_f s_m}) - ds_f(q_{l_f s_f}) = 0.$$

If, for example, the male's health declines ($ds_m > 0$) and the female's health is held constant ($ds_f = 0$)[6] we obtain the following expressions for the changes in h_m and h_f:

$$\frac{dh_m}{ds_m} = \frac{p_{l_m s_m}(w_f^2 r_{cc} + q_{l_f l_f}) - q_{l_f s_m} w_m w_f r_{cc}}{w_f^2 r_{cc} p_{l_m l_m} + p_{l_m l_m} q_{l_f l_f} + w_m^2 r_{cc} q_{l_f l_f}}, \tag{3}$$

$$\frac{dh_f}{ds_m} = \frac{q_{l_f s_m}(w_m^2 r_{cc} + p_{l_m l_m}) - p_{l_m s_m} w_m w_f r_{cc}}{w^2 r_{cc} p_{l_m l_m} + p_{l_m l_m} q_{l_f l_f} + w_r^2 u_{cc} q_{l_f l_f}}. \tag{4}$$

The sign of expressions (3) and (4) will determine if the labour supply of each spouse increases or decreases when the male faces an increase in s_m.[7] Since the denominator is positive by the second order conditions, it follows that the numerator will rule the direction of the change.

[6]The solution is analogous when $ds_f > 0$ and $ds_m = 0$.

[7]So far, I have not included medical expenditures. If I were to include them in the budget constraint they would reduce the household's income as a result of poor health. For simplicity, if I were to include medical expenditure I would consider it as exogenous, and it would enter into the analysis through the utility of consumption. This simplification would not affect the qualitative results, since the effect of a major health shock for the spouse on the household members' labour supply would still be determined by (3) and (4).

Another important assumption in this model is that there is no household production. It could be argued that the spouses can allocate time in leisure, the labour market and household production, hence the sign of (3) and (4) could be different. In Appendix A, a version of the model with household production is presented. The results of the comparative statics of the latter model are similar to those of the model without household production. For this reason and the lack of household production data I use the model presented in this section to derive the empirical specification.

Parsons (1977), Riphahn (1999) and Siegel (2006) suggest that a negative shock to an individual's own health reduces the time dedicated to market work. To my knowledge there are no earlier empirical studies suggesting that an individual would work more when becoming ill. Therefore, I focus on the case where the marginal utility of leisure of the ill person (the male) increases with respect to a major health shock, $p_{l_m s_m} > 0$.[8] It follows that the sign of the numerator in (3) and (4) will depend on the sign of the cross effect between the marginal utility of the female's leisure and her spouse's major health shock ($q_{l_f s_m}$).

First, a negative female's cross effect ($q_{l_f s_m} < 0$) is a sufficient condition for the male to work less ($\frac{dh_m}{ds_m} < 0$) and the female to work more ($\frac{dh_f}{ds_m} > 0$). Second, a positive female's cross effect ($q_{l_f s_m} > 0$), and the following constraint:

$$\frac{w_m w_f r_{cc}}{w_m^2 r_{cc} + p_{l_m l_m}} < \frac{q_{l_f s_m}}{p_{l_m s_m}} < \frac{w_f^2 r_{cc} + q_{l_f l_f}}{w_m w_f r_{cc}}, \tag{5}$$

are sufficient conditions for both the male ($\frac{dh_m}{ds_m} < 0$) and female to work less ($\frac{dh_f}{ds_m} < 0$).[9]

The sufficient conditions previously mentioned show that when the male becomes ill

[8] The model also includes the case where the marginal utility of the ill member's leisure decreases when he or she faces a major health shock $p_{l_m s_m} < 0 \quad (q_{l_f s_f} < 0)$. Under this assumption the ill person works more when becoming ill. This case is less likely to happen based on the empirical evidence referred to above.

[9] For $\frac{dh_m}{ds_m} < 0$ the following inequality must hold

$$\begin{aligned} &p_{l_m s_m}(w_f^2 r_{cc} + q_{l_f l_f}) - q_{l_f s_m}(w_m w_f r_{cc}) < 0 \\ &\Longrightarrow \frac{q_{l_f s_m}}{p_{l_m s_m}} < \frac{w_f^2 r_{cc} + q_{l_f l_f}}{w_m w_f r_{cc}}. \end{aligned} \tag{F1}$$

Likewise, for $\frac{dh_f}{ds_m} < 0$ the following inequality must hold

$$\begin{aligned} &q_{l_f s_m}(w_m^2 r_{cc} + p_{l_m l_m}) - p_{l_m s_m}(w_m w_f r_{cc}) < 0 \\ &\Longrightarrow \frac{q_{l_f s_m}}{p_{l_m s_m}} > \frac{w_m w_r u_{cc}}{w_m^2 r_{cc} + p_{l_m l_m}}. \end{aligned} \tag{F2}$$

Inequalities $(F1)$ and $(F2)$ together imply constraint (5),

$$\frac{w_m w_f r_{cc}}{w_m^2 r_{cc} + p_{l_m l_m}} < \frac{q_{l_f s_m}}{p_{l_m s_m}} < \frac{w_f^2 r_{cc} + q_{l_f l_f}}{w_m w_f r_{cc}}.$$

he spends more time on leisure, perhaps to rest and recover. They also show that when the female's cross effect is negative she spends less time on leisure, and works more since perhaps being with an unhealthy partner gives her less utility than when her partner was healthy. Alternatively, when the female's cross effect is positive she works less when her spouse faces an increase in the major health shock; hence, it might be the case that she spends her leisure time taking care of her partner (Siegel 2006).

2.2 Parametric Model.

The functional form of the utility function in the following maximization programme is a special case of the theoretical model discussed in section 2.1. I follow Ngai and Pissarides (2011), and Ragan (2013), and use a quasi-linear functional form to derive expressions that can be used to estimate the probability for the labour force participation of the female and male in the household, respectively. It follows that the latter specification only depends on major health shocks and wage determinants. Hence, the effects of health shocks can be estimated; first on the household's decision to work, and second on the household's decision to work part-time or full-time. The household's decision problem takes the form

$$\begin{aligned}\max_{\{c,h_m,h_f\}} u(c,h_m,h_f,s_m,s_f) = \quad & c + \frac{A}{1-\gamma_m} e^{\eta_m}\left(\frac{T-h_m}{e^{\eta_m}}\right)^{1-\gamma_m} + \frac{B}{1-\gamma_f} e^{\eta_f}\left(\frac{T-h_f}{e^{\eta_f}}\right)^{1-\gamma_f} \\ & s.t. \quad w_m h_m + w_f h_f = c, \\ & h_m \geq 0, h_f \geq 0,\end{aligned}$$

Furthermore,

$$(5) \implies \frac{w_m w_f r_{cc}}{w_m^2 r_{cc} + p_{l_m l_m}} < \frac{w_f^2 r_{cc} + q_{l_f l_f}}{w_m w_f r_{cc}} \qquad \implies w_f^2 r_{cc} p_{l_m l_m} + p_{l_m l_m} q_{l_f l_f} + w_m^2 r_{cc} q_{l_f l_f} > 0.$$

The latter is the denominator in expressions (3) and (4) and is positive by the second order conditions.

where the parameters γ_m and γ_f are positive, η_i is the unobservable characteristics of family member i and has zero mean and constant variance σ_η^2 conditional on s_m and s_f. Let $A = e^{-(\alpha_0+\alpha_m s_m+\alpha_f s_f)}$ and $B = e^{-(\beta_0+\beta_m s_m+\beta_f s_f)}$, where $\alpha_m, \beta_f > 0$. Therefore, the major health shocks affect the utility function through the components A and B. In this way, major health shocks directly affect the spouses' demand for leisure, while the household consumption is affected indirectly via h_m and h_f.

From the maximization programme the first order condition for an interior solution with respect to h_m and h_f are,

$$A(\frac{T-h_m}{e^{\eta_m}})^{-\gamma_m} = w_m, \tag{6}$$

$$B(\frac{T-h_f}{e^{\eta_f}})^{-\gamma_f} = w_f, \tag{7}$$

In this model, the left hand side of equations (6) and (7) represents the marginal rate of substitution of l_m for c, and l_f for c, respectively.

2.2.1 Labour Force Participation

The prior setting allows the household members to choose zero hours of work. The marginal rate of substitution between the male's leisure and the household consumption, given that the male does not work, is given by:

$$A\left(\frac{T}{e^{\eta_m}}\right)^{-\gamma_m}. \tag{8}$$

Therefore, the male chooses to participate if (8) falls short of the wage rate, i.e.

$$\ln A\left(\frac{T}{e^{\eta_m}}\right)^{-\gamma_m} < \ln w_m. \tag{9}$$

I assume that $\ln(w_m) = X_m\zeta_m + \epsilon_m$, where X_m is a vector of observable and exogenous determinants of the male's wage, and the disturbance ϵ_m is assumed to be uncorrelated with X_m having zero mean and constant variance σ_ϵ^2 conditioned on X_m. Equation (9) can be rewritten as follows:

$$\ln A - \gamma_m \ln(T) + \gamma_m \eta_m < X_m \zeta_m + \epsilon_m. \tag{10}$$

With the previous information we can find the male's probability of participating in the labour force

$$\begin{aligned} &Pr(\ln A - \gamma_m \ln(T) + \gamma_m \eta_m < X_m \zeta_m + \epsilon_m | s_m, s_f, X_m) \\ =&Pr(\alpha_0 + \alpha_m s_m + \alpha_f s_f + \gamma_m \ln(T) + X_m \zeta_m + > \gamma_m \eta_m - \epsilon_m | s_m, s_f, X_m). \end{aligned}$$

Define $\Upsilon_{m0} = \alpha_0 + \gamma_m \ln(T)$, and let Y_m be an indicator variable such that $Y_m = 1$ if the individual participates, and $Y_m = 0$ otherwise. Define $\mu = \gamma_m \eta_m - \epsilon_m$; since $\gamma_m \eta_m \sim \mathcal{N}(0, \gamma_m^2 \sigma_\eta^2)$ and $\epsilon_m \sim \mathcal{N}(0, \sigma_\epsilon^2)$ it follows that $\mu \sim \mathcal{N}(0, \gamma_m^2 \sigma_\eta^2 + \sigma_\epsilon^2)$. Finally, if μ is normalized such that $\mu \sim \mathcal{N}(0, 1)$ then a probit model of the following form is obtained:

$$Pr(Y_m = 1 | s_m, s_f, X_m) = \Phi(\Upsilon_{m0} + \alpha_m s_m + \alpha_f s_f + X_m \zeta_m), \tag{11}$$

where $\Phi_m(\cdot)$ is the standard normal distribution function. Likewise, the female's probability of joining the labour force is:

$$Pr(Y_f = 1 | s_m, s_f, X_f) = \Phi(\Upsilon_{f0} + \beta_m s_m + \beta_f s_f + X_f \zeta_f). \tag{12}$$

Expressions (11) and (12) can be used to estimate probit specifications for the probability of labour force participation of the male and female, respectively.

2.2.2 Part-time and Full-time Participation

The previous analysis is limited since it only considers the cases where the individual participates or does not participate in the labour force, but does not distinguish between individuals working full-time or part-time. It could be the case that a negative health-shock influences the decision of whether to work part-time or full-time. If this is true, these effects cannot be observed in the previous model.

I disentangle the extensive and intensive margins of labour force participation to test if there is an effect of a major health shock on the probabilities of working part-time and full-time. To do this I extend the model in section 2.2.1 by redefining the indicator variable Y_i to equal 1 if the member decides not to participate, 2 if the member participates part-time and 3 if the member participates full-time. I follow Zabalza et al. (1980) and define a convex opportunity set that includes the three labour force participation states mentioned before.[10] Each state is associated with a level of consumption c, and leisure time $T - h$, such that $h_1 < h_2 < h_3$, and $c_1 < c_2 < c_3$. For any number of hours of work of the partner, h_j, each household member behaves as if he/she chooses the opportunity that gives the highest utility for the household.

$$Y_i = \begin{cases} 1 \quad if \quad U(c_1, T-h_1, T-h_j) \geq U(c_2, T-h_2, T-h_j), \\ 2 \quad if \quad U(c_2, T-h_2, T-h_j) > U(c_1, T-h_1, T-h_j) \\ \quad and \quad U(c_2, T-h_2, T-h_j) \geq U(c_3, T-h_3, T-h_j), \\ 3 \quad if \quad U(c_3, T-h_3, T-h_j) > U(c_2, T-h_2, T-h_j). \end{cases}$$

For example, the member chooses not to participate in the labour force ($Y_i = 1$) if the utility of not participating is greater than or equal to the utility of participating part-time; chooses part-time if the utility attached to the part-time option exceeds the utilities associated with non-participation and full-time, respectively; and chooses full-time if this choice leads to higher utility than working part-time.

I use the parametric model in section 2.2. Given the spouse's choices, I derive the probability corresponding to each of the three labour market states for, e.g. the male. As in

[10]In the convex opportunity set if an individual prefers state 3 over 2 then he/she will prefer state 3 over 1.

the probit specification of labour force participation μ is normalized such that $\mu \sim \mathcal{N}(0,1)$ (see the Appendix for a mathematical derivation), but in this case the resulting specification is an ordered probit model where the male's probabilities of not participating, participating part-time and participating full-time conditional on the labour participation choice of the other member become

$$Pr(Y_m = 1|\nu_m) = \Phi(\theta_{m1} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m), \tag{14}$$

$$Pr(Y_m = 2|\nu_m) = \Phi(\theta_{m2} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m) - \Phi(\theta_{m1} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m), \tag{15}$$

$$Pr(Y_m = 3|\nu_m) = 1 - \Phi(\theta_{m2} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m), \tag{16}$$

where $\nu_m \equiv (s_m, s_f, X_m)$,
$\theta_{m1} = -\ln(h_2) + \gamma_m \ln(T) - \ln(1-\gamma_m) + \ln\left[\frac{T^{1-\gamma_m} - (T-h_2)^{1-\gamma_m}}{1-\gamma_m}\right]$,
$\theta_{m2} = -\ln(h_3 - h_2) + \gamma_m \ln(T) - \ln(1-\gamma_m) + \ln\left[\frac{(T-h_2)^{1-\gamma_m} - (T-h_3)^{1-\gamma_m}}{1-\gamma_m}\right]$.

From (14) to (16) the empirical specification for the male's probability of choosing between not participating, participating part-time and participating full-time, respectively, is obtained:

$$Pr(Y_m = k|s_m, s_f, X_m) = \Phi(\theta_{mk} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m) - \Phi(\theta_{mk-1} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m), \tag{17}$$

where k equals 1 for the male's probability of choosing to not participate, 2 to participate part-time, and 3 to participate full-time. Likewise, the female's probability of choosing between not participating, participating part-time and participating full-time, respectively, is:

$$Pr(Y_f = k|s_m, s_f, X_f) = \Phi(\theta_{fk} - \Upsilon_{f0} - \beta_m s_m - \beta_f s_f - X_f\zeta_f) - \Phi(\theta_{fk-1} - \Upsilon_{f0} - \beta_m s_m - \beta_f s_f - X_f\zeta_f). \tag{18}$$

3 Data

Individual and household data are obtained from the Mexican Health and Aging Study (MHAS). The MHAS was a collabourative effort among researchers from the Universities of Pennsylvania, Maryland, and Wisconsin in the U.S., and the Instituto Nacional de Estadística, Geografía e Informática (INEGI) in Mexico. The baseline survey, in 2001, included a nationally representative sample of Mexicans aged 50 and above and their spouse or partner. The survey had follow-ups in 2003, 2012 and 2015. In each wave, I use data for couples where both spouses are aged 50 to 75 years old. Finally, the sample has an unbalanced panel structure and its size is 7,270 female observations and 8,632 male observations.[11]

Tables 1 and 2 show the mean values of the variables used in this study for females and males, respectively.[12] The mean values are measured over all the four waves and are presented by the labour force participation states: not in the labour force, in the labour force, part-time, and full-time. Figures 1 and 2 in Appendix D show that a small fraction of females and males work seven to eight hours per day, and that there are spikes at eight working hours per day for both females and males. Hence, I consider that the household member works full-time if he or she works eight or more hours per day, and if the individual works more than zero and less than eight hours per day I consider that he or she works part-time.[13]

The mean age was higher for males that did not take part in the labour force, 66.7 years old, than for males that did participate, 62.1 years old. Females were, on average, younger than males. Females that did not participate in the labour force had a mean age

[11]The number of male and female observations differs due to missing values in the female data.

[12]See Table 5 in the Appendix for variable definitions.

[13]Furthermore, in Mexico Art. 61 in the Federal Law of labour states that the maximum duration of a work day is eight hours. The law also states that a work day can be extended, and that the worker does not receive extra payment if the extension does not exceed three hours per day and three times per week. Otherwise the worker must be compensated (Art. 66 and 68 in the Federal Law of labour).

of 62.2 years and females that did participate had a mean age of 58.6 years.

The variable major health shock indicates if the individual has had a new cancer diagnosis, a heart attack or a stroke in the last two years previous to the corresponding wave of the survey. Table 1 shows that 2% of all females, 3% of females that were not participating, and 2% of participating females experienced a major health shock. Table 2 shows that 3% of all males reported a major health shock. The proportion of males that reported a major health shock was greater for those that were not in the labour force, 5%, than for those that were part of it, 2%.

Table 2 shows that 3% of the spouses of males experienced a major health shock. The incidence of a major health shock for the spouse was 3% when the male was participating and when he was not participating. In contrast, there is a larger variation in Table 1; when females were participating 2% of their spouses faced a major health shock while 3% of their spouses faced a major health shock when females did not participate in the labour force.

Table 1 shows that the incidence of a major health shock for females who worked part-time and females who worked full-time was 2%. The incidence of a major health shock was higher for males who worked part-time than for those who worked full-time; for instance, Table 2 shows that 2% of males who worked part-time reported a major health shock while only 1% of males who worked full-time did so.

The statistics of a major health shock for the spouse show a different pattern. The incidence of a major health shock for the spouse was the same for females who worked full-time and females who worked part-time, 2% in both cases. In contrast, the incidence of a major health shock for the spouse was higher for males who worked part-time than for males who worked full-time, 3% versus 2%.

The data also show that the household members participated more when their partners also participated in the labour force, for example, 28% of females participate in the labour force when their spouses were part of the labour force. In contrast, 19% of females participate in the labour force when their spouses did not participate in the labour force.

Similarly, the labour force participation rate among males was 73% when their partner participated, and 62% when the partner did not.

Table 1 shows that 71% of the spouses of females who worked part-time participated in the labour force. When the females worked full-time 76% of the spouses worked. In Table 2 the proportion of females participating in the labour force was similar in both cases when their spouses worked part-time and full-time, female participation was approximately 26%.

Table 1: Summary Statistics (Female)

	Total	**Not In labour Force**	**In labour Force**	**Part-time**	**Full-time**
Age	61.2	62.2	58.6	58.9	58.1
	(6.35)	(6.29)	(5.72)	(5.79)	(5.56)
Major Health Shock	0.02	0.03	0.02	0.02	0.02
	(0.16)	(0.17)	(0.15)	(0.15)	(0.14)
Other Health Problem*	0.64	0.66	0.56	0.57	0.55
(Diabetes,hypertension and/or arthritis)	(0.48)	(0.47)	(0.49)	(0.49)	(0.49)
Years of Education	4.86	4.38	6.03	6.14	5.83
	(4.20)	(3.88)	(4.69)	(4.78)	(4.55)
Work Experience	10.5	8.46	15.7	15.2	16.6
	(11.4)	(10.6)	(11.8)	(11.9)	(11.7)
Spouse labour Force	0.65	0.62	0.73	0.71	0.76
Participation	(0.47)	(0.48)	(0.44)	(0.45)	(0.45)
Spouse Major Health Shock	0.03	0.03	0.02	0.02	0.02
	(0.17)	(0.18)	(0.15)	(0.14)	(0.14)
Spouse Other Health Problem*	0.59	0.61	0.53	0.52	0.54
(Diabetes,hypertension and/or arthritis)	(0.49)	(0.49)	(0.49)	(0.49)	(0.49)
Urban Area	0.72	0.71	0.76	0.76	0.75
	(0.44)	(0.46)	(0.42)	(0.42)	(0.43)
Number of Children	5.15	5.41	4.52	4.37	4.76
	(2.81)	(2.82)	(2.66)	(2.61)	(2.74)
N	7,270	5,161	2,109	1,491	618

Standard deviations in parenthesis. Source: http://www.mhasweb.org/Data.aspx, Mexican Health and Aging Study (MHAS). Note. On average, each couple is observed twice in the sample. *The subsample for other health problems consists of 6,905 observations. 29% of all females were in the labour force, of these 61% worked part-time and 39% worked full-time.

Table 2: Summary Statistics (Male)

	Total	**Not In labour Force**	**In labour Force**	**Part-time**	**Full-time**
Age	63.5	66.7	62.1	63.9	61.1
	(6.41)	(5.47)	(6.26)	(5.95)	(6.19)
Major Health Shock	0.03	0.05	0.02	0.02	0.01
	(0.17)	(0.22)	(0.16)	(0.16)	(0.13)
Other Health Problem*	0.47	0.57	0.42	0.45	0.41
(Diabetes,hypertension and/or arthritis)	(0.49)	(0.49)	(0.49)	(0.49)	(0.49)
Years of Education	5.67	5.46	5.77	6.07	5.47
	(4.90)	(4.86)	(5.00)	(5.19)	(4.79)
Work Experience	27.4	27.5	27.4	27.6	4.79
	(12.3)	(11.3)	(12.8)	(13.3)	(12.5)
Spouse labour Force	0.25	0.19	0.28	0.26	0.27
Participation	(0.43)	(0.39)	(0.45)	(0.44)	(0.44)
Spouse Major Health Shock	0.03	0.03	0.02	0.03	0.02
	(0.17)	(0.17)	(0.17)	(0.18)	(0.15)
Spouse Other Health Problem	0.52	0.56	0.51	0.51	0.52
(Diabetes, hypertension and/or arthritis)	(0.49)	(0.49)	(0.49)	(0.50)	(0.49)
Urban Area	0.72	0.79	0.69	0.70	0.70
	(0.44)	(0.40)	(0.45)	(0.45)	(0.45)
Number of Children	5.11	5.31	5.01	4.94	5.09
	(2.75)	(2.78)	(2.73)	(2.72)	(2.81)
N	8,632	3,105	5,527	3,241	2,286

Standard deviations in parenthesis. Source: http://www.mhasweb.org/Data.aspx, Mexican Health and Aging Study (MHAS). Note. On average, each couple is observed twice in the sample. *The subsample for other health problems consists of 8,158 observations. 67% of all males were in the labour force, of these 59% worked part-time and 41% worked full-time.

4 Empirical Specification

Two models are estimated. First, a model where each household member chooses between participating or not in the labour force. Equations (11) and (12) in section 2.2.1 suggest that I can estimate separate probit specifications of the male's and the female's labour force participation, respectively:

$$Pr(Y_{mt} = 1|s_{mt}, s_{ft}, X_{mt}) = \Phi(\Upsilon_{m0} + \alpha_m s_{mt} + \alpha_f s_{ft} + X_{mt}\zeta_m),$$
$$Pr(Y_{ft} = 1|s_{mt}, s_{ft}, X_{ft}) = \Phi(\Upsilon_{f0} + \beta_m s_{mt} + \beta_f s_{ft} + X_{ft}\zeta_f),$$

where the dependent variable is a binary variable indicating if the individual participates in the labour force or not, s_m and s_f are dummy variables for whether the male and the female have experienced a major health shock during the last two years prior to the wave. X is a vector of explanatory variables that include the individual's age, age squared, years of education, working experience, the square of working experience, the number of children, a dummy variable indicating whether the household is located in an urban area or not, and wave dummies.

Second, equations (17) and (18) in section 2.2.2 suggest that I can use an ordered probit model (Maddala, 1986) to estimate the female's and male's probabilities of not participating, participating part-time and participating full-time, respectively, in the labour force. I consider the same explanatory variables as in the binary probit specification and estimate separate ordered probit models for males and females, respectively.

$$\begin{aligned} Pr(Y_{mt} = k | s_{mt}, s_{ft}, X_{mt}) &= \Phi(\theta_{mk} - \Upsilon_{m0} - \alpha_m s_{mt} - \alpha_f s_{ft} - X_{mt}\zeta_m) \\ &\quad - \Phi(\theta_{mk-1} - \Upsilon_{m0} - \alpha_m s_{mt} - \alpha_f s_{ft} - X_{mt}\zeta_m), \\ Pr(Y_{ft} = k | s_{mt}, s_{ft}, X_{ft}) &= \Phi(\theta_{fk} - \Upsilon_{f0} - \beta_m s_{mt} - \beta_f s_{ft} - X_{ft}\zeta_f) \\ &\quad - \Phi(\theta_{fk-1} - \Upsilon_{f0} - \beta_m s_{mt} - \beta_f s_{ft} - X_{ft}\zeta_f), \end{aligned}$$

One concern that arises is that major health shocks may, nevertheless, be endogenous. For example, working in some industries can increase the risks of cancer due to exposure to chemicals or radiation. However, Dwyer and Mitchell (1999) instrument objective health measures such as cancer diagnoses, heart attacks, and strokes, and find little evidence of endogeneity.[14] In the case of the present study, the variables that could be used as

[14]The instruments used by the authors are parents' health and mortality, a dummy variable indicating if the parents were ill prior to death, a dummy variable indicating if the parents needed help with basic daily activities, the respondent's weight/height ratio and nights spent at the hospital. The authors estimate OLS and IV regressions of retirement models where the explanatory variables include a measure of objective health. Following these models, Dwyer and Mitchell (1999) carry out the Hausman-Wu Test for exogeneity

instruments such as parents' health and mortality, the respondent's weight/height ratio and nights spent at the hospital (Dwyer and Mitchell, 1999) were not available or had a large number of missing observations in the data.

Another concern arises with the estimation method, since I would need to estimate probit and ordered probit models where the major health shock would be a binary endogenous explanatory variable. Some common approaches that are used to estimate the latter are linear probability methods, maximum likelihood and control functions. But all of these estimation methods have limitations; linear probability methods produce predicted probabilities that are close to zero or one; maximum likelihood methods for endogenous variables require full specification of the first stage model; and control functions require the endogenous regressors to be continuous (Lewbel, Dong and Yang, 2012). Given the data in this study there is not enough information to consider major health shocks to be endogenous. Therefore, I follow the empirical evidence provided by Dwyer and Mitchell (1999), and assume like Blau (1998), Blau and Riphahn (1998), and Berger and Pelkowski (2004) that major health shocks are exogenous.

5 Results

5.1 Major Health Shocks

Tables 3 and 4 present the average semi-elasticities of the female and male specifications, respectively. The semi-elasticities are calculated at the observation level and are then averaged. These semi-elasticities estimate the percentage change in the probability of participation when the individual or the spouse faces a major health shock (an absolute change from 0 to 1). Specification Ia considers the labour force participation model, while specification IIa accounts for part-time and full-time labour force participation.[15]

and cannot reject the null hypothesis that estimated effects are equal when using OLS and IV, respectively.

[15]The parameter estimates of all specifications as well as the marginal effects of these can be found in the Appendix C. Specification b includes a binary variable named other health problem that indicates whether

Table 3: Female's Semi-elasticities and Elasticities of Probit and Ordered Probit

	Ia	**IIa**	**IIa**	**Ib**	**IIb**	**IIb**
	LFP	**Part-time**	**Full-time**	**LFP**	**Part-time**	**Full-time**
Semi-elasticities						
Major Health Shock	-0.245†	-0.251*	-0.465*	-0.232	-0.238†	-0.443*
	(0.147)	(0.119)	(0.213)	(0.153)	(0.124)	(0.222)
Male Major Health Shock	-0.378**	-0.185	-0.347†	-0.343*	-0.146	-0.276
	(0.138)	(0.113)	(0.206)	(0.142)	(0.116)	(0.213)
Other Health Problem				-0.157**	-0.125**	-0.245**
				(0.046)	(0.037)	(0.073)
Male Other Health Problem				-0.042	-0.042	-0.083
				(0.043)	(0.037)	(0.072)
Urban	0.031	0.041	0.078	0.027	0.033	0.064
	(0.060)	(0.043)	(0.084)	(0.062)	(0.046)	(0.089)
Elasticities						
Age	-7.050***	-4.417***	-8.118***	-7.262***	-4.623***	-8.420***
	(0.474)	(0.349)	(0.549)	(0.512)	(0.380)	(0.595)
Years of Education	0.046	0.015	0.032	0.037	0.009	0.020
	(0.030)	(0.017)	(0.037)	(0.031)	(0.018)	(0.039)
Work Experience	0.253***	0.159***	0.344***	0.246***	0.158***	0.338***
	(0.024)	(0.015)	(0.036)	(0.025)	(0.016)	(0.038)
Number of Kids	-0.072	-0.055	-0.103	-0.087	-0.057	-0.106
	(0.060)	(0.041)	(0.077)	(0.063)	(0.044)	(0.081)
Observations	7270	7270	7270	6905	6905	6905

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. LFP: labour Force Participation. Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command. All specifications include age squared, working experience squared, and year controls. The semi-elasticity is the percentage change in the probability of participating in the labour force relative to an absolute change in the factor variable in the labour force relative to an absolute change in the factor variable (The discrete change from 0 to 1). The quantities in the table need to be multiplied by 100 in order to be transformed into percentages.

the individual and her/his spouse had diabetes, arthritis and/or hypertension. This specification will be discussed in the next subsection.

Table 4: Male's Semi-elasticities and Elasticities of Probit and Ordered Probit

	Ia	IIa	IIa	Ib	IIb	IIb
	LFP	**Part-time**	**Full-time**	**LFP**	**Part-time**	**Full-time**
Semi-elasticities						
Major Health Shock	-0.321***	-0.171***	-0.746***	-0.233***	-0.114**	-0.571***
	(0.066)	(0.042)	(0.128)	(0.063)	(0.036)	(0.127)
Female Negative Major Health Shock	0.001	-0.006	-0.055	0.019	-0.001	-0.010
	(0.049)	(0.012)	(0.095)	(0.049)	(0.010)	(0.097)
Other Health Problem				-0.138***	-0.028***	-0.273***
				(0.019)	(0.004)	(0.037)
Female Other Health Problem				-0.028	-0.005	-0.053
				(0.017)	(0.003)	(0.036)
Urban	-0.104***	-0.011***	-0.129**	-0.107**	-0.011**	-0.124**
	(0.023)	(0.003)	(0.039)	(0.024)	(0.003)	(0.041)
Elasticities						
Age	-3.888***	-0.824***	-6.147***	-3.833***	-0.779***	-5.983***
	(0.185)	(0.133)	(0.277)	(0.193)	(0.134)	(0.284)
Years of Education	0.008	-0.004†	-0.043*	0.008	-0.004*	-0.049*
	(0.013)	(0.002)	(0.021)	(0.013)	(0.002)	(0.022)
Work Experience	0.347***	0.037***	0.492***	0.347***	0.031***	0.500***
	(0.023)	(0.005)	(0.044)	(0.023)	(0.005)	(0.045)
Number of Kids	0.008	-0.001	-0.009	0.009	-0.002	-0.014
	(0.021)	(0.004)	(0.035)	(0.021)	(0.004)	(0.036)
Observations	8632	8632	8632	8158	8158	8158

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. LFP: labour Force Participation. Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command. All specifications include age squared, working experience squared, and year controls. The semi-elasticity is the percentage change in the probability of participating in the labour force relative to an absolute change in the factor variable in the labour force relative to an absolute change in the factor variable (The discrete change from 0 to 1). The quantities in the table need to be multiplied by 100 in order to be transformed into percentages.

Table 3 shows that a major shock to the female's own health is estimated to reduce her likelihood of participating in the labour force by approximately 24.5%. Both the likelihood to work part-time and full-time are also reduced by a major health shock but the effect on full-time is larger in absolute value than the effect on part-time, the reduction corresponds to 46.5% and 25.1%, respectively. The onset of a spouse major health shock decreases the female's probability of participating in the labour force by 37.8%, the probability of participating full-time by 34.7%, and there is not a statistically significant effect on her probability of participating part-time.

The results for males are shown in Table 4. The effect of the male's own major health shock is negative and significant in all specifications. The male's probability of participating in the labour force is reduced by 32.1% and 23.3% in specifications Ia and Ib, respectively. There is a larger effect of a major health shock on the full-time than the part-time regime. For example, in specification IIa the probability of participating part-time is reduced by 17.1% while the probability of participating full-time is reduced by 74.6%. The estimated effect of a major health shock for the female is insignificant and close to zero. Therefore, a negative health shock for women does not seem to affect their husband's labour supply behaviour.

According to the empirical evidence in this paper and the theoretical model in section 2.1, the results suggest that the marginal utility of the female's non-working time increases when her husband faces a major health shock; as a result the female works less, perhaps to take care of the ill spouse. In contrast, the marginal utility of the male's non-working time does not seem to change with respect to a major health shock to his spouse.

Work experience has a statistically significant effect in this analysis (Table 3 and Table 4). In the case of the female, a 1% increase in years of work experience increases her probability of participating in the labour force by 0.25% in specification Ia. In the case of the male, his probability of participating in the labour force increases by 0.34% when there is a 1% increase in his years of work experience in specifications Ia and Ib, respectively.

5.2 Other Health Problems

While a major health shock is unexpected and may affect the individual in a severe way, other health problems such as diabetes, arthritis and hypertension are not a significant impediment to the individual when treated and controlled (Testa and Simonson 1998). These health problems should have a lesser effect than major health problems.

I define the variable 'other health problem' as a binary variable that takes the value 1 if the individual has diabetes, hypertension and/or arthritis; otherwise the variable takes the value 0. The same definition is used for the spouse's other health problems. I include other health problem and spouse other health problem in specification Ib (the probit model) and specification IIb (the ordered probit model), together with the explanatory variables used in specification Ia and IIa.

Table 3 shows that other health problems do decrease the female's probability of being in the labour force, but in a smaller magnitude than a major health shock. For example, in specification IIb, a major health problem significantly decreases the female probability of participating full-time by 44.3% while other health problems decrease it by 24.5%. The results show that the estimates of the effects of a female's major health problem do not change much when including other health problems in the specification.

Other health problems in specification IIa reduce the female's probability of working full-time by 46.5% and reduce her probability of working part-time by 25.1%. Furthermore, women's labour force participation behaviour is not significantly affected by the other health problems facing the spouse. Since conditions such as diabetes are not a significant impediment to work when treated and controlled, it can be the case that the female is not largely affected by her spouse's condition when making her labour force participation decision.

In response to another health problem, the male's probability of participating decreases by a smaller amount than when he faces a major health problem. Table 4 shows that other health problems have a negative effect on the probability of participating in the labour force. For instance, the full-time probability decreases by 27.3% while the part-time probability decreases by 2.8%.

5.3 Sensitivity Analysis

I considered two upper age limits other than the 75 years old limit, namely 70 and 80 years old. The results are similar to those obtained in the main analysis, i.e. in the regressions with the 70 years limit (Table 10 in Appendix C), a major health shock to the male reduces the female's probability of participating in the labour force by 39.6% in specification Ia. In specification IIa, the probability of participating part-time is reduced by 26.8%, while the probability of participating full-time is reduced by 51.2%. When the upper limit is defined as 80 years old (Table 11 in Appendix C), specification Ia shows that the female's probability of participating in the labour force is reduced approximately by 37.5%, while specification IIa shows that the probability of participating part-time and full-time are reduced by 24.3% and 45.1%, respectively. The results are also similar for own major health shock, own other health problems, and the male's other health problems. The male's probability of participating in the labour force, part-time and full-time, remains unaffected by his spouse facing a major health shock when the age upper limits are changed to 70 or to 80 years old (Table 12 and 13 in Appendix C).

I also consider another specification with two interaction effects; one between the individual's major health shock and the individual's age, and a second one between the spouse's major health shock and the spouse's age. If the female faces the major health shock then her probability of participating in the labour force is reduced by 41.9%, but if it is her spouse facing this type of shock then her likelihood of participating in the labour force is reduced by 33.8% (Table 14 in Appendix C). Similarly, her probability of participating part-time and full-time is reduced by 21.8%, and 40.8%, respectively when her husband faces a major health shock. The semi-elasticity of the own major health shock is larger in absolute value when compared to the model in Table 3 (24.5%), while the semi-elasticity of the spouse major health problem is smaller in absolute value than in Table 3 (37.8%). The female's interaction effect between the major health shock and age is negative and statistically significant while the spouse's interaction effect is not. In the male's case the interaction effects are not statistically significant. The male's probability of participating is reduced by 32.1% when he faces a major health shock and as before there is not a statistically significant effect when his spouse faces this kind of shock (Table 15 in Appendix C).

The results seem to be similar, except for the female own negative health shock, to those in Tables 3 and 4. In the case of the female facing a major health shock, the result might suggest that the effect of major health shocks differ at different ages, and that it might be more severe the older the female is. When her husband faces a major health shock, the results are similar to those presented in subsection 5.1.

6 Conclusions

In this paper, I find that spouses' major health shocks such as new cancer diagnoses, heart attacks, and strokes do not have a significant effect on the labour supply decisions of males. In contrast, the effect is large and significant for females. They are 37.8% less likely to participate in the labour force if their partner faces a major health shock. Likewise, when I classify labour force participation into part-time and full-time there is a reduction of 34.7% in the full-time regime, while there is not a statistically significant effect in the part-time regime.

As outlined by the theoretical model in section 2.1, the results suggest that the average response of Mexican females between 50 and 75 years old to a spousal major health shock is described by an increase in the marginal utility of the females' leisure, which is a sufficient condition to decrease their labour force participation. Perhaps females feel that spending time with their ill spouse is more rewarding than working, hence females will reduce work time and spend leisure time taking care of their partner. The results also suggest that there is no significant change in the males' labour force participation when their spouse becomes ill.

The findings of the present paper may be of particular interest to policy makers, as they seem to suggest that females act as caregivers when males fall ill. If the females' labour income is an important component of the household labour income, and if they have to reduce their work time or even stop working to take the roll of caregivers, households would experience a loss in labour income. This loss, and having a sick or ill member, which potentially would add medical expenditures to the household, might create an adverse environment for the household's well-being. Therefore, to protect, maintain, and even increase women's labour supply and consequently the household's well-being, it is important for policy makers to design strategies for making formal caregiving more accessible to Mexican households. The latter is of concern since Mexico is one of three OECD countries without universal health coverage and it is also one of OECD countries with higher out of pocket medical expenditure; in 2016 it represented nearly 50% of total health spending (OECD, 2016).

The estimates of the effects of major health shocks reported in this paper are larger than those obtained by Coile (2004) but smaller than the estimates reported by Siegel (2006). Coile (2004) finds that females in the US decrease labour force participation by approximately 6 percentage-points when their spouse faces major health shocks that are accompanied by a large decrease in functioning. Similarly, using US data Siegel (2006) finds that wives of men who had heart problems reduce their labour force participation by approximately 11 percentage-points. In this study females reduce their labour force participation by 7.3 percentage-points (I mention the average marginal effect and not the semi-elasticity to make the estimates comparable). This difference in results may in part be due to differences between the access and cost to medical attention in Mexico and the US.

Future research can be enriched by taking into account the access to medical attention and formal caregiving for Mexican households. Furthermore, for a more extensive analysis, the estimation of the household member's decision to participate in the labour force could use some of the partner's characteristics as explanatory variables, i.e. partner's labour force participation, partner's wage, partner's age, etc. Finally, it is also important to revisit the theory and consider a theoretical framework that includes medical expenses where major health shocks are endogenous.

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Appendix A A model with household production.

Following Ngai and Pissarides (2011), and Ragan (2013), I assume that the household members behave as if they maximize a household utility function with respect to household consumption of goods produced in the market and at home, male's leisure, and female's leisure. The household members maximize the household utility function subject to a family budget constraint. Let l_m and l_f denote leisure, t_m and t_f time dedicated to home work, and h_m and h_f denote market work hours, respectively for the male and female. Furthermore, let c denote the household consumption of market goods, and n denote consumption of the home produced good. Let s_m and s_f be an indicator of illness (or sickness), i.e., the larger s_m, the poorer is the health of the male, and similarly for s_f. The exogenous wage rate is defined by w. Household behavior is thus the solution to the following programme:

$$
\begin{aligned}
\max_{\{c,n,l_m,h_m,t_m,l_f,h_f,t_f\}} \quad & U = r(c) + g(n) + p(l_m, s_m, s_f) + q(l_f, s_m, s_f) \\
s.t. \quad & w_m h_m + w_f h_f \geq c, \\
& n = z(t_m, t_f), \\
& l_i \geq 0, h_i \geq 0, t_i \geq 0 \quad i \in \{m, f\}, \\
& T = l_i + h_i + t_i \quad i \in \{m, f\}.
\end{aligned}
$$

The utility function is increasing, strictly quasi-concave, and twice differentiable in c, n, l_m, and l_f. Furthermore, p is decreasing in s_m, and q is decreasing in s_f. Likewise, the home production function z is increasing in t_m and t_f, respectively. Additionally, z exhibits diminishing marginal productivity. Thus, using the budget, time and home production function constraints, the household chooses $\{h_m, t_m, h_f, t_f\}$ to maximize

$$
r(w_m h_m + w_f h_f) + g(z(t_m, t_f)) + p(T - h_m - t_m, s_m, s_f) + q(T - h_f - t_f, s_m, s_f).
$$

The first order conditions with respect to h_m, h_f, t_m, and t_f for a maximum are

$$w_m r_c - p_{l_m} = 0, \tag{A1}$$

$$w_f r_c - q_{l_f} = 0, \tag{A2}$$

$$g_n z_{t_m} - p_{l_m} = 0, \tag{A3}$$

$$g_n z_{t_f} - q_{l_f} = 0, \tag{A4}$$

where r_c is the marginal utility of consumption, and g_n is the marginal utility of consumption of the home produced good. Furthermore, z_{t_m} and z_{t_f} are the marginal productivity of the male's and female's home work, respectively. Finally, p_{l_m} is the marginal utility of the male's leisure and q_{l_f} is the marginal utility of the female's leisure . The issue is how a negative health shock, a decrease in s_m or s_f, affects the labour supply of the members in the household.

The expressions for the change in labour supply with respect to health are obtained by totally differentiating the first order conditions (A3)-(A4) with respect to $h_m, h_f, t_m, t_f, s_m, s_f$

$$dh_m(w_m^2 r_{cc} + p_{l_m l_m}) + dh_f(w_m w_f r_{cc}) + dt_m(p_{l_m l_m}) - ds_m(p_{l_m s_m}) - ds_f(p_{l_m s_f}) = 0, \tag{A5}$$

$$dh_m(w_f w_m r_{cc}) + dh_f(w_f^2 r_{cc} + q_{l_f l_f}) + dt_f(q_{l_f l_f}) - ds_m(q_{l_f s_m}) - ds_f(q_{l_f s_f}) = 0, \tag{A6}$$

$$\begin{aligned} &dh_m(p_{l_m l_m}) + dt_m(g_{nn} z_{t_m}^2 + g_n z_{t_m t_m} + p_{l_m l_m}) + dt_f(g_{nn} z_{t_f} z_{t_m} + g_n z_{t_m t_f}) \\ &- ds_m p_{l_m s_m} - ds_f p_{l_m s_f} = 0, \end{aligned} \tag{A7}$$

$$\begin{aligned} &dh_f(q_{l_f l_f}) + dt_f(g_{nn} z_{t_f}^2 + g_n z_{t_f t_f} + q_{l_f l_f}) + dt_m(g_{nn} z_{t_m} z_{t_f} + g_n z_{t_f t_m}) \\ &- ds_m q_{l_f s_m} - ds_f q_{l_f s_f} = 0. \end{aligned} \tag{A8}$$

If for example, there is a negative shock to the male's health ($ds_m < 0$), the female's health is held constant ($ds_f = 0$)[16], and dividing (A5)-(A8) by ds_m we obtain the following expressions:

[16]The solution is analogous when $ds_f < 0$ and $ds_m = 0$.

$$\frac{dh_m}{ds_m}(w_m^2 r_{cc} + p_{l_m l_m}) + \frac{dh_f}{ds_m}(w_m w_f r_{cc}) + \frac{dt_m}{ds_m}(p_{l_m l_m}) = p_{l_m s_m}, \tag{A9}$$

$$\frac{dh_m}{ds_m}(w_f w_m r_{cc}) + \frac{dh_f}{ds_m}(w_f^2 r_{cc} + q_{l_f l_f}) + \frac{dt_f}{ds_m}(q_{l_f l_f}) = q_{l_f s_m}, \tag{A10}$$

$$\frac{dh_m}{ds_m}(p_{l_m l_m}) + \frac{dt_m}{ds_m}(g_{nn} z_{t_m}^2 + g_n z_{t_m t_m} + p_{l_m l_m}) + \frac{dt_f}{ds_m}(g_{nn} z_{t_f} z_{t_m} + g_n z_{t_m t_f}) = p_{l_m s_m}, \tag{A11}$$

$$\frac{dh_f}{ds_m}(q_{l_f l_f}) + \frac{dt_f}{ds_m}(g_{nn} z_{t_f}^2 + g_n z_{t_f t_f} + q_{l_f l_f}) + \frac{dt_m}{ds_m}(g_{nn} z_{t_m} z_{t_f} + g_n z_{t_f t_m}) = q_{l_f s_m}. \tag{A12}$$

The system of equations (A9)-(A12) can be solved to obtain the change in h_m and h_f with respect to s_m:

$$\frac{dh_m}{ds_m} = \frac{p_{l_m s_m}[(w_f^2 r_{cc} + q_{l_f l_f})(ab - x^2) - q_{l_f l_f}^2 a] - q_{l_f s_m}[w_f w_m r_{cc}(ab - x^2) + p_{l_m l_m} q_{l_f l_f} x]}{\Gamma}, \tag{A13}$$

$$\frac{dh_f}{ds_m} = \frac{q_{l_f s_m}[(w_m^2 r_{cc} + p_{l_m l_m})(ab - x^2) - p_{l_m l_m}^2 b] - p_{l_m s_m}[w_m w_f r_{cc}(ab - x^2) + q_{l_f l_f} p_{l_m l_m} x]}{\Gamma}, \tag{A14}$$

where

$$a = g_{nn} z_{t_m}^2 + g_n z_{t_m t_m} + p_{l_m l_m} < 0, \tag{A15}$$

$$b = g_{nn} z_{t_f}^2 + g_n z_{t_f t_f} + q_{l_f l_f} < 0, \tag{A16}$$

$$x = g_{nn} z_{t_m} z_{t_f} + g_n z_{t_f t_m} < 0, \tag{A17}$$

$$ab - x^2 > 0. \tag{A18}$$

$$\begin{aligned}\Gamma &= (w_m^2 r_{cc} + p_{l_m l_m})[(w_f^2 r_{cc} + q_{l_f l_f})(ab - x^2) - q_{l_f l_f}^2 a] - (w_m w_f r_{cc})^2 ab - \\ &(w_m w_f r_{cc}) q_{l_f l_f} p_{l_m l_m} x - p_{l_m l_m}[w_f w_m r_{cc} q_{l_f l_f} x + (w_f^2 r_{cc} + q_{l_f l_f}) p_{l_m l_m} b + q_{l_f l_f} p_{l_m l_m}],\end{aligned} \tag{A19}$$

where $\Gamma > 0$. The expressions for the model without household production are

$$\frac{dh_m}{ds_m} = \frac{p_{l_m s_m}(w_f^2 r_{cc} + q_{l_f l_f}) - q_{l_f s_m}(w_m w_f r_{cc})}{w_f^2 r_{cc} p_{l_m l_m} + p_{l_m l_m} q_{l_f l_f} + w_m^2 r_{cc} q_{l_f l_f}}, \tag{A20}$$

$$\frac{dh_f}{ds_m} = \frac{q_{l_f s_m}(w_m^2 r_{cc} + p_{l_m l_m}) - p_{l_m s_m}(w_m w_f r_{cc})}{w_f^2 r_{cc} p_{l_m l_m} + p_{l_m l_m} q_{l_f l_f} + w_m^2 r_{cc} q_{l_f l_f}}. \tag{A21}$$

The comparative statics in both models look similar, though in the model with home production the change in work hours will also depend on the following three expressions:

$$(w_f^2 r_{cc} + q_{l_f l_f})(ab - x^2) - q_{l_f l_f}^2 a < 0, \tag{A22}$$

$$w_f w_m r_{cc}(ab - x^2) + p_{l_m l_m} q_{l_f l_f} x < 0, \tag{A23}$$

$$(w_m^2 r_{cc} + p_{l_m l_m})(ab - x^2) - p_{l_m l_m} b < 0. \tag{A24}$$

Since (A22)-(A24) are negative then (A20) and (A21) will only differ from (A13) and (A14), respectively, in magnitude and not in direction. Hence, if $p_{l_m l_m} < 0$ and $q_{l_f s_m} > 0$ then $\frac{dh_m}{ds_m} > 0$ and $\frac{dh_f}{ds_m} < 0$, as in the model with out household production.

If it is the case that $p_{l_m s_m} < 0$, and $q_{l_f s_m} < 0$, it must also be true that

$$\left|p_{l_m s_m}[(w_f^2 r_{cc} + q_{l_f l_f})(ab - x^2) - q_{l_f l_f}^2 a]\right| > \left|q_{l_f s_m}[w_f w_m r_{cc}(ab - x^2) + p_{l_m l_m} q_{l_f l_f} x]\right|,$$

and

$$\left|q_{l_f s_m}[(w_m^2 r_{cc} + p_{l_m l_m})(ab - x^2) - p_{l_m l_m}^2 b]\right| > \left|p_{l_m s_m}[w_m w_f r_{cc}(ab - x^2) + q_{l_f l_f} p_{l_m l_m} x]\right|,$$

to obtain $\frac{dh_m}{ds_m} > 0$ and $\frac{dh_f}{ds_m} > 0$. The results of the comparative statics of the model with household production are similar to those to the reduced model with out household production, the difference is in magnitude and not in the direction of the change.

Appendix B Variable Definitions

Table 5: Variable Definitions

Age	Age of the household member at the time of the survey's wave.
Major Health Shock	Equals to one if the respondent had have a new cancer diagnose, a heart attack and/or a stroke two years previous to the survey's wave, and it equals zero otherwise.
Other Health Problem	Equals to one if the respondent had diabetes, hypertension or arthritis at the time of the survey's wave, and it equals zero otherwise.
labour Force Participation	Equals to one if the respondent had a job during the week prior to the survey's wave, and zero otherwise.
Years of Education	The respondent's years of schooling at the time of the survey's wave.
Work Experience	The respondent's accumulated number of years devoted to the activities of the main job, throughout life at the time of the survey's wave.
Spouse major Health Shock	Equals to one if the spouse of the respondent had have a new cancer diagnose, a heart attack and/or a stroke two years previous to the survey's wave, and it equals zero otherwise.
Spouse Other Health Problem	Equals to one if the spouse of the respondent had diabetes, hypertension or arthritis at the time of the survey's wave, and it equals zero otherwise.
Spouse labour Force Participation	Equals to one if the spouse of the respondent had a job during the week prior to the survey's wave, and zero otherwise.
Urban Area	Equals to one if the household was located in an urban area at the time of the survey's wave, and it equals zero otherwise.
Number of Children	The household's number of children who were alive at the time of the survey's wave.

Data source: http://www.mhasweb.org/Data.aspx, Mexican Health and Aging Study (MHAS).

Appendix C Estimates

Table 6: Female's Estimates of Probit and Ordered Probit Models

	Probit		Ordered Probit	
	Ia	**Ib**	**IIa**	**IIb**
Major Health Shock	-0.301†	-0.285	-0.216*	-0.203*
	(0.172)	(0.180)	(0.095)	(0.098)
Male Major Health Shock	-0.454 **	-0.414**	-0.163†	-0.129
	(0.157)	(0.163)	(0.094)	(0.097)
Other Health Problem		-0.205***		-0.118**
		(0.061)		(0.035)
Male Other Health Problem		-0.055		-0.039
		(0.055)		(0.034)
Age	0.010	0.043	0.068	0.104†
	(0.094)	(0.100)	(0.055)	(0.058)
Years of Education	0.014	0.011	0.003	0.002
	(0.009)	(0.009)	(0.004)	(0.004)
Work Experience	0.123***	0.126***	0.067***	0.069***
	(0.009)	(0.010)	(0.004)	(0.004)
Number of Children	-0.017	-0.020	-0.009	-0.009
	(0.013)	(0.014)	(0.006)	(0.007)
Urban	0.040	0.035	0.038	0.030
	(0.077)	(0.080)	(0.040)	(0.042)
Constant	1.681	0.912		
	(2.834)	(3.008)		
Observations	7270	6905	7270	6905
Wald chi^2	521.03*	496.85$^\diamond$	1346.33*	1325.24$^\diamond$
lnsigma2u	1.300	1.314		
rho	0.628	0.633		

*12 degrees of freedom, $^\diamond$14 degrees of freedom.
† $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. LFP: labour Force Participation.
Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command
All specifications include age squared, working experience squared, and year controls.

Table 7: Female's Average Marginal Effects of Probit and Ordered Probit

	Ia	**Ib**	**IIa**	**IIa**	**IIb**	**IIb**
	LFP	**LFP**	**Part-time**	**Full-time**	**Part-time**	**Full-time**
Major Health Shock‡	-0.049†	-0.046	-0.033*	-0.025**	-0.032*	-0.023*
	(0.027)	(0.028)	(0.014)	(0.009)	(0.015)	(0.009)
Male Major Health Shock‡	-0.073**	-0.066**	-0.025†	-0.019†	-0.020	-0.015
	(0.023)	(0.025)	(0.014)	(0.010)	(0.015)	(0.010)
Other Health Problem‡		-0.034***			-0.018**	-0.015**
		(0.010)			(0.005)	(0.004)
Male Other Health Problem‡		-0.009			-0.006	-0.005
		(0.009)			(0.005)	(0.004)
Urban‡	0.006	0.005	0.006	0.005	0.004	0.004
	(0.013)	(0.013)	(0.006)	(0.005)	(0.006)	(0.005)
Age	-0.023***	-0.022***	-0.009***	-0.007***	-0.009***	-0.006***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Years of Education	0.002	0.001	0.001	0.001	0.000	0.000
	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)
Work Experience	0.013***	0.012***	0.007***	0.004***	0.007***	0.004***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Number of Children	-0.002	-0.003	-0.001	-0.001	-0.001	-0.001
	(0.002)	(0.013)	(0.001)	(0.001)	(0.001)	(0.000)
Observations	7270	6905	7270	7270	6905	6905

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
LFP: labour Force Participation. Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command. All specifications include age squared, working experience squared, and year controls. ‡ The average marginal effects for factor variables is the discrete change from 0 to 1. The quantities in the table need to be multiplied by 100 to be transformed into percentage points.

Table 8: Male's Estimates of Probit and Ordered Probit Models

	Probit		Ordered Probit	
	Ia	**Ib**	**IIa**	**IIb**
Major Health Shock	-0.749***	-0.571***	-0.497***	-0.386***
	(0.133)	(0.136)	(0.076)	(0.078)
Female Negative Major Health Shock	0.003	0.056	-0.041	-0.007
	(0.136)	(0.142)	(0.070)	(0.072)
Other Health Problem		-0.389***		-0.203***
		(0.053)		(0.027)
Female Other Health Problem		-0.079		-0.040
		(0.050)		(0.027)
Age	-0.025	-0.012	-0.059	-0.060
	(0.085)	(0.089)	(0.040)	(0.042)
Years of Education	0.004	0.004	-0.005*	-0.006*
	(0.006)	(0.007)	(0.002)	(0.002)
Work Experience	-0.095***	-0.099***	-0.035***	-0.036***
	(0.009)	(0.010)	(0.004)	(0.004)
Number of Children	0.004	0.004	-0.001	-0.001
	(0.011)	(0.011)	(0.005)	(0.005)
Urban	-0.302***	-0.314***	-0.099**	-0.094**
	(0.071)	(0.074)	(0.030)	(0.031)
Constant	7.295	7.117		
	(2.729)	(2.864)		
Observations	8632	8158	8632	8158
Wald chi^2	597.57*	588.76$^\diamond$	1575.06*	1549.54$^\diamond$
lnsigma2u	0.576	1.318		
rho	0.640	0.634		

*12 degrees of freedom, $^\diamond$14 degrees of freedom. Robust standard errors in parentheses.
† $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. LFP: labour Force Participation.
Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command
All specifications include age squared, working experience squared, and year controls.

Table 9: Male's Average Marginal Effects of Probit and Ordered Probit

	Ia	**Ib**	**IIa**	**IIa**	**IIb**	**IIb**
	LFP	**LFP**	**Part-time**	**Full-time**	**Part-time**	**Full-time**
Major Health Shock‡	-0.149***	-0.233***	-0.171***	-0.746***	-0.114**	-0.571***
	(0.027)	(0.063)	(0.042)	(0.128)	(0.036)	(0.127)
Female Negative Major Health Shock‡	0.000	0.019	-0.006	-0.055	-0.001	-0.010
	(0.026)	(0.049)	(0.012)	(0.095)	(0.010)	(0.097)
Other Health Problem‡		-0.138***			-0.028***	-0.273***
		(0.019)			(0.004)	(0.037)
Female Other Health Problem‡		-0.028			-0.005	-0.053
		(0.017)			(0.003)	(0.036)
Urban‡	-0.056***	-0.107***	-0.011***	-0.129**	-0.010**	-0.124*
	(0.013)	(0.024)	(0.003)	(0.039)	(0.003)	(0.040)
Age	-0.028***	-0.057***	-0.009***	-0.093***	-0.009***	-0.091***
	(0.000)	(0.002)	(0.002)	(0.003)	(0.002)	(0.004)
Years of Education	0.000	0.001	-0.000*	-0.007*	-0.000*	-0.008*
	(0.001)	(0.002)	(0.000)	(0.003)	(0.000)	(0.003)
Work Experience	0.003***	0.005***	0.001***	0.007***	0.001***	0.007***
	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)	(0.000)
Number of Children	0.000	0.001	-0.001	-0.001	-0.000	-0.002
	(0.002)	(0.004)	(0.006)	(0.006)	(0.000)	(0.007)
Observations	8632	8158	8632	8632	8158	8158

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
LFP: labour Force Participation. Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command. All specifications include age squared, working experience squared, and year controls. ‡ The average marginal effects for factor variables is the discrete change from 0 to 1. The quantities in the table need to be multiplied by 100 to be transformed into percentage points.

Table 10: Female's Semi-elasticities and Elasticities of Probit and Ordered Probit

Sample 50-70 years old	**Ia**	**Ib**	**IIa**	**IIa**	**IIb**	**IIb**
	LFP	**LFP**	**Part-time**	**Full-time**	**Part-time**	**Full-time**
Semi-elasticities						
Major Health Shock	-0.185	-0.169	-0.142	-0.277	-0.134	-0.262
	(0.144)	(0.151)	(0.110)	(0.209)	(0.116)	(0.220)
Male Major Health Shock	-0.396**	-0.362**	-0.268*	-0.512*	-0.243*	-0.466*
	(0.140)	(0.144)	(0.113)	(0.206)	(0.117)	(0.216)
Other Health Problem		-0.131**			-0.077*	-0.156*
		(0.046)			(0.035)	(0.071)
Male Other Health Problem		-0.030			-0.042	-0.085
		(0.042)			(0.032)	(0.065)
Urban	0.008	-0.003	0.013	0.026	-0.006	-0.013
	(0.059)	(0.061)	(0.045)	(0.091)	(0.048)	(0.096)
Elasticities						
Age	-7.525***	-8.036***	-4.989***	-9.356***	-5.481***	-10.089***
	(0.507)	(0.560)	(0.404)	(0.643)	(0.451)	(0.712)
Years of Education	0.048	0.043	0.008	0.019	0.007	0.016
	(0.031)	(0.032)	(0.020)	(0.045)	(0.021)	(0.048)
Work Experience	0.248***	0.241***	0.151***	0.339***	0.148***	0.331***
	(0.024)	(0.025)	(0.015)	(0.040)	(0.017)	(0.042)
Number of Children	-0.059	-0.077	-0.045	-0.087	-0.057	-0.109
	(0.060)	(0.063)	(0.046)	(0.089)	(0.050)	(0.095)
Observations	6706	6370	6706	6706	6370	6370

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
LFP: labour Force Participation. Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command. All specifications include age squared, working experience squared, and year controls. The semi-elasticity is the percentage change in the probability of participating in the labour force relative to an absolute change in the factor variable (The discrete change from 0 to 1).
The quantities in the table need to be multiplied by 100 to be transformed into percentages.

Table 11: Female's Semi-elasticities and Elasticities of Probit and Ordered Probit

Sample 50-80 years old	**Ia**	**Ib**	**IIa**	**IIa**	**IIb**	**IIb**
	LFP	**LFP**	**Part-time**	**Full-time**	**Part-time**	**Full-time**
Semi-elasticities						
Major Health Shock	-0.263†	-0.251	-0.211†	-0.393†	-0.209†	-0.389†
	(0.149)	(0.156)	(0.117)	(0.211)	(0.123)	(0.222)
Male Major Health Shock	-0.375**	-0.342*	-0.243*	-0.451*	-0.215†	-0.401†
	(0.138)	(0.141)	(0.112)	(0.201)	(0.116)	(0.209)
Other Health Problem		-0.158**			-0.099**	-0.193**
		(0.047)			(0.035)	(0.070)
Male Other Health Problem		-0.048			-0.052	-0.102
		(0.043)			(0.033)	(0.065)
Urban	0.025	0.021	0.025	0.050	0.011	0.022
	(0.061)	(0.062)	(0.047)	(0.091)	(0.050)	(0.096)
Elasticities						
Age	-6.695***	-6.833***	-4.578***	-8.527***	-4.758***	-8.789***
	(0.466)	(0.503)	(0.383)	(0.588)	(0.422)	(0.643)
Years of Education	0.041†	0.031	0.004	0.009	0.001	0.001
	(0.030)	(0.031)	(0.021)	(0.045)	(0.021)	(0.046)
Work Experience	0.293***	0.291***	0.177***	0.408***	0.178***	0.405***
	(0.025)	(0.026)	(0.015)	(0.041)	(0.016)	(0.044)
Number of Children	-0.090	-0.108†	-0.064	-0.119	-0.077	-0.142
	(0.061)	(0.063)	(0.048)	(0.088)	(0.051)	(0.094)
Observations	7449	7084	7449	7449	7084	7084

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
LFP: labour Force Participation. Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command. All specifications include age squared, working experience squared, and year controls. The semi-elasticity is the percentage change in the probability of participating in the labour force relative to an absolute change in the factor variable (The discrete change from 0 to 1).
The quantities in the table need to be multiplied by 100 to be transformed into percentages.

Table 12: Male's Semi-elasticities and Elasticities of Probit and Ordered Probit

Sample 50-70 years old	**Ia**	**Ib**	**IIa**	**IIa**	**IIb**	**IIb**
	LFP	**LFP**	**Part-time**	**Full-time**	**Part-time**	**Full-time**
Semi-elasticities						
Major Health Shock	-0.258***	-0.198**	-0.093**	-0.597***	-0.064*	-0.479***
	(0.068)	(0.065)	(0.035)	(0.137)	(0.030)	(0.135)
Female Major Health Shock	0.008	0.032	-0.000	-0.008	0.001	0.025
	(0.048)	(0.048)	(0.003)	(0.097)	(0.002)	(0.099)
Other Health Problem		-0.104***			-0.007**	-0.193***
		(0.018)			(0.002)	(0.037)
Female Other Health Problem		-0.025			-0.001	-0.051
		(0.017)			(0.001)	(0.035)
Urban‡	-0.113***	-0.110***	-0.001	-0.136**	-0.002	-0.130**
	(0.021)	(0.022)	(0.001)	(0.044)	(0.001)	(0.046)
Elasticities						
Age	-3.542***	-3.467***	-0.423**	-5.817***	-0.393**	-5.692***
	(0.198)	(0.204)	(0.144)	(0.320)	(0.147)	(0.332)
Years of Education	0.008	0.005	-0.001	-0.036	-0.001	-0.042
	(0.013)	(0.013)	(0.001)	(0.026)	(0.001)	(0.027)
Work Experience	0.280***	0.276***	-0.001	0.376***	-0.003	0.379***
	(0.021)	(0.021)	(0.007)	(0.051)	(0.007)	(0.052)
Number of Children	0.010	-0.006	-0.000	-0.002	-0.001	-0.013
	(0.020)	(0.021)	(0.001)	(0.041)	(0.001)	(0.042)
Observations	7153	6814	7153	7153	6814	6814

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
LFP: labour Force Participation. Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command. All specifications include age squared, working experience squared, and year controls. The semi-elasticity is the percentage change in the probability of participating in the labour force relative to an absolute change in the factor variable (The discrete change from 0 to 1).
The quantities in the table need to be multiplied by 100 to be transformed into percentages.

Table 13: Male's Semi-elasticities and Elasticities of Probit and Ordered Probit

Sample 50-80 years old	**Ia**	**Ib**	**IIa**	**IIa**	**IIb**	**IIb**
	LFP	**LFP**	**Part-time**	**Full-time**	**Part-time**	**Full-time**
Semi-elasticities						
Major Health Shock	-0.331***	-0.243***	-0.165***	-0.668***	-0.118**	-0.528***
	(0.067)	(0.064)	(0.039)	(0.117)	(0.035)	(0.118)
Female Major Health Shock	0.010	0.024	-0.001	-0.011	0.002	0.022
	(0.051)	(0.051)	(0.012)	(0.091)	(0.011)	(0.094)
Other Health Problem		-0.143***			-0.032***	-0.234***
		(0.019)			(0.005)	(0.036)
Female Other Health Problem		-0.033†			-0.008†	-0.060†
		(0.018)			(0.004)	(0.033)
Urban	-0.118***	-0.121***	-0.015***	-0.128***	-0.016***	-0.133***
	(0.024)	(0.024)	(0.004)	(0.042)	(0.006)	(0.044)
Elasticities						
Age	-4.345***	-4.339***	-0.745***	-4.467***	-0.719***	-4.393***
	(0.193)	(0.201)	(0.459)	(2.756)	(0.147)	(2.875)
Years of Education	0.008	0.008	-0.004	-0.035	-0.004	-0.037
	(0.013)	(0.013)	(0.003)	(0.024)	(0.003)	(0.025)
Work Experience	0.374***	0.378***	0.055***	0.560***	0.051***	0.563***
	(0.023)	(0.023)	(0.006)	(0.045)	(0.006)	(0.046)
Number of Children	0.022	0.025	0.002	0.019	0.002	0.019
	(0.021)	(0.022)	(0.004)	(0.034)	(0.005)	(0.036)
Observations	8979	8505	8979	8979	8505	8505

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
LFP: labour Force Participation. Specifications Ib and IIb have fewer observations due to missing observations for other health problems. Specifications Ia and Ib are estimated using the stata xtprobit command, while IIa and IIb are estimated using the stata xtoprobit command. All specifications include age squared, working experience squared, and year controls. The semi-elasticity is the percentage change in the probability of participating in the labour force relative to an absolute change in the factor variable (The discrete change from 0 to 1).
The quantities in the table need to be multiplied by 100 to be transformed into percentages.

Table 14: Female's Semi-elasticities and Elasticities of Probit and Ordered Probit specifications with interactions between Major Health Shocks and Age

	Ia	**IIa**	**IIa**
	LFP	**Part-time**	**Full-time**
Semi-elasticities			
Major Health Shock	-0.419*	-0.369*	-0.576*
	(0.207)	(0.167)	(0.271)
Male Major Health Shock	-0.338*	-0.218†	-0.408*
	(0.140)	(0.116)	(0.206)
Urban	0.030	0.028	0.056
	(0.061)	(0.047)	(0.092)
Elasticities			
Age	-6.756***	-4.479***	-8.322***
	(0.561)	(0.442)	(0.726)
Male Age	-1.032*	-0.872*	-1.509*
	(0.461)	(0.368)	(0.647)
Years of Education	0.046	0.008	0.018
	(0.031)	(0.021)	(0.045)
Work Experience	0.251***	0.154***	0.342***
	(0.025)	(0.016)	(0.040)
Number of Children	-0.052	-0.036	-0.068
	(0.061)	(0.048)	(0.090)
Observations	7088	7088	7088

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. LFP: labour Force Participation. Specifications Ia and Ib use the stata xtprobit command, while IIa and IIb use the stata xtoprobit command. The specifications include age squared, spouse's age squared work experience squared, year controls, as well as two interaction terms; one between Major Health Shock and Age, and another one between Female Major Health Shock and Female Age. The semi-elasticity is the percentage change in the probability of participating in the labour force relative to an absolute change in the factor variable (The discrete change from 0 to 1).

Table 15: Male's Semi-elasticities and Elasticities of Probit and Ordered Probit specifications with interactions between Major Health Shocks and Age

	Ia	**IIa**	**IIa**
	LFP	**Part-time**	**Full-time**
Semi-elasticities			
Major Health Shock	-0.321***	-0.166**	-0.673***
	(0.068)	(0.049)	(0.128)
Female Major Health Shock	0.007	0.012	-0.008
	(0.048)	(0.022)	(0.091)
Urban	-0.102***	-0.011*	-0.107*
	(0.023)	(0.004)	(0.045)
Elasticities			
Age	-3.395***	-0.794***	-5.697***
	(0.216)	(0.155)	(0.372)
Female Age	-0.616***	-0.343**	-0.951**
	(0.163)	(0.105)	(0.294)
Years of Education	0.006	-0.005	-0.045†
	(0.013)	(0.003)	(0.025)
Work Experience	0.348***	0.049***	0.507***
	(0.023)	(0.006)	(0.051)
Number of Children	0.004	-0.001	-0.009
	(0.021)	(0.005)	(0.041)
Observations	8583	8583	8583

Robust standard errors in parentheses. † $p < 0.1$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$. LFP: labour Force Participation. Specifications Ia and Ib use the stata xtprobit command, while IIa and IIb use the stata xtoprobit command. The specifications include age squared, spouse's age squared work experience squared, year controls, as well as two interaction terms; one between Major Health Shock and Age, and another one between Female Major Health Shock and Female Age. The semi-elasticity is the percentage change in the probability of participating in the labour force relative to an absolute change in the factor variable (The discrete change from 0 to 1).

Appendix D Figures

Figure 1: Female's daily hours of work.

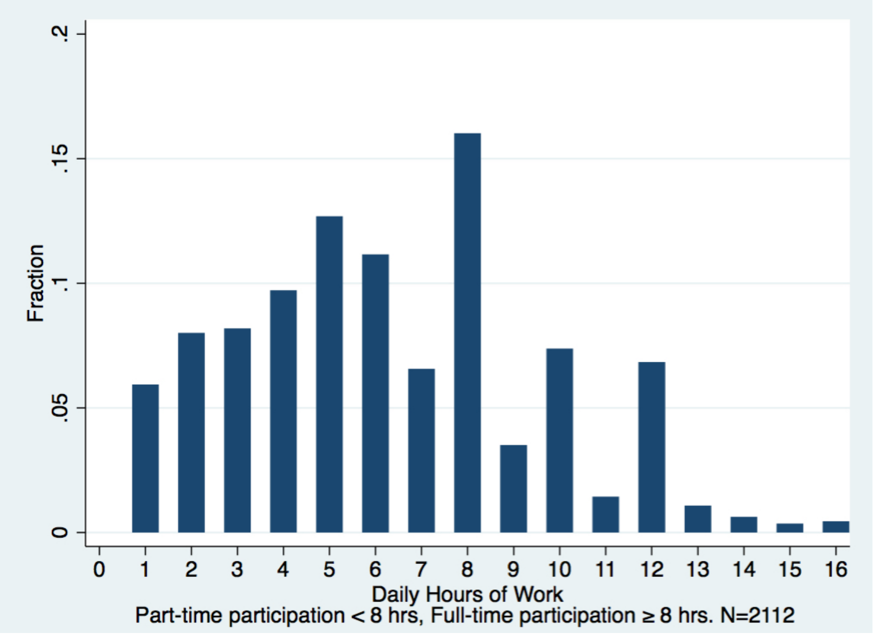
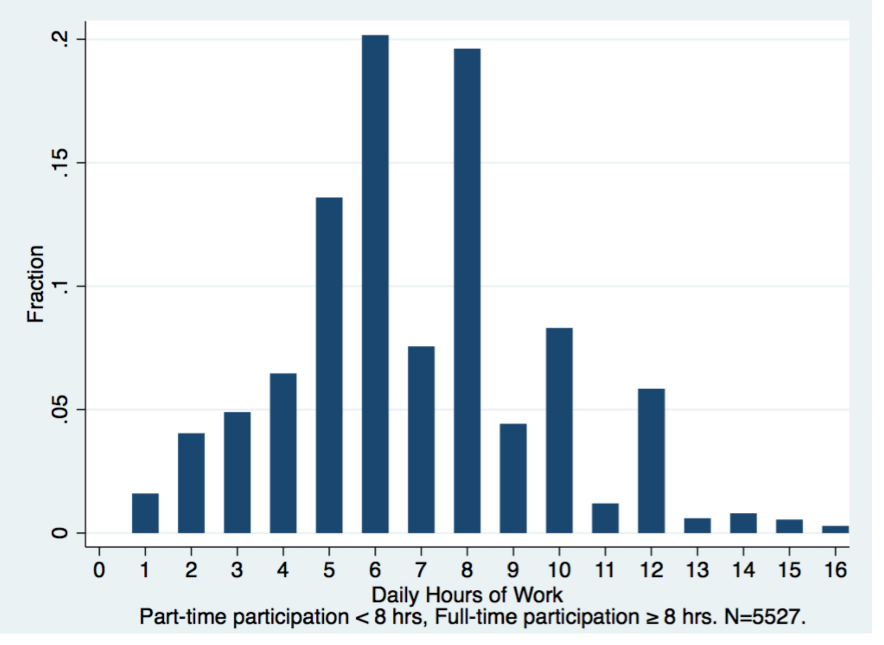


Figure 2: Male's daily hours of work.



Appendix E Mathematical Appendix

Following the budget constraint in section 2.2 and given the females's choices and holding the female's major health shock constant, I define consumption in each labour force state as

$$c_1 = \pi_0 + w_f h_f,$$
$$c_2 = \pi_0 + w_m h_2 + w_f h_f,$$
$$c_3 = \pi_0 + w_m h_3 + w_f h_f,$$

where π_0 is an initial level of consumption. I use the functional form in section 2.2 and develop the male's choice to participate in the labour force. The male does not work if $U(c_1, 0, T - h_f) \geq U(c_2, T - h_2, T - h_f)$

$$\Rightarrow \pi_0 + w_f h_f + \frac{A}{1-\gamma_m} e^{\gamma_m \eta_m} T^{1-\gamma_m} \geq \pi_0 + w_f h_f + w_m h_2 + \frac{A}{1-\gamma_m} e^{\gamma_m \eta_m} (T - h_2)^{1-\gamma_m}$$
$$\Rightarrow A e^{\gamma_m \eta_m} \Big[\frac{T^{1-\gamma_m} - (T - h_2)^{1-\gamma_m}}{1-\gamma_m}\Big] \geq w_m h_2$$
$$\Rightarrow \ln(A) + \gamma_m \eta_m + \ln \Big[\frac{T^{1-\gamma_m} - (T - h_2)^{1-\gamma_m}}{1-\gamma_m}\Big] \geq \ln(w_m) + \ln(h_2) + \gamma_m \ln(T) - \gamma_m \ln(T)$$
$$\Rightarrow \quad \theta_{m1} \geq \Upsilon_{m0} + \alpha_m s_m + \alpha_f s_f + X_m \zeta_m + \epsilon_m - \gamma_m \eta_m$$
$$\therefore \quad \theta_{m1} \geq y_m^*,$$

where
$\theta_{m1} = -\ln(h_2) + \gamma_m \ln(T) - \ln(1 - \gamma_m) + \ln \big[\frac{T^{1-\gamma_m} - (T-h_2)^{1-\gamma_m}}{1-\gamma_m}\big]$,
$\Upsilon_{m0} = \alpha_0 + \gamma_m ln(T)$,
$y_m^* = \Upsilon_{m0} + \alpha_m s_m + \alpha_f s_f + X_m \zeta_m + \epsilon_m - \gamma_m \eta_m$, and $\mu_m = \gamma_m \eta_m - \epsilon_m$; μ_m is normalized such that $\mu_m \sim \mathcal{N}(0, 1)$.
The male works part-time if $U(c_2, T - h_2, T - h_f) \geq U(c_1, 0, T - h_f)$ and $U(c_2, T - h_2, T -$

$h_f) \geq U(c_3, T-h_3, T-h_f)$

$$\Rightarrow \Upsilon_0 + \alpha_m s_m + \alpha_f s_f + X_m\zeta_m + \epsilon_m - \gamma_m\eta_m \geq \theta_1 \quad and \quad \theta_2 \geq \Upsilon_0 + \alpha_m s_m + \alpha_f s_f + X_m\zeta_m + \epsilon_m - \gamma_m\eta_m$$

$$\therefore \quad \theta_1 \leq y_m^* \leq \theta_2,$$

where $\theta_{m2} = -\ln(h_3 - h_2) + \gamma_m \ln(T) - \ln(1-\gamma_m) + \ln\left[\frac{(T-h_2)^{1-\gamma_m} - (T-h_3)^{1-\gamma_m}}{1-\gamma_m}\right]$, and $\mu_m \sim \mathcal{N}(0,1)$.

The male works full-time if $\quad U(c_3, T-h_3, T-h_f) \geq U(c_2, T-h_2, T-h_f)$

$$\Rightarrow \Upsilon_{m0} + \alpha_m s_m + \alpha_f s_f + X_m\zeta_m + \epsilon_m - \gamma_m\eta_m \geq \theta_{m2}$$

$$\therefore \theta_{m2} \leq y_m^*.$$

The male's probabilities of not participating, participating part-time and participating full-time, respectively, are

$$Pr(Y_m = 1|\nu_m) = \Phi(\theta_{m1} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m), \tag{MA1}$$

$$Pr(Y_m = 2|\nu_m) = \Phi(\theta_{m2} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m) - \Phi(\theta_{m1} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m), \tag{MA2}$$

$$Pr(Y_m = 3|\nu_m) = 1 - \Phi(\theta_{m2} - \Upsilon_{m0} - \alpha_m s_m - \alpha_f s_f - X_m\zeta_m), \tag{MA3}$$

where $\nu_m \equiv (s_m, s_f, X_m)$, and Φ is the standard normal distribution function. Likewise, for the female

$$Pr(Y_f = 1|\nu_f) = \Phi(\theta_{f1} - \Upsilon_{f0} - \beta_m s_m - \beta_f s_f - X_f\zeta_f), \tag{MA4}$$

$$Pr(Y_f = 2|\nu_f) = \Phi(\theta_{f2} - \Upsilon_{f0} - \beta_m s_m - \beta_f s_f - X_f\zeta_f) - \Phi(\theta_{f1} - \Upsilon_{f0} - \beta_m s_m - \beta_f s_f - X_f\zeta_f), \tag{MA5}$$

$$Pr(Y_f = 3|\nu_f) = 1 - \Phi(\theta_{f2} - \Upsilon_{f0} - \beta_m s_m - \beta_f s_f - X_f\zeta_f), \tag{MA6}$$

where $\nu_f \equiv (s_m, s_f, X_f)$, and Φ is the standard normal distribution function.

2

Health and the share of wealth held in risky assets

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Abstract

This paper focuses on the correlation between negative health shocks and the households' share of wealth held in risky assets. By using U.S. data from the Health and Retirement Study, we try to establish a link between negative health shocks and financial outcomes such as a household's probability of owning risky assets and share of risky assets held. In our definition of a recent negative health shock, we include: cancer or malignant tumor diagnoses, stroke or transient ischemic attack, and heart attack, coronary heart disease, angina, congestive heart failure, or other heart problems. We find an important negative correlation between a negative health shock to the female and her household's probability of owning risky assets and share of risky assets held, respectively. In contrast, we do not find a corresponding statistically significant correlation for males.

JEL code: G50, G11, I10

Keywords: aging couples, health shock, household portfolio choice, share of risky assets, risky financial assets, Health and Retirement Study

*We would like to thank Thomas Aronsson, Gauthier Lanot, David Granlund, Katharina Jenderny and Magnus Wikström for valuable comments.

1 Introduction

The purpose of this study is to analyze the correlation between negative health shocks and the households' ownership of risky financial assets and share of risky financial assets in their portfolios. For our analysis, we use a sample of aging couples in the U.S. derived from 11 waves (1994-2014) of the Health and Retirement Study (HRS). The HRS provides a very rich set of data that includes information on recent negative health shocks as diagnosed by a medical doctor, the households' financial wealth, as well as various demographic and economic variables.

Exploring the factors that affect the financial decisions of aging couples is particularly interesting since this segment of the population holds a substantial amount of the total U.S. household net worth (Coile and Milligan 2009). For example, in 2004, households where the spouses were 65 years old or older had a median net worth of 190,100 USD, which lies well above the 93,100 USD median net worth of all households in the U.S. (Bucks et al. 2006). Moreover, we would expect aging couples to be more exposed to negative health shocks than younger counterparts, which motivates us to investigate whether there is an effect of such shocks on their financial decisions. Finally, it is also interesting to investigate this in light of an aging population; we can expect that in the future, there will be a larger number of older people holding a considerable amount of wealth in the U.S.

According to Rosen and Wu (2004) and Edwards (2008), we should expect a lower share of wealth invested in risky financial assets among households that have experienced a recent negative health shock. One reason for this change might be that these households tend to sell risky assets in order to cover medical expenses. However, unlike Rosen and Wu 2004, who use self-perceived health ratings, and Edwards 2008, who use expectations regarding health, we choose to use an objective measure of health, in this case conditions that have been diagnosed by a medical doctor. That way, we avoid the subjectivity that is inevitable in any type of self-rated variable. Furthermore, compared with Wu (2003), who examined the same measure of negative health shock as we do, we focus on an older age group and use more waves of the HRS, i.e., from 1994 to 2014.

Moreover, we contribute to the existing literature by presenting a theoretical model that helps explain the household's investment choice in risky assets when taking into account the spouses' health. Earlier studies (Wu 2003, Rosen and Wu 2004, and Berkowirtz and Qiu 2006) focus their work only on empirical modelling and estimation and do not model the households' choice of consumption and investment in risky assets. We go further and present a theoretical model, which serves as a background to the empirical analysis. That way, our study provides further understanding of how negative health shocks, as a source of background risk, might affect households' ownership of and share of wealth held in risky assets.

We also present a simple theoretical model that helps us analyze the effect of a recent negative health shock to one of the spouses on their household's investment in risky assets. This model is important since it provides a description of the mechanisms behind this relationship. The mechanisms show that there are no theoretical reasons to expect households to change the amount invested in risky assets in a particular fashion in response to negative health shocks. Therefore, we perform a numerical analysis based on the model's comparative statics and find that the direction of the change in the amount invested in risky assets will depend on the functional form of the household's utility function and that risk aversion does not necessarily induce the households in question to reduce their investment in risky assets.

The insights of the theoretical model suggest that it remains an empirical question to determine the direction of the change in the households' share of wealth held in risky assets. We address this issue and find that a negative health shock to the female is associated with a 1.2 percentage-point reduction in a household's probability of owning risky assets, and with a 2 percentage-point reduction in its share of wealth held in such assets.[1] We do not find any statistically significant correlation between negative health shocks to the male and outcome variables.[2]

Finally, we reconcile the theoretical and empirical models and suggest that, on average, the households insure themselves against the cost of the female's negative health shock by reducing the share of risky assets. This result is important because it suggests that households are to some extent poorly insured against the cost of a negative health shock to the female. The results also suggest that households are better insured against the male facing a negative health shock, which implies that the households' consumption is less affected by such a shock. The share of wealth held in risky assets is less affected as well.

The rest of this paper is structured as follows. Previous literature is reviewed in Section 2. Section 3 presents the theoretical model. Section 4 discusses the data, and Section 5 specifies the empirical models. The results are presented in Section 6. Finally, conclusions are drawn in Section 7.

2 Literature review

Our study builds upon and aims to further develop existing literature that examines the effect of health events on household wealth and portfolio decisions. It has already been documented that negative health

[1]Even though we are mostly interested in the risky share and whether it increases or decreases after a negative health shock, we also believe it is important to investigate the association between the probability of owning risky assets, i.e., stock market participation, and negative health shocks.

[2]We do not claim that we have established a casual effect between health shocks and the specific economic decisions we study even though we have controlled for covariates that previous literature has suggested play a role in these decisions. We believe there could still be other factors unaccounted for in our analysis that may affect a households' decision to own risky assets and share of risky assets held.

shocks affect the economic status of married couples approaching retirement. In our case, we focus on older households where both spouses are 65 years old or older, while the literature mentioned in this section that uses the US HRS focuses on households where the survey's main respondent was 51 years old or older. Specifically, Wu (2003) looks into the effect of severe exogenous health shocks on household wealth, income, and consumption by using data on couples from the first two waves of the HRS. He estimates quantile regressions where the dependent variable is change in household wealth between periods, and finds that new severe diagnoses of diseases or health conditions reduce household wealth, yet the effect is asymmetric between females and males. More precisely, there is a statistically significant drop in household wealth associated with a negative health shock facing the female, an effect that persists even after controlling for initial health condition and changes in retirement decisions. For males, on the other hand, the impact of negative health shocks on household wealth disappears.

Apart from the effect of health on wealth, the existing literature has also documented how negative health shocks influence households' ownership of various types of assets. Rosen and Wu (2004) use four waves of the HRS and focus on the impact of changes in the individual's perception of health on financial assets. They classify assets into four categories consisting of safe assets (checking and savings accounts, money market funds, CDs, government savings bonds, and T-bills), bonds (corporate, municipal and foreign bonds, and bond funds), risky assets (stocks and mutual funds), and retirement accounts (IRAs and Keoghs[3]). They estimate a probit model for each asset category in their classification, and their results show that poor health is associated with a lower probability of holding financial assets, a larger share of financial wealth held in safe assets, and a smaller share allocated to other asset classes.[4]

Rosen and Wu (2004) argue that the effect of poor health on portfolio decisions can operate through three different channels: risk aversion, planning horizons, and/or health insurance. To investigate whether the effect of self-reported health status on asset allocation operates through increasing the respondents' risk aversion, through decreasing their planning horizon, or through their access to health insurance, they create binary variables for each channel and include those as independent variables in their original model. Specifically, they include a binary variable that takes the value one if the individual answers affirmatively to a question designed to provide information about the respondents' being risk averse, and zero otherwise.[5] They create dichotomous variables for planning horizon by using the answers provided to

[3]Keogh plans: Keogh accounts allow self-employed individuals to invest in a tax-deferred retirement plan, where the maximum permitted contribution is considerably higher than for IRAs (Hira et al. 2009).

[4]The explanatory variables they use include a dichotomous variable for poor health and controls for total wealth, income and other individual characteristics.

[5]The question being whether the respondent would take a job that would double their income with a 50 percent probability and reduce it by half with a 50 percent probability (Rosen and Wu 2004).

questions designed to measure the respondents' financial planning periods.[6] They also use a dichotomous variable taking a value of one when the household has health insurance and zero otherwise. The authors show that including risk aversion, planning horizon, and health insurance variables in the model does not affect the relationship between health shocks and financial outcomes, and argue that this indicates that the impact of poor health on portfolio decisions does not operate through any of these channels. Finally, they suggest that there is a robust relationship between health status and portfolio decisions, and that the channels through which poor health operates are not entirely clear.

Berkowitz and Qiu (2006), regarding negative health shocks, distinguish between the effect on financial and non-financial wealth and show that new severe diagnoses, such as a heart problem, stroke, cancer, lung disease, and diabetes, have a greater impact on financial wealth than on non-financial wealth for single households. For married couples, the results indicate that a negative health shock facing the female affects negatively household financial and non-financial assets while a negative health shock facing the male does not have a statistically significant effect. They mention that the latter could be attributed to a liquidity effect and eligibility requirements for public insurance.

By using data from the first six waves of the HRS, Coile and Milligan (2009) document that ownership of different types of assets (principal residence, vehicle, real estate, and business) declines while ownership of liquid assets, time deposits, IRAs[7], stocks, and bonds increases with negative health shocks such as a stroke, a heart attack, or a diagnosis of a new chronic disease. Finally, following these results, they suggest that households have larger reductions in the share of risky than of liquid assets.

Edwards (2008) focuses solely on elderly individuals and examines how self-perceived risky health affects portfolio allocation.[8] He uses individual-level data from the Study of Assets and Health Dynamics Among the Oldest Old in the US (AHEAD), individuals aged 70 or older, which reveals that health risk prompts safer investment. He finds that elderly singles respond the most to health risk and that poor health may explain 20 percent of the age-related decline in financial risk taking after retirement. He suggests that retired individuals tend to view their health as risky, and they appear to decrease their exposure to financial risk by hedging against it. Because health tends to deteriorate with age, the presence

[6] For example, they create the variable plan1, which takes the value 1 if "the next few months" are the most important when the household plans its saving and spending, and zero otherwise. They create five possible planning types in total. Plan 5 takes the value 1 if "more than 10 years" is the most important time horizon for the household.

[7] IRAs: IRAs stands for Individual Retirement Accounts. IRAs were established in 1974 as part of the Employee Retirement Income Security Act to encourage retirement savings by employees without private pension plans. The availability of IRAs was extended to all employees and the contribution limit was raised by the Economic Recovery Tax Act of 1981 (Wise 1987).

[8] In the AHEAD surveys, respondents were asked to report the probability that medical expenses will use up all of the household's savings in the next five years. Edwards (2008) uses these answers for the self-perceived risky health variable.

of undiversifiable health risk may explain why investors on average decrease their financial risk with age after retirement (Edwards 2008).

Our study is positioned within the literature that explores the effect of health on portfolio decisions. However, as mentioned previously, it differs from previous studies in the field by the large number of HRS waves that we have included in our analysis. Particularly, to our knowledge, we are the first ones to use data from 11 waves of the RAND HRS Longitudinal File 2014, i.e., from 1994 to 2014, to explore the potential link between actual health decline (i.e., recent negative health shocks as diagnosed by a medical doctor) and the probability of owning, and the share of a household's wealth allocated to, risky assets. We use data for households where both spouses are at least 65 years old, while previous literature that uses the HRS considers households where the main respondent is 51 years old or older. Furthermore, compared with previous studies that only focus on empirical results (Wu 2003, Rosen and Wu 2004, and Berkowirtz and Qiu 2006), we also present a theoretical model in order to provide a foundation for our empirical analysis.

3 Theoretical framework

We extend the framework presented by Eeckhoudt et al. (2005) by considering a household that invests an amount of its wealth in a risky asset. Moreover, the household comprises two spouses, each of whom faces a probability of enjoying good health.[9] In this way we can capture both health risk and uncertainty in the financial market within a simple framework. To begin with, we consider a model where the household derives utility solely from its consumption, c. The utility function $u(c)$ is increasing in its argument and strictly concave. The male's probability of enjoying good health is p^m and his probability of experiencing poor health is $1 - p^m$. Likewise, the female experiences good health with probability p^f and poor health with probability $1 - p^f$. This means that the household faces four possible states: both spouses are in good health, gg; the female is in good health and the male in poor health, gb; the female is in poor health and the male in good health, bg; and both spouses are in poor health, bb. Second, the household invests an amount α of its total wealth A in a risky asset with an associated uncertain risky rate r. The remainder of A is invested in a risk-free asset at a risk-free rate ρ. There is also a monetary cost for the household when either the female or the male face poor health, k_f and k_m, respectively. The household must make its consumption and portfolio choice before the health status of the spouses and the rate of return of the risky asset materialize. Hence, the household maximizes the expected utility function, $Eu(\cdot)$, with respect

[9] In Eeckhoudt et al. (2005, pp. 142-145) the agent can invest in preventive measures that reduce the probability of a certain outcome occurring. In contrast, we assume that the probability of good health is not affected by the household decisions. We do this to maintain a certain degree of simplicity in our model.

to c, and α,

$$\max_{c,\alpha} \quad p^f p^m Eu(c^{gg}) + p^f(1-p^m)Eu(c^{gb}) + \\ (1-p^f)p^m Eu(c^{bg}) + (1-p^f)(1-p^m)Eu(c^{bb}) \tag{1}$$

$$s.t. \quad c^{ij} = A(1+\rho) + \alpha(r-\rho) - k_f^i - k_m^j, \quad i \in \{b,g\}, \quad j \in \{b,g\} \tag{2}$$

$$A > 0,$$

$$k_f^g = k_m^g = 0.$$

Note that the consumption in each state is a random variable, since the risky rate of return is random. We substitute equation (2) into equation (1), solve an unconstrained maximization problem in α, and obtain the following first-order condition:

$$p^f p^m E[u'^{gg}\delta] + p^f(1-p^m)E[u'^{gb}\delta] + (1-p^f)p^m E[u'^{bg}\delta] + (1-p^f)(1-p^m)E[u'^{bb}\delta] = 0, \tag{3}$$

where u' indicates the marginal utility of consumption in each state. The superscripts indicate that both spouses face good health, u'^{gg}; the female faces good health and the male faces poor health, u'^{gb}; the female faces poor health and the male faces good health, u'^{bg}; and both are face poor health, u'^{bb}. $\delta = r - \rho$ is the excess return of the risky asset. Equation (3) is the condition for optimal investment and implies that the sum of the expected products of the marginal utility of consumption in each state and the difference in return equals zero. We use the first-order condition to compute the change in α, the amount invested in the risky asset, when p^f or p^m changes, respectively.[10] The change in the amount invested in the risky asset when the female's probability of good health changes is given by

$$\frac{d\alpha}{dp^f} = -\frac{1}{D}\big\{p^m\big(E[u'^{gg}]\hat{\delta} - E[u'^{bg}]\hat{\delta} + cov[u'^{gg},\delta] - cov[u'^{bg},\delta]\big) \\ + (1-p^m)\big(E[u'^{gb}]\hat{\delta} - E[u'^{bb}]\hat{\delta} + cov[u'^{gb},\delta] - cov[u'^{bb},\delta]\big)\big\}. \tag{4}$$

Similarly, the expression for the change in the amount invested in the risky asset when the male's probability of good health changes can be written as

$$\frac{d\alpha}{dp^m} = -\frac{1}{D}\big\{p^f\big(E[u'^{gg}]\hat{\delta} - E[u'^{gb}]\hat{\delta} + cov[u'^{gg},\delta] - cov[u'^{gb},\delta]\big) \\ + (1-p^f)\big(E[u'^{bg}]\hat{\delta} - E[u'^{bb}]\hat{\delta} + cov[u'^{bg},\delta] - cov[u'^{bb},\delta]\big)\big\}. \tag{5}$$

In equations (4) and (5), $\hat{\delta} = E[\delta] > 0$, and $D < 0$ by the second-order conditions.[11] The sign of these equations is undetermined. For example, on the right-hand side of equation (4), there are two differences:

[10] See the Appendix for the mathematical derivation.

[11] $D = p^f p^m E[u''^{gg}\delta^2] + p^f(1-p^m)E[u''^{gb}\delta^2] + (1-p^f)p^m E[u''^{bg}\delta^2] + (1-p^f)(1-p^m)E[u''^{bb}\delta^2]$.

the difference between the products of expectations and $\hat{\delta}$, and the difference between covariances. What we know from this model is that the difference between the products of expectations and $\hat{\delta}$ is negative because the marginal utility of consumption in a state with poorer health is higher than the marginal utility of consumption in a state with better health.

We also know that the covariance between the marginal utility of consumption and δ is negative; however, the difference between covariances can be either positive or negative. If it is negative, then the right-hand side of equation (4) will be negative and the household will decrease its investment in risky assets when the probability of good health increases. If it is positive and larger in absolute value than the difference between the products of expectations and δ, then the right-hand side of equation (4) will be positive. Hence, the household increases the amount invested in risky assets when the probability of good health increases. As suggested by the numerical analysis carried out later, the sign of the difference in covariances will be determined by the functional form of the household's utility function.

To go further we assume that the household faces a monetary cost related to the female's poor health and that the household is fully insured against the male facing poor health ($k_f^b > 0$, and $k_m^b = 0$).[12] It follows that the household's consumption when the male is ill equals the household's consumption when both spouses are healthy. Thus, $u'^{gg} = u'^{gb} < u'^{bg} = u'^{bb}$. If we substitute these in equations (4) and (5), respectively, we obtain the following expressions:

$$\frac{d\alpha}{dp^f} = -\frac{1}{D}\left[(E(u'^{gg})\hat{\delta} - E(u'^{bb})\hat{\delta}) + cov(u'^{gg}, \delta) - cov(u'^{bb}, \delta)\right] \tag{6}$$

$$\frac{d\alpha}{dp^m} = 0, \tag{7}$$

where $E(u'^{gg}) - E(u'^{bb}) < 0$. Even in this case, equation (6) can be either positive or negative depending on the sign and magnitude of the difference between the covariances. If this difference is positive and larger in absolute value than the difference between expectations, then equation (6) is positive, otherwise it is negative. Furthermore, we have also shown in equation (7) that if the household is fully insured against the male facing poor health the amount invested in the risky asset does not change when his probability of good health changes.

Now we assume that the cost of the female's poor health and the cost of the male's poor health are the same ($\bar{k} = k_f^b = k_m^b > 0$) to show that, in this case, the changes in the amount invested in the risky asset when the probabilities of good health change will not be symmetrical. It follows that $u'^{gg} < u'^{gb} = u'^{bg} < u'^{bb}$. We substitute the latter in (4) and (5) and obtain the following expressions for

[12] The analysis is symmetrical if we consider that the household is fully insured against the female facing poor health.

the female and male, respectively:

$$\frac{d\alpha}{dp^f} = -\frac{1}{D}\{p^m\big(E[u'^{gg}]\hat{\delta} - E[u'^{bg}]\hat{\delta} + cov(u'^{gg},\delta) - cov(u'^{bg},\delta)]$$

$$+ (1-p^m)\big(E[u'^{bg}]\hat{\delta} - E[u'^{bb}]\hat{\delta} + cov(u'^{bg},\delta) - cov(u'^{bb},\delta)\big)\} \tag{8}$$

$$\frac{d\alpha}{dp^m} = -\frac{1}{D}\{p^f\big(E[u'^{gg}]\hat{\delta} - E[u'^{bg}]\hat{\delta} + cov(u'^{gg},\delta) - cov(u'^{bg},\delta)]$$

$$+ (1-p^f)\big(E[u'^{bg}]\hat{\delta} - E[u'^{bb}]\hat{\delta} + cov(u'^{bg},\delta) - cov(u'^{bb},\delta)\big)\} \tag{9}$$

Equations (8) and (9) will differ unless $p^f = p^m$. The probabilities of good health may differ between the female and the male, since health depends on many factors, e.g., health history, and genetic background. Henceforth, the model suggests that changes in the probability of good health facing the female and male, respectively, may not have a symmetrical effect on the household's decision to invest in risky assets even if the household faces the same monetary costs of the spouses' health.

3.1 Numerical analysis

Since we cannot sign equations (4) and (5) without further assumptions, we choose two different utility functions and do a numerical analysis to illustrate that the right-hand side of equation (4) can be positive or negative. For each utility function we use the same values for the parameters of interest: $p^f = 0.7, p^m = 0.5, A = 5, k_m = 1$ and $k_f = 1$.[13] We also use a uniform distribution between $[-0.05, 0.05]$ for δ. With this information, we calculate the optimal value of α, and the sign of the right-hand side of equation (4).

First, we assume that the utility function is logarithmic $u(c) = ln(c)$. In this case, we find that the two differences between covariances are positive, i.e., $cov(u'^{gg},\delta) - cov(u'^{bg},\delta) > 0$ and $cov(u'^{gb},\delta) - cov(u'^{bb},\delta) > 0$. We also find that the difference between covariances dominates the difference between the products of expectations and $\hat{\delta}$, and thus the right-hand side of equation (4) is positive. It follows that the amount invested in risky assets increases with an increase in the probability of good health.

Second, we assume a cubic utility function of the following form $u(c) = c - \frac{1}{1000}\frac{c^2}{2} - \frac{1}{10000}\frac{c^3}{3}, \quad c \in (0, 95)$.[14] In this case we find the opposite result. The covariance between the household's marginal utility of consumption and the excess return is smaller in absolute value when one of the spouses faces poor health than when both spouses face good health. It follows that all differences between covariances are negative, and the right-hand side of equation (4) is negative. The amount invested in risky assets decreases with an

[13] We chose a higher probability for the female in this case since the data shows that females face fewer health shocks than males. We also performed the analyses with different values for the parameters, and obtained the same qualitative results for the sign of the right hand side of equation (4). For example, in one of the analyses we chose $p^f = 0.4, p^m = 0.6, A = 2, k_f = 1$ and $k_m = 0$, while in another we chose $p^f = 0.5, p^m = 0.3, A = 5, k_f = 2$ and $k_m = 1$.

[14] $c \in (0, 95)$ in order for $u'(c) > 0$.

increase in the probability of good health. Furthermore, for low (high) levels of consumption a decrease in the probability of good health will have a larger (smaller) effect on the marginal utility of consumption for prudent households than for imprudent ones. We choose these utility functions to show that prudence may play an important role for the household's decision to invest in risky assets when the probability of being healthy increases or decreases, and that risk aversion does not necessarily induce the household to reduce the investment in risky assets when the probability of being healthy decreases.[15] Prudence is represented by the utility function's third derivative. If the third derivative is positive then the household is prudent, and if it is negative then the household is imprudent. It follows that a household with logarithmic preferences is prudent while a household with the cubic preferences in our example is imprudent.

If we think of a negative health shock as a decrease in the probability of good health, it follows that the prudent household prefers to invest less in the risky asset as an insurance against loss in consumption due to facing poor health. On the other hand, the imprudent household chooses to invest more in the risky asset when having experienced a negative health shock.

3.2 Adding health capital

In this subsection, we add health capital to the model, in such a way that it influences the utility of the household. We assume that the health capital stock varies exogenously over the health states. Specifically, s_f, and s_m represent the female's and male's health capital, respectively. The household solves the following program:

$$\begin{aligned}
\max_{c,\alpha} \quad & p^f p^m Eu(c^{gg}, s_f^g, s_m^g) + p^f(1-p^m)Eu(c^{gb}, s_f^g, s_m^b) + \\
& (1-p^f)p^m Eu(c^{bg}, s_f^b, s_m^g) + (1-p^f)(1-p^m)Eu(c^{bb}, s_f^b, s_m^b) \qquad (10) \\
s.t. \quad & c^{ij} = A(1+\rho) + \alpha(r-\rho) - k_f^i - k_m^j, \quad i \in \{b,g\}, \quad j \in \{b,g\}, \\
& A > 0, \\
& k_f^g = k_m^g = 0, \\
& s_i^g > s_i^b \quad i \in \{f,m\}.
\end{aligned}$$

The utility function u has the same properties with respect to consumption as before, and is, in this case, also increasing in s_f, and s_m. The superscripts g and b indicate good and poor health, respectively. The first-order condition with respect to α takes the same form as before. The change in the amount invested in the risky asset when the female's probability of good health changes takes the same form as equation

[15] Note that in our example, both utility functions are concave. Their first derivative with respect to c is positive and the second is negative.

(4), i.e.,

$$\begin{aligned}\frac{d\alpha}{dp^f} = -\frac{1}{D}\big\{ & p^m\big(\hat{\delta}(E[u'^{gg}] - E[u'^{bg}]) + cov[u'^{gg}, \delta] - cov[u'^{bg}, \delta]\big) \\ & + (1-p^m)\big(\hat{\delta}(E[u'^{gb}] - E[u'^{bb}]) + cov[u'^{gb}, \delta] - cov[u'^{bb}, \delta]\big)\big\}. \end{aligned} \tag{11}$$

The sign of the right-hand side of equation (11) will, as discussed in the previous section, depend on the differences between the products of expected marginal utilities and $\hat{\delta}$ and the differences between covariances, respectively. However, the difference in expected marginal utilities of consumption now depends on two distinct effects. First, for given levels of health capital, the household consumes more when the spouses are healthy than when they are not, hence the marginal utility of consumption will be smaller when the household is healthy (consumption effect). Second, if consumption and health are complements, more health capital contributes to increasing the marginal utility of consumption, ceteris paribus (health effect).

The consumption effect contributes to making the differences $E[u'^{gg}] - E[u'^{bg}]$ and $E[u'^{gb}] - E[u'^{bb}]$ negative, whereas the health effect contributes to making them positive. The sign of the difference between the expectations will depend on which effect dominates. If the consumption effect dominates the difference is negative, and if the health effect dominates the difference is positive. Furthermore, if the effects offset each other the difference may be close to zero. The interaction between the health and consumption effects constitutes another mechanism that may drive a positive relationship between p^f and α.

Now we consider the case where the household is fully insured against the cost of the male facing poor health, $k_m^b = 0$. In contrast to the model without health capital in the utility function, the change in the amount invested in the risky asset when the male's probability of good health changes is different from zero since the marginal utility of consumption also depends on the health capital. It follows that $E[u'^{gg}] \neq E[u'^{bg}]$ and $E[u'^{gb}] \neq E[u'^{bb}]$ and therefore $\frac{d\alpha}{dp^m} \neq 0$. Even in this case where there is no monetary cost of the male's poor health, the amount invested in the risky asset can change positively or negatively.

The theory, even in a simple setting as the one presented in this section, does not provide a clear answer to the question of the direction of the change in the amount invested in the risky asset due to a change in one of the spouse's probability of enjoying good health without making additional assumptions about the household's preferences. Yet, the model highlights some of the important underlying mechanisms that may drive the behavioral change. However, it still remains an empirical investigation to show in which direction the household's investment in risky assets changes when one of the spouses faces a negative health shock. We address the aforementioned question in the empirical specification and link its results with the theoretical framework.

4 Data

The study uses data from the University of Michigan Health and Retirement Study (HRS). The HRS is supported by the National Institute on Aging and the Social Security Administration. It is a biennial nationally representative panel that follows approximately 7,000 households in the U.S. over time. In the first wave (1992), the main respondent is 51 years old or older.

In our study, we choose to use data retrieved from the RAND HRS Longitudinal File 2014 (V2)[16]. The file contains health-related information for the respondent and the respondent's spouse, economic and demographic variables, and spousal counterparts of most individual-level variables. Of particular interest for our study is that the dataset provides information about whether the spouses had experienced a negative health shock in the two years prior to the wave. In our study, as defined by the RAND HRS, any or a combination of the following diagnoses constitutes a negative health shock: cancer or a malignant tumor of any kind except skin cancer; heart attack, coronary heart disease, angina, congestive heart failure, or other heart problems; and stroke or transient ischemic attack (TIA). These health shocks can be interpreted in terms of an increase (decrease) in the likelihood of poor health (good health), ceteris paribus, and in this way we can relate our results to the theoretical model.

The HRS also provides information on the respondent's health insurance plans and the household's ownership of retirement accounts, safe and risky financial assets, labor and non-labor income, total wealth, and various individual characteristics important for our analysis.

The respondent's government health insurance plans incorporate coverage by any of the following federal government health insurance programs: Medicare, Medicaid, VA/CHAMPUS, CHAMP-VA, or any other government-provided health insurance.

The household's financial assets reported in the RAND HRS file which we include in our definition of the household's financial portfolio are: checking, savings, and money market accounts; CDs, government savings bonds, and T-bills; corporate, municipal and foreign bonds and bond funds; stocks, mutual funds, investment trusts, and IRA and Keogh accounts. We define the household's risky assets as the sum of the net value of stocks, mutual funds and investment trusts, and calculate the household's share of wealth held in risky assets by dividing this by the sum of the net value of the household's financial portfolio. The household's non-earned income is defined as the sum of the households' capital income and the spouses' income from employer pension or annuities, social security, unemployment or workers' compensation, and all other government transfers..[17] Our total wealth measure is the sum of the net value of the primary residence, real estate (not primary residence), vehicles, businesses, financial wealth, and all other savings.

[16] http://hrsonline.isr.umich.edu/data /index.html.

[17] All types of taxable income are reported as pre-tax income.

We focus our study on heterosexual couples where both spouses are 65 years old or older and use 11 waves of the RAND HRS Longitudinal File 2014, i.e., from 1994 to 2014. We choose to focus on this age group because, as mentioned before, members of this group own a large amount of wealth, which suggests that they face lower economic barriers to investing in the stock market. Moreover, individuals who are 65 years old or older are more prone to experience health shocks than younger individuals. Occurrence of health shocks is a precondition so that we can investigate the association between these and the probability of owning risky assets and share of household wealth held in such assets, respectively. Last but not least, using this age group instead of a younger one provides a less heterogeneous sample, reducing the number of factors we have to consider in our model.

The first wave of the HRS was conducted in 1992, but we do not use this wave since it does not offer any observations where both spouses are 65 years old or older.[18] While most existing studies use a smaller number of waves (mostly 2-4 waves), we choose to include as much data as possible in order to increase the number of observations used in our analysis.

Furthermore, an advantage of focusing on these waves of the RAND HRS file is that they, in a sense, constitute an unexplored area of study. To our knowledge, these waves have not been used in previous studies that analyze the relationship between negative health shocks and the share of risky assets in financial portfolios because they were not available at that time.

Our final sample has an unbalanced panel structure, where each household is observed on average four times. The sample consists of 44,303 observations of U.S. households where both spouses are aged 65 years or older. We use 11 waves of the HRS from 1994 to 2014. Table 1 shows that 40 percent of the households own risky assets and allocate 17.7 percent of their financial wealth in risky assets. The average incidence of a negative health shock occurring differs largely between males and females in the sample: 37 percent of females had experienced such a shock, while the figure was 56 percent for males. Both numbers may seem large, but remember that the negative health shock variable includes diverse shocks like cancer diagnoses, heart problems and strokes. Heart problems account for the majority of the shocks, i.e., 39 percent in males and 23 percent in females, while 9 percent of males and 6 percent of females had had a stroke.[19] Cancer diagnoses were also higher for males than for females, i.e., 22 percent of males vs. 16 percent of females.[20]

[18]In this wave, the average age is 57 for males and 53 for females.

[19]The American Heart Association reported that from 1999 to 2012, 69.1 percent of males and 67.9 percent of females 60-79 years old experienced a cardiovascular disease (Mozaffarian et al. 2016). Using data from 1999, the American Stroke Association reported that 6.1 percent of males and 5.2 percent of females 60-79 years old had suffered a stroke at some point in life (Mozaffarian et al. 2016).

[20]According to the U.S. National Cancer Institute, approximately 38.4 percent of males and females will be diagnosed with cancer at some point during their lifetime. Source: https://www.cancer.gov/about-cancer/understanding/statistics.

Table 1: Summary Statistics

	Mean	Standard Deviation
Household Risky Assets Ownership	0.401	
Household Share in Risky Assets	0.177	(0.289)
Male's Negative Health Shock	0.561	
Female's Negative Health Shock	0.375	
Male's Age	75.702	(6.129)
Female's Age	73.073	(5.827)
Male's Years of Education	12.612	(3.275)
Female's Years of Education	12.637	(2.583)
Number of Children	3.291	(2.033)
Household Total Wealth*	722,133	(1,521,183)
Household Earned Income*	8,681	(3,200)
Household Non Earned Income*	63,056	(93,589)
Male's Government Health Insurance	0.981	
Male's Medicare	0.978	
Male's Medicaid	0.023	
Male's VA/CHAMPUS	0.075	
Female's Government Health Insurance	0.978	
Female's Medicare	0.976	
Female's Medicaid	0.022	
Female's VA/CHAMPUS	0.045	
Male's Private Health Insurance**	0.653	
Female's Private Health Insurance**	0.673	
Male's Out of Pocket Medical Expenditure*	2,916	(6,959)
Female's Out of Pocket Medical Expenditure*	3,007	(6,144)
N		44,303

Standard deviations in parenthesis. Source: waves 1994-2014 of the RAND HRS Longitudinal File 2014 (V2). *Monetary variables are measured in 2014 U.S. dollars. **The number of observations is 44,143 due to missing data. Note. See the Appendix for a description of the variables.

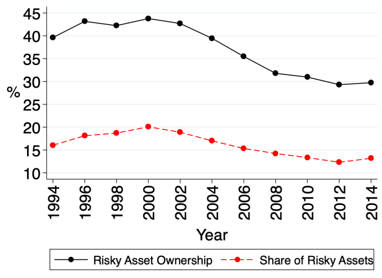
The average age of females is 73 years, while the average for males is about 3 years higher, i.e., 76 years. Years of education are on average about the same for males and females: 12 years. We can also observe that households had on average more than 3 children. Approximately 98 percent of the males and females were covered by Medicare, and approximately 2.2 percent by Medicaid.[21] The latter may imply that there is a low percentage of males and females with limited income and resources in the data. Males were also more likely to have VA/Champus insurance. Table 1 also shows that the average household earned income in our sample was about $8,681, non-earned income $63,056.[22], and total wealth $722,133. Males had slightly lower out-of-pocket medical expenditures despite exhibiting a higher incidence of negative health shocks. This may suggest that they are better insured than females.

[21]Not all individuals were covered by Medicare since some may have preferred to have only private insurance. It can also be the case that some did not qualify for both Medicare and Medicaid.

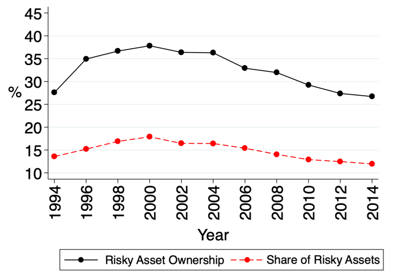
[22]Non-earned income is higher since approximately 75 percent of the individuals in our data reported being retired.

Figure 1: Risky Asset Ownership, Share of Risky Assets

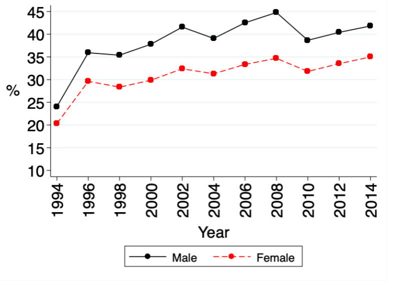
and Negative Health Shocks.



(a) Risky asset ownership and share of risky assets of healthy households



(b) Risky asset ownership and share of risky assets of unhealthy households



(c) Negative health shocks

Source: waves 1994-2014 of the RAND HRS Longitudinal File 2014 (V2). There is a total of 44,303 observations where 12,753 are healthy households and 31,550 are unhealthy households. A healthy household is a household where both spouses were healthy while an unhealthy one is a household where at least one of the spouses faced a negative health shock.

Panel (a) of Figure 1 shows the risky asset ownership and share of wealth held in risky assets among households where both spouses are healthy. Panel (a) shows that the percentage of healthy households that owned risky assets increased from 1994 to 2000 and then started to decline. In 1994, approximately 40 percent of such households owned risky assets. By the year 2000, the proportion had increased to 44 percent, and then it started to decrease. Finally, in 2014, approximately 30 percent of healthy households owned risky assets. Panel (a) also shows the share of wealth held in risky assets among healthy households.

We can see that the risky asset share has a similar trajectory as risky asset ownership. In 1994, healthy households held approximately 15 percent of their portfolios in risky assets. The share peaked in the year 2000 at 20 percent and then started to fall. By the year 2014, the households held, on average, 14 percent of their portfolio in risky assets. Panel (b) shows the risky asset ownership and share of wealth held in risky assets among households where at least one of the spouses faced a negative health shock. Panel (b) shows that in general, unhealthy households had an ownership percentage lower than healthy households. For example, in 2002, 43 percent of healthy households and 36 percent of their unhealthy counterparts owned risky assets. The same is true for share of wealth held in risky assets, i.e., healthy households seem to exhibit a larger share than unhealthy households, although the difference is small.

Panel (c) shows the trajectory of negative health shocks for males and females. The rate of negative health shocks increases more for males than females, i.e., in 2014, approximately 42 percent of males reported having experienced a negative health shock, compared with only 35 percent for females. The increase in negative health shock incidence over the years can be explained by aging spouses in the data, since the HRS follows households across time. This can be seen in Table 2, i.e., older people are more likely to experience negative health shocks. Finally, Table 2 also shows that the incidence differs between males and females even when we group them by age cohorts.

Table 2: Age and Negative Health Shocks

Age	**Male's Negative Health Shock**	**Female's Negative Health Shock**
65-70	43.7%	31.4%
N	9,618	17,287
71-75	52.9%	37.0%
N	13,882	13,193
76-80	61.1%	42.9%
N	11,006	8,404
81 and older	66.9%	49.7%
N	9,797	5,419

Source: waves 1994-2014 of the RAND HRS Longitudinal File 2014 (V2). There are 44,303 observations in total.

Table 3 presents the percentage of households owning risky assets and the share of their wealth held in that asset class conditioned on the households' health state and the male's age. There are four health states: households where both spouses are healthy, households where the female is healthy and the male is unhealthy, vice versa, and households where both spouses are unhealthy. We define the individual in the household as unhealthy if they had faced a negative health shock and as healthy otherwise. We also consider four age groups: 65-70, 71-75, 76-80, and 81 and older. We observe that the first two age groups, respectively, present larger variation across household health states than the last two. For example, in the first age group, the percentage of households owning risky assets is approximately the same, 41 percent, when both spouses are healthy and when the female is healthy and the male unhealthy, yet it is lower for households where the female is unhealthy, 35.2 percent, and both spouses unhealthy, 31.6 percent. There is a similar pattern in households' share of wealth held in risky assets: households with two healthy spouses and households with an unhealthy male (17.6 percent and 16.8 percent, respectively) have larger shares of risky assets in their portfolios than households where either the female or both spouses are unhealthy (13.6 percent and 12.7 percent, respectively). We observe a similar relation in the second age group as in the first one, but the differences between health states are smaller. Finally, the differences between groups are very small in the last two age categories. This may suggest that older households in our data hold a similar share of their wealth in risky assets irrespective of their health status.

Table 3: Risky Assets Ownership and share of risky assets by Health States and Age Group of the Household

Household health states by age group	**Households owning risky assets**	**Share of risky assets**	**Households within age group**
65-70 years old			
Both Healthy	41.3%	17.6%	38.9%
Healthy Female and Unhealthy Male	41.7%	16.8%	29.2%
Unhealthy Female and Healthy Male	35.2%	13.6%	17.3%
Both Unhealthy	31.6%	12.7%	14.6%
N			9,618
71-75 years old			
Both Healthy	41.5%	17.3%	31.9%
Healthy Female and Unhealthy Male	40.5%	16.1%	33.1%
Unhealthy Female and Healthy Male	37.3%	15.2%	15.1%
Both Unhealthy	36.0%	14.1%	19.9%
N			13,882
76-80 years old			
Both Healthy	39.1%	17.7%	24.5%
Healthy Female and Unhealthy Male	41.3%	18.5%	36.9%
Unhealthy Female and Healthy Male	42.5%	19.1%	14.4%
Both Unhealthy	40.6%	18.9%	24.2%
N			11,006
81 and older			
Both Healthy	40.5%	21.1%	18.9%
Healthy Female and Unhealthy Male	41.3%	20.3%	35.3%
Unhealthy Female and Healthy Male	41.9%	21.2%	14.1%
Both Unhealthy	41.2%	21.2%	31.7%
N			9,797
N Total			44,303

Source: waves 1994-2014 of the RAND HRS Longitudinal File 2014 (V2).
An unhealthy individual has faced a negative health shock while a healthy one has not.

5 Empirical specification

We follow the model presented in Section 3 and assume that the share of wealth held in risky assets depends on negative health shocks facing the female, S_f, and the male, S_m.[23] The share of risky assets also depends on a vector of the male's individual characteristics X_m, a vector of the female's individual characteristics X_f, the household's wealth W, the household's labor income I, and the household's non-labor income NI. Taking this into account, we estimate empirical specifications for the household's probability of owning risky assets and portfolio share of risky assets, respectively.[24] Although our main interest lies in the specification for the portfolio share of risky assets, we think that it is also important to look into the household's probability of owning risky assets as a first step, i.e., the extensive margin. We start by examining whether the probability of participating in the stock market is positively or negatively associated with a health shock for either spouse. We would like to see whether households were less likely to own risky assets if one of the spouses had faced a negative health shock.

Our first empirical specification estimates the following random effects probit model for the household's probability of owning risky assets at time t.[25]

$$Y_{it}^{*} = \beta_0 + \beta_1 S_{mit} + \beta_2 S_{fit} + \beta_3 X_{mit} + \beta_4 X_{fit} + \beta_5 W_{it} + \beta_6 I_{it} + \beta_7 NI_{it} + \tau_t + v_i + \epsilon_{it}, \tag{12}$$

$$Y_{it} = \begin{cases} 1 & \text{if} \quad Y_{it}^{*} > 0 \\ 0 & \text{otherwise} \end{cases} .$$

In equation (12), τ_t is a time-specific effect, v_i is the household-specific effect, and ϵ_{it} is an error term that is normally distributed with zero mean and variance σ_{ϵ}^{2}. S_{mit} and S_{fit} are dummies for whether the male and the female, respectively, have experienced a negative health shock during in the last two years prior to the year of observation, t. The vectors X of individual characteristics for the male and female include, for each of them, explanatory variables that are available in our dataset and that the previous literature has found to have statistically significant effects on the probability of owning risky assets and on portfolio share of risky assets, i.e., age (Bertaut and Starr-McCluer 2002, Rosen and Wu 2004), education (King

[23] In section 3 we studied the amount, and not the share, of wealth invested in the risky asset and not the share of risky assets, but this does not modify the qualitative results of our analysis since the share of wealth held in risky assets would be defined as the amount invested in risky assets, α, divided by total wealth A (a positive number).

[24] Similar specifications were estimated by Berkowitz and Qiu (2006). Compared with our estimations, they use the first six waves of the HRS, and also include younger individuals in their sample.

[25] We present the estimates of random effects models rather than fixed effects models since this strategy is more practical in a short panel with relatively little time variation (Edwards 2008). Furthermore, the estimation of probit and tobit models with fixed effects results in estimates that are not consistent, i.e., the estimates of the probit case are not sufficient statistics. There may be other statistics in the sample that provide additional information as to the value of the parameter (Greene 2002).

and Leape 1998, Rosen and Wu 2004, Berkowitz and Qiu 2006), and the household's number of children (Rosen and Wu 2004).[26] To account for wealth W, we include the household's total wealth as defined in the data section. Finally, we include the household's earned income I and non-earned income NI (similar to previous studies such as Rosen and Wu [2004] that include variables for household income and net worth in their model, and Berkowitz and Qiu [2006], who include labor income as well as financial and non-financial assets).

We extend equation (12) and estimate a second model that includes dummies for private medical insurance and federal government health insurance, respectively. The latter equals one if the respondent had access to any of the following federal government health insurance programs: Medicare, Medicaid, TRI-CARE, CHAMP-VA or any other government health care plan. It equals zero otherwise. We also add the male's and female's out-of-pocket medical expenditure, respectively, to the explanatory variables. We include the insurance and out-of-pocket variables since expenditure on own health that is not covered by a medical insurance can affect the amount allocated to risky assets.

To address the determination of the share of risky assets, we estimate a Tobit model as follows:

$$\alpha_{it} = \begin{cases} \alpha^*_{it} \quad if \quad \alpha^*_{it} > 0 \\ 0 \quad if \quad \alpha^*_{it} \leq 0 \end{cases} \quad ,$$

where the latent portfolio share of risky assets is

$$\alpha^*_{it} = \pi_0 + \pi_1 S_{mit} + \pi_2 S_{fit} + \pi_3 X_{fit} + \pi_4 X_{mit} + \pi_5 W_{it} + +\pi_6 I_{it} + \pi_7 NI_{it} + \mu_t + \kappa_i + \psi_{it}. \qquad (13)$$

In equation (13), μ_t is a time-specific effect, κ_i is the household-specific effect, and ψ_{it} is an error term that is normally distributed with zero mean and variance σ^2_ψ. We use the same explanatory variables as in (12) to estimate (13).

Finally, we relate the theoretical framework in Section 3 to the empirical specification (13) in the following way: equation (4) and equation (5) relate to π_1 and π_2, respectively. In equation (13), π_1 and π_2 can be understood as the change in the latent share of risky assets when the male's health and female's health decreases (they face a negative health shock), respectively, everything else constant. Although there are two main differences between the theoretical model and the empirical model, these do not affect the qualitative results of our analysis. First, in the theoretical model, we focus on the amount, and not share, of wealth invested in risky assets. This does not affect the sign of the right-hand side of equations (4) and (5) since the share of risky assets equals the amount of wealth invested in the risky assets divided by total wealth, which we have restricted to be positive in both the theoretical and empirical models. Second, in the theoretical model, we calculate the change in the investment in risky assets due to a change

[26] Not necessarily living in the household.

in the probability of enjoying good health. Hence, we can think of a negative health shock as a decrease in the probability of being in good health.

6 Results

6.1 Ownership probabilities

Table 4 presents the average marginal effects for the probability of owning risky assets. More specifically, we show that in all model specifications, the average marginal effect of a negative health shock facing the female is statistically significant at the 5 percent level, and it is associated with an approximately 1.2 percentage-point reduction in the household's probability of owning risky assets.[27] Based on the first model and the numerical analysis in Section 3, this result suggests that households show prudence and are more likely to own less risky assets when the female faces a negative health shock. In this way, the household insures itself against the female facing a health shock. There is no similar relationship between the probability of holding risky assets and a negative health shock facing the male: in all specifications, the effect of a negative health shock facing the male is statistically insignificant. Based on the theoretical model, this result suggests that prudent households are better insured against the male facing a negative health shock.

Table 4 also shows that the probability of owning risky assets is positively correlated with years of education. Every additional year of education attained by the male is associated with a 3.1 percentage-point increase in that probability. For females, the corresponding percentage-point increase is 2.3 in all specifications.[28] Rosen and Wu (2004) and Berkowitz and Qiu (2006) also estimated a positive effect of the male's education level on the probability of owning risky assets. This may suggest that individuals with more education may understand the financial market better and be more able to face financial risk than less educated individuals. The estimates of the average marginal effects for number of children are also statistically significant and negative in all specifications. One more child is associated with a 0.4 percentage-point reduction in the probability of owning risky assets. Hence, we can think that households with children consume more than households without children, and thus, they prefer to hold less risky assets. It can also be the case that households are closer to the age of leaving a bequest, which increases

[27]We also estimated specifications with interaction effects between the negative health shock and age, but the estimates of these interaction effects were not statistically significant and those of the other parameters remained robust compared with the specification without interaction effects.

[28]We estimated a specification with two dummy variables: 1) high school and 2) more than high school as highest attained education. The estimates for these dummies were positive and statistically significant, and the estimates of the effects of covariates remained robust.

the households' preferences for safe over risky assets. Furthermore, our results indicate that increases in household wealth are associated with increases in the probability of owning risky assets, suggesting that the households become less risk averse when wealth increases. This result is also supported by Berkowitz and Qiu (2006), who estimate the effect of wealth by including a variable for financial and a variable for non-financial wealth. Both variables seem to have a positive and statistically significant effect on the probability of owning risky assets in their study. Unlike Berkowitz and Qiu (2006), who document a positive effect of labor income on the ownership probability, although they include individuals aged 50 and older, our results suggest that non-earned income is positively associated with ownership probability. We find no statistically significant correlation between earned income and ownership probability. This seems reasonable since nearly 75 percent of the individuals in our sample reported being retired and therefore had non-earned income as their only source of income.

We can also see that government medical insurance does not seem to be significantly correlated with the probability of owning risky assets in specifications 2 and 3. Hence, we separate such insurance into different types in specifications 4 and 5 to see whether they affect the ownership probability in different directions. We find that there are statistically significant correlations between some of the government insurances and the probability of owning risky assets. Specifically, the female's MEDICARE estimate is associated with a 2.7 percentage-point increase in the probability of owning risky assets, and the male's and female's MEDICAID estimates are associated with a 3.6 and 11.6 percentage-point decrease, respectively. Finally, the estimate for the male's private medical insurance is associated with a 3.6 percentage-point increase in the ownership probability, and the female's private medical insurance estimate is associated with a 1.8 percentage-point increase. These results suggest that households where the female had access to MEDICAID are also associated with a lower probability of owning risky assets. Since this type of insurance is for those whose resources are insufficient to pay for health care, it may pick up characteristics that are not captured by the income and wealth variables that affect the household's investment in risky assets.

Table 4: Average Marginal Effects of Risky Asset Ownership Probit Model.

	(1)	(2)	(3)	(4)	(5)
Male's Negative Health Shock	-0.241	-0.313	-0.303	-0.318	-0.301
	(0.497)	(0.500)	(0.501)	(0.503)	(0.504)
Female's Negative Health Shock	-1.183*	-1.245*	-1.277*	1.215*	-1.248*
	(0.521)	(0.522)	(0.523)	(0.527)	(0.528)
Male's Age	0.137*	0.145	0.143	0.146	0.145
	(0.082)	(0.082)	(0.082)	(0.082)	(0.083)
Female's Age	0.211*	0.207*	0.208*	0.203*	0.204**
	(0.085)	(0.085)	(0.085)	(0.086)	(0.086)
Male's Years of Education	3.173***	3.137***	3.138***	3.107***	3.107***
	(0.123)	(0.124)	(0.124)	(0.125)	(0.125)
Female's Years of Education	2.356***	2.264***	2.263***	2.197***	2.196***
	(0.157)	(0.157)	(0.158)	(0.159)	(0.159)
Number of children	-0.462**	-0.417**	-0.416**	-0.377*	-0.377*
	(0.164)	(0.164)	(0.164)	(0.165)	(0.165)
Household Wealth/(10^5)	0.542***	0.546***	0.546***	0.544***	0.544***
	(0.022)	(0.022)	(0.022)	(0.022)	(0.022)
Household Earned Income/(10^4)	-0.106	-0.121	-0.121	-0.123	-0.123
	(0.064)	(0.065)	(0.065)	(0.065)	(0.065)
Household Non Earned Income/(10^4)	0.291***	0.284***	0.281***	0.281***	0.278***
	(0.028)	(0.028)	(0.028)	(0.028)	(0.028)
Male's Government Insurance		1.261	1.267		
		(1.658)	(1.657)		
Males's Medicare				1.140	1.144
				(1.603)	(1.603)
Males's Medicaid				-3.639*	-3.638**
				(1.642)	(1.643)
Males's VA/CHAMPUS				-0.355	-0.350
				(0.951)	(0.951)
Female's Governement Insurance		1.865	1.867		
		(1.451)	(1.451)		
Female's Medicare				2.789*	2.791*
				(1.400)	(0.014)
Female's Medicaid				-11.621***	-11.612***
				(1.771)	(1.772)
Female's VA/CHAMPUS				0.297	0.333
				(1.450)	(1.451)
Males's Private Insurance		3.687***	3.683***	3.628***	3.624***
		(0.506)	(0.506)	(0.512)	(0.512)
Female's Private Insurance		2.108**	2.121**	1.823**	1.840**
		(0.516)	(0.517)	(0.523)	(0.524)
Males's Out of Pocket Med Exp/(10^4)			-0.042		-0.115
			(0.241)		(0.249)
Female's Out of Pocket Med Exp/(10^4)			0.326		0.340
			(0.282)		(0.288)
N	44, 303	44, 143	44, 143	43,717	43,717

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$. Standard errors in parenthesis. Source: waves 1994-2014 of the HRS. All specifications include year controls. The estimates of the xtprobit regression in Stata for all specifications are presented in the Appendix.

In Table 5, we divide the sample into age groups conditional on the male's age. The results show that the effect of a negative health shock to the female is statistically significant only in the group where the male is 65-70 years old, and it is associated with a 4.8 percentage-point decrease in the probability of owning risky assets. For this age group, negative health shocks are least frequent (as shown in Table 2) and might therefore be the most unexpected, which can make the effect of such a shock largest in this age group.

The estimate of the effect corresponding to the male's private insurance is statistically significant and positive in all specifications, and is associated with an increase in the probability of owning risky assets. In contrast, the estimate of the female having a private insurance is only statistically significant in specifications (2) and (3), and is also associated with an increase in the ownership probability. This suggests that the household owns more assets when it is insured against the cost of poor health. Hence, the private insurance reduces the risk of having monetary costs related to poor health.

Table 5: Average Marginal Effects of Asset Ownership Probit Model Conditioned on Male Age Groups.

	Male's Age			
	65-70	**71-75**	**76-80**	**81 and older**
	(1)	**(2)**	**(3)**	**(4)**
Male's Negative Health Shock	-0.908	0.024	0.095	-1.076
	(1.057)	(0.862)	(0.964)	(1.081)
Female's Negative Health Shock	-4.801***	-1.142	0.440	-1.584
	(1.130)	(0.917)	(0.991)	(1.047)
Male's Age	0.178	-0.044	-0.544*	-0.102
	(0.311)	(0.251)	(0.271)	(0.201)
Female's Age	-0.226	0.133	0.411**	0.182
	(0.253)	(0.151)	(0.135)	(0.141)
Male's Years of Education	2.545***	2.939***	2.833***	2.748***
	(0.228)	(0.180)	(0.193)	(0.211)
Female's Years of Education	2.291***	1.664***	1.931***	2.094***
	(0.276)	(0.228)	(0.255)	(0.287)
Number of children	-1.167***	-0.579*	-0.511*	-0.162
	(0.304)	(0.239)	(0.259)	(0.302)
Household Wealth/(10^5)	0.316***	1.008***	1.291***	1.133***
	(0.032)	(0.050)	(0.061)	(0.067)
Household Earned Income/(10^4)	-0.070	-0.307**	-0.132	-0.481
	(0.105)	(0.108)	(0.207)	(0.311)
Household Non Earned Income/(10^5)	0.651***	0.301***	0.322***	0.074
	(0.076)	(0.053)	(0.069)	(0.050)
Males's Medicare	0.231	-0.308	1.285**	-0.323
	(2.584)	(2.895)	(3.845)	(4.645)
Female's Medicare	1.513	1.052***	-0.551	-4.329
	(2.302)	(2.426)	(4.009)	(4.124)
Males's Medicaid	-1.233	-7.197*	-7.215*	-7.416
	(3.682)	(3.364)	(3.025)	(3.984)
Female's Medicaid	-13.144**	-8.991*	-15.309***	-12.359**
	(4.374)	(3.629)	(3.254)	(3.813)
Males's VA/CHAMPUS	-4.193	-1.179	0.693	1.342
	(2.456)	(1.884)	(1.833)	(1.987)
Female's VA/CHAMPUS	5.786	4.351	-4.527	-0.835
	(3.326)	(2.738)	(2.643)	(2.910)
Male's Private Insurance	6.108***	4.456***	2.747**	5.455***
	(1.182)	(0.974)	(1.058)	(1.092)
Female's Private Insurance	1.230	4.707***	4.284***	0.152
	(1.215)	(1.011)	(1.079)	(1.105)
Male's Out of Pocket Med Exp/(10^4)	-0.141	-1.075*	0.621	1.318*
	(0.887)	(0.487)	(0.705)	(0.624)
Female's Out of Pocket Med Exp/(10^4)	0.040	-0.714	-0.384	0.811
	(0.784)	(0.694)	(0.610)	(0.540)
N	9,533	13,731	10,864	9,589

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$. Standard errors in parenthesis. Source: waves 1994-2014 of the HRS. All specifications include year controls. The estimates of the xtprobit regression in Stata for all specifications are presented in the Appendix.

6.2 Share of wealth held in risky assets

Table 6 presents the estimates of the tobit model for the household's share of wealth held in risky assets. We are mostly interested in the correlation between negative health shocks and this share. Our results show that the effect of a negative health shock to the male is not statistically significant while a negative health shock to the female is statistically significantly correlated with the share of wealth held in risky assets at the 5 percent level, and corresponds to a decrease of approximately 2 percentage-points. The theoretical framework suggests that the household experiences a loss in consumption due to the monetary cost of the female facing a health shock. Moreover, if we assume that the spouses' health and household consumption are complements, a reduction in consumption will also follow after a negative health shock to the female. Hence, the household has an incentive to reduce the share of risky assets held in order to insure itself against the female facing a health shock. Finally, as discussed before, the results in this section also suggest that the household is better insured against a negative health shock facing the male, which implies that the household's consumption is less affected by this.

In all specifications, we estimate that the share of risky assets held increases by approximately 2.5 percentage-points with the male's age, by 0.3 percentage-points with household wealth, and by 1.4 percentage-points with household non-earned income. As in the previous section, these results may suggest that households become less risk averse when their wealth increases. The associations between the male's and female's private insurance, respectively, and the share of risky assets held are also statistically significant and imply an increase of approximately 4.5 and 2 percentage-points. Finally, only the estimate of the effect of the female's out-of-pocket medical expenditure is statistically significant and corresponds to a 0.8 percentage-point increase in the share of wealth held in risky assets. The latter could reflect the households' wealth characteristics, i.e., wealthier households can afford to cover medical expenditures and invest in the stock market at the same time. Finally, both the male's and female's MEDICAID are statistically significant and associated with a reduction in the share of risky assets

Table 6: Estimates of Share of Risky Assets Tobit Model.

	(1)	(2)	(3)	(4)	(5)
Male's Negative Health Shock	-0.218	-0.242	-0.287	-0.283	-0.321
	(0.748)	(0.748)	(0.748)	(0.750)	(0.751)
Female's Negative Health Shock	-1.892*	-2.057**	-2.142**	-1.896*	-1.977*
	(0.794)	(0.794)	(0.795)	(0.797)	(0.798)
Male's Age	2.517*	2.346	2.467*	2.104	2.213
	(1.215)	(1.217)	(1.218)	(1.224)	(1.224)
Female's Age	-2.035	-1.938	-1.914	-1.949	-1.921
	(1.291)	(1.296)	(1.296)	(1.303)	(1.303)
Male's Years of Education	5.329***	5.251***	5.248***	5.179***	5.176***
	(2.144)	(0.213)	(0.213)	(0.214)	(0.214)
Female's Years of Education	4.022***	3.904***	3.889***	3.746***	3.733***
	(0.261)	(0.261)	(0.261)	(0.261)	(0.261)
Number of children	-0.352	-0.295	-0.293	-0.232	-0.232
	(0.265)	(0.265)	(0.265)	(0.265)	(0.265)
Household Wealth/(10^5)	0.360**	0.361***	0.360***	0.359***	0.358***
	(0.018)	(0.018)	(0.018)	(0.018)	(0.018)
Household Earned Income/(10^4)	-0.093	-0.105	-0.106	-0.101	-0.101
	(0.088)	(0.090)	(0.090)	(0.089)	(0.089)
Household Non Earned Income/(10^5)	0.146***	0.142***	0.139***	0.142***	0.139***
	(0.024)	(0.024)	(0.024)	(0.024)	(0.024)
Males's Government Insurance		0.307	0.295		
		(2.485)	(2.484)		
Males's Medicare				1.183	1.178
				(2.408)	(2.408)
Males's Medicaid				-6.637*	-6.628*
				(2.654)	(2.655)
Males's VA/CHAMPUS				1.319	1.393
				(1.431)	(1.431)
Female's Government Insurance		3.010	3.006		
		(2.192)	(2.196)		
Female's Medicare				4.159	4.152
				(2.145)	(2.144)
Female's Medicaid				-22.935***	-22.874***
				(3.228)	(3.227)
Female's VA/CHAMPUS				-2.774	-2.609
				(2.208)	(2.208)
Males's Private Insurance		4.737***	4.737***	4.539***	4.541***
		(0.759)	(0.759)	(0.765)	(0.765)
Female's Private Insurance		2.494***	2.518***	2.043**	2.077**
		(0.769)	(0.769)	(0.776)	(0.776)
Males's Out of Pocket Med Exp/(10^4)			0.513		0.438
			(0.328)		(0.341)
Female's Out of Pocket Med Exp/(10^4)			0.815*		0.806*
			(0.394)		(0.399)
Constant	-1.951***	-1.993***	-2.043***	-1.879***	-1.925***
	(0.389)	(0.391)	(0.005)	(0.394)	(0.394)
N	44, 303	44, 143	44, 143	43, 717	43, 717

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$. Standard errors in parenthesis. Source: waves 1994-2014 of the HRS. The table shows the estimates of a random effects tobit regressions in Stata(xttobit command) for all specifications. All specifications include the square of the male's age, the square of the female's age, and year controls.

In Table 7, once again we divide the data into age groups conditional on the male's age. The results show that the correlation between the female's negative health shock and the share of wealth held in risky assets is statistically significant for the groups where the male is 65-70 and 71-75 years old, with it being larger for the first group. It could be the case that for households with younger females, the negative health shock is more unexpected, and thus, this type of household might be less insured against the female's poor health and face higher costs. If this is true, households with younger females might reduce their share of wealth held in risky assets more than households with older females.

Table 7: Estimates of Share of Risky Assets Tobit Model
Conditioned on Male Age Groups.

	Male's Age			
	65-70	**71-75**	**76-80**	**81 and older**
	(1)	**(2)**	**(3)**	**(4)**
Male's Negative Health Shock	-1.402	-0.983	1.342	-2.110
	(1.517)	(1.211)	(1.449)	(1.696)
Female's Negative Health Shock	-6.668***	-2.825*	0.729	-2.917
	(1.651)	(1.299)	(1.505)	(1.695)
Male's Age	26.818	-3.671	-20.989	-3.427
	(28.026)	(3.385)	(43.011)	(6.382)
Female's Age	-4.521	4.741	17.398***	-0.239
	(5.687)	(4.163)	(4.288)	(2.871)
Male's Years of Education	3.742***	4.632***	4.780***	5.201***
	(0.346)	(0.273)	(0.314)	(0.383)
Female's Years of Education	3.584***	2.714***	3.224***	4.054***
	(0.411)	(0.332)	(0.398)	(0.497)
Number of children	-1.477***	-0.588	-0.617	-0.158
	(0.449)	(0.345)	(0.401)	(0.513)
Household Wealth/(10^5)	0.245***	0.479***	0.846***	0.543***
	(0.026)	(0.033)	(0.051)	(0.045)
Household Earned Income/(10^4)	-0.078	-0.118	0.020	-0.890*
	(0.136)	(0.132)	(0.272)	(0.414)
Household Non Earned Income/(10^5)	0.289***	0.291***	0.037	0.165*
	(0.046)	(0.051)	(0.044)	(0.067)
Males's Medicare	-3.446	4.953	15.797*	5.159
	(3.591)	(4.092)	(6.262)	(7.471)
Female's Medicare	3.714	14.895***	0.536	-8.949
	(3.184)	(3.907)	(5.974)	(6.583)
Males's Medicaid	-2.113	-6.361	-1.932***	-9.101
	(5.303)	(4.934)	(5.232)	(7.377)
Female's Medicaid	-22.847**	-15.963*	-2.427***	-33.843***
	(7.481)	(5.639)	(6.362)	(7.766)
Males's VA/CHAMPUS	-3.927	1.322	0.917	4.517
	(3.757)	(2.624)	(2.696)	(3.073)
Female's VA/CHAMPUS	1.619	0.058	-9.216*	-3.331
	(4.685)	(3.710)	(4.233)	(4.726)
Male's Private Insurance	7.163***	4.352***	4.905**	7.744***
	(1.656)	(1.370)	(1.606)	(1.712)
Female's Private Insurance	1.516	7.571***	5.154**	0.155
	(1.719)	(1.395)	(1.634)	(1.715)
Male's Out of Pocket Med Exp/(10^4)	0.458	-0.778	0.132	1.336
	(1.155)	(0.572)	(0.781)	(0.716)
Female's Out of Pocket Med Exp/(10^4)	0.135	0.186	-0.487	0.903
	(1.055)	(0.891)	(0.831)	(0.717)
Constant	-8.868	-1.979	0.069	-0.029
	(9.671)	(12.333)	(16.731)	(2.742)
N	9,533	13,731	10,864	9,589

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$. Standard errors in parenthesis. Source: waves 1994-2014 of the HRS. The table shows the estimates of a random effects tobit regressions in Stata (xttobit command) for all specifications. All specifications include the square of the male's age, the square of the female's age, and year controls.

6.3 Sensitivity analysis

We estimated two more specifications: one with an "initial health" variable and a second with a lagged health shock variable. Initial health indicates whether the individual had experienced a health shock before their first observation in our data, while the lagged variable indicates whether the individual had a health shock in the previous wave. We also include interaction effects between these variables and the actual health shock. We included these variables and interaction effects since a recent health shock may be more expected for households that have already experienced a health shock in the past, thus they might have changed their share of risky assets in the past already. The results are in line with our previous specifications. The male's negative health shock is not statistically significant, while the female's is. Specifically, when we control for initial health, the estimates for the female's negative health shock are negative and statistically significant when the male is 65-70 years old and 71-75 years old (see Table 12 in the Appendix). When we control for the lagged health shock, the female negative health shock is significant when the male is 65-70 years old (see Table 13 in the Appendix). Finally, we find that the estimates of the interaction effects mentioned above and the estimates of the effects of the initial health and lagged health shock are not statistically significant.

7 Conclusions

In this paper, we examine the relationship between recent negative health shocks and the probability of owning risky assets and the risk exposure of household portfolios, respectively, among aging couples in the U.S. We define negative health shocks as one or more of the following diagnoses by a medical doctor: cancer or a malignant tumor of any kind except skin cancer, stroke or transient ischemic attack, and heart attack, coronary heart disease, angina, congestive heart failure, or other heart problems.

More specifically, we observe that a major negative health shock to the female is associated with a decrease in the probability of owning risky assets. However, a similar health shock to the male does not seem to have an effect on the probability of owning risky assets.

Our results are in line with the previous studies of Rosen and Wu (2004) or Berkowitz and Qiu (2006), who find that a negative health shock to the female is negatively associated with the probability of owning risky assets, while a negative health shock to the male does not significantly affect this probability.

Furthermore, we find that a negative health shock to the female is associated with a decrease in the share of household wealth invested in risky assets, in particular when the male is 65-70 years old and 71-75 years old. We do not find a statistically significant correlation for the male facing a negative health shock.

The theoretical model suggests that the results could be explained by households being better insured against the male facing poor health, which implies that the household's consumption is less affected by a negative health shock to the male. In turn, the share of household wealth held in risky assets is less affected as well. Our theoretical analysis also suggests that households show prudence and prefer to reduce the share of wealth held in risky assets to insure themselves against the loss in consumption due to the cost of the female's negative health shock.

Future research will need to consider the household's intertemporal decisions regarding consumption and investment in risky assets. Theoretical analyses should take into account that households consists of two spouses who may have different preferences, and should therefore use collective household models. Finally, it is also important to incorporate more channels that may affect the investment in risky assets, i.e., life expectancy, mortality, and bequest motives, into both the theoretical and empirical analysis.

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Appendices

A Mathematical Appendix

By totally differentiating (3) with respect to α, p^f, and p^m, we obtain the following equation:

$$\begin{aligned}
&\big[p^f p^m E[u''^{gg}\delta^2] + p^f(1-p^m)E[u''^{gb}\delta^2] + (1-p^f)p^m E[u''^{bg}\delta^2] + (1-p^f)(1-p^m)E[u''^{bb}\delta^2]\big]\,d\alpha \\
+&\big[p^m E[u'^{gg}\delta] + (1-p^m)E[u'^{gb}\delta] - p^m E[u'^{bg}\delta] - (1-p^m)E[u'^{bb}\delta]\big]\,dp^f \\
+&\big[p^f E[u'^{gg}\delta] - p^f E[u'^{gb}\delta] + (1-p^f)E[u'^{bg}\delta] - (1-p^f)E[u'^{bb}\delta]\big]\,dp^m = 0. \qquad \text{(A1)}
\end{aligned}$$

If we assume that the female's probability of good health changes $dp^f \neq 0$ and the male's probability of good health does not change $dp^m = 0$, then from equation (A1) we obtain the change in the expression for the change in the amount invested in the risky asset due to a change in the female's probability of good health,

$$\frac{d\alpha}{dp^f} = -\frac{1}{D}\big\{p^m\big(E[u'^{bb}\delta] - E[u'^{bg}\delta]\big) + (1-p^m)\big(E[u'^{gb}\delta] - E[u'^{bb}\delta]\big)\big\}$$

If we decompose the expected value into the sum of the product of the expectations and the covariance, $E[XY] = E[X]E[Y] + cov[X,Y]$, we obtain the following equation:

$$\begin{aligned}
\frac{d\alpha}{dp^f} = -\frac{1}{D}\big\{&p^m\big(E[u'^{gg}]\hat{\delta} - E[u'^{bg}]\hat{\delta} + cov[u'^{gg},\delta] - cov[u'^{bg},\delta]\big) \\
&+ (1-p^m)\big(E[u'^{gb}]\hat{\delta} - E[u'^{bb}]\hat{\delta} + cov[u'^{gb},\delta] - cov[u'^{bb},\delta]\big)\big\},
\end{aligned}$$

where

$$D = p^f p^m E[u''^{gg}\delta^2] + p^f(1-p^m)E[u''^{gb}\delta^2] + (1-p^f)p^m E[u''^{bg}\delta^2] + (1-p^f)(1-p^m)E[u''^{bb}\delta^2],$$

and $\hat{\delta} = E[\delta]$. Similarly, the expression when the male's probability of good health changes, ceteris paribus, is

$$\frac{d\alpha}{dp^m} = -\frac{1}{D}\big\{p^f\big(E[u'^{gg}\delta] - E[u'^{gb}\delta]\big) + (1-p^f)\big(E[u'^{gb}\delta] - E[u'^{bb}\delta]\big)\big\}.$$

Again, by decomposing into expectations and covariances we obtain:

$$\begin{aligned}
\frac{d\alpha}{dp^m} = -\frac{1}{D}\big\{&p^f\big(E[u'^{gg}]\hat{\delta} - E[u'^{gb}]\hat{\delta} + cov[u'^{gg},\delta] - cov[u'^{gb},\delta]\big) \\
&+ (1-p^f)\big(E[u'^{bg}]\hat{\delta} - E[u'^{bb}]\hat{\delta} + cov[u'^{gb},\delta] - cov[u'^{bb},\delta]\big)\big\}.
\end{aligned}$$

B Notes and Tables

Note to Table 1.

Number of children is the number of the spouses' living children. The household's total wealth is the sum of the net value of real estate (including primary residence), vehicles, and businesses and the value of the household's financial assets and all other savings. The household's financial assets include: checking, savings and money market accounts, CDs (Certificates of Deposit), government savings and T-bills, corporate, municipal and foreign bonds and bond funds, stocks, mutual funds, and IRA and Keogh accounts. The spouses' health insurance variables are dummies that indicate if the spouses' health insurance plans incorporate coverage by any of the following federal government health insurance programs: Medicare, Medicaid, VA/CHAMPUS (Civilian Health and Medical Program of the Uniformed Services), CHAMP-VA (The Civilian Health and Medical Program of the Department of Veterans Affairs) or any Medicare, Medicaid, VA/CHAMPUS, CHAMP-VA or any other government-provided health insurance. Private insurance indicates whether the spouse is privately insured. The household's earned income is the sum of the spouses' wage/salary income, commissions/tips and earnings from professional trade or practice. The household's non-earned income is defined as the sum of the households' capital income and the income from employer pensions or annuities, social security, unemployment or workers compensation for injuries, and all other government transfers. The negative health shock takes the value one if the spouse has experienced cancer or a malignant tumor of any kind except skin cancer, heart attack, coronary heart disease, angina, congestive heart failure, or other heart problems, or stroke or transient ischemic attack, and zero otherwise.

Table 8: Estimates of Asset Ownership Probit Model.

	(1)	(2)	(3)	(4)	(5)
Male's Negative Health Shock	-0.014	-0.018	-0.017	-0.018	-0.017
	(0.029)	(0.029)	(0.029)	(0.029)	(0.029)
Female's Negative Health Shock	-0.069*	-0.072*	-0.074*	-0.071*	-0.072*
	(0.030)	(0.030)	(0.030)	(0.031)	(0.031)
Male's Age	0.064	0.060	0.061	0.052	0.054
	(0.047)	(0.047)	(0.047)	(0.048)	(0.048)
Female's Age	-0.007	-0.012	-0.012	-0.012	-0.012
	(0.051)	(0.051)	(0.051)	(0.051)	(0.051)
Male's Age^2	-0.000	-0.000	-0.000	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Female's Age^2	0.000	0.000	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Male's Years of Education	0.185***	0.182***	0.182***	0.180***	0.180***
	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)
Female's Years of Education	0.137***	0.131***	0.131***	0.127***	0.127***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Number of children	-0.027**	-0.024*	-0.024*	-0.021*	-0.022*
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Household Wealth/(10^5)	0.032***	0.032***	0.032***	0.031***	0.031***
	(0.001)	(0.001)	(0.001)	(0.012)	(0.001)
Household Earned Income/(10^4)	-0.006	-0.007	-0.007	-0.007	-0.007
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Household Non Earned Income/(10^4)	0.017***	0.017***	0.016***	0.016***	0.016***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Male's Government Insurance		0.073	0.074		
		(0.097)	(0.097)		
Males's Medicare				0.066	0.066
				(0.093)	(0.093)
Males's Medicaid				-0.214*	-0.214*
				(0.098)	(0.098)
Males's VA/CHAMPUS				-0.021	-0.020
				(0.055)	(0.055)
Female's Government Insurance		0.109	0.109		
		(0.085)	(0.085)		
Female's Medicare				0.163*	0.163*
				(0.083)	(0.083)
Female's Medicaid				-0.719***	-0.719***
				(0.119)	(0.119)
Female's VA/CHAMPUS				0.017	0.019
				(0.083)	(0.083)
Male's Private Insurance		0.214***	0.214***	0.211***	0.211***
		(0.029)	(0.029)	(0.029)	(0.029)
Female's Private Insurance		0.122***	0.123***	0.105***	0.106***
		(0.030)	(0.030)	(0.030)	(0.030)
Males's Out of Pocket Med Exp/(10^4)			-0.002		-0.006
			(0.001)		(0.014)
Female's Out of Pocket Med Exp/(10^4)			0.001		0.019
			(0.001)		(0.016)
Constant	-7.991***	-8.005***	-8.006***	-7.662***	-7.737***
	(1.573)	(1.577)	(1.577)	(1.596)	(1.597)
N	44, 303	44, 143	44, 143	43, 717	43, 717
rho	0.735	0.731	0.731	0.731	0.731
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)

$^*p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$. Standard errors in parenthesis. Source: waves 1994-2014 of the HRS. The table shows the estimates of a random effects probit regression in Stata (xtprobit command) for all specifications. All specifications include year controls.

Table 9: Estimates of Asset Ownership Probit Model

Conditioned on Male Age Group.

	Male's Age			
	65-70	**71-75**	**76-80**	**81 and older**
	(1)	**(2)**	**(3)**	**(4)**
Male's Negative Health Shock	-0.054	0.001	0.006	-0.070
	(0.063)	(0.057)	(0.066)	(0.070)
Female's Negative Health Shock	-0.291***	-0.076	0.030	-0.103
	(0.069)	(0.061)	(0.068)	(0.068)
Male's Age	1.744	1.139	-0.797	-0.052
	(1.256)	(1.718)	(2.074)	(0.272)
Female's Age	-0.096	0.124	0.973***	-0.090
	(0.236)	(0.205)	(0.200)	(0.121)
Male's Age^2	-0.012	-0.007	0.004	0.000
	(0.009)	(0.011)	(0.013)	(0.001)
Female's Age^2	0.000	-0.001	-0.006***	0.000
	(0.001)	(0.001)	(0.001)	(0.000)
Male's Years of Education	0.153***	0.196***	0.195***	0.179***
	(0.014)	(0.013)	(0.015)	(0.015)
Female's Years of Education	0.137***	0.111***	0.133***	0.136***
	(0.017)	(0.015)	(0.018)	(0.019)
Number of children	-0.070***	-0.038*	-0.035*	-0.011
	(0.018)	(0.016)	(0.018)	(0.019)
Household Wealth/(10^5)	0.019***	0.067***	0.089***	0.073***
	(0.018)	(0.003)	(0.005)	(0.004)
Household Earned Income/(10^4)	-0.004	-0.021**	-0.009	-0.031
	(0.006)	(0.007)	(0.014)	(0.020)
Household Non Earned Income/(10^5)	0.039***	0.020***	0.022***	0.004
	(0.004)	(0.003)	(0.005)	(0.003)
Males's Medicare	0.013	-0.021	0.952**	-0.021
	(0.155)	(0.192)	(0.313)	(0.301)
Female's Medicare	0.091	0.743***	-0.037	-0.278
	(0.140)	(0.185)	(0.275)	(0.261)
Males's Medicaid	-0.074	-0.497*	-0.514*	-0.500
	(0.223)	(0.242)	(0.224)	(0.279)
Female's Medicaid	-0.853**	-0.628*	-1.155***	-0.860**
	(0.316)	(0.269)	(0.281)	(0.288)
Males's VA/CHAMPUS	-0.256	-0.079	0.047	0.087
	(0.152)	(0.126)	(0.126)	(0.128)
Female's VA/CHAMPUS	0.342	0.286	-0.318	-0.054
	(0.193)	(0.177)	(0.190)	(0.191)
Male's Private Insurance	0.368***	0.297***	0.189**	0.355***
	(0.072)	(0.065)	(0.073)	(0.071)
Female's Private Insurance	0.074	0.314***	0.296***	0.009
	(0.073)	(0.068)	(0.075)	(0.072)
Male's Out of Pocket Med Exp/(10^4)	-0.008	-0.071*	0.042	0.085*
	(0.053)	(0.032)	(0.048)	(0.041)
Female's Out of Pocket Med Exp/(10^4)	0.002	0.047	-0.026	0.052
	(0.047)	(0.046)	(0.042)	(0.035)
Constant	-60.956	-52.541	-11.642	0.367
	(43.299)	(62.684)	(80.677)	(11.752)
N	9,533	13,731	10,864	9,589
rho	0.753	0.789	0.794	0.779
	(0.015)	(0.011)	(0.012)	(0.013)

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$. Standard errors in parenthesis. Source: waves 1994-2014 of the HRS. The table shows the estimates of a random effects probit regressions in Stata (xtprobit command) for all specifications. All specifications include year controls.

Table 10: Sensitivity Analysis 1: Estimates of Risky Share Tobit Model Conditioned on Male Age Groups.

	Male's Age			
	65-70	**71-75**	**76-80**	**81 and older**
	(1)	**(2)**	**(3)**	**(4)**
Female's Negative Health Shock	-7.929***	-8.478**	-0.653	-4.008
	(2.249)	(1.934)	(2.591)	(4.246)
Female's Initial Health	14.274	5.533	-5.089	-31.384
	(13.017)	(12.964)	(15.892)	(18.496)
N	6,568	7,621	4,569	1,827

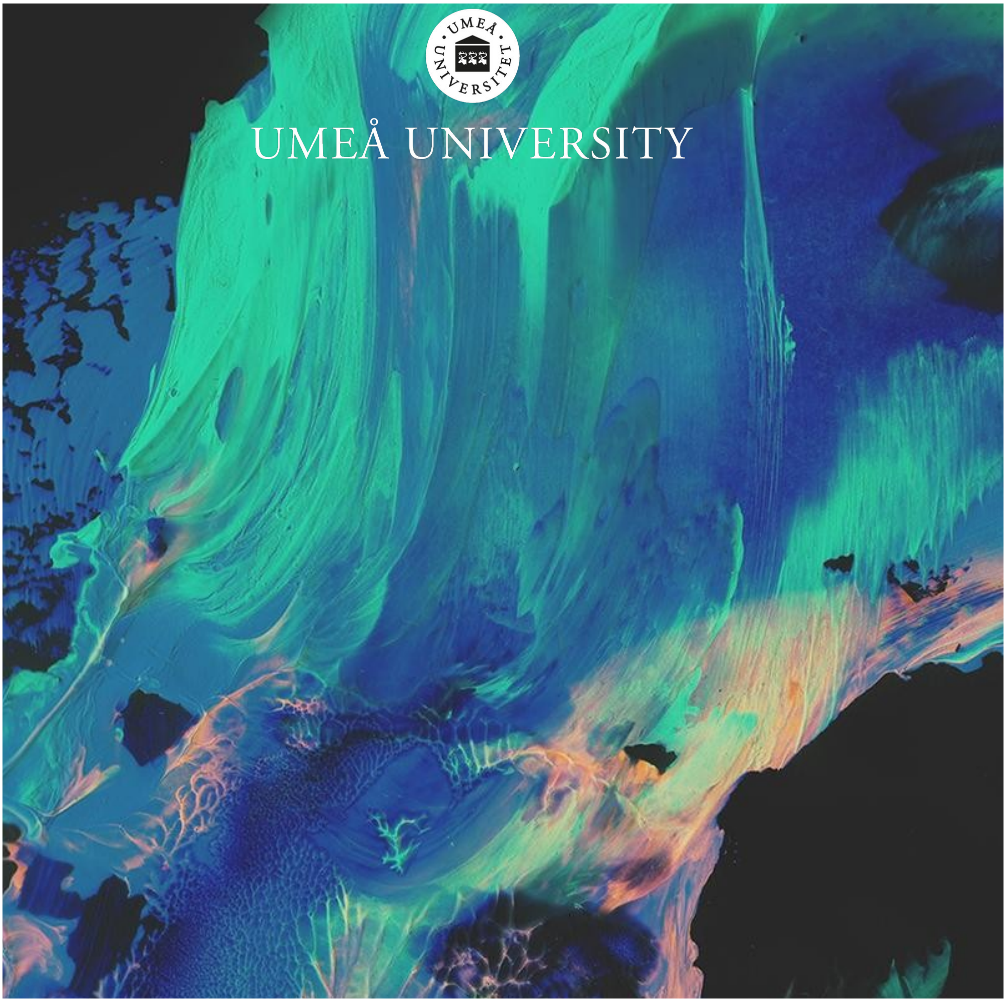
*$p < 0.05$, **$p < 0.01$, ***$p < 0.001$. Standard errors in parenthesis. Source: waves 1994-2014 of the HRS. Initial health indicates if the female has ever had a health shock before her first observation. All specifications include year controls, the same covariates as in Table 7, and an interaction effect between the negative health shock and initial health.

Table 11: Sensitivity Analysis 2: Estimates of Risky Share Tobit Model Conditioned on Male Age Groups.

	Male's Age			
	65-70	**71-75**	**76-80**	**81 and older**
	(1)	**(2)**	**(3)**	**(4)**
Female's Negative Health Shock	-9.285**	0.264	2.427	-3.217
	(2.725)	(2.064)	(2.296)	(2.418)
Female's Lagged Negative Health Shock	2.304	-5.384	-2.395	-10.184
	(10.484)	(6.664)	(5.794)	(5.244)
N	8,378	11,281	9,679	8,539

*$p < 0.05$, **$p < 0.01$, ***$p < 0.001$. Standard errors in parenthesis. Source: waves 1994-2014 of the HRS. Lagged negative health shock indicates if the female had a health shock in the previous wave. All specifications include year controls, the same covariates as in Table 7, and an interaction effect between the negative health shock and the lagged negative health shock.

3



The effects of involuntary job losses on depressive symptoms among couples

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CEDAR Working Papers 2021:5
Centre for Demographic and Ageing Research

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Abstract

The way job loss affects mental health within couples attracts increasing attention. While previous studies focused on the working-age population, this paper investigates how job loss affects spouses' depressive symptoms in later life. We focus on spouses aged 50 or older, who represent a vulnerable segment of the labor market, and who also have a high prevalence of depressive symptoms.
We use data from the Health and Retirement Study, which provides longitudinal information about changes in labor market status and mental health outcomes among respondents and their spouses in the United States. To deal with potential reverse causality problems, we utilize data on job loss resulting from business closures. We find that job loss may affect the levels of depressive symptoms among the spouses of older adults who experience these adverse life course events. The effects are gendered, as women are negatively affected by job losses experienced by their husbands, but we do not observe such harmful effects among men whose wives lose their jobs. We also show how the effects of job loss vary across couples with differing levels of economic resources and health care needs, as well as differential access to health care.

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The effects of involuntary job losses on depressive symptoms among couples.

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This study has received funding from the European Research Council (ERC) under the European Union's Horizon 2020 research and innovation program under grant agreement No 802631.

1. Introduction

Losing a job carries social stigma, undermines personal status and identity, lowers self-esteem, and triggers stress and anxiety (Jahoda 1981; Pearlin et al. 1981). While many studies provide valuable insights on the effects of unemployment on mental health (Krug and Prechsl 2020; Mousteri et al. 2018; Paul and Moser 2009; Pearlman 2015; Wanberg 2012), these studies mostly focus on the individuals who experienced job losses. However, a job loss may affect not only the individual who experiences this negative life course event, but also the people in his or her family (Brand 2015; Ström 2003). Especially spouses, who are emotionally close and also economically interdependent, may be indirectly affected by job losses within families.

The effects of job losses on mental health among couples remains an understudied topic. Previous research provides some evidence for the working-age population. Specifically, Bubonya et al. (2017); Marcus (2013) and Mendolia (2014) as well as Jolly (2020), show that men's job losses have negative effects on their female spouses' mental health, while job loss among women has no significant effect on the mental health of their husbands[1]. However, we need more knowledge about the consequences of job losses on mental health among older couples. The only study which addresses this topic specifically was carried out by Siegel et al. (2003), who found no evidence of spillover effects, possibly due to a small sample size[2]. In the context of population aging, it is crucial to understand the interrelatedness of the working lives and health of people approaching retirement. This is particularly true in the United States, where more than 20 percent of the total population will be over the age of 65 within a decade (Colby and Ortman 2014). Older workers with less assets can be viewed as particularly vulnerable because job losses have more significant financial consequences later

[1] Related literature that focused on a different outcome, namely, happiness or life satisfaction, observed gendered patterns in the spillover effects of unemployment among couples of working age (Inanc 2018; Kim and Do 2013; Nikolova and Ayhan 2019).

[2] Siegel et al. (2003) examined mental health outcomes among the wives of 38 husbands who lost their jobs. Our study uses a larger sample, hence we were able to use more sophisticated methods of analysis that provide stronger conclusions.

in life (Couch et al. 2009). Moreover, re-employment becomes less likely for older workers (Chan and Stevens 2001), meaning that employability, which moderates the negative consequences of missed earnings in the working-age population (Green 2011; Silla et al. 2009), can no longer shield older people.

The theoretical framework in this paper brings together insights from different disciplines to explain the mechanisms of the impact of job loss on mental health among couples, as well as to discuss potential buffering mechanisms. Then, in the empirical analysis, we compare the magnitude of the effects of spousal job losses on mental health across age, employment status, and wealth, to see whether economic resources may play a protective role. We also examine whether these effects differ across couples with diverging levels of health needs and access to health insurance. The empirical analysis uses longitudinal data from the Health and Retirement Study (HRS), which provides data on the labor market careers and the mental health dynamics of individuals and their spouses. By focusing on job losses resulting from business closures, we reduce the potential reverse causality bias. Previous research on the health effects of unemployment stressed that workers with poorer health face a higher risk of job loss, and a lower chance of reemployment (Stewart 2001; Strully 2009), and highlighted the need to take this selection bias into account. It is plausible that similar selectivity arises with respect to spousal mental health, which may impose additional caregiving duties and limit individual productivity (Benazon and Coyne 2000; Revenson et al. 2016; Van Wijngaarden et al. 2004), limiting the labor market opportunities of spouses (García-Gómez et al. 2013; Siegel 2006). However, when an entire business closes, it is unlikely that the mental health of a worker's spouse caused the ensuing job loss. Thus, by focusing on job losses that result from business closures, we take this potential source of bias into account.

The following section discusses the theoretical framework. Section 3 presents the data used in the analysis. Section 4 provides the methodology used in the empirical analysis, while Section 5 presents the results. Finally, Section 6 concludes the paper.

2. Theoretical background

The mental health effects of job loss on spouses can be understood from the perspective of family stress models developed within psychology and family sociology (Lavee 2013). According to this literature, adverse life course events such as the loss of a job, may lead to a cascade of additional stressors such as an abrupt change in financial situation, a lifestyle change, or relocation to less expensive housing. Loss of a job may also require the substantial re-organization of a couple's private and professional lives. This in turn may lead to emotional reactions and behaviors that place a burden on spouses (Rook et al. 1991). This pile-up of stressors may negatively affect spouses' mental health (Maitoza 2019). Hence, the negative impact of a job loss experienced by one of the spouses may go beyond the directly affected individual and extend to the other spouse, possibly impairing his or her mental health.

While theories developed in psychology and family sociology focused on the negative effects of job losses among family members, economic theory often emphasizes health as something that results from individual—rather than family-specific—processes. However, economic health capital models present a different perspective and view families as units that jointly manage resources to improve the health of family members (Bolin et al. 2001). This perspective emphasizes that families do not invest the same amount of time and resources in the health of different family members (Jacobson 2000). Adverse life course events such as job losses shape the re-allocations of investments in health capital within the family, which are the outcome of a non-cooperative bargaining game between spouses (Bolin et al. 2001).[3]

[3] In the non-cooperative game, each spouse is assumed to know the equilibrium strategies of the other spouse, and no spouse has anything to gain by changing only his/her strategy.

In these models, each spouse's utility function increases with the health of both spouses and in the consumption of goods. Also, the marginal utility of consumption for each spouse increases with each spouse's health capital. Furthermore, the assumptions within a theoretical model posit that when one spouse loses a job, the distribution of health capital between them will be affected by the possibility of divorce and by specialization within the household. For example, if the wife loses her job, this leads to a loss of family income that reduces both spouses' consumption and hence their utility. It follows that the utility for the husband of divorce will increase, compared with the utility of staying married, since he has the advantage of being employed, and therefore retains bargaining power. There are two ways for a family to compensate for lost income and avoid divorce in this example. One is by investing in the husband's health, and the other is by investing in his wife's health, since the utility of each spouse increases with the health of both spouses. By investing in the husband's health, the family makes the husband derive more utility from the consumption good. Assuming that the husband attaches a higher value to his own health than to his wife's health, it follows that the family will reallocate the wife's lost market time into health-enhancing time for her husband. Hence, the Nash-bargaining process will reallocate time and resources in order to increase the husband's utility within the marriage and it will decrease the wife's utility. This result becomes stronger if traditional gender roles are considered. In this case, the husband would have a higher level of market-related human capital than his wife, who would have invested more in household-related capital. For the husband, being healthy has a higher value in the labor market than in household production activities. Hence, when one spouse loses a job, a couple may use the time gained in this way to redistribute resources in order to improve the health of the other spouse. These relocations may take different forms, including additional efforts that the jobless spouse may undertake to improve their spouse's wellbeing, such as additional caregiving or joint leisure activities. While implying greater economic inequalities within couples, these re-allocations of resources can be considered compensation strategies that buffer family members against the spillover effects of job loss. Therefore, based on this theoretical model, one can expect that the mechanisms behind the job loss of one spouse may positively affect the mental health of the other spouse.

Both the family-stress literature and the health capital models recognize that the effects of job loss on spousal mental health may vary within and across couples. Couples may have differential possibilities for reallocating resources after job losses. Regarding differences within couples, individuals differ in terms of their knowledge and capabilities to improve health (Tipper 2010). Due to gender differences in socialization processes, women are more likely to assume caregiving roles than men, and as a result may be more likely to develop skills related to improving the health and wellbeing of other family members (Eagly and Wood 2016). Women are also more likely than men to feel responsible for meeting the needs of family members (Kessler and McLeod 1984). Hence, job loss experienced by a female spouse may be more easily followed by re-allocations of resources, compared to when a male spouse loses their job. These gendered differences may also result from specialization within the family. If the partner’s labor income plays an important role in covering household consumption costs, the female may specialize in household production, and invest in her husband’s health capital, to keep him participating in the labor market. Therefore, we may expect gender differences in the effects of job losses on spousal health, with more negative effects on mental health observed among women whose husbands lose jobs, and less negative or even positive effects observed on men’s mental health when their wives lose jobs.

The effects of job losses on spousal mental health could also vary across age and spousal health status. The relative importance of labor income for covering household consumption costs varies across life stages. As people approach retirement age, pensions may make them less dependent on labor income, and hence job losses may lead to weaker effects among older couples. In addition, preferences for joint leisure activities as well as healthcare needs may change across variations in age and health status. When a spouse who lost a job reallocates time to care for his or her partner, any beneficial effects may be stronger among older couples, or couples where one spouse suffers from poor health.

Some couples may experience greater disruption in functioning, and higher levels of distress, following a job loss, because of differential access to economic resources. Previous research argues that missed earnings, the loss of employer-provided health insurance, and financial hardship, play an important role in mediating the effects of unemployment on spousal health (Salm 2009). Therefore, especially in the context of the United States, we can expect job loss to produce stronger effects on couples where the lost job results in a lack of earnings or health insurance, as compared to when one spouse remains employed, and his or her income can protect the couple from economic hardship. Likewise, wealth, in the form of economic resources accumulated over a lifetime, may cushion couples from both the financial and health consequences of job loss, especially later in life (Berkowitz and Qiu 2006; Christelis et al. 2005; Vega and Velli 2020). Some studies suggest that wealth is a more relevant measure of economic standing among older people than earnings (Pollack et al. 2007; Semyonov et al. 2013), especially because this population subgroup holds a substantial amount of the total U.S. household net worth (Coile and Milligan 2009).

3. Data

We use longitudinal data from the Health and Retirement Study (HRS). The HRS is supported by the National Institute on Aging and the Social Security Administration. It is a biennial, nationally representative panel that follows across time approximately 7,000 households in the U.S. We use 11 waves from 1994 to 2014[4], focusing on couples where both the respondent and his/her spouse are 50 to 65 years old. Our final sample consists of 10,338 couples where the female reported the depressive symptoms, and 9,598 couples where the male did so. The unique feature of this survey is that the data contain health-related information, and economic and demographic variables for both the respondent, and the respondent's spouse. In the subsequent part of the paper, we use the term 'spouses' to refer to the partners of respondents who experienced job losses (our treatment group), or partners of the respondents who did not experience job losses (our control group).

Of particular interest for our study is that the data set provides information about the spouses' depressive symptoms measured by Center for Epidemiologic Studies Depression scale (CES-D). The CES-D inventory was constructed to identify populations at risk of developing depressive disorders, and it has a high construct validity and internal consistency (Radloff 1977). This inventory is used for screening purposes and it should not be used as a diagnostic tool. The reliability of the inventory has been confirmed across genders and nationalities (Van de Velde et al. 2010). It measures whether spouses identified with one or more of the following feelings all or most of the time: 'depressed,' 'have sleeping problems,' 'everything is an effort,' 'feel alone,' 'feel sad,' 'cannot get going,' 'feel happy,' and 'enjoy life.' Every sentiment takes a value of one if it was experienced by the individual, and a value of zero otherwise. The CES-D score is obtained by adding the value of every 'negative' sentiment, the first six sentiments in the list, plus two minus the total value of the 'positive' sentiments. The summated rating ranges from zero to eight, with higher scores indicating higher levels of depressive symptoms.

[4] The first wave of the HRS was conducted in 1992, but we do not use it because CES-D scores were not reported in this wave.

The HRS also provides information about respondents' employment history. In each wave of the study, the respondents are asked whether they are currently working for pay. If the answer is no, the respondents are asked to indicate reasons why they are not working. One reason provided by the HRS questionnaire is, due to lay-offs resulting from a business closure. With this information, we can determine whether the respondent faced an involuntary job loss due to a business closure between waves of the HRS. Thus, our measure of involuntary job loss includes 264 observations who answered that their previous employers' business had closed. We focus on those individuals who remained unemployed between waves. By focusing on involuntary job loss as the reason for the respondent's unemployment, we deal with the potential issue of reverse causation, which could result in biased estimates.

Since our dependent variable measures a spouse's depressive symptoms, our set of covariates includes characteristics that may affect spouse's mental health: age, age squared, race, years of education, number of marriages, number of children, and a variable indicating the year when the survey was carried out. Age is measured as the year of the survey's wave, minus the year when the spouse was born. The covariate for race indicates if the spouse belongs to one of the following ethnic groups: white/Caucasian, black/African American, or other ethnic group. Number of marriages denotes the number of times that the individual has been married, including his/her current marriage. Number of children denotes the number of living children of the respondent and spouse.

Apart from controlling for spousal characteristics, we also include health and social capital characteristics of a respondent that may affect the spouses' marital choices, employment dynamics, and mental health. Such variables are the respondent's early life health, i.e., the number of conditions infections experienced during childhood, as well as the spouse's parents' years of education. The number of conditions indicates whether the respondent experienced one or more of the following conditions before the age of sixteen: vision problems, asthma, diabetes, respiratory problems, speech problems, allergy, heart trouble, ear problems, epilepsy, headache, stomach problems, high blood pressure, depression, drug problems, or other psychological problems. The number of infections that the respondent experienced before the age of sixteen includes measles, mumps, and chicken pox. We create and include dummy variables for those respondents with missing observations on the number of illnesses, as well as on the number of infections during childhood, and include them as covariates in our analyses.

Following insights from previous research showing that macroeconomic changes, and especially major recessions, may affect the risk for both job loss and mental health, we control for such period effects. To this end, the wave of the survey is included in the analysis.

Table 1 shows the means of the variables used in this paper for females and males, respectively. The table shows that the CES-D score is larger for females whose spouse experienced a job loss as compared to females whose spouses did not experience a job loss. Females whose spouse remained employed report a CES-D score of approximately 1.27, while those whose spouse became unemployed report a score of 2. The table shows that for males the opposite is true; the CES-D score of males whose spouse became unemployed is larger than for those whose spouse remained working. The table also shows that the CES-D differences across job-loss categories are significant. These differences suggest that there may be gendered differences in the effects of job loss on the spouses' CES-D scores.

Table 1. Summary Statistics

	Respondent did not experience job loss (Untreated)	Respondent experienced job loss (Treated)	Respondent did not experience job loss (Untreated)	Respondent experienced job loss (Treated)
	Female		Male	
Spouse's CES-D	1.27	2.00***	1.00	0.79*
	(1.86)	(2.15)	(1.61)	(1.44)
Age	56.03	57.97***	58.31	59.91***
	(3.77)	(3.87)	(3.81)	(3.37)
White	83.59%	77.41%*	85.18%	87.4%
Black/Afro-American	9.94%	15.06%*	11.67%	5.88%
Other ethnicity	6.47%	7.53%*	6.15%	6.72%
Years of education	13.01	12.01***	13.46	12.73*
	(2.89)	(3.23)	(2.83)	(3.51)
Number of marriages	1.37	1.31	1.37	1.411
	(0.69)	(0.71)	(0.65)	(0.76)
Number of children	3.09	3.17	3.02	3.00
	(1.85)	(1.74)	(1.79)	(1.84)
# of infections during childhood	1.21	1.21	1.22	1.14
	(1.32)	(1.31)	(1.29)	(1.25)
# of conditions during childhood	0.31	0.43*	0.41	0.35
	(0.71)	(0.92)	(0.92)	(0.75)
Mother's years of education	10.31	9.48**	10.56	9.69***
	(3.66)	(4.08)	(3.49)	(3.55)
Father's years of education	9.92	9.26*	10.29	9.87
	(4.11)	(4.47)	(3.92)	(3.79)
Observations	10,192	146	9,479	119

Source: HRS data. The table shows the means and the standard deviations in parentheses. The table also shows the proportion of ethnic groups in the sample. *p<0.1,**p<0.05,***p<0.01, significance levels for differences between the means for the untreated and for the treated, for females and males respectively. The differences were estimated by two sample t-tests with unequal variances.

4. Methodology

Inference about the impact of the respondent's involuntary job loss on his/her spouse's CES-D score involves an assumption about how the spouse would have performed had the respondent not experienced a job loss. To address this, we use a binary treatment where the treatment indicator Di equals one if respondent i experienced an involuntary job loss due to a business closure, and zero otherwise. There are two potential outcomes: Yi1 and Yi0. Yi1 is the spouse's CES-D when the respondent experienced an involuntary job loss, while Yi0 is the outcome when the respondent kept his/her job. It follows that the treatment effect for the treated spouse i can be written as

$$\tau_i = [Y_i^1 - Y_i^0 | D_i = 1] \quad (1)$$

where i=1,..., N. Since only one of the potential outcomes is observed for each individual i, estimating the individual treatment effect for the treated τi is not possible. Therefore, we have to concentrate on the average treatment effects. We would like to focus on the effects of those who experienced a job loss, in other words, on the average treatment effect on the treated (ATET). The ATET is given by

$$\tau_{ATET} = E[\tau|D=1] = E[Y^1|D=1] - E[Y^0|D=1] \quad (2)$$

The ATET is defined as the difference between expected outcome values with and without treatment for those whose spouse experienced a job loss. As the counterfactual mean for those being treated -E[Y 0|D = 1]- is not observed, we have to choose a proper substitute for it in order to estimate the ATET. Using the mean outcome of untreated individuals E[Y0|D = 0] is not a good option, since it is likely that components which determine the treatment decision also determine the outcome variable.

Thus leading to selection bias:

$$E[Y^1|D=1] - E[Y^0|D=0] = \tau_{ATET} + E[Y^0|D=1] - E[Y^0|D=0]. \quad (3)$$

where $E[Y0|D = 1]-E[Y0|D = 0]$ is the bias attributed to non-random selection. The true parameter $\tau ATET$ is only identified if the bias equals zero.[5]

4.1 Estimation strategy

We follow recent literature on matching methods, which suggests the use of multivariate distance matching (MDM) (Gu and Rosenbaum, 1993) in order to identify τ_{ATET}, over other matching methods that have been used in the past, such as propensity score matching (King and Nielsen, 2019). Part of the argument in favor of MDM is that PSM approximates complete randomization, while MDM approximates fully blocked randomization.[6] The estimates obtained from fully blocked randomization are more efficient than those obtained by complete randomization (King and Nielsen, 2019). Another issue with PSM is that complete randomization gives place to random pruning, which increases the variance by deleting observations at random. The latter increases imbalance within confounding factors and model dependence. MDM does not lead to random pruning and therefore dominates PSM in terms of imbalance and model dependence.[7]

MDM creates a sample of individuals who experienced a job loss resulting from the closure of a business, which is comparable on all observed covariates X to a sample of individuals who remained working. MDM conditions on X, and for each observation in the treatment group, match on a metric distance (MD) that measures the proximity between observations in the

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[5] The methodology also takes into account the conditional independence assumption. This assumption requires that the common variables that affect treatment assignment and treatment-specific outcomes should be observable. The dependence between treatment assignment and treatment-specific outcomes can be removed by conditioning on these observable variables.

[6] The blocks method consists of splitting the observations into groups (blocks) that are similar to each other, separated by smaller sub-blocks, so that the groups can be considered almost independent. This method randomizes observations within blocks such that an equal number are assigned to each treatment and will tend to be uniformly distributed by their related characteristics. A complete randomization may allocate a different number of observations into treatment groups, the latter may reduce the statistical power to reject the null hypothesis (Efird, 2011).

[7] In the Appendix, we show the results when using PSM. We find that the qualitative results found with MDM hold even when using PSM, although some of the values for the ATET differ slightly.

multivariate space of X. The Y values of these matching observations are then used to compute the counterfactual outcome of the observation at hand. Finally, an estimate for the average causal effect can be obtained as the mean of the differences between the observed values and the computed counterfactual values over all observations. We use a Kernel matching algorithm which gives a larger weight to controls for which the metric distance is small. The ATET in this case is:

$$\tau^{MDM} = E\,\{E[\tau|D=1] = E[Y(1)|D=1, MD(X)] - E[Y(0)|D=0, MD(X)]\} \quad (4)$$

MD(X) is the Mahalanobis metric distance.[8] X is a vector that includes the covariates listed earlier in the data section. We use the kmatch md command in Stata for the estimations (Ben, 2017). The command uses a bandwidth algorithm similar to the one proposed by Huber et al. (2013, 2015) to determine whether the treated and control observations are a potential match.[9] If the distance between these two observations is larger than the specified bandwidth, then these are not matched. In most of our estimations, all treated observations were within the bandwidth and were matched.

4.2 Sample Balancing and Matching Quality

We examine if the matching procedure is able to balance the distribution between the covariates in the control and treatment groups, and produce good-quality matching. To do this, we use the covariates' standard mean difference and variance ratio. An ideal balance is when the mean and variance are the same between the treated and the control group. Nevertheless, a standard mean difference close to zero, and a variance ratio below two, is

[8] We use the kmatch command in Stata which uses the Mahalanobis (Rubin, 1980) metric distance: $MD(X_i,X_j)= \sqrt{(Xi-Xj)'\Sigma-1(Xi-Xj,)}$ where Σ is the covariance matrix of X.

[9] The algorithm sets the bandwidth to f times the q-quantile of the distribution of (non-zero) distances between observations in one-to-one matching (Ben, 2017).

generally acceptable for a good match (Zhang et al. 2017). The figures section in the Appendix show that the mean difference of most covariates is zero or very close to zero. The covariate that has the largest difference in absolute value is age, with an approximate value of 0.2. The variance ratio is also similar across specifications; it is smaller than 1.8, and for most covariates it equals or approaches one. In most of the specifications, the covariates with the largest ratio are the number of marriages and the father's years of education. Overall, the values for the covariates' mean difference and variance ratio suggest that the sample balancing and matching quality is generally good.

4.3 Stratified analyses

Following previous research, which emphasizes gender differences in the health effects of job losses, we estimate models for male and female spouses separately. To see whether the effects of job losses on spousal health vary across population subgroups, with differential preferences for leisure activities and companionship and different needs for personal care, we stratify samples by age and health status. Age is divided into three categories: 50-55, 55-60, and 60-65. Health status is based on whether spouses have experienced at least one of the following health shocks between waves in the survey: heart attack, stroke, or new cancer diagnosis. To see whether social and economic resources constitute a buffer against the effects of job loss, we also estimate the specification mentioned above, across the spouse's employment status, the household's level of wealth, and the spouse's access to medical insurance. We consider whether the spouses have access to a government or private insurance plan.[10] We classify households as 'low wealth level' if their wealth level is below the sample's median household wealth (USD 441,378.90), and otherwise as high-wealth level households. Household wealth includes both financial and non-financial assets.

[10] The government health insurance plans incorporate coverage by any of the following federal government health insurance programs: Medicare, Medicaid, VA/CHAMPUS, CHAMP-VA, or any other government-provided health insurance.

5. Results

In order to address our question regarding the effects of job losses on spousal mental health, we examine the effects of job losses using MDM. Table 2 shows the ATET by age group for female and male spouses, respectively. The results suggest that the impact of a job loss experienced by a respondent affects his/her spouse in different ways depending on gender. Specifically, the ATET increases levels of depressive symptoms among females, while it decreases levels of depressive symptoms among males. The CES-D score for females increases by an average of 0.74 points when their husband loses his job. At the same time, the results show that the CES-D score for males aged 56 to 60 decreases by an average of 0.51 (MDM) points when wives experience a job loss.

We expected that the effects of job losses on spousal mental health may differ across age. As people approach retirement age, pensions may make them less dependent on labor income, and job losses may have stronger effects on spousal mental health among older couples. However, the results presented in Table 2 do not support this idea. We do not observe any specific age pattern. We find that the ATET is statistically significant. Increases in depressive symptoms are observed for female spouses in all groups, but the increase is larger for females aged 50-55; somewhat smaller and not significant in the 56-60 age group; then again rather large (over 0.8 CES-D score) in the age group 60-65. Similarly, we do not observe any specific patterns of decreasing effects across groups among male spouses. While the analysis of the pooled sample revealed that the male spouses of respondents who experienced job losses reported fewer depressive symptoms, this result is clearly not driven by the effects observed in the oldest age group, which most likely has access to old-age pensions and other non-labor income sources.

Table 2. Treatment effects by age group

	Age Group			
	50-65	50-55	56-60	61-65
Female spouse				
	0.742***	1.234***	0.465*	0.926***
	(0.189)	(0.411)	(0.247)	(0.333)
Obs	10,338	4,951	3,875	1,512
Male spouse				
	-0.186	-0.079	-0.510***	0.046
	(0.138)	(0.352)	(0.131)	(0.227)
Obs	9,598	2,478	4,052	3,068

Source: HRS data. *p<0.1,**p<0.05,***p<0.01. Standard errors in parenthesis. The dependent variable is the spouse's CES-D.

We hypothesized that preferences for companionship and healthcare needs may change across health status, diminishing any negative effects of job losses on mental health among spouses. Following argumentation by Bolin et al. (2001), one could argue that when a respondent who loses a job reallocates the time gained, after ceasing paid work, to provide care for his or her spouse, the beneficial effects of this re-location of resources may offset the negative effects of job losses. According to the results presented in Table 3, we find a statistically significant effect of job loss among female spouses, who did not experience a health shock. Among women, who did experience a health shock, the effect was smaller and it is not statistically significant. However, male spouses of respondents who experienced job losses do not report fewer depressive symptoms if they have poorer health. In fact, the positive effects of job losses on mental health of male spouses observed earlier turn out to be mainly driven by the subgroup of men who have not experienced any health shocks.

Table 3. Treatment Effects by health shock

	Health Shock	
	No	Yes
Female spouse		
	0.816***	0.719*
	(0.215)	(0.415)
Obs	8,306	1,839
Male spouse		
	-0.343***	0.129
	(0.108)	(0.414)
Obs	7,491	1,983

Source: HRS data. *p<0.1,**p<0.05,***p<0.01. Standard errors in parenthesis. The dependent variable is the spouse's CES-D. ATET stands for average treatment effects on the treated. Health shock equals yes if the spouse has experienced one or more of the following between waves in the survey: heart attack, stroke, or new cancer diagnosis.

In the next step, we examined whether economic resources may constitute a buffer against the effects of job loss. In Table 4, spouses age 50 to 65 are classified by employment status, household wealth level, and access to health insurance. The results show that the ATET is statistically significant for both non-working and working females, and the estimates are larger for non-working females. The latter may suggest that there might be a buffering effect of employment among women. Furthermore, the ATET for female spouses in low-wealth households is larger than for members of wealthier households. The latter may suggest that for female spouses, household wealth may constitute a buffer against their husbands' job losses. Table 4 also shows that the ATET is smaller but still statistically significant for female spouses who do have health insurance than for those who are uninsured. All in all, we find some evidence for the moderating role of economic resources, but it mainly concerns employment status and accumulated resources such as wealth, whereas the evidence is less clear when it comes to health insurance.

The results for male spouses reveal a completely different pattern. The effects of job losses on CES-D scores are negative and stronger among non-working male spouses as compared to working male spouses. That is, the number of depressive symptoms reported by men decreases when their wives face a job loss, and it does so more strongly among men who do not have earnings that could potentially serve as a buffer for financial consequences of job losses experienced by their wives. Similarly, the effects of job loss are stronger among male spouses in less affluent households. However, we do find that for male spouses the ATET is statistically significant only for those who have access to health insurance. Overall, we observe that if female respondents do not work, male spouses report fewer depressive symptoms, and these effects are not contingent on access to economic resources.

Table 4. Treatment effects by access to economic resources

	Employment status		Wealth level		Health insurance	
	Unemployed	Employed	Low	High	No	Yes
Female spouse						
	0.946***	0.698***	0.995***	0.416**	0.667	0.495**
	(0.353)	(0.197)	(0.288)	(0.211)	(0.421)	(0.207)
Obs	3,332	6,997	4,386	5,952	800	8,570
Male spouse						
	-0.671**	-0.086	-0.449**	-0.141	-0.276	-0.282**
	(0.270)	(0.158)	(0.179)	(0.155)	(0.343)	(0.135)
Obs	2,224	7,360	3,766	5,832	649	8,160

Source: HRS data. *p<0.1,**p<0.05,***p<0.01. Standard errors in parenthesis. The dependent variable is the spouse's CES-D. ATET stands for average treatment effects on the treated. We classify households as low-wealth househoulds if their wealth level is below the sample's median household wealth (USD 441,378.90), and otherwise as high-wealth households. The household's wealth includes both financial and non-financial assets. Health insurance equals yes if the spouse has access to a government or private insurance plan. The government health insurance plans incorporate coverage by any of the following federal government health insurance programs: Medicare, Medicaid, VA/CHAMPUS, CHAMP-VA, or any other government- provided health insurance.

5.1 Sensitivity Analysis

We do a sensitivity analysis by controlling for pre-existing mental health. Although this is not part of our main strategy, since we lose a significant number of observations, we would like to see if the results remain robust when we add the following covariates to the specifications: the spouses' CES-D score from the previous wave, the respondent's CES-D score in the last wave, the CES-D score change of the respondent between waves, and an interaction term between the latter and job loss. We find that adding these covariates to the specifications do not affect the main results of our analyses.

One question that we would like to address is how response to the treatment differs across spouses. One reason for heterogenous effects could be, for example, the respondent's differential risk of job loss. Individuals whose spouses are more likely to become unemployed could respond with mental health outcomes that differ from those married to individuals with a lower probability of job loss.[11] To address how the effects of job loss differ across spouses of respondents with diverging employability, we use a stratification method to estimate heterogenous treatment effects similar to Xie et al. (2012). We construct balanced propensity score strata such that the average value of the propensity score and covariates do not significantly differ between spouses of respondents who lost a job and those who did not. By doing this, we assume that individuals in the same stratum have a similar response to their spouses losing a job. We estimate the MDM specifications given the propensity score stratum. We only estimate the base-line specification for the 50-65 group since stratification makes the number of treated observations very small or zero when trying to follow our previous set-up.[12]

[11] One way to do this is to control for industry or sector. We could not control for these in our specifications because we did not have enough observations in the data.

[12] Xie et al (2012) recommend using at least 20 treated and untreated observations per strata.

Table 5 shows the results of the heterogeneous treatment effects estimation by spousal gender. The table shows that the effects for female spouses are positive and significant while they are not statistically significant for male spouses. The estimates for female spouses range from a 0.65 increase in the CES-D score for spouses of respondents with the lowest propensity to lose a job, to a 0.78 and 0.74 increase in stratums 2 and 3, respectively. The results are in range with the estimate shown in Table 2 (an increase of 0.56 in the CES-D score).

Table 5. Heterogenous effects of job loss on spouses' CES-D

	ATET	Standard Error	Obs.
Female			
P-score Stratum [0-0.012)	0.644*	(0.331)	6,413
P-score Stratum [0-12.025)	0.778**	(0.357)	2,336
P-score Stratum [0.025-1]	0.738***	(0.280)	1,525
Male			
P-score Stratum [0-0.012)	-0.277*	(0.157)	7,211
P-score Stratum [0-12.025)	-0.187	(0.282)	1,674
P-score Stratum [0.025-1]	0.311	(0.470)	662

Source: HRS data. *p<0.1,**p<0.05,***p<0.01. Standard errors in parenthesis. The dependent variable is the spouse's CES-D. ATET stands for average treatment effects on the treated. Propensity scores were generated by a probit regression model.

6. Discussion

How losing a job affects a spouse has remained an understudied topic in the literature on the health effects of unemployment (Brand 2015; Ström 2003), but has received increasing attention. Recent studies have shown that the effects of job losses may extend to spouses among couples of working age (Bubonya et al. 2017; Marcus 2013; Mendolia 2014), pointing also to the gendered nature of these spillover effects. In this paper, we provide evidence for how job loss affects spouses' mental health among older couples, using data from the HRS. This paper provides also insights into how the mental health effects of job losses for spouses vary across population subgroups, and shows what factors magnify such effects or protect older couples.

Our results confirm earlier findings on the gendered patterns of the spillover effects of job losses, showing that the differences are also pronounced among older couples. The levels of depressive symptoms of female spouses of respondents who lost jobs increases, which is consistent with the family stress literature (Lavee 2013). We also find that among women, the negative effects of job loss on mental health are stronger if they are not employed, and if couples have less accumulated wealth and lack access to health insurance, suggesting the buffering role of economic resources. In contrast, for male spouses of respondents who lost jobs, depressive symptoms decrease, which is consistent with predictions from the economic health capital models mentioned before. Our analysis revealed also that respondents’ job losses do not affect their male spouse’s mental health negatively, even if these men are not involved in paid work, if they live in less affluent households, or if they lack access to health insurance. Overall, our results suggest that when women lose jobs, this may trigger a reallocation of intrahousehold resources towards improving male spouses’ wellbeing, as suggested by theoretical predictions from Bolin et al. (2001). The same relocation of resources towards a female spouse does not necessarily take place when the respondent losing a job is a man.

Gendered differences in how job losses resonate within couples may stem from the diverging bargaining power of husbands and wives. Indeed, the assumptions in the health capital model formulated by Bolin et al. (2001) are related to how bargaining power within couples is changed by adverse life course events. However, gendered socialization processes, making women responsible for improving the welfare of others (Eagly and Wood 2016), may also contribute to the varying response to job losses.

While the gendered patterns are robust when we condition on economic resources, it is striking that the evidence for the buffering role of these resources can be observed when we consider accumulative resources such as wealth. But the results are less clear when we consider employment status or health insurance. Wealth is built up during years of working, and, among older couples, wealth may be a more relevant measure of economic wellbeing than contemporaneous employment status or income (Semyonov et al. 2013). The lack of clear evidence supporting the hypothesis of the buffering effects of employment status may also be related to a strong variation in the level of wages as well as working conditions in the United States (Kalleberg et al. 2000).

Our analysis has a number of limitations. First, it is restricted by the low incidence of involuntary job loss for individuals in our sample. However, the fact that we are able to identify the effects despite limited statistical power means that these effects could be even stronger in larger datasets. Second, we do not have full knowledge about the exact time when the job loss, or the onset of depressive symptoms, took place, nor can we examine the lagged or anticipated effects of job losses. In that way we cannot control for habit formation or adaptability, the possibility that the effect of the job loss becomes smaller as time passes. However, our analysis assures that the treatment, i.e., the involuntary job loss, took place before our measures of depressive symptoms were taken. Finally, while our research design, focused on involuntary job losses, reduces the problem of reverse causality, the internal validity comes at the cost of external validity. Plant closures may be more likely in some specific industries, occupations, or geographic areas, restricting generalizability to the overall workforce (Brand 2015). Considering the well-established patterns of marital homogamy (Lichter et al. 1995), this also means that spouses of these individuals are not a random draw of all the spouses of workers who lost their jobs. While data limitations prevent us from more detailed analysis across industries, occupations, and other characteristics that may give result to different responses to spousal job losses, we address this issue by estimating heterogeneous treatment effects across different population strata with varying levels of propensity scores. We find that the effects of job losses on depressive symptoms remain robust. Finally, while theoretical models proposed in Bolin et al. (2001) suggest that divorce constitutes a threat in couples' bargaining processes, our analyses do not follow couples after the divorce, and we focus instead on the effects of job losses on spousal health among couples that stayed together. From the empirical perspective, the divorce can be seen as a (probably non-random) dropout. Given the recent developments in methods for handling non-ignorable dropout in longitudinal data on health (Josefsson and Daniels 2019; Josefsson et al. 2016), future research could benefit from analyses that examine partnership dynamics and mental health within couples.

Despite these limitations, our results contribute to the recently emerging field of research on contagion of ill health within families (Umberson and Thomeer 2020). We also add to research on the interplay of employment and health within older couples. Previous studies have shown that men and women are differently exposed to stress over their lives, and they respond to individual experiences differently (Hagestad and Dykstra 2016; Simon 2014). There is also a growing understanding that men's and women's responses to events affecting family members also vary. For example, an acute health shock experienced by a spouse has a strong negative effect on women's employment, when women give up paid work to provide care, but men do not make similar adjustments when female partners experience a health shock (Vega 2019). Our paper adds to the evidence of gender differences in responses to adverse events by showing that a male spouses' job loss affects women's mental health, but the reverse does not hold when female spouses lose their jobs. Our results corroborate earlier findings of better mental health among men whose wives don't work, suggesting stronger preferences for joint leisure in this population group (Szinovacz and Davey 2004).

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Appendix

Table A1. Treatment effects by age group

	Age Group			
	50-65	50-55	56-60	61-65
Female				
MDM				
	0.742***	1.234***	0.465*	0.926***
	(0.189)	(0.411)	(0.247)	(0.333)
PSM				
	0.442*	1.318***	0.317	0.842**
	(0.199)	(0.431)	(0.251)	(0.364)
Obs	10,338	4,951	3,875	1,512
Male				
MDM				
	-0.186	-0.079	-0.510***	0.046
	(0.138)	(0.352)	(0.131)	(0.227)
PSM				
	-0.114	-0.313	-0.769***	-0.184
	(0.152)	(0.349)	(0.189)	(0.284)
Obs	9,598	2,478	4,052	3,068

*p<0.1,**p<0.05,***p<0.01. Standard errors in parenthesis. The dependent variable is the spouse's CES-D.

Table A2. Treatment Effects by health shock

	Health Shock	
	No	Yes
Female		
MDM		
	0.816***	0.719*
	(0.215)	(0.415)
PSM		
	0.630***	0.228
	(0.228)	(0.444)
Obs	8,306	1,839
Male		
MDM		
	-0.343***	0.129
	(0.108)	(0.414)

PSM		
	-0.343***	0.213
	(0.122)	(0.411)
Obs	7,491	1,983

*p<0.1,**p<0.05,***p<0.01. Standard errors in parenthesis. The dependent variable is the spouse's CES-D. Health shock equals yes if the spouse has experienced one or more of the following between waves in the survey: heart attack, stroke, or a new cancer diagnosis.

Table A3. Treatment effects

	Employment status		Wealth level		Health insurance	
	Unemployed	Employed	Low	High	No	Yes
Female						
MDM						
	0.946***	0.698***	0.995***	0.416**	0.667	0.495**
	(0.353)	(0.197)	(0.288)	(0.211)	(0.421)	(0.207)
PSM						
	0.867**	0.505**	0.859***	0.271	0.896*	0.346
	(0.380)	(0.218)	(0.316)	(0.229)	(0.457)	(0.216)
Obs	3,332	6,997	4,386	5,952	800	8,570
Male						
MDM						
	-0.671**	-0.086	-0.449**	-0.141	-0.276	-0.282**
	(0.270)	(0.158)	(0.179)	(0.155)	(0.343)	(0.135)
PSM						
	-0.431*	-0.116	-0.325	-0.117	-0.215	-0.371**
	(0.241)	(0.167)	(0.204)	(0.199)	(0.400)	(0.155)
Obs	2,224	7,360	3,766	5,832	649	8,160

*p<0.1,**p<0.05,***p<0.01. Standard errors in parenthesis. The dependent variable is the spouse's CES-D. ATET stands for average treatment effects on the treated. We classify households as low wealth if their wealth level is below the sample's median household wealth (USD 441,378.90), and otherwise as high-wealth households. Household wealth includes both financial and non-financial assets. Health insurance equals yes if the spouse has access to either a government or private insurance plan. The government health insurance plans incorporate coverage by any of the following federal government health insurance programs: Medicare, Medicaid, VA/CHAMPUS, CHAMP-VA, or any other government- provided health insurance.

Figures

FIGURE 1. FEMALE SPOUSE (50-65)

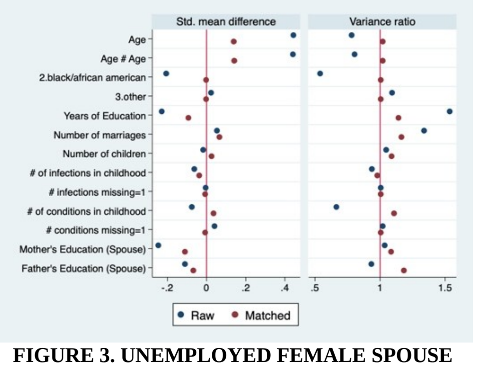


FIGURE 2. MALE SPOUSE (50-65)

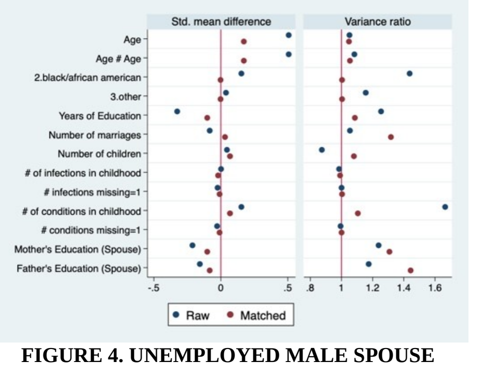


FIGURE 3. UNEMPLOYED FEMALE SPOUSE

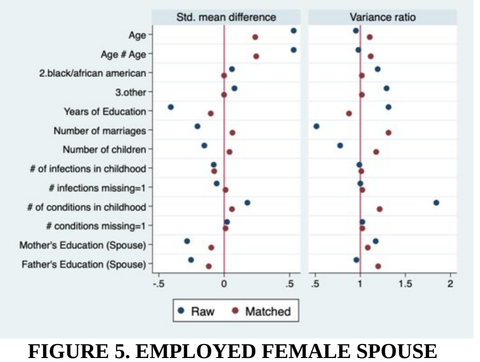


FIGURE 4. UNEMPLOYED MALE SPOUSE

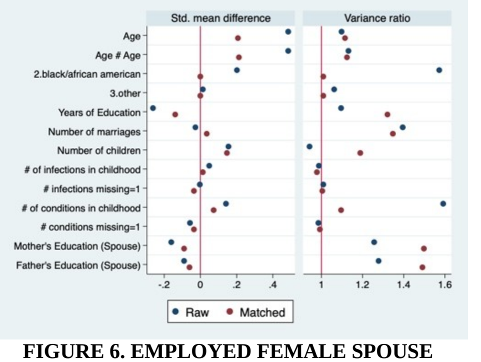


FIGURE 5. EMPLOYED FEMALE SPOUSE

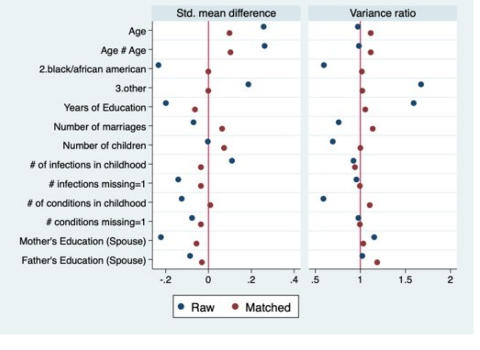


FIGURE 6. EMPLOYED FEMALE SPOUSE

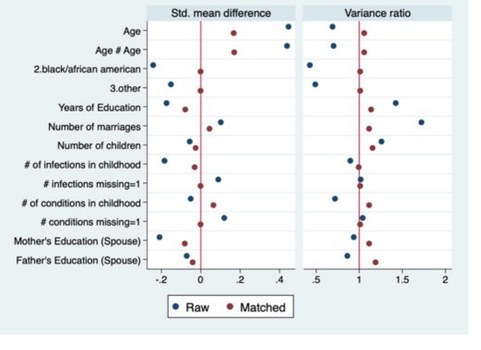


FIGURE 7. FEMALE SPOUSE IN GOOD HEALTH

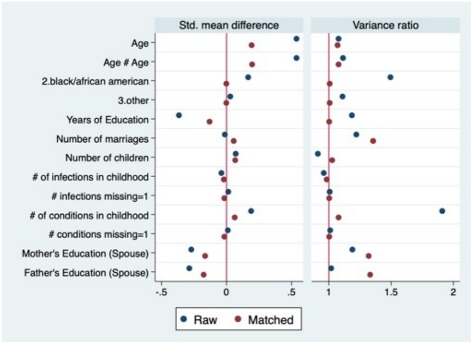


FIGURE 8. MALE SPOUSE IN GOOD HEALTH

FIGURE 9. FEMALE SPOUSE IN BAD HEALTH

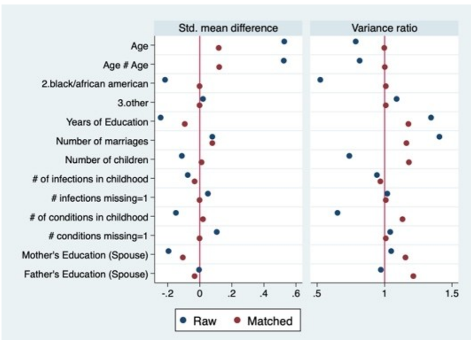


FIGURE 10. MALE SPOUSE IN BAD HEALTH

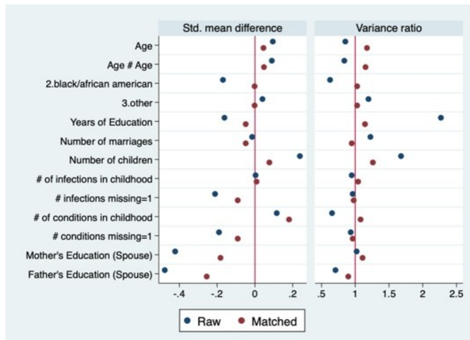


FIGURE 11. FEMALE SPOUSE DOES NOT HAVE MEDICAL INSURANCE

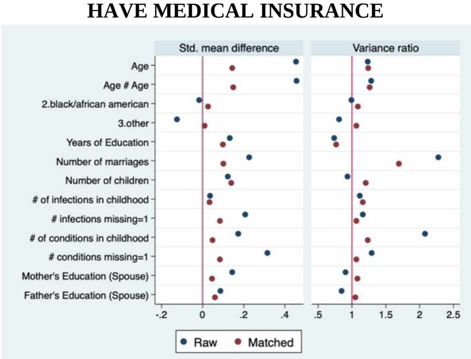


FIGURE 12. MALE SPOUSE DOES NOT HAVE MEDICAL INSURANCE

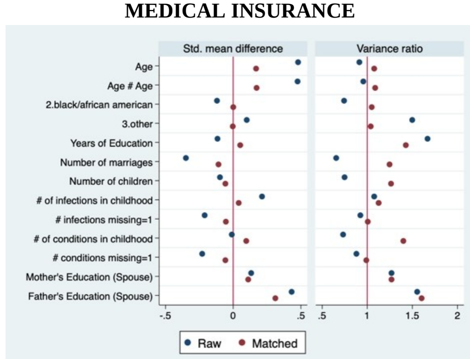
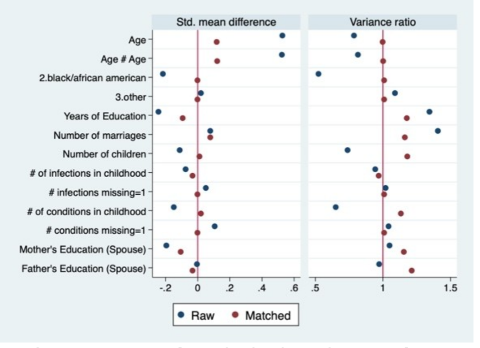


FIGURE 13. FEMALE SPOUSE HAS MEDICAL INSURANCE

FIGURE 14. MALE SPOUSE HAS MEDICAL INSURANCE



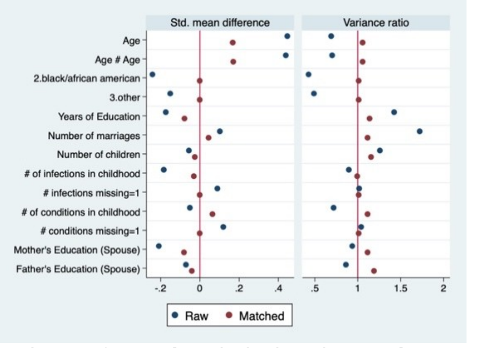
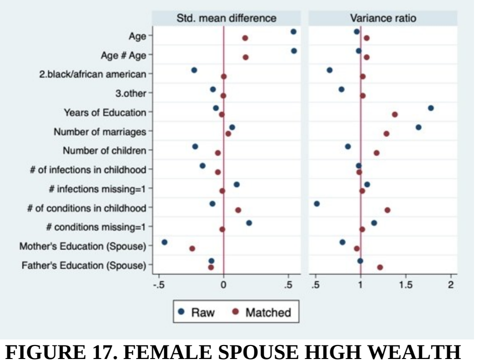


FIGURE 15. FEMALE SPOUSE LOW WEALTH LEVEL

FIGURE 16. FEMALE SPOUSE LOW WEALTH LEVEL



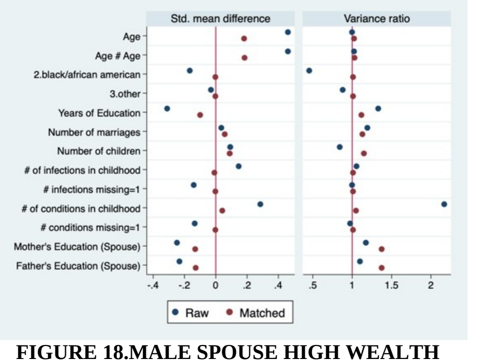
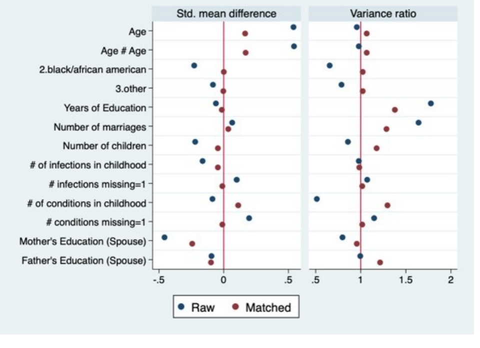
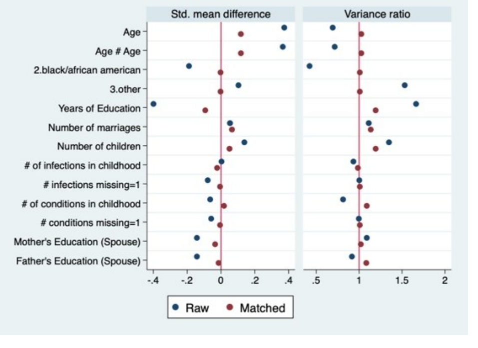


FIGURE 17. FEMALE SPOUSE HIGH WEALTH LEVEL

FIGURE 18.MALE SPOUSE HIGH WEALTH LEVEL





4

The Labor Supply of Elderly Mexican Women

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Abstract

This paper estimates the labor supply response to an increase in the marginal wage rate among middle-aged to elderly Mexican women. Using data from the National Survey of Occupation and Employment, I find that an increase in the marginal wage rate is associated with an increase in worked hours. The results suggest that the marginal wage rate elasticities are greater for older women than for their younger counterparts.

JEL code: I10 J01

Keywords: labor supply, marginal wage, women, Mexico.

1 Introduction

The purpose of this study is to estimate the labor supply response among Mexican middle-aged to elderly women to changes in the marginal wage rate. This is of particular interest in the context of an ageing population, lower fertility rates, and increased life expectancy. In Mexico, the share of women aged 50–70 increased from 13% in 2010 to approximately 17% in 2018. This increasing trend is expected to continue in coming years, and the share of women in this age group is expected to be close to 24% t by 2050 (CONAPO, 2017). Furthermore, the population aged ≥65 represented 7.2% of the total population in 2018 and is predicted to increase up to 10.3% by 2030 and to 16.8% by 2050 (CONAPO, 2017).

Some stylized facts are of importance when analyzing the labor supply of Mexican women. First, like in other Latin American countries, there has been an increase in female labor force participation in Mexico since the late 1960s, where a relevant factor is a strong negative correlation between fertility rates and labor force participation (Gasparini and Marchionni 2015). The birth rate in Mexico decreased from 20.2 births per thousand people (btp) in 2010 to 17.1btp in 2017, while women's participation in the labor force increased from 55% to 61% in the same period. The decline in btp is expected to continue and to fall to 11.3btp by 2050[2]. However, despite increased labor force participation, there has been virtually no changes to weekly hours worked; Mexican women between 50 and 70 years old have worked approximately 21 hours per week on average between 2004 and 2017.[3] In 2017, female life expectancy in Mexico was 75 years and is expected to increase to 83 years by 2050 (CONAPO, 2017).

These facts are likely to restrain social security spending in the future, since a larger share of elderly persons will require increased expenditures on pension benefits or transfers from the government to those who are not eligible for a pension. For example, in 2017, pension expenditure represented 3.12% of Mexico's GDP and if the current pension system prevails, pension expenditure is expected to increase to 5.2% of the GDP by 2030 (Villareal and Macias,

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[1] I would like to thank Thomas Aronsson, Gauthier Lanot, David Granlund, Johan Holmberg and Magnus Wikström for valuable comments.

[2] https://www.macrotrends.net/countries/MEX/mexico/birth-rate

[3] This average reflects both labor force participants and non-participants. When only participants are considered, the average number of weekly hours worked is approximately 36.

2020). Vogel et al. (2017) suggest that labor market policies that focus on the extensive margin by increasing the retirement age imply strong welfare gains for all future generations.

In this paper, I analyze how middle-aged to elderly women change their participation in the workforce in response to an increase in the marginal (net-of-tax) wage rate. It is important to understand how the labor supply of women at older ages reacts to economic incentives that could eventually make them increase the time they spend in the labor market after they are eligible for retirement. If women have incentives to join or remain in the labor force, and if those who are already working would work more hours later in life, tax collection could increase and thus alleviate some of the restraints on social security. In the data used below, approximately 39% of women aged 50–70 do not participate in the labor market, suggesting a substantial labor supply reserve.

While a large number of studies have examined the labor supply behavior of males (Pencavel 2002, Devereux 2004) and women (Eissa and Hoynes 2004, Bargain 2005, Blau and Kahn 2007), these studies mostly focus on the prime-age to pre-retirement age population in the US (Blau and Kahn 2007) and in Europe (Blomquist 1996, van Soest et al. 2002).[4] Attanasio et al. (2018) studied the labor supply of women aged 25 to 60 in the US and showed that there is substantial heterogeneity in their labor supply elasticities. The study also shows that the response to wage changes declines with age, and that women with the lowest wage levels have larger wage elasticities.

To my knowledge, there are only a few studies addressing the labor supply behavior of women in Mexico. Gómez and Campos-Vázquez (2010) estimated the wage elasticity of the labor supply of married women who are between 25 and 55 years old using 1990 and 2000 census data and found that women with children <5 years of age are less attached to the labor force than the average married woman. They also found that the labor supply elasticity ranged from 1.52–3.32 in 1990 and from 0.65–1.18 in 2000. The least amount of elasticity in 1990 was reported for women in the 45–54-year-old group, while in 2000, the elasticity for this group was the greatest. These elasticities are large when compared to corresponding elasticities for women aged 20–60 in the US. For example, Kuma and Liang (2005) show that for the year 1990, the elasticity was no larger than 0.6, while in 2000 they estimated an elasticity below 0.2.

Some studies focus on the labor supply of elderly Mexicans, although they do so without reporting or addressing the relationship between labor supply and wages. Juárez (2010) focuses on a so-called 'nutrition transfer' for individuals aged ≥70 years and finds that the transfer reduces the labor force participation of single men aged ≥70 by 46%. For women, the point estimate is negative, but not statistically significant. Avila-Parra and Escamilla-Guerrero (2017) found that expanding the pension program did not reduce the labor force participation of Mexican adults between 66 and 69 years of age.

The contribution of this paper is an estimation of labor supply elasticity with respect to the marginal wage rate for Mexican women aged 50–70, as well as examine how this elasticity varies between age-groups in this interval. To this end, I use the ENOE surveys from 2005 to 2017. First, I follow a two-step methodology (Blundell and MaCurdy 1999) to control for selection bias, and afterwards I use an instrumental variable tobit model to estimate the labor

[4]A common finding is that the wage elasticity of labor supply for women is larger than the corresponding elasticity for men, and that the latter is similar across countries (Evers et al. 2008). For extensive reviews of the literature, see Pencavel (1986), Blundell and MaCurdy (1999), and Evers et al. (2008).

supply of the whole sample of women and of women that participate in the labor force, respectively. The results suggest that cohorts of older women have larger marginal wage rate elasticities than cohorts of younger women. The reported marginal wage rate elasticities for women who participate in the labor force range from 0.02–0.10. The elasticities estimated for the entire sample range from 0.29–0.66. I also estimate a labor supply model that controls for random effects. The model is estimated for women participating in the labor force; the qualitative results do not differ from the results of the instrumental variable model tobit model. In this case, the elasticities are smaller and range from 0.01–0.08. The results are in line with studies for developed countries that find that older individuals have larger elasticities than their younger counterparts (Mroz 1987; French 2005; Keane and Rogerson 2015; Attanasio et al. 2018).

The rest of the paper is structured as follows. Section 2 discusses the data and section 3 outlines the empirical model. The results are presented in section 4. Finally, conclusions are drawn in section 5.

2 Data

The data were obtained from the National Occupation and Employment Survey (Encuesta Nacional de Ocupación y Empleo or ENOE). The ENOE has been collected quarterly since 2005 by the National Institute of Geographic Statistics and Information and is the consolidation and fusion of the National Urban Employment Survey and the National Employment Survey. As a result, the ENOE includes urban and rural observations. The ENOE sample includes 120,260 households. The survey contains demographic information for individuals aged ≥12 years, as well as detailed information on employment and unemployment status.

The design of the ENOE includes a five-quarter rotating panel of households. Each household is interviewed every 3 months for a period of 15 months, so there are 5 observations for each household (and consequently for all the individuals in said household). The design is such that in any given cross section of the ENOE, 20% of the households are in their first interview, 20% are in their second interview, etc. I use data for women aged 50–70 years old from the first quarter of 2005 to the last quarter of 2017.[5]

Although the ENOE reports variables that are relevant to the study of labor supply, it does not collect information about non-labor income. Non-labor income may affect the number of hours a person works, since increased non-labor income means that an individual can afford to work fewer hours. To control for (at least some of) the effects of non-labor income, I include women’s marital status and the number of children. These variables can act as proxies of non-labor income, since Mexican women are likely to receive financial support from their children and husband in the form of money transfers (Gomes, 2007).[6]

In this paper, I study both participants and non-participants in the labor market. Furthermore, I exclude women working in the agricultural sector, women who are self-employed, women who work in the informal sector, and women who have two jobs.[7] These groups are excluded

[5]This was the last quarter available for the data at the time of completing this study.

[6] I also estimated a specification of the labor supply for those aged 65–70 that includes a dummy variable indicating whether the individual is receiving pension-based income. I chose this age group since the retirement age is 65. The wage elasticities of the labor supply derived from this estimation were similar to those shown in Table 6 in section 4.3.

[7] From the ENOE total sample, approximately 34.5% of women work in the informal sector, 6.8% are self-employed, and 2.7% work in the agriculture sector. I do not address non-participants who work in the informal sector. One way to deal with the complications of work in the informal sector would be to consider that the individual has three choices when deciding to participate in the labor market: she can decide to participate in the formal sector, the informal sector, or not to participate

to avoid difficulties relating to measuring their hours and earnings, not to mention that people working in the informal sector do not pay taxes. The final sample is an unbalanced panel with 125,985 distinct women, where each woman is observed in different quarters within a year.[8] Women are observed on average 2.7 times, which gives a total of 242,066 observations in the sample. Among these, 136,533 participated in the labor force while 105,533 did not. Only a total of 37,310 women are observed over five quarters.

Table 1 shows the mean values and standard deviations of the variables used in this study. We observe that participants on average work 37 hours per week. The wages are presented in real terms, and we observe that the hourly gross wage rate is approximately MXN48 (USD2.37), while the hourly marginal wage rate net-of-tax is approximately MXN41 (USD2.02).[9] The marginal tax rate is on average 11%, but is smaller for older women who, on average, earn less. For example, the hourly gross wage rate for women aged 50–55 years is on average MXN49 (USD2.42), while for women aged 66–70 years, hourly gross wage averages MXN39 (USD1.91). The average hourly marginal wage rate is also lower for older women; it is 34MXN (1.68USD) for the oldest cohort, while it is approximately 42MXN (2.07USD) for the youngest. The data suggests, on average, older women work fewer hours and earn less, although differences between groups are not large. For example, if the youngest group is compared to the oldest group, the difference in working hours is approximately 5 hours per week, while the difference in the hourly marginal wage rate is approximately 8MXN (0.39USD).[10]

Compared to non-participants, those who participate in the labor force average more years in education, fewer children, and a smaller percentage of them are married. For example, when looking at the 50–70 group, participants have on average 3 more years of education and 2 fewer children than non-participants. Finally, in this same age group, the share of married women is approximately 70% for participants and 86% for non-participants.

Among those who participate, the share of married women as well as the hourly wage rates are smaller for the older cohorts.

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at all (Pradhan and van Soest, 1995). I exclude women in the agricultural sector, since their incentives to work may differ from those characterizing other sectors. For example, the ENOE shows that most households that depend on the agricultural sector own the land they farm and use it to procure their own food and to sell their harvest. Furthermore, the decision to work and the hours of work are dependent on the seasons; for example, more hours of work and more participation might be needed when preparing the land and during harvest, while less work might be needed in between these periods.

[8] The number of women is larger than the number of households because there are households with more than one generation of women. For example, take a grandmother, daughter, and granddaughter living in the same household.

[9] In the dataset we don't see the hourly wage. Instead, we see monthly earnings (E) and monthly worked hours (h). Both are measured with error. Suppose the true values for a person are (E, h), and the observed values are:

$$E^0 = E + e_1$$

$$h^0 = h + e_2$$

The hourly wage 'wage' is constructed as $w^0 = E^0/h^0$. Note that when e_2 is positive, hours of work are overestimated, and the observed wage is lower. Likewise, when e_2 is negative, observed hours are lower than they were, and the wage is higher. This is called the problem of 'division bias': there is negative correlation between hours and the hourly wage caused by measurement error in hours (Borjas, 1980). I do not test for division bias, since I do not find a negative correlation between hours of work and the marginal wage rate, but I encourage future research to consider this issue in more detail.

The marginal wage rate is defined as the gross wage rate times one minus the marginal tax rate. 'MXN' stands for Mexican Pesos; the average exchange rate in July 2018 was MXN20.22 per 1USD. The 2018 minimum wage in Mexico was MXN88.36 (USD4.36) per day. The wages are presented in 2018 CPI adjusted MXN (Mexican Pesos). Source: www.banxico.org.mx

[10] The data is used later to impute wages according to the methodology section, the description of the summary statistics of the imputed hourly wages are presented in the appendix.

Table 1: Summary Statistics.

Age Group	50–70	50–55	56–60	61–65	66–70
Participants					
Weekly Worked Hours	36.82	37.62	36.62	34.73	32.47
	(14.35)	(13.81)	(14.56)	(15.39)	(16.13)
Hourly Gross Wage Rate (w)	48.27	49.73	47.55	45.97	38.79
	(78.03)	(66.73)	(97.13)	(88.85)	(55.15)
Hourly Marginal Wage Rate (w^n)	40.78	41.88	40.24	38.94	33.73
	(54.61)	(47.45)	(66.81)	(61.98)	(40.08)
Marginal Tax Rate	11.37%	11.71%	11.06%	10.83%	10.23%
Age	55.56	52.41	57.68	62.60	67.56
	(4.57)	(1.61)	(1.39)	(1.39)	(1.41)
Education years	9.41	10.12	8.91	7.88	6.45
	(5.12)	(4.91)	(5.11)	(5.29)	(5.12)
Number of Children	2.88	2.71	3.01	3.21	3.50
	(2.13)	(1.85)	(2.23)	(2.65)	(3.16)
Married	69.81%	71.91%	69.04%	64.98%	58.01%
Observations	136,533	79,757	35,485	15,471	5,820
Non-Participants					
Age	55.23	52.59	57.88	62.89	67.81
	(5.66)	(1.62)	(1.42)	(1.43)	(1.42)
Education years	6.05	7.18	6.12	5.07	4.21
	(4.49)	(4.55)	(4.43)	(4.20)	(3.91)
Number of Children	4.41	3.78	4.34	4.94	5.54
	(2.99)	(2.46)	(2.86)	(3.27)	(3.64)
Married	86.12%	86.93%	86.82%	85.49%	83.47%
Observations	**105,533**	**40,929**	**28,096**	**21,374**	**15,134**

Standard deviations in parenthesis. The wages are presented in 2018 CPI adjusted MXN (Mexican Pesos). An observation is defined as a woman in a particular quarter at a particular year, and each woman is observed on average 2 times in the data set. There is a total of 242, 066 observations. The observations correspond to 125,985 women. Of the total number of observations, 25.22% are observed once, 23.42% twice, 19.06% three times, 17.36% four times, and 15.09% five times.

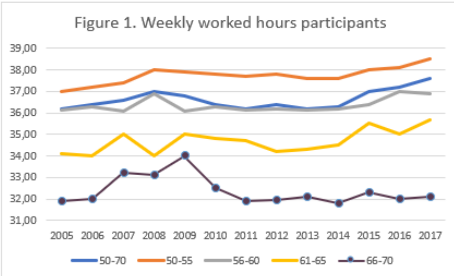

Figure 1. Weekly worked hours participants

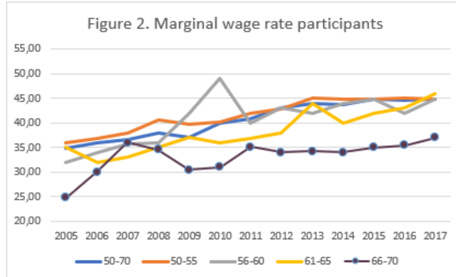

Figure 2. Marginal wage rate participants

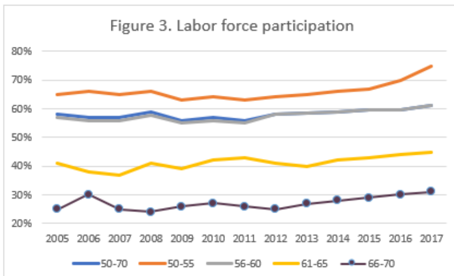

Figure 3. Labor force participation

Figures 1 and 2 show the weekly worked hours and marginal wage rates for participants. We can see that hours worked per week were relatively steady until 2014 for the age groups 50–55 and 56–60, as well for the 50–70 cohort; after this period, they were characterized by a slightly increasing trend. This may have been a result of small increases in the real minimum wage. The peak in the weekly marginal wage rate in 2010 for women aged 56–60 is explained by women with large earnings entering the survey in that year.

Figure 3 shows the share of women participating in the labor force from 2005–2017 divided into five age groups. For the first three groups (blue, gray, and orange lines) there was an increase in participation between 2013–2017 of approximately 6%, while for the two oldest cohorts (yellow and dark blue lines) the increase was approximately 4% and 7%, respectively. The increase in participation in recent years has been larger than the increase in weekly worked hours for those who participate. For example, in the period between 2013 and 2017, weekly worked hours increased by approximately 2.4%, 2.3%, and 0.16% for women aged 50–70, 56–60 and 66–70, respectively. As in other Latin American countries, more women have joined the labor force in recent years, but at the same time, weekly worked hours have only increased modestly (Gasparini and Marchionni, 2015). The weekly worked hours have increased at most 2 hours since 2005 in most of the cohorts, even when the real minimum wage has increased and after reforms were made to the tax system.

The income tax schedule had two major reforms during this period, in 2008 and again in 2014 (see Table 2). The first reform added three new brackets at the top of the income range while simultaneously reducing the taxable income range for the remaining brackets. For example, in 2007, the tax rate for those with a monthly labor income higher than MXN8601.5 (USD425.4) was 28%, while in 2008 it changed to 17.92%. The 2014 reform increased the top tax rate by creating 3 new brackets with a tax rate of 32%, 34%, and 35% for those earning more than MXN363.38 (USD17.97), MXN484.51 (USD23.96), and MXN1453.5 (USD71.88), respectively.[11]

[11] Taxation is defined in terms of yearly income; the monthly schedules of the Mexican tax code are calculated by dividing the yearly income by 12. For example, according to the tax code, an individual with yearly earnings of 120,000 MXN has a monthly income of 10,000 MXN and is assigned a 30% income tax.

Table 2. The Income Tax Schedule in Mexico 2005-2017*.

Monthly Earnings MXN lower limit	upper limit	2005	2006	2007	2008	2009	2010	2011	2012	2013	2014	2015	2016	2017
0.01	496,07	3.0%	3.0%	3.0%	1.9%	1.9%	1.9%	1.9%	1.9%	1.9%	1.9%	1.9%	1.9%	1.9%
496,08	4,206,41	10.0%	10.0%	10.0%	6.4%	6.4%	6.4%	644%	6.4%	6.4%	6.4%	6.4%	6.4%	6.4%
4,206,42	7 399,42	17.0%	17.0%	17.0%	10.9%	10.9%	10.9%	10.9%	10.9%	10.9%	10.9%	10.9%	10.9%	10.9%
7,399,43	8,601,5	25.0%	25.0%	25.0%	16.0%	16.0%	16.0%	16.0%	16.0%	16.0%	16.0%	16.0%	16.0%	16.0%
8,601,51	10,298,35	30.0%	29.0%	28.0%	17.9%	17.9%	17.9%	17.9%	17.9%	17.9%	17.9%	17.9%	17.9%	17.9%
10,298,36	20,770,29	30.0%	29.0%	28.0%	19.9%	19.9%	21.4%	21.4%	21.4%	21.4%	21.4%	21.4%	21.4%	21.4%
20,770,91	32,736,83	30.0%	29.0%	28.0%	22.0%	22.0%	23.5%	23.5%	23.5%	23.5%	23.5%	23.5%	23.5%	23.5%
32,736,84	62,500	30.0%	29.0%	28.0%	28.0%	28.0%	30.0%	30.0%	30.0%	30.0%	30.0%	30.0%	30.0%	30.0%
62,500,01	83,333,33	30.0%	29.0%	28.0%	28.0%	28.0%	30.0%	30.0%	30.0%	30.0%	32.0%	32.0%	32.0%	32.0%
83,333,34	250,000	30.0%	29.0%	28.0%	28.0%	28.0%	30.0%	30.0%	30.0%	30.0%	34.0%	34.0%	34.0%	34.0%
250,000,01	∞	30.0%	29.0%	28.0%	28.0%	28.0%	30.0%	30.0%	30.0%	30.0%	35.0%	35.0%	35.0%	35.0%

*Mexican income tax code. The official gazette of the Mexican Federation: https://www.dof.gob.mx/nota_detalle.php?codigo=5640505&fecha=12/01/2022. The table shows the marginal tax per year and income bracket.

3 Empirical Specification

Identifying the response of the labor supply to wages is not an easy task, since several issues are simultaneously at play. Selection and endogeneity issues, as well as a lack of data are among the obstacles in the empirical identification of the causal links between the female labor supply and its covariates.

The selection issue—which results from the decision of whether to work/participate in the labor market—implies that gross wage rates are unobserved for non-participants. If only the observations that report positive hours of work and positive wages were considered, the labor supply estimation would suffer from selection bias. Women with strong preferences toward work will tend to participate more in the labor market regardless of their own wage. To account for selection in the estimation, I follow a two-step methodology (Blundell and MaCurdy 1999). First, I estimate the participation equation, followed by the estimation of a gross wage equation. In these estimations I consider the panel structure of the data and use a random effects specification for the unobserved components. The labor force participation equation is estimated by a probit model. The estimates of the participation equation are then used to calculate the inverse Mills ratio (imr), which is included as an additional explanatory variable in the log wage estimation.[12]

Following Triest (1990), I exclude education from the labor supply equation, and include it in both the labor force and wage imputation estimations. I do not consider health, non-labor income, or ethnicity (Triest, 1990) in the estimations, because the ENOE does not report these variables. I also exclude higher-order terms for education or age as well as interaction terms between these, because adding them to the estimations produces multicollinearity problems. The variables used in each estimation are presented in more detail in the following paragraphs.

The participation equation includes variables describing women's preferences to participate in the labor force such as her age, education, number of children, and marital status. These variables are chosen because they are identified by economic theory as well as by earlier empirical studies as relevant determinants undergirding a woman's choice to participate in the labor force (Killingsworth and Heckman, 1986), and they are commonly included in the ENOE. The participation equation is as follows:

$$p_{it} = X_{it}\alpha_1 + Yr_t\alpha_2 + Q_t\alpha_3 + v_i + e_{it} \tag{1}$$

where p_{it} takes the value 1 if individual *i* participates in the labor force, and takes the value 0 otherwise. The subscript *i* refers to the individual, and *t*, for simplicity, to the survey's year and quarter. X_{it} is a vector of individual characteristics such as age, education, children, and marital status. Age is defined as the year of the survey minus the year of birth, while the education variable denotes the individual's years of schooling. The variable for children indicates the

[12] The inverse Mill's ratio is obtained from the participation equation (1); it represents the marginal probability of a woman being employed at a specific wage rate, given her individual characteristics and prevailing market conditions. The ratio reflects the idea that women with large negative wage errors are not working, so the expected value of the wage error is no longer zero for some of the women who do work.

number of children that are alive at the time of the survey, and marital status indicates whether the individual is married. *Yr* is a vector of year dummies and *Q* is a vector of quarter dummies which capture time-specific effects and indicates the time when the survey took place. Finally, v_i is the individual-specific random effect and is normally distributed with mean 0 and variance σ^2_v. The idiosyncratic component is denoted by e_{it} and is independent and identically distributed (iid) with mean 0 and variance 1, independent of v_i.[13]

The second step is to estimate a log-gross wage equation for participants which corrects for the selection effect using the results from the estimation of the participation equation. The results from estimating this wage equation are useful for imputing log gross wages to all individuals in the sample, whether or not they are participating. For simplicity's sake, I assume that the log-gross wage equation takes a standard human capital form, where the log gross wage depends on the individual's age and education (Blundell et al. 2001), as well as year and quarter dummies and the inverse mills ratio to control for selection.[14] The log gross wage equation is as follows:

$$logw_{it} = S_{it}\beta_1 + imr_{it}\beta_2 + Yr_t\beta_3 + Q_t\beta_4 + z_i + u_{it} \tag{2}$$

where w_{it} denotes the hourly wage rate for individual *i* in a specific year and quarter, S_{it} is a vector containing the variables age and education, and imr_{it} is the inverse Mills ratio. The equation is estimated as a random-effects specification, as mentioned before, where z_i is the individual-specific random effect and is normally distributed with mean 0 and variance σ^2_z, u_{it} is the idiosyncratic effect, and is iid with mean 0 and variance 1, independently of z_i. The log gross wage equation is estimated for those who work, and log-gross wages are imputed for all individuals in the dataset. The marginal tax rate is identified based on the earnings and the log-marginal wage rate is calculated as follows[15];

$$logw^n_{it} = logw_{it} + \log(1 - \tau_{it}) \tag{3}$$

where w^n is the marginal net wage rate, $log\ w$ is the imputed log gross wage rate, and τ is the marginal tax rate. The imputed wages are presented and discussed in section 4.

The second issue is the endogeneity of the marginal wage rate. To address this problem and include non-participants, I estimated two regressions using an instrumental variables tobit model[16], according to equation (4) for each age group. The first model uses the log-marginal wage rate if an individual worked the mean weekly hours as the instrument for the log-marginal rate, and the second model uses the log-marginal wage rate if an individual worked 40 hours per week (which corresponds to a full-time working week) as instruments.[17] These instruments are correlated with the marginal wage rate, but are not affected by the

[13] The results for the participation equation are presented in Table A1 in the appendix.

[14] The gross wage equation does not include working experience, since this variable is not reported by the ENOE.

[15] The marginal tax rate for non-participants is estimated considering the weekly ISR table published by the federation. It is assumed that if the non-participating individual were to enter the labor market, she would work 5 days per week and 8 hours per day, which corresponds to a full-time workday. This way, the weekly income is divided by 40 hours to obtain an approximation of the ISR table per hour. With this table, it is possible to assign the marginal tax rate to non-participants. Similar results were obtained when the ISR table per hour was estimated considering a 7-day week, as well as 5-hour workdays.

[16] For a description of the tobit model see Bierens (2004), and for the instrumental variables tobit model, see Newey (1987), Miranda and Rabe-Hesketh (2006), and Finlay and Magnusson (2009).

[17] The gross wage rate is calculated by dividing weekly earnings by the number of weekly worked hours; 40 hours is considered as a full-time working week according to Mexican law, or, alternatively 8 working hours per day.

individual's choice of hours of work.[18] Both models allow for intragroup correlation (Papke and Wooldrige 2008) to account for the panel structure of the data. A semi-log specification that allows for zero hours of work is adopted to estimate the labor supply defined as weekly hours worked h_i^* (Blundell et al. 2001),[19]

$$h_i^* = \begin{cases} h_i & if h_i > 0 \\ 0 & otherwise \end{cases}, \tag{4}$$

where

$$h_i = C_i\gamma_1 + logw_i^n\gamma_2 + Yr_i\gamma_3 + Q_i\gamma_4 + b_i \tag{5}$$

Here C_i is a vector containing the variables age, children, and marital status and b_i is assumed to be normally distributed with mean 0 and variance.

Even if equation (5) considers the panel structure of the data, it does not estimate a random effects model. Due to the difficulty of estimating a censored model with panel data and endogeneity (Honoré et al., 2004) an alternative is to estimate the parameters of a random effects model using the sub-sample of individuals who report positive hours. In this case, the labor supply equation has the following form,

$$h_{it} = C_{it}\eta + logw_{it}^o\gamma + Yr_i\psi + Q_i\varsigma + imr_{it}\lambda + f_i + g_{it}, \tag{6}$$

The logarithm of the observed marginal wage rate is denoted by $logw_{it}^o$ and is instrumented via the log-marginal wage rate if an individual worked the mean weekly hours, while the components f_i and g_{it} capture random specific and idiosyncratic effects.

[18] To test the exclusion restriction (which states that the instrument should not directly affect the dependent variable) I include one of the two instruments in the main regression (the log-marginal wage rate if an individual worked the mean weekly hours) while using the remaining regression as the instrument (the log-marginal wage rate if an individual worked 40 hours) and vice versa. If the "instrument" (now as a covariate in the labor supply regression) is not statistically significant in the labor supply regression, it supports to some degree the notion that it does not have a direct effect on the hours of work (Bound et al. 1995, Staiger and Stock 1997, Stock and Yogo 2005). The latter must be interpreted with caution since there could be other factors such as measurement error or sample size limitations that prevent detecting a small direct effect. The results of these regressions showed that none of the instruments were statistically significant when added as covariates in the labor supply estimation (i.e., this supports the exclusion of the instruments from the substantive equation), while the estimated effects of the marginal wage rate—the endogenous regressor—were similar to those presented below (i.e., this suggests that there is little evidence of overidentification).

[19] In this case an observation "i" is an individual in a particular year and quarter; i.e., the same individual can be present in the data more than one time. The standard errors are clustered by the individual's identification number to allow for intragroup correlation.

4 Results
4.1 Imputed wages

Table 3 shows the mean of the imputed log gross wages obtained from equation (2) as well as the log marginal wage rates calculated for participants and the whole sample, respectively. (See the Appendix for the results of the estimation of equation (2)). Table 3 shows that both log gross and log marginal wage rates are larger for younger cohorts. The imputed log gross wages for women that participate in the labor force are on average 1.9% larger than their observed log gross wages. The Table also shows that log-imputed wages for participants are approximately 6% larger than those for non-participant women.[20]

Table 3. Summary Statistics: mean imputed gross and marginal wage rates.

Age Group	50–70	50–55	56–60	61–65	66–70
All women					
Imputed log w	3.86	3.91	3.84	3.78	3.61
	(3.25)	(3.28)	(3.27)	(3.20)	(2.97)
Imputed log w^n	3.49	3.54	3.47	3.41	3.23
	(2.86)	(2.89)	(2.88)	(2.81)	(2.7)
Marginal tax rate	27.67%	29.85%	28.39%	27.57%	28.15%
Observations	242,066	120,686	63,581	36,845	20,954
Participants					
Observed log w	3.78	3.81	3.76	3.73	3.56
	(4.13)	(3.97)	(4.35)	(4.26)	(3.78)
Observed log w^n	3.62	3.65	3.61	3.58	3.43
	(3.79)	(3.65)	(4.00)	(3.92)	(3.49)
Imputed log w	3.99	4.01	3.97	3.94	3.79
	(3.35)	(3.35)	(3.37)	(3.34)	(3.16)
Imputed log w^n	3.62	3.65	3.61	3.58	3.42
	(3.93)	(4.11)	(3.90)	(4.31)	(4.16)
Marginal tax rate (imputed)	26.58%	27.88%	27.45%	26.43%	25.47%
Observations	136,533	79,757	35,485	15,471	5,820
Non-Participants					
Imputed log w	3.74	3.78	3.74	3.71	3.59
	(3.11)	(3.14)	(3.14)	(3.08)	(2.89)
Imputed log w^n	3.50	3.53	3.48	3.46	3.30
	(2.71)	(2.74)	(2.74)	(2.68)	(2.49)
Marginal tax rate	21.25%	23.87%	20.89%	20.15%	19.31%
Observations	105,533	40,929	28,096	21,374	15,134

Standard deviations in parenthesis. The table shows the mean of the imputed gross wage rates and marginal wage rates by age group in 2018 MXN (Mexican Pesos). The log-imputed wages are first estimated according to equation (2) and then marginal wage rates are calculated by applying the exponential to the log-imputed wages.

4.2 Estimates

Table 4 presents the results of the estimation of equations (4) and (5) by age group for participants and for all individuals, respectively, when the instrument is the imputed log-marginal wage rate corresponding to the mean weekly worked hours.[21] The first stage results are presented in Table A3 in the appendix and show that the instruments are significantly correlated with the log-marginal wage rate. For participants in the labor market, the estimated parameter for the logarithm of the marginal wage rate is statistically significant and positive

[20] This is the average of the differences between participants and non-participants age groups.

[21] The results for the instrument estimated at 40 hours are presented in the Appendix. I do not present these results in this section since they were similar for the two instruments. This is due to the mean hours being close to 40 hours.

for all groups, except the 50–55 group. Hence, the results suggest that an increase in the marginal wage rate increases the number of weekly worked hours. The estimates are larger for older groups. Furthermore, the estimates are positive and statistically significant for all groups when non-participants are included in the regressions. The estimates are larger for older cohorts, which suggests that the labor supply response to the marginal wage rate is larger for older cohorts, even when non-participants are included. In addition, the estimates are much larger compared to those for working women.

The rest of Table 4 shows that the estimated effects of age are statistically significant and negative for all age groups in both subsamples. The latter suggests that women spend less time in the labor market as they get older, even within small age intervals. The estimated effects of the number of children suggest that women work fewer hours per week the more children they have. The estimated effects of children are also larger when all individuals are considered in the estimations when compared to the results of participants. We can also see a negative correlation between being married and weekly worked hours. The correlation suggests that participants that are married work less than their single counterparts. For example, a married woman in the 56–60 group works approximately 2.6 hours less per week than their unmarried counterparts. This difference becomes even larger when all women are included in the estimation, as a married female in the 56–60 group works approximately 10 hours less per week than a single female.

Mexican families, especially families with members in these age intervals, often have a traditional structure, where gender roles are well-defined and adult children support their parents (Gomes, 2007). Women spend more time in household production activities and may receive non-labor income from their husband and adult children. The larger the number of children, the larger economic and social support she may receive from them at older ages, thus preferring to spend less time in the labor market. It is also expected that married women spend less time at the labor market than their single counterparts (Schmitz and Diefenthaler 1998).

Table 4: Estimated Parameters of the Labor Supply model (equation 4).

Age Group	50–70	50–55	56–60	61–65	66–70
Participants					
$logw_i^n$	0.791***	–0.099	1.178***	1.919***	3.254***
	(0.058)	(0.081)	(0.116)	(0.179)	(0.384)
Age	–0.269***	–0.116***	–0.232***	–0.467***	–0.353**
	(0.008)	(0.031)	(0.054)	(0.087)	(0.148)
Children	–0.409***	–0.432***	–0.405***	–0.391***	–0.331***
	(0.021)	(0.029)	(0.039)	(0.053)	(0.079)
Married	–2.431***	–2.586***	–2.418***	–2.188***	–1.791***
	(0.094)	(0.121)	(0.187)	(0.296)	(0.499)
Constant	51.746***	46.701***	49.076***	60.096***	48.451***
	(0.543)	(1.618)	(3.193)	(5.568)	(10.216)
Prob > Chi2	0.0000	0.0000	0.0000	0.0000	0.0000
Observations	136,533	79,757	35,485	15,471	5,820
All women					
$logw_i^n$	8.067***	7.907***	8.741***	10.374***	13.779***
	(0.103)	(0.142)	(0.215)	(0.343)	(0.731)
Age	–1.559***	–0.752***	–1.578***	–2.753***	–2.639**
	(0.014)	(0.053)	(0.098)	(0.158)	(0.264)
Children	–2.455***	–2.508***	–2.402***	–2.472***	–2.099***
	(0.032)	(0.047)	(0.064)	(0.089)	(0.133)
Married	–12.501***	–10.377***	–13.896***	–18.243***	–22.476***
	(0.177)	(0.225)	(0.361)	(0.579)	(0.959)
Constant	93.011***	48.981***	95.418***	161.741***	142.584***
	(0.911)	(2.871)	(5.775)	(10.014)	(18.156)
Observations	**242,066**	**120,686**	**63,581**	**36,845**	**20,954**

*p<0.1, **p<0.05, ***p<0.01. Standard errors in parentheses. All specifications include year and quarter controls. The table shows the estimation of the iv tobit model in Stata. The Wald test was conducted in tobit and not iv tobit regressions to see whether the presence of endogeneity could be rejected. The Wald test examines whether the error terms in the structural equation and the reduced-form equation for the endogenous explanatory variable are correlated (Wooldridge, 2002). If the test is statistically significant, the null hypothesis of exogeneity is rejected. The Wald tests for most regressions were statistically significant, implying the presence of endogeneity.

Table 5 shows the results of estimating equation (6). We can see that the estimated effects of the logarithm of the observed marginal wage rate are positive and statistically significant for all age groups. As before, the estimates are smaller for the younger age groups. In contrast to the results of equation (5), the estimates are smaller, except for the 50–55 age group, and the estimate for the 50–55 age group is statistically significant in this case. The fraction of variance due to f_i (rho) is different from zero and indicates that the panel-variance component is important, and that this estimator differs from the pooled estimator. Although the qualitative results of both equations for participants are similar, the magnitudes of the estimates differ depending on the choice of model.

Table 5: Estimated Parameters of the Labor Supply model (equation 6).

Age Group	50–70	50–55	56–60	61–65	66–70
$logw_{it}^{o}$	0.593***	0.562**	0.857***	1.253***	2.816***
	(0.063)	(0.095)	(0.122)	(0.201)	(0.391)
Age	–0.119***	–0.096**	–0.174***	–0.375***	–0.253***
	(0.011)	(0.062)	(0.074)	(0.097)	(0.156)
Children	–0.382***	–0.394***	–0.377***	–0.325***	–0.297***
	(0.032)	(0.042)	(0.067)	(0.078)	(0.082)
Married	–1.897***	–1.911***	–2.165***	–1.754***	–1.236***
	(0.097)	(0.168)	(0.192)	(0.321)	(0.501)
Inverse Mills Ratio	0.054***	0.125***	0.037***	0.009***	0.924***
	(0.007)	(0.007)	(0.007)	(0.005)	(0.015)
Constant	36.125***	26.245***	35.682***	45.342***	31.928***
	(0.452)	(1.437)	(2.911)	(4.857)	(9.873)
Prob>Chi2	0.0000	0.0000	0.0000	0.0000	0.0000
R^2	0.328	0.365	0.287	0.301	0.292
Sigma_f	0.235	0.168	0.193	0.211	0.257
Sigma_g	0.279	0.232	0.365	0.382	0.395
rho	0.415	0.344	0.219	0.234	0.297
Observations	136,533	79,757	35,485	15,471	5,820

*p<0.1, **p<0.05, ***p<0.01. Standard errors in parentheses. All specifications include year and quarter controls. The table shows the estimates of a random effects regression with endogenous covariates in Stata (xtivreg). Rho is the fraction of variance due to f_i, and is different from zero, suggesting that the panel-variance component is important, and that the estimator is different from the pooled estimator.

4.3 Elasticities

Table 6 shows the marginal wage elasticity by age group for participants and all individuals, respectively. The elasticity derived from equation 5 for participants aged 50–70 shows that a 1% increase in the marginal wage rate is associated with a 0.021% increase in weekly hours worked. The table shows that the elasticities are smaller for younger women compared to their older counterparts.[22] Thus, the elasticity is larger for women close to retirement, but it is even larger for those in the post-retirement age group. The elasticities for all women (participants and non-participants) are larger, and suggest that the increase in working hours varies from 0.34% to 0.65% when the marginal wage rate increases by 1%.[23] The elasticities for participants derived from equation (6) show similar properties, although they are smaller than those from equation (5) (except for the 50–55 group), and the elasticity for the 50–55 group is statistically significant. These results suggest that older females react more to changes in the marginal wage rate than their younger counterparts.

A possible interpretation of the results, in line with earlier research (Eissa and Liebman, 1996; Meyer and Rosenbaum, 2001; Blundell, 2001), is that the labor supply response at the extensive margin (labor force participation) is more sensitive than the intensive margin (hours of work.

[22] The t-test showed that these differences were statistically significant.

[23] It is possible for women to belong to two different age groups in the data—in some cases, a woman that is 55 years old at the beginning of the survey will be 56 by the end of it. A total of 5,467 women—4.3% of all women in the sample—belong to two different groups. The results presented here includes women belonging to two different age groups. I also estimated two more specifications. In the first, I placed the intersecting observations in the lower age-group while in the second specification these were placed in the upper age-group. The elasticity results of these two specifications did not differ extensively from the baseline specification.

Table 6: Marginal Wage Elasticities.

Age Group	50–70	50–55	56–60	61–65	66–70
Participants (equation 5)	0.021***	-0.002	0.032***	0.055**	0.101***
	(0.001)	(0.002)	(0.003)	(0.005)	(0.012)
All women (equation 5)	0.342***	0.286***	0.351***	0.414***	0.656***
	(0.001)	(0.001)	(0.002)	(0.003)	(0.007)
Participants (equation 6)	0.014***	0.012***	0.021***	0.037***	0.082***
	(0.002)	(0.001)	(0.004)	(0.001)	(0.005)

*p<0.1, **p<0.05, ***p<0.01. Standard errors in parentheses. The first row shows the marginal wage rate elasticity for women that work while the second row shows the elasticity for working and non-working women.

5 Conclusions

Earlier studies on the labor supply of middle-aged to elderly women have been mostly focused on the US and on European countries and to a much lesser extent on Latin American countries. Hence, the purpose of this study is to analyze the labor supply response of this population to changes in the marginal wage rate in Mexico. The results suggest that older women in these age cohorts react more to a change in the marginal wage, which is in line with studies based in developed countries (see Mroz 1987; French 2005), suggesting that older individuals have larger elasticities than their younger counterparts. Furthermore, elasticities seem to decrease when compared to the elasticities of prime-aged women in Mexico. Gómez and Campos-Vázquez (2010) report labor supply elasticities for Mexican women in their prime years that range between 0.65 and 1.52. The elasticities reported in the present study are smaller and assume values between 0.34 and 0.65.

I also find that the wage elasticity of labor supply is larger for older women than for middle-aged women. The results also show that the choice of model matters for the magnitude of the elasticities measured among the participants, although the qualitative results are quite similar. In addition, the elasticities suggest that the participation decision is more sensitive to changes in the marginal wage rate than the hours of work among the participants.[24] The latter is interesting in the sense that this part of the population can be seen as a labor supply reserve, which can contribute to alleviating some of the pressure on the social welfare system in Mexico. Policy-makers could tailor policies to increase the marginal wage rate of older women to incentivize non-participants to join the labor market. Eventually, tax collection could increase and assist Mexican social welfare in depending less on government debt.

The results in this study should be read with caution due to the limitations of the data. First, the data does not report non-labor income, and the severity of this omitted variable problem is difficult to assess. I have tried to reduce this problem by including variables that are likely to correlate with non-labor income such as age, marital status, and the number of children. The data also lacks information on work experience, which is potentially useful in estimating the wage equation. In the future, richer datasets may be available to solve these issues.

[24] This interpretation suggests that changes in the marginal wage rate could increase the extensive margin of the labor supply and incentivize women to work past the age of retirement. It is important to note that this interpretation comes from a static labor supply model; however, a dynamic model could also suggest that favorable marginal tax treatment of older workers delays retirement and increases life-cycle labor supply (Gustafsson, 2021).

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Appendix

Table A1: Labor Force Participation.

Age Group	50–70	50–55	56–60	61–65	66–70
Age	–0.188***	–0.109***	–0.211***	–0.298***	–0.222***
	(0.008)	(0.014)	(0.031)	(0.086)	(0.049)
Education years	0.286***	0.288***	0.301***	0.332***	0.175***
	(0.014)	(0.019)	(0.015)	(0.106)	(0.054)
Children	–0.181***	–0.215***	–0.198***	–0.207***	–0.103***
	(0.011)	(0.027)	(0.029)	(0.081)	(0.011)
Married	–2.133***	–1.856***	–2.601***	–7.954***	–2.149*
	(0.104)	(0.097)	(0.303)	(2.816)	(1.204)
Constant	10.767***	6.498***	12.631***	19.163***	12.246***
	(0.485)	(0.789)	(1.783)	(5.461)	(4.257)
Prob > Chi[2]	0.0000	0.0000	0.0000	0.0362	0.0000
Sigma_u	1,426	1,264	1,624	1,795	1,821
Rho	0,671	0,615	0,725	0,763	0,768
Observations	242,066	120,686	63,581	36,845	20,954

*p<0.1, **p<0.05, ***p<0.01. Standard errors in parentheses. All specifications include year and quarter controls. The table shows the estimates of a random effects probit regression in Stata (xtprobit). The table also shows that rho is different from zero, which indicates that the panel-variance component is important and that the panel estimator is different from the pooled estimator.

Table A1 presents the results of the participation equation (1). The table shows that all estimates are statistically significant. The estimated effects of Age, Children, and Married have a negative sign. Therefore, the probability of a female participating in the labor force decreases with age, the number of children, and being married. On the other hand, the estimated effects of education suggest that having more years in education increases the probability of joining the labor force.

Table A2: Log-gross wage Regression.

Age Group	50–70	50–55	56–60	61–65	66–70
Age	–0.004***	–0.002	–0.001	–0.002	–0.012
	(0.000)	(0.002)	(0.003)	(0.004)	(0.007)
Education years	0.102***	0.109***	0.098***	0.090***	0.085***
	(0.000)	(0.008)	(0.001)	(0.001)	(0.003)
Inverse Mills	0.082***	0.142***	0.065***	0.016***	1.023**
Ratio	(0.003)	(0.005)	(0.005)	(0.002)	(0.010)
Constant	2.400***	2.196**	2.252	2.451	3.013
	(0.033)	(0.088)	(0.172)	(0.293)	(0.521)
R^2	0.367	0.378	0.353	0.351	0.297
Sigma_z	0.249	0.181	0.196	0.198	0.261
Sigma_u	0.281	0.242	0.252	0.269	0.301
Rho	0.412	0.375	0.381	0.401	0.435
Observations	136,533	76,757	35,485	15,471	5,820

*p<0.1, **p<0.05, ***p<0.01. Standard errors in parentheses. All specifications include year and quarter controls. The table shows the estimates of a random effects regression in Stata (xtreg). The table also shows that rho is different from zero, which indicates that the panel-variance component is important and that the panel estimator is different from the pooled estimator.

Table A2 presents the results of the log-gross wage equation (2). Due to the data limitations, the equation follows a simple human capital model, where wage is regressed based on an individual's age and education years (Blundell et al. 2001). The equation also controls for selection by including the inverse Mill's ratio. The inverse Mill's ratio is obtained from the participation equation (1), and the ratio represents the covariance between the error terms in the participation and wage equations. Table A2 shows that the effect of Age is only statistically significant for the 50–70 group. On the other hand, the effect of Education is statistically significant and positive. The log gross wage increases with education years, but the increase is smaller for older cohorts. The effect of the inverse Mills ratio is also positive and statistically significant, implying that women with the greatest earning potential are relatively more likely to be observed in the labor market. The latter suggests the presence of selection bias if non-participants are not included in the labor supply model. Hence, I follow the estimation methodology presented in section 3.

Table A3: First stage regression for the marginal wage rate

Age Group	**50–70**	**50–55**	**56–60**	**61–65**	**66–70**
Participants					
Log W^M	0.915***	0.907***	0.916***	0.918***	0.904***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)
Age	0.000**	-0.000	0.000***	0.001***	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Children	0.002***	0.002***	0.003***	0.002***	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Married	–0.006***	–0.000***	–0.005***	0.003***	0.004***
	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)
Constant	0.042***	0.089***	0.028***	–0.008	0.085***
	(0.001)	(0.004)	(0.007)	(0.013)	(0.019)
Observations	136,533	76,757	35,485	15,471	5,820
Prob > Chi2	0.0000	0.0000	0.0000	0.0000	0.0000
All women					
Log W^M	0.918***	0.908***	0.916***	0.919***	0.915***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Age	0.000***	0.000	0.000***	0.001***	0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Children	0.002***	0.003***	0.003***	0.002***	0.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Married	–0.003***	0.000***	–0.002***	0.009***	0.007***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Constant	0.036***	0.084***	0.008	–0.058***	0.071***
	(0.001)	(0.003)	(0.005)	(0.007)	(0.009)
Observations	242,066	120,686	63,581	36,845	20,954
Prob > Chi^2	0.0000	0.0000	0.0000	0.0000	0.0000

*p<0.1, **p<0.05, ***p<0.01. Standard errors in parentheses. All specifications include year and quarter controls. The table shows the estimates of the first stage of the iv tobit regression in Stata (iv tobit) with clustered standard errors. Log W^M denotes the log-marginal wage rate had an individual worked the mean weekly hours.

Table A3 presents the first-stage results of equation (5) when the instrument is the log-marginal wage at means. The table shows that the instrument has a statistically significant effect on the log-marginal wage rate for all cohorts in both groups: all participants and all women.

Table A4 presents the first-stage results of equation (5) when the instrument is the log-marginal wage at 40 hours. The table shows that the instrument has a statistically significant effect on the log-marginal wage rate for all cohorts in both groups: all participants and all women.

Table A4: Labor Supply IV at 40 hours.

Age Group	50-70	50-55	56-60	61-65	66-70
Participants					
Log W^{40}	0.819***	–0.074	1.191***	1.928***	3.316***
	(0.058)	(0.081)	(0.116)	(0.179)	(0.384)
Age	–0.269***	–0.116***	–0.232***	–0.467***	–0.351**
	(0.008)	(0.031)	(0.054)	(0.087)	(0.148)
Children	–0.406***	–0.431***	–0.404***	–0.391***	–0.327***
	(0.021)	(0.029)	(0.039)	(0.053)	(0.079)
Married	–2.441***	–2.595***	–2.422***	–2.192***	–1.806***
	(0.094)	(0.121)	(0.187)	(0.296)	(0.499)
Constant	51.638***	46.604***	49.036***	60.066***	48.176***
	(0.543)	(1.618)	(3.193)	(5.568)	(10.216)
Prob > Chi^2	0.0000	0.0000	0.0000	0.0000	0.0000
Observations	136,533	79,757	35,485	15,471	5,820
All women					
Log W^{40}	8.101***	7.927***	8.768***	10.434***	14.087***
	(0.103)	(0.142)	(0.215)	(0.344)	(0.733)
Age	–1.556***	–0.0752***	–1.574***	–2.738***	–2.647***
	(0.014)	(0.053)	(0.098)	(0.158)	(0.264)
Children	–2.453***	–2.507***	–2.403***	–2.492***	–2.115***
	(0.032)	(0.047)	(0.064)	(0.089)	(0.133)
Married	–12.506***	–10.384***	–13.901***	–18.149***	–22.435***
	(0.177)	(0.225)	(0.361)	(0.579)	(0.961)
Constant	92.709***	48.926***	95.077***	160.729***	142.103***
	(0.911)	(2.871)	(5.775)	(10.017)	(18.167)
Prob > Chi^2	0.0000	0.0000	0.0000	0.0000	0.0000
Observations	242,066	120,686	63,581	36,845	20,954

*p<0.1, **p<0.05, ***p<0.01. Standard errors in parentheses. All specifications include year and quarter controls. The table shows the estimates of the first stage of the ivtobit regression in Stata (ivtobit) with clustered standard errors. Log W^{40} denotes the log-marginal wage rate had an individual worked the mean weekly hours.

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