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共同富裕与现代化国家建设

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北大赛瑟(CCISSR)论坛文集·*2022*

北京大学中国保险与
社会保障研究中心(CCISSR) 编

# 致 谢

北京大学中国保险与社会保障研究中心（CCISSR）衷心感谢以下理事会员单位（2021-2022 年度）的大力支持！

# 编者简介

北京大学中国保险与社会保障研究中心（CCISSR）成立于 2003 年 9 月，是一家致力于保险与社会保障研究的非营利学术机构，其宗旨是充分发挥北京大学经济学院在风险管理与保险学教育和研究方面的优势，借助中外风险管理与保险以及社会保障领域的重要资源，加强学界、监管部门和业界的密切合作，促进保险与社会保障的理论研究、知识传播与实际应用。中心成立十九年来，在学术研究、国内合作与国际交流等方面积极开展工作，在国内外产生了较大的影响。

一年一度的"北大赛瑟（CCISSR）论坛"是中心的一项重要学术活动。"赛瑟"二字源自中心英文简称 CCISSR 的谐音，"赛"意指科学、竞争与交流，"瑟"取庄重、严谨、明洁之意，集中体现论坛汇集国内外学界专家、业界精英及政界高层人士、打造学术争鸣与信息共享的绝佳平台、提供思想碰撞和科学研讨之良机的目标。

# 内容简介

第十九届"北大赛瑟（CCISSR）论坛"于 2022 年 10 月 28 日–29 日在北京大学召开。本届论坛的主题是"共同富裕与现代化国家建设"。来自国内外的学界、政界和业界的专家、学者在论坛上发表了精彩的演讲，数十位论文入选作者和与会者分享了自己的研究成果。

本书收录了大会主题演讲人的发言及具有较高学术价值或创新理论的论坛入选论文，以飨读者。该书适合保险与社会保障领域的学界、业界和政府部门相关人士阅读，是读者了解中国保险与社会保障领域的理论前沿、政策动态和业界发展的理想参考读物。

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# 1

# 共同富裕与现代化
# 国家建设

# 推动中国式保险业现代化 为实现共同富裕保驾护航

梁涛 1

保险是现代经济的重要产业和风险管理的基本手段，具有经济补偿的"补血"功能和资金融通的"造血"功能，是实现现代化和经济富裕过程中重要的助推器。同时，保险的本源产生于互帮互助、扶危济困，自带强烈的社会属性，具有社会管理的"活血"功能，发挥着服务社会、保障民生的稳定器作用。因此，保险与共同富裕具有天然的联系，是实现共同富裕的有效途径和重要保障，在助力共同富裕过程中不可或缺。党的十八大以来，保险业在脱贫攻坚、实现全面小康中发挥了重要作用。在实现中华民族伟大复兴的新征程中，保险业亦大有可为。

保险业可以通过以下路径和方式助力中国式现代化和实现共同富裕：

第一，服务实体经济的高质量发展。要实现中国式现代化和共同富裕，需要大量高附加值的先进制造和专利技术创新。近年来，保险业通过科技保险、专利保险等产品支持国家创新发展战略，支持产业链提升。

第二，助力完善社会分配制度。完善分配制度是实现共同富裕的关键。商业健康保险参与三次分配，更具有针对性和公平性，能够在时间和空间上平滑风险带来的损失；保险通过资金融通参与要素分配，建立"供血"机制，有效缓解民生改造项目资金短缺问题，助推协调发展。

第三，支持构建民生保障体系。例如，疫情防控保险为中小企业的持续经营和员工的稳定生活提供了重要支持；农业保险、大病保险等政策性、普惠性民生保险增强了低收入群体的抗风险能力。

第四，促进社会和谐稳定。除了生活富裕富足外，环境宜居宜业，

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社会和谐和睦，公共服务普及普惠也是共同富裕的重要目标。保险机制能够较好地将"有效市场"与"有为政府"对接起来，已经成为国家社会治理体系的有效补充。

第五，保障人的全面发展。共同富裕更加需要和关注人的全面发展。医疗保险、养老保险助力"病有所医"、"老有所依"，有力支持健康中国战略的实施。

从当前来看，保险业服务共同富裕和中国式现代化仍有较大空间。一是保险业发展水平还不充分；二是保险服务内容有待扩展；三是保险保障仍有较大缺口。

在新的阶段，要积极推动中国式保险业现代化，为实现共同富裕保驾护航。

一是坚持保险服务的政治性和人民性。坚持以人民为中心的发展思想，追求让改革发展成果更多、更公平地惠及全体人民。

二是持续推进供给侧结构性改革。要完善不断创新的多元化保险产品体系，达到需求牵引供给、供给创造需求的更高水平动态平衡。

三是坚持服务实体经济。切实找准保险行业在实现中国式现代化中的职能定位，通过深化改革，融入服务经济社会大盘发展之中。

四是推动保险业的高质量发展。保险公司要坚持守正创新，提高自身专业化、精细化和集约化发展水平，增强金融科技应用，实现降本增效，为共同富裕和中国式现代化提供优质、高效的保险保障。

# 关于共同富裕与高质量发展的几点思考

陈文辉 [1]

第一，实现共同富裕必须坚持高质量发展。党的二十大报告明确提出"中国式现代化是全体人民共同富裕的现代化"。"共同"要求分好蛋糕，"富裕"则要做大蛋糕，做大蛋糕是分好蛋糕的前提。实现共同富裕，既要完善公平分配机制，更要大力发展。发展经济不能简单以生产总值增长率论英雄，要把重点放在推动产业结构转型升级上，要让创新成为第一动力、协调成为内生特点、绿色成为普遍形态、开放成为必由之路、共享成为根本目的，这样才能成就高质量发展。党的二十大报告强调，高质量发展是全面建设社会主义现代化国家的首要任务。只有坚持高质量发展，通过全国人民共同奋斗把"蛋糕"做大做好，才能厚植共同富裕基础，最终实现共同富裕。

第二，高质量发展的核心是夯实实体经济根基。实体经济是我国发展的本钱，也是构筑未来发展战略优势的重要支撑。高质量发展的核心是用人才红利替代人口红利，利用科技创新不断开辟新领域、新赛道，形成新的经济增长极，引导传统企业转型升级，夯实实体经济根基。近年来，我国在多个领域高质量发展取得突破，例如新能源汽车及其产业链，以及光伏行业。我国新能源汽车企业利用超大规模国内市场和工程师红利，不断迭代创新，抓住历史机遇，取得技术和市场的同步领先，同时也带动了产业链上下游的企业快速成长。通过大力发展光伏等新能源产业，为实体经济长期稳定发展提供清洁廉价的能源，既符合中国式现代化提出的人与自然和谐共生的要求，也为我国突破能源瓶颈提供了解决方案，为中华民族伟大复兴打下坚实基础。此外，高质量发展对扩大和稳定就业具有重要意义。社保基金会始终坚持服务实体经济，关注国家战略产业发展，下一步，社保基金会将

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[1] 陈文辉，全国社会保障基金理事会副理事长。本文根据会场速记稿整理，未经本人确认。

坚持把发展经济的着力点放在实体经济上，不断提升实业投资管理运营水平，为服务构建新发展格局，助力国家重大战略实施，推动高质量发展贡献更大的力量。

第三，实现高质量发展需要相应的保障措施。一是充分发挥市场在资源配置中的决定作用。政府立足国家长远发展立场，将经济活动中的外部性内生化，而具体的资源配置则应交给市场，政府不应过多干预，否则会影响公平和效率。二是大力推动高水平对外开放。一方面，只有继续加大开放力度，充分发挥我国比较优势，用好全球资源，才能不断提升我国竞争力，实现经济高质量可持续发展。另一方面，高水平对外开放对树立我国良好的国际形象，增强文化软实力和国际影响力具有重要作用，是中国式现代化的必然选择。三是建立覆盖范围更宽、更广的社会保障体系。在推动高质量发展过程中，经济社会面临的不确定性增加，政府应不断完善社会保障体系，进一步扩大社会保险覆盖范围，适当提高保障标准，大力发展商业保险，为高质量发展保驾护航。

# 共同富裕议题下我国的养老金水平

胡晓义 [1]

我国已经建成世界上最大的社会保障体系，以企业职工基本养老保险为例，虽然只靠基本养老金，企业退休人员不可能达到富裕程度，还需要拓展第二、第三支柱以及其他收入渠道；但有可靠的基本收入托底，富裕才有基础，而且基本养老保险公平、普惠的特质，与"共同富裕"的命题更是密切相关。

第一，衡量养老金水平有明确的国际标准。国际劳工组织（ILO）有 3 个正式文件涉及养老金标准，一是 1952 年第 102 号《社会保障（最低标准）公约》，其中列示"本人、达到领养恤金年龄的妻子"老龄津贴的最低标准为规定工资的 40%；二是 1967 年第 128 号《残疾、老年和遗属津贴公约》，其中列示"有一个达到领抚恤金年龄的妻子的男子"老龄津贴的最低标准为规定工资的 45%；三是 1967 年第 131 号《残疾、老年和遗属津贴建议书》，在"津贴 22"的建议中，提出第 128 号公约"附表列举的百分比应提高十个百分点"，即 55%。根据《国际劳工组织章程》第十九条，公约"应送交各会员国以备批准"；建议书由"会员国考虑通过国家立法或其他方式予以实施"。我国基本养老金水平持续提高是基本事实。但近几年社会上对国际标准也有一些误传，比如将非强制性标准当作了经会员国批准后具有法律效力的国际公约标准，而且忽略两个国际公约中都规定了缴费满 30 年的前提以及该标准是含配偶的 2 人合计待遇水平等。对这种容易误导社会大众认知的议论，专业领域人士有责任正确阐释，防止我国基本社会保障水平跨越"尽力而为"的界限。

第二，评估养老金水平有多个维度。目前评估养老金水平的最基

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本指标是替代率，这一指标可以从多个维度来评估，比如个人或群体、静态或动态等。大多数人在面对多种坐标对照时，通常倾向于选择自己获利更多的标准。国际劳工组织在《世界社会保障报告 2017-2019》中表示，养老金待遇给付充足性的评估维度复杂且具有挑战性，会随着社会、文化、人口和经济条件的变化而变化。这样的评论视角，为我们客观研判我国基本养老金水平以更多的参考。

第三，应多角度观察我国基本养老金水平。首先，转入社会主义市场经济体制后，多种所有制经济共同发展，局限于"非私单位在岗职工工资"确定和评估基本养老金水平，已经脱离实际。其次，加入私营单位工资权重，厘定全社会加权平均工资是一大进步，但仍需形成全国统一评估标准。最后，在我国目前的治理水平和能力条件下，以实际缴费工资作基数评估养老金水平，可能比工资统计指标更准确，而且体现权利与义务对应原则，有鼓励多缴长缴多得的导向功能。

第四，传统的替代率评价需与时俱进。传统的替代率指标是工业社会产物，目前已缺乏代表性，我们可以对替代率及其变化有新的观察和评估视角，从基本养老保险保障基本生活需求的定位出发，综合平衡退休群体与缴费群体的收入、财富积累及消费结构。
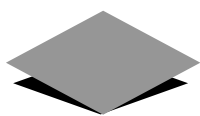
# 2

保险业务经营

# 家庭债务、金融素养对商业寿险消费的影响

唐宏凯、许敏敏 [1]

**摘要:**

我国家庭负债率长期保持一个较高的水平，高负债隐含着大量风险，而商业寿险能够在家庭主要收入来源者遭遇意外时，确保债务能按时偿还，为家庭成员的生活提供保障。为研究二者的关系，本文将用中国家庭金融调查(CHFS)的研究数据，构建 Probit 和 Tobit 模型，探究其影响过程，并引入金融素养分析调节效应。研究发现：家庭债务能够促进商业寿险消费；金融素养能够削弱家庭债务对商业寿险消费的正向刺激，减少家庭过度负债的倾向；金融素养的调节作用在过度负债压力的家庭显著，对于低收入家庭的影响也更加显著。基于此本文提出建议：一是出台配套措施缓解居民家庭高负债压力；二是普及金融知识，增强保险意识，发挥其调节作用；三是保险公司需及时推出针对性的保险产品。

**关键词:** 家庭债务，金融素养，调节作用，商业保险，寿险

## 一、引言

近十年来，我国家庭的负债率与增速都大幅上涨。根据国际清算银行（BIS）2008 年-2020 年第二季度的统计数据，中国家庭负债率由 2008 年的 17.9%升至 2018 年的 59.1%，在十年间大幅上升了 31.2 个百分点。中国的家庭负债现象较为普遍，负债率也在以较快的速度增加。负债是把"双刃剑"。适度负债有助于增强工作动力，释放消费需求，促进生活质量的提升，而过度负债不仅会削弱即期消费能力，还会强化信贷约束，摧毁家庭信用，掣肘未来消费张力，诱发金融风险，损害经济增长[1]。家庭高负债的存在隐含较多风险：家庭主要收入来源者一旦遭遇意外，整个家庭可能陷入困境，无法偿还到期债务，家庭成员的生活无以为续或较难保

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持之前的生活质量。没有适当的商业保险，高负债家庭将较为脆弱，并影响整个社会的稳定。商业寿险能缓和中国家庭高负债所带来的类似问题。在家庭主要收入来源者遭遇意外等时，确保债务能按时偿还且家庭成员的生活能有一定的保障。

在 2021 年《政府工作报告》中商业保险被明确在形成保险民生保障网、促进经济增长中具有重要地位。根据中国银保监会的数据显示，2020年，中国保险保费中，寿险占比达 52.99%，可见我国寿险业务在保险业务中起着比较重要的作用。根据国家统计局数据显示，我国 2012-2017 年寿险保费收入增长迅速，且增速较快，2017 年增幅达 23.3%，但 2017 年后寿险保费收入增长速度与之前相比较为缓慢。

与此同时，我国居民的金融素养水平不高，中国人民银行发布的《消费者金融素养调查分析报告（2021）》全国消费者金融素养指数为 66.81，这说明还需要有进一步提升的空间。金融素养对消费有着两方面的作用，一方面，金融素养的提高有助于提高金融市场参与度，优化配置家庭资产，促进财富稳健增长[2]；另一方面，金融素养的提高也有助于降低过度负债存在的可能性[3]。那么，在对商业寿险支出上，金融素养具有怎样的影响？金融素养是否会调节家庭债务对商业寿险消费的影响？对于不同负债压力、不同家庭收入的家庭，这种调节效应是否会有差异？

因此，本文利用数据分析家庭债务对商业寿险消费影响，并且引入金融素养，分析其作用机制。希望能够为提升我国商业保险市场发展、缓解高负债压力下居民家庭的脆弱性。

## 二、文献综述与研究假说

家庭负债的增加，往往能够起到平滑跨期消费的作用，进而产生对未来良好的预期。VGuerrieri 等（2017）认为，家庭会将其所有资产平滑到生命周期的不同阶段，进行消费或储蓄的跨期决策，以实现效用最优，因而家庭债决策是因平滑跨期消费的需要而产生的[4]；PhilippeBacchetta 等（1997）表示，可预期的消费贷款增长与消费支出之间存在显著正相关关系，而且这种关系对于耐用品和服务的消费更为敏感[5]。张雅淋等（2019）认为，尽管支撑居民家庭做出消费决策的重要决定因素是其在消费当期的可支配收入，但如若消费者的预期收入是稳定的，同时又基于某种消费需求，那么消费者也将更有可能做出借贷的决策，将所借款项包含在当期期末的可支配收入，形成"负债性消费"行为[6]。商业寿险作为家庭消费的一部分，在家庭对未来收入预期向好的情况下，"负债性消费"可能也会作用在商业寿险上面，因此，本文提出假设：

H1:在其他因素不变的情况下，家庭债务能够促进商业寿险消费。

江静琳等（2019）提出：金融素养会影响经济行为主体的资产配置决

策，金融素养高的家庭对于其当前借贷需求和未来还款能力有更好的认知，对于不同渠道和不同种类借款的可得性、利率水平、还款期限等各项信息也都有着更加充分的认识和了解[7]，（吴卫星、孟德锋、尹志超等）认为该类群体持有债务的可能性相对较高，借助有效的金融知识，选择合适的借款渠道、负债规模等，有效缓解当期的信贷约束，并对今后各期的消费和投资行为进行合理规划[2][3][8]。同时金融素养较高的个体会有更高的金融市场参与度、更合理的资产配置结构，并最终通过家庭财富的稳健增长对消费起到促进作用[9]。与此同时，KalantarieTM 等(2013)提出具有较高水平金融素养的行为主体通常不会进行过度负债，在债务偿还期内也基本都会保留较为充足的资金以便按时还款，鲜有因此产生的财务焦虑等现象[11]。

因此，一方面，金融素养的提高会对家庭的信贷约束起到削弱的作用，增加对于包括商业寿险在内的消费水平；另一方面，金融素养越高的家庭，往往会较少过度负债的出现，而是追求债务资产结构的合理性，发挥其调节作用。据此，本文提出如下假设：

H2：在其他因素不变的情况下，金融素养能够提高商业寿险的消费规模

H3：金融素养在家庭债务对于商业寿险消费影响过程中具有调节作用，减少居民家庭对于过度负债的依赖程度。

## 三、研究设计

（一）数据来源与处理

本文使用的数据来源于中国家庭金融调查数据库（CHFS）2017 年的调查数据。该数据样本分布全国 29 个省（市、自治区），样本规模为 40011户。调查问卷的对象为社区、家庭、居民个人，内容包括家庭成员的年龄、工作、婚姻等个人信息，家庭层面的负债、消费、保险保障等信息，基本能够比较全面的反映所调查家庭的基本信息特征。对于 2019 年的最新数据，与本文所选取的 2017 年数据相比，在金融素养相关问题、保险保障方面的数据缺失值较多，难以满足本文的研究分析，所以本文选取 2017年的数据。

本文研究中的地区变量：人均国内生产总值和房价的经济变量来源于中国统计年鉴。

在进行实证分析时，为了保证数据的有效性，本文对数据做了预处理：第一，剔除数据中的缺失值，根据问卷的信息对部分数据的缺失值以 0 值代替，以保证样本数据的连续性；第二，家庭数据中只保留户主的信息，以减弱不同家庭之间数据的对比难度，便于开展研究。最终，得到 9713户样本数据进行研究。

（二）变量选取与样本描述

1.变量选取

(1)被解释变量

本文主要研究家庭负债、金融素养对商业人寿保险消费的影响。本文所使用的数据使用了调查问卷，问卷将商业保险的回答分为了商业人寿保险、商业健康保险、其他商业保险、都没有四个选项，选择商业人寿保险的样本被赋值 1（包括多选的选项），其余选项赋值为 0，构建虚拟变量：商业人寿保险是否支出。

同时，借鉴樊纲治、王宏扬(2015)的研究，本文还选择了另一个变量：家庭商业人寿保险总支出[12]，由于寿险一般采用均衡保费，所以本文利用问卷中去年缴纳的保费×已经缴纳该保险年限得出此变量。

（2）核心解释变量

a.负有债务

借鉴以往对于家庭负债的研究，本文选取负有债务这一虚拟变量。主要根据数据库中关于家庭总负债这一变量进行赋值，若此家庭有家庭总负债数值，即赋值为 1，否则赋值 0。

b.债务资产比

借鉴张雅淋等（2019）的研究，本文选取债务资产比这一相对值变量从另一方面研究家庭债务对于家庭商业人寿保险消费的影响，在此选用相对量指标而非绝对量指标主要是因为，在我国，正规金融机构发放信贷均需以家庭资产或个人信用作为抵押或担保，即使是民间借贷，也多会综合考虑贷款人的信用资质等，因而采用相对量更易反映出家庭部门债务水平的高低状况[6]。

c.金融素养

在调查问卷中，有关金融知识共设置了 15 道题目，根据罗文颖、梁建英（2020）的研究，结合实际情况，本文选取了其中 11 个问题降维处理。这 11 个问题包含了主客观的信息，是一种比较全面的测度方式[13]。进行主成分分析，据此得出金融素养得分。具体选择相关问题以及得分处理见下图

表 1：金融素养指标的相关问题及得分处理

| 问卷题号 | 赋予变量名 | 题目内容 | 得分处理 |
|---|---|---|---|
| h3101 | x1 | 受访者对经济金融信息关注度 | 将五个选项赋分处理,1—5 分关注度逐渐升高 |
| h3103 | x2 | 受益风险识别 | 选项 1 正确，赋值为 1，若选择其他选项则赋值为 0 |
| h3105 | x3 | 利息计算 | 选项 2 正确，赋值为 1，若选择 |

| | | | 其他选项则赋值为 0 |
|---|---|---|---|
| h3106 | x4 | 通货膨胀识别 | 选项 1 正确，赋值为 1，若选择其他选项则赋值为 0 |
| h3107 | x5 | 受访者彩票选择 | 选项 2 正确，赋值为 1，若选择其他选项则赋值为 0 |
| **h3110** | **x6** | 对股票、债券、基金的整体知晓程度 | 按了解程度对选项赋分，1—5 逐渐升高 |
| h3111 | x7 | 股票和基金风险研判 | 选项 1 正确，赋值为 1，若选择其他选项则赋值为 0 |
| h3112 | x8 | 主板和创业板风险判断 | 选项 2 正确，赋值为 1，若选择其他选项则赋值为 0 |
| h3113 | x9 | 偏股型基金和偏债型基金风险判断 | 选项 1 正确，赋值为 1，若选择其他选项则赋值为 0 |
| h3114 | x10 | 国债和公司债风险判断 | 选项 2 正确，赋值为 1，若选择其他选项则赋值为 0 |
| h3115 | x11 | 金融资产风险识别 | 选项 1 正确，赋值为 1，若选择其他选项则赋值为 0 |

基于此，将 11 个变量投入到 STATA 软件中进行主成分分析，首先进行变量间的相关系数检验，虽然适合做主成分分析，但是存在异常值。[1]剔除四个异常值以后，进行 KMO 和巴特利特球形检验。正如表 2 所示，剔除后的样本依然满足显著性要求，p 值显著，并且 KMO 值大于 0.6,适合做主成分分析。

表 2：剩余变量的 KMO 和巴特利特球形检验结果

| KMO 和巴特利特球形检验结果 | |
| --- | --- |
| Det | 0.152 |
| Chi-square | 24851.850 |
| Degreesoffreedom | 21 |
| p-value | 0.000 |
| H0:variablesarenotintercorrelated | |
| 抽样适度测定值 | |
| KMO | 0.816 |

表3：主成分贡献分布表

| Component | Eigenvalue | Difference | Proportion | Cumulative |
| --- | --- | --- | --- | --- |
| Comp1 | 3.001 | 1.769 | 0.429 | 0.429 |
| Comp2 | 1.231 | 0.455 | 0.176 | 0.605 |
| Comp3 | 0.777 | 0.056 | 0.111 | 0.716 |
| Comp4 | 0.721 | 0.210 | 0.103 | 0.819 |
| Comp5 | 0.511 | 0.111 | 0.073 | 0.891 |
| Comp6 | 0.400 | 0.040 | 0.057 | 0.949 |
| Comp7 | 0.360 | . | 0.051 | 1.000 |

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[1] 由于篇幅限制，此处相关系数表以及分析异常值过程没有展示。
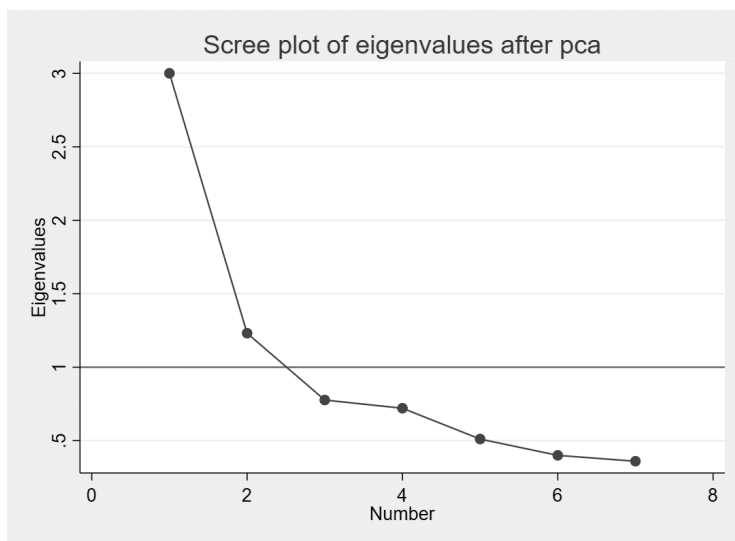
图 1:碎石图

      下一步，我们进行主成分分析，确定要提取的变量。如表 3 所示，特征值大于 1 的变量有两个并且他们的累计贡献率达到了 60%以上，满足主成分分析的要求。再结合图 1 所报告的碎石图，其中靠前的两个因子方差较大，解释力强，而后面平坦的部分解释力较弱。综合比较，决定选取前两个值作为因子来计算金融素养的得分。此外，表 4 的因子载荷矩阵也表明，剔除后无异常值的出现。

表 4：剔除变量后的因子载荷矩阵

| Variable | Factor1 | Factor2 | Uniqueness |
|----------|---------|---------|------------|
| x2 | 0.118 | 0.787 | 0.366 |
| x3 | 0.650 | −0.102 | 0.567 |
| x6 | 0.846 | −0.070 | 0.279 |
| x7 | 0.185 | 0.767 | 0.377 |
| x8 | 0.788 | −0.044 | 0.377 |
| x9 | 0.733 | −0.068 | 0.459 |
| x10 | 0.810 | −0.031 | 0.343 |

表 5：成分得分系数矩阵

| Variable | Factor1 | Factor2 |
|----------|---------|---------|
| x2 | 0.039 | 0.639 |
| x3 | 0.217 | -0.083 |
| x6 | 0.282 | -0.057 |
| x7 | 0.062 | 0.623 |
| x8 | 0.263 | -0.036 |
| x9 | 0.244 | -0.055 |
| x10 | 0.270 | -0.025 |

最后，计算主成分得分，如表 5 所示，将提取出的两个变量作为主成分，并计算所得系数。然后以特征值作为权重计算得出金融素养得分：

金融素养=(0.4287*受益风险识别+利息计算)/0.6046

（3）控制变量

根据以往对于商业保险以及家庭金融的研究经验，本文选取三类控制变量：第一类个体特征变量：性别、年龄、年龄的平方项、受教育年限、婚姻状态、是否创业、是否工作、自评身体状况；第二类为家庭特征变量：家庭总收入、家庭净资产；第三类为地区特征变量：所在省份的人均国内生产总值、所在省份的房价。

表 6：主要变量定义

| 维度 | 变量 | 定义 |
|------|------|------|
| 被解释变量 | 商业人寿保险支出与否 | 虚拟变量，家庭有商业人寿保险支出为 1，否则为 0 |
| | 商业人寿保险总支出 | 商业人寿保险年支出金额和已交年限相乘所得 |
| 解释变量 | 负有债务 | 虚拟变量，家庭负有债务为 1，否则为零 |
| | 债务资产比 | 家庭总债务与总资产的比值（单位：%） |
| | 金融素养 | 受访者金融素养水平高低，数值越大表示金融素养水平越高，范围 1-10 分 |
| 控制变量 | 性别 | 虚拟变量，男性赋值为 1，女性赋值为 0 |
| | 年龄 | 家庭户主年龄 |
| | 年龄的平方 | 家庭户主年龄的平方项 |

| 项 | |
|---|---|
| 受教育年限 | 小学以下为 0 年，小学为 6 年，初中为 9 年，高中、职业高中、中专、技校为 12 年，大专、高职为 15 年，大学本科为 16 年，硕士研究生为 19 年。博士研究生为 22 年 |
| 婚姻状态 | 虚拟变量，已婚或同居为 1，否则为 0 |
| 是否创业 | 虚拟变量，问卷中家中从事工商业赋值为 1，否则为 0 |
| 是否工作 | 虚拟变量，当前处于工作状态为 1，否则为 0 |
| 自评身体状况 | 赋值 1~5，数值越大表示越健康 |
| 家庭总收入 | 包括工资性收入、农业经营收入、工商业经营收入、转移性收入和投资性收入（单位：元） |
| 家庭净资产 | 家庭总资产减去家庭总负债所得（单位：元） |
| 人均生产总值 | 家庭所在省份的人均生产总值（单位：元） |
| 房价 | 家庭所在省份的住房价格（单位：元） |

2.样本描述

从被解释变量来看，此次调查中的家庭样本购买商业人寿保险的比率仅占 5%左右，从商业人寿保险支出的规模来看，全部样本的总支出平均值为 3392 元，从中可以看出，家庭商业人寿保险整体参与率偏低，一定程度上反映出我国商业人寿保险市场还有较大的发展空间。

从解释变量来看，样本中负有债务的家庭占比 38%，并且债务资产比达到 52%，这反映出，样本中的家庭负有债务还是比较普遍，不仅如此，家庭负有的债务还占到家庭净资产的一半左右，家庭负债压力比较大，一定程度上也反映出当前我国家庭普遍负债压力大的现象；从金融素养来看，家庭金融素养得分基本维持在较低的水平，金融知识普及度不高。

从控制变量来看，在个体特征变量中，性别均值为 0.794，说明家庭户主男性比例偏高，在两个子样本中，也是男性比例偏高；家庭户主年龄均值为 56 岁左右；受教育年限均值为 9 年左右，整体学历不高，可能也是影响金融素养偏低的原因之一；从均值来看，大部分户主为已婚或同居状态；从创业和工作状态来看，家庭户主较多处于工作状态，有一部分同时有自营工商业，并且对于负有债务的户主来说，

其处于工作状态和自营工商业的比率更高，这也能体现出负债的压力下，会促进人们进行工作或创业以获取资金；身体状况均值为 3.5 左右，身体状况良好。家庭特征变量：家庭总收入和家庭净资产具有很大的标准差，说明样本家庭财富间的差异非常明显；地区特征变量：家庭所在省份的人均 GDP 和房价标准差较大，说明样本家庭间地区差异也比较明显。

表 7：主要变量描述性统计

| 维度 | 变量 | 全样本 | 负债样本 | 无负债样本 |
|---|---|---|---|---|
| 被解释变量 | 商业人寿保险支出与否 | 0.0546<br>(0.227) | 0.0636612<br>(0.2441818) | 0.0490666<br>(0.2160249) |
| | 商业人寿保险总支出 | 3,392<br>(38,139) | 4444.92<br>(44183.34) | 2755.572<br>(33955.45) |
| 解释变量 | 负有债务 | 0.377<br>(0.485) | / | / |
| | 债务资产比 | 52.3<br>(927.3) | 1.387156<br>(15.06732) | / |
| | 金融素养 | 2.204<br>(0.802) | 2.241151<br>(0.8842078) | 2.181266<br>(0.7467739) |
| 控制变量 | 性别 | 0.794<br>(0.404) | 0.8005464<br>(0.3996442) | 0.7901867<br>(0.4072089) |
| | 年龄 | 55.74<br>(14.12) | 51.593<br>(13.09475) | 58.24483<br>(14.12357) |
| | 年龄平方项 | 3,306<br>(1,605) | 2833.263<br>(1407.909) | 3591.903<br>(1649.507) |
| | 受教育年限 | 9.903<br>(3.965) | 10.08169<br>(4.203051) | 9.795143<br>(3.810836) |
| | 婚姻状态 | 0.868<br>(3.965) | 0.8721311<br>(0.3339894) | 0.8658516<br>(0.3408398) |
| | 是否创业 | 0.190<br>(0.392) | 0.2459016<br>(0.4306793) | 0.1556253<br>(0.3625297) |
| | 是否工作 | 0.676<br>(0.468) | 0.7571038<br>(0.4288915) | 0.6264662<br>(0.4837819) |

| 自评身体状况 | 3.564<br>(0.991) | 3.490984<br>(1.06096) | 3.608459<br>(0.9442531) |
|---|---|---|---|
| 家庭总收入 | 118,350<br>(251,990) | 137495.9<br>(304043.7) | 106773.4<br>(213668.6) |
| 家庭净资产 | 1372360<br>(2570524) | 1374451<br>(2754165) | 1371096<br>(2453058) |
| 人均生产总值 | 70,477<br>(28,892) | 65155.25<br>(26524.12) | 73695.52<br>(29778.33) |
| 房价 | 10,108<br>(6,536) | 9196.726<br>(5717.556) | 10658.78<br>(6926.598) |

为了更为直观地展示家庭负债对商业人寿保险消费的影响，本文绘制了图 2。将全样本分为负债样本和无负债样本，柱状图表示不同样本下商业人寿保险总支出，折线图表示不同样本下，商业人寿保险支出所占比率。从图中可以很明显地看出，负债家庭购买商业人寿保险的比率更高并且购买商业人寿保险的总支出也更高。
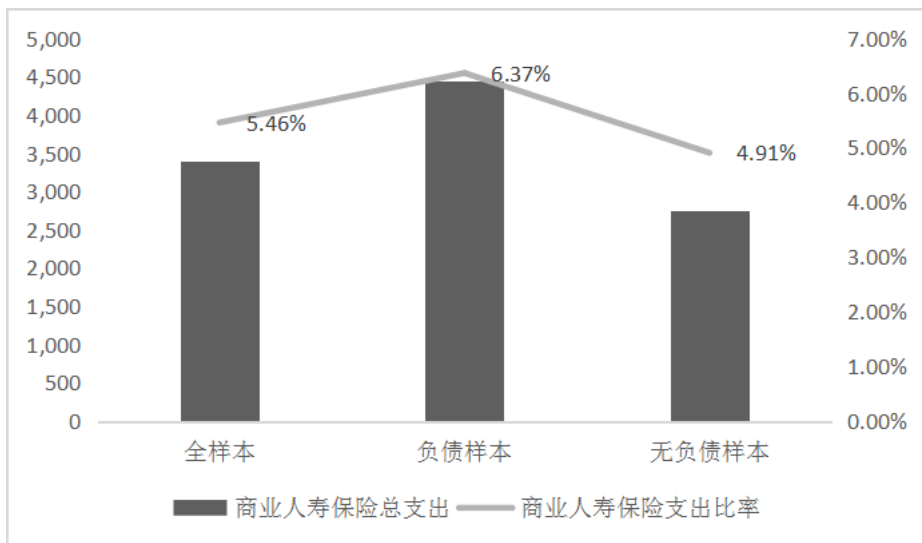


图 2：分样本支出比率及总支出

（三）模型建立

本文拟通过实证分析回答如下问题：第一，家庭负债和金融素养

分别对家庭的商业人寿保险是否支出有何影响？第二，家庭负债和金融素养分别对家庭商业人寿保险支出水平有何影响?第三,金融素养在家庭负债影响商业寿险消费的过程中起到怎样的调节作用？

1.Probit 模型

如表 8 所示，相关系数展示了本文主要变量间的相关系数关系，由于是否负债和金融素养两个变量间存在显著的相关性关系，为了排除共线性的影响，同时根据张雅淋、孙聪等（2019）的研究，债务资产比能够更好的排除家庭其他特征所带来的影响，同时也能够反映出家庭负债的特征，故下文模型中，选取债务资产比这一解释变量构建模型。

表 8：变量间相关系数表

|  | 是否负债 | 债务资产比 | 金融素养 |
|---|---|---|---|
| 是否负债 | 1 | | |
| 债务资产比 | 0.072*** | 1 | |
| 金融素养 | 0.036*** | −0.0150 | 1 |

为了考察家庭负债和金融素养分别对商业人寿保险消费有何影响，并且考虑到商业人寿保险支出与否为虚拟变量，本文使用 Probit 模型，具体模型如下：

$$Probit(Y=1) = \alpha + \beta_1 debt\_asset\_r_{ij} + \beta_2 fin\_literacy_{ij} + \beta_3 I + \beta_4 F + \beta_5 R + \varepsilon_{ij} \#(1)$$

其中 i 和 j 代表家庭个体和家庭所在地，被解释变量 Y 表示商业人寿保险是否支出，购买商业人寿保险记为"1"，否则记为"0"；解释变量 debt_asset_r 为资产负债比，fin__literacy 为金融素养。同时，借鉴已有研究本文也控制了影响家庭消费的其他因素，主要包括个体特征向量组 I,家庭特征变量组 F 和地区特征向量组 R。其中，为了减弱异方差性，本文在后续回归中，对家庭净资产以及人均 GDP 和房价进行取对数处理，模型最后 $\varepsilon$ 为随机扰动项。

2.Tobit 模型

考虑到目前取对数的商业人寿保险总支出变量存在较多的 0 值，所以本文选取左截断点为 0 的 Tobit 模型进行回归：

$$Tobit y^* = \alpha + \beta_1 debt\_asset\_r_{ij} + \beta_2 fin\_literacy_{ij} + \beta_3 I + \beta_4 F + \beta_5 R + \varepsilon_{ij} \#(2)$$

其中 y*表示的是商业人寿保险总支出，并且做了取对数处理，其他控制变量与（1）式相同。

3.调节效应的 Tobit 模型

本文借鉴姚玲珍、张雅琳（2020）的研究，构建金融素养和债务指标的交乘项，与本文而言，就是金融素养与资产负债比的乘积。通过建立 Tobit 模型以研究金融素养在家庭负债影响商业人寿保险支出过程中所起到的调节作用。具体模型如下：

$$Tobity^* = \alpha + \beta_0 debt_asset_{r_{ij}} \times fin_literacy_{ij} + \beta_1 debt_asset_{r_{ij}} + \beta_2 fin_literacy_{ij}$$

$$+ \beta_3 I + \beta_4 F + \beta_5 R + \varepsilon_{ij} \#(3)$$

（四）稳健性检验

为了确保本文研究内容结构的可靠性，根据已有研究，实证研究内生性主要来源于遗漏变量、互为因果以及样本选择偏差。针对本研究，由于本文所选择的控制变量基于以往研究所得，基本涵盖了可以选择的控制变量，可以认为由遗漏变量产生的内生性问题不会对模型产生太大影响；针对互为因果问题，已有研究认为家庭消费会影响借贷决策[14]，但是就本文来说，从常识来分析，家庭购买寿险不会对家庭的借贷决策产生太大影响，也就是说，大部分家庭不太可能因为购买寿险而选择去借贷消费；而针对可能因为样本选择偏误带来的影响，本文选择倾向得分匹配（PSM）一对一近邻匹配的方法，按照是否负有债务分别定义为处理组和控制组，构建一个新的子样本以此缓解样本偏误的问题。

（五）异质性分析

为例进一步考察家庭负债、金融素养对商业人寿保险消费的影响，本文将群体进行以下分组进行回归：第一，在资产负债比这一变量的2/3 处插入分位点，分为一般负债压力和过度负债压力进行考察。第二，在家庭总收入变量 1/3 和 2/3 处分组，将负债家庭的样本分为三组。

# 四、实证结果分析

（一）家庭负债与金融素养对商业人寿保险支出与否的影响

解释变量：表 9 报告了家庭负债和金融素养对商业人寿保险是否支出的影响。表中第一列为基于模型(1)的 Probit 回归结果，可以表明，家庭负债对商业人寿保险支出与否的影响显著为正，统计意义上具有 5%的显著性水平，即债务资产比上升 1%，会导致商业人寿保险支出的可能性增加，正如前述所言，"负债性商业寿险消费"现象是存在的。这可能源于当前借贷消费体系日渐完善，居民家庭通过借贷形式实现了对预期可支配收入的进一步优化分配[15]，居民可实现借贷的渠道也增多，居民通过这些渠道提高了预期收入，同时一些分期付款的方式也减轻了当前消费的支出门槛，能够促进消费。另一方面，随着人们保险意识的提高，商业人寿保险作为一种预防未来风险的有效方式，已经成为了较为必要的消费支出。所以，家庭债务刺激了商业人寿保险的消费。金融素养也在 5%的显著性水平上为正，这表明金融素养的提高也会促进商业寿险的支出可能性。金融素养高的家庭能获取到更多的保险知识，所以他们去购买商业人寿保险的可能性也越高。

控制变量：在个体特征变量方面，性别与是否支出有负效应，这可能是因为相对于女性来说，男性对自身身体健康、未来保障的意识没有那么强，所以他们购买商业人寿保险的意愿相对较少；结合户主年龄的两个变量，户主的年龄可能与商业人寿保险是否支出呈现倒 U型的关系；受教育年限显著为正，表明受教育程度越高，接受的保险知识越多，从未增加支出的可能性；婚姻状态显著为负，这可能是因为未婚状态的受访者会对增加对未来风险的担忧，因为担心未来没有相伴的人照顾自己和家庭，从而增加购买寿险的可能；是否工作显著为正，这可能是因为拥有工作的户主，享受到了社保的待遇，保险意识更强，更愿意进行商业寿险的消费。家庭特征变量方面，家庭净资产显著为正。因为家庭资产越高，购买力越强，能够获取的信息资源也相对较多，从而增加寿险消费。地区特征变量方面，结果不显著，这可能是因为样本量不足，影响了显著性。

表 9 第二列展示了稳健性检验所得结果，主要思想是使得两组变量中每一户主尽可能匹配到具有相近家庭情况的受访者，再进行回归。回归结果显示,资产负债比以及金融素养两个核心解释变量依然显著,且系数方向和大小并未发生明显变化。所以，可以认为本文模型（1）的回归结果具有一定的可靠性。

表 9：家庭负债与金融素养对商业人寿保险是否支出的影响

被解释变量：商业人寿保险支出与否

| 变量 | 模型(1) | PSM匹配后样本 |
|---|---|---|
| 债务资产比 | 0.349** | 0.401** |
| | (0.162) | (0.172) |
| 金融素养 | 0.065*** | 0.063** |
| | (0.023) | (0.026) |
| 性别 | −0.150*** | −0.099 |
| | (0.055) | (0.066) |
| 年龄 | 0.046*** | 0.047*** |
| | (0.013) | (0.018) |
| 年龄的平方 | −0.001*** | −0.001*** |
| | (0.000) | (0.000) |
| 受教育年限 | 0.022*** | 0.021** |
| | (0.007) | (0.008) |
| 婚姻状态 | −0.236*** | −0.298*** |
| | (0.071) | (0.085) |
| 是否创业 | 0.074 | 0.099 |
| | (0.054) | (0.061) |
| 是否工作 | 0.151** | 0.145* |
| | (0.063) | (0.079) |
| 自评身体状况 | 0.010 | −0.004 |
| | (0.026) | (0.032) |
| ln(家庭净资产) | 0.163*** | 0.167*** |
| | (0.018) | (0.022) |
| ln(人均生产总值) | −0.034 | −0.145 |
| | (0.109) | (0.130) |
| ln(房价) | −0.006 | 0.096 |
| | (0.085) | (0.104) |
| _cons | −4.538*** | −4.203*** |
| | (0.763) | (0.924) |
| N | 9374.000 | 6168.000 |

Standard errors in parentheses

$^{*}p<0.1$, $^{**}p<0.05$, $^{***}p<0.01$

（二）家庭负债和金融素养对商业人寿保险支出水平的影响

解释变量：债务资产比对于商业人寿保险支出水平显著为正，不

仅进一步证明家庭债务能够促进商业寿险的消费，还表明了促进消费增长的幅度。即资产负债比增加 1%，商业寿险支出费用增加 38%，从需求端考虑，这可能是因为负债的压力使得受访者（户主）更倾向于选择寿险以应对未来因去世而无法偿还贷款的风险，比如房屋的主要供款人，因为房贷的压力，促使其选择购买寿险，如果未来不幸去世，还可以用保险金偿还房贷，避免断供。金融素养提高 1%也会带来商业寿险支出增长 11.2%，金融素养高的家庭更愿意去购买商业寿险作为自身权益的保障。

控制变量：个体特征变量方面：性别的负向效应影响较大，能够减少 17.%的商业寿险支出，这可能还是和男性女性的思维差异有关，男性往往对于自身健康状况，未来受保障的程度没有那么重视；年龄因素影响效应相对较少，而且仍然呈现倒 U 型关系；受教育年限的正向效应达到 2.3%，低于金融素养的正向效应，这可能是因为学生受教育的专业不同，教学内容也有很大差异，总体上对于金融知识的了解自然也会相对来说低一些；是否工作能够带来的正向效应为 18.1%，对商业人寿保险的支出的影响也相对大一些，可见，随着参加工作，缴纳社保，对商业寿险的了解程度和支付意愿也会较多的提高。家庭特征变量方面：家庭净资产的正向效应也比较强，可见家庭的资产较多，可支配收入也会相应多一些，从而也有着更高的商业寿险支付意愿。

表格第二列展示了稳健性检验的结果，可以看出，再 PSM 匹配后，债务资产比的正向效应更加强，金融素养也保持了一致的显著性，可以充分说明家庭负债和金融素养对商业保险支出水平都有着促进作用，回归结果可靠。

表 10：家庭负债与金融素养对商业人寿保险支出水平的影响

被解释变量：ln(商业人寿保险支出水平)

| 变量 | 模型(2) | PSM匹配后样本 |
| --- | --- | --- |
| 债务资产比 | 0.381** | 0.522*** |
| | (0.159) | (0.184) |
| 金融素养 | 0.112*** | 0.106*** |
| | (0.029) | (0.037) |
| 性别 | −0.175*** | −0.157** |
| | (0.059) | (0.079) |

| | | |
|---|---|---|
| 年龄 | 0.011 | 0.014 |
| | (0.011) | (0.017) |
| 年龄的平方 | -0.000* | -0.000 |
| | (0.000) | (0.000) |
| 受教育年限 | 0.023*** | 0.021** |
| | (0.007) | (0.009) |
| 婚姻状态 | -0.142* | -0.225** |
| | (0.074) | (0.101) |
| 是否创业 | 0.118* | 0.131* |
| | (0.061) | (0.075) |
| 是否工作 | 0.181*** | 0.208** |
| | (0.059) | (0.082) |
| 自评身体状况 | 0.006 | -0.017 |
| | (0.025) | (0.035) |
| ln(家庭净资产) | 0.139*** | 0.178*** |
| | (0.015) | (0.022) |
| ln(人均生产总值) | -0.053 | -0.208 |
| | (0.107) | (0.143) |
| ln(房价) | 0.042 | 0.224* |
| | (0.086) | (0.120) |
| _cons | -1.619** | -1.939* |
| | (0.743) | (1.026) |
| / | | |
| var(e.ln_premium) | 4.845*** | 5.566*** |
| | (0.071) | (0.100) |
| N | 9374.000 | 6168.000 |

$^*p<0.1, ^{**}p<0.05, ^{***}p<0.01$

（三）金融素养的调节作用

表 11 展示了金融素养在家庭负债对商业寿险支出水平影响过程中的调节作用，从表中可以看到，资产负债比金融素养具有显著正效应，而金融素养与资产负债比的交互项具有强显著性，系数为负。这说明在家庭债务对商业寿险的影响过程中，金融素养起到了削弱或缓解的调节效应。如前所述，家庭负债对商业寿险的支出具有促进作用，即在商业寿险上，存在"负债性消费"的现象。但是，随着家庭金融

素养的提高,家庭有了更多的金融知识以及合理配置资产结构的意识,家庭就会更追求更合理的借贷途径以及合理的负债结构,不会使得家庭一味增加借贷,家庭对于过度负债的意愿减少,即减少商业寿险的消费支出。

表11:金融素养的调节作用回归结果

| | 被解释变量:ln(商业人寿保险支出水平) |
|---|---|
| 金融素养×资产负债比 | −0.023*** |
| | (0.007) |
| 资产负债比 | 0.049*** |
| | (0.014) |
| 金融素养 | 0.129*** |
| | (0.029) |
| 性别 | −0.166*** |
| | (0.058) |
| 年龄 | 0.017 |
| | (0.011) |
| 年龄的平方 | −0.000** |
| | (0.000) |
| 受教育年限 | 0.032*** |
| | (0.007) |
| 婚姻状态 | −0.112 |
| | (0.073) |
| 是否创业 | 0.209*** |
| | (0.059) |
| 是否工作 | 0.145** |
| | (0.058) |
| 自评身体状况 | 0.023 |
| | (0.025) |
| ln(家庭净资产) | 0.092*** |
| | (0.017) |
| ln(人均生产总值) | 0.099 |
| | (0.085) |
| ln(房价) | −0.041 |

|  |  | (0.105) |
| _cons |  | −1.797** |
|  |  | (0.728) |

| / |  |  |
| var(e.ln_premium) |  | 4.786*** |
|  |  | (0.069) |

| N |  | 9533.000 |

（四）金融素养调节作用的异质性分析

为了更好的衡量金融素养的调节作用在不同负债压力下有何不同，本文在资产负债比变量中插入一个2/3位置的分位点，将低于2/3位置资产负债比的数据定义为一般负债压力，而大于该分位点的数据定义为过度负债压力，进行分组回归。

表12:不同负债压力下金融素养的调节作用

被解释变量：ln(商业人寿保险支出水平)

| 变量 | 一般负债压力 | 过度负债压力 |
| --- | --- | --- |
| 金融素养×资产负债比 | −29.813 | −0.024*** |
|  | (20.942) | (0.008) |
| 资产负债比 | 66.594 | 0.052*** |
|  | (47.861) | (0.016) |
| 金融素养 | 0.136*** | 0.135*** |
|  | (0.037) | (0.050) |
| 性别 | −0.176*** | −0.156 |
|  | (0.068) | (0.111) |
| 年龄 | 0.009 | 0.039* |
|  | (0.012) | (0.023) |
| 年龄的平方 | −0.000 | −0.000* |
|  | (0.000) | (0.000) |
| 受教育年限 | 0.029*** | 0.035*** |
|  | (0.008) | (0.012) |
| 婚姻状态 | −0.163* | −0.004 |

| | | |
|---|---|---|
| | (0.085) | (0.137) |
| 是否创业 | 0.148** | 0.274*** |
| | (0.075) | (0.099) |
| 是否工作 | 0.105 | 0.199* |
| | (0.067) | (0.111) |
| 自评身体状况 | 0.013 | 0.037 |
| | (0.030) | (0.045) |
| ln(家庭净资产) | 0.063*** | 0.136*** |
| | (0.021) | (0.031) |
| ln(人均生产总值) | 0.023 | 0.314* |
| | (0.098) | (0.170) |
| ln(房价) | 0.030 | −0.196 |
| | (0.124) | (0.198) |
| _cons | −1.198 | −3.382** |
| | (0.859) | (1.423) |
| / | | |
| var(e.ln_premium) | 4.441*** | 5.454*** |
| | (0.079) | (0.137) |
| N | 6380.000 | 3153.000 |

Standard errors in parentheses

$^*p<0.1, ^{**}p<0.05, ^{***}p<0.01$

通过表 12 可以看出，金融素养仅在过度负债压力的家庭中产生缓解或削弱效应，并且在统计意义上具有 1%的显著性水平，而在一般负债压力的家庭中，金融素养的调节作用并不显著。正如前文所述，金融素养的调节作用体现在其能够调节家庭的负债结构，防止过度负债的出现，从而起到削弱或抑制作用。此部分的结果正好与之相对应，金融素养的调节作用在过度负债的家庭中发挥了显著的作用，而在一般负债压力的家庭中，其负债结构相对来说比较良好，所以金融素养调节作用不显著。

不同的家庭收入下，家庭负债对商业寿险支出水平的影响也会有所不同。表 13 展示了不同家庭收入下金融素养的调节作用，从中可以看出，金融素养的调节项、资产负债比和金融素养对商业寿险支出水平的影响都显著为正，这与前文的论述一致。但是这三个变量在中高收入的家庭中，影响都不显著。这是因为，对于低收入的家庭，其应

对风险的能力更弱，对于预防未来不可预测的生命风险的需求更大。因此当低收入的家庭背负着债务时，若是主要还款人遭遇不测，就需要商业寿险给予保险费以防止中断还贷。其中，对于低收入的家庭来说，它的金融素养相对来说较低，那么增加1%的金融素养，所带来的边际效用以及调节作用也会更显著，而中高收入家庭，则并不显著。因此，家庭负债对商业寿险支出水平的影响，主要体现在低收入家庭上。

表 13:不同家庭收入金融素养的调节作用

被解释变量：ln(商业人寿保险支出水平)

| | 低收入家庭 | 中收入家庭 | 高收入家庭 |
|---|---|---|---|
| 金融素养×资产负债比 | −0.023*** | −0.251 | 0.051 |
| | (0.005) | (0.347) | (0.472) |
| 资产负债比 | 0.048*** | 0.690 | −0.356 |
| | (0.010) | (0.839) | (1.359) |
| 金融素养 | 0.128*** | 0.030 | 0.173*** |
| | (0.047) | (0.050) | (0.062) |
| 性别 | −0.083 | −0.051 | −0.280** |
| | (0.072) | (0.092) | (0.125) |
| 年龄 | −0.006 | 0.032* | 0.021 |
| | (0.012) | (0.017) | (0.026) |
| 年龄的平方 | 0.000 | −0.000** | −0.000 |
| | (0.000) | (0.000) | (0.000) |
| 受教育年限 | 0.009 | 0.010 | 0.046*** |
| | (0.008) | (0.011) | (0.015) |
| 婚姻状态 | −0.065 | −0.240** | −0.123 |
| | (0.077) | (0.120) | (0.179) |
| 是否创业 | 0.098 | 0.031 | 0.317** |
| | (0.086) | (0.088) | (0.128) |
| 是否工作 | 0.077 | 0.130 | 0.168 |
| | (0.065) | (0.090) | (0.140) |
| 自评身体状况 | −0.034 | 0.060 | 0.072 |
| | (0.027) | (0.038) | (0.063) |
| ln(家庭净资产) | 0.030 | 0.047 | 0.140*** |

|  | (0.019) | (0.029) | (0.047) |
|---|---|---|---|
| ln(人均生产总值) | -0.104 | -0.125 | 0.241 |
|  | (0.106) | (0.145) | (0.183) |
| ln(房价) | 0.057 | -0.015 | -0.168 |
|  | (0.125) | (0.157) | (0.252) |
| _cons | 0.384 | 0.302 | -2.408 |
|  | (0.912) | (1.197) | (1.858) |
| / |  |  |  |
| var(e.ln_premium) | 2.253*** | 3.714*** | 8.211*** |
|  | (0.057) | (0.093) | (0.206) |
| N | 3154.000 | 3196.000 | 3183.000 |

$^*p<0.1, ^{**}p<0.05, ^{***}p<0.01$

## 五、研究结论及启示

商业人寿保险的发展对提振国家经济运行有着重大作用。研究商业人寿保险消费影响因素的文献大多集中在个体特征、家庭经济特征、主观态度以及社会因素方面，研究家庭债务对其影响的文献很少。同时，在居民家庭负债率持续攀升的背景下，为了促进商业寿险的发展，政府应该提供怎样的负债政策？保险公司应该如何开发出合适的产品？基于此，本文利用 CHFS 的数据进行了实证检验。研究结果首先证明了家庭债务的增加会促进商业寿险的消费。同时，研究发现了金融素养能够通过调节家庭债务资产结构的合理化，削弱家庭债务对商业寿险消费的正向刺激，减少家庭过度负债的倾向。此外，研究发现，金融素养的调节作用在存在过度负债压力的家庭是显著的，进一步证明了上述结论，并且调节作用对于低收入家庭的影响也更加显著，对于中高收入家庭影响较小。

完善商业寿险体制机制，建立促进商业寿险消费的长效机制能够有效发挥其对于经济的"稳定器"作用。而家庭负债是其中不可忽略的一环，结合本文研究结论，为了更好的缓解当前我国居民家庭高负债的脆弱性和不稳定性，更好地发挥商业寿险的功能。本文提出如下建议：1、相关部门需及时缓解当前高负债家庭的压力。这需要相关部门依据不同的家庭收入、负债压力等因素综合考虑，制定出针对化，差异化的政策，尤其是目前我国居民住房债务负担较大，对于住房债务以及高负债压力的家庭需要注重缓解其负担,避免一味增大杠杆;2、通过建立多样化的学习平台、普及金融知识，提高居民整体的金融素

养水平，促进居民合理调整债务结构，减少居民过度负债倾向。三是保险公司需要及时开发出相适应的保险产品，以满足因为家庭债务而产生的对于商业寿险的需求。例如，推出债务促进型寿险，对于债务压力较高的家庭减少保费；或者推出"保供型"寿险，对于投保人中途断供，可以提供保障等产品；与此同时，还可以在保单中提供相关增值业务，以帮助投保人管理家庭债务等。

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# 基于流动性风险的存款保险定价模型研究

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**摘要：**

　　流动性风险主要产生于银行无法应对因负债下降或资产增加而导致的流动性困难。当商业银行流动性不足时，它无法以合理的成本迅速减少负债或变现资产获取足够的资金，极端情况下，流动性不足能导致银行资不抵债而倒闭。文章基于流动性风险研究了存款保险定价问题，进一步分析了定期存款和活期存款在存款保险上的差别风险。使用 Gap 期权定价模型，纳入清算折扣因子，给出了不考虑技术性破产和考虑技术性破产情况下定期存款和活期存款的存款保险费率；其次根据中国 25家上市银行的数据进行实证测算，分别计算不同情况下的存款保险费率，并对实证的结果进行比较。测算结果表明，考虑银行的流动性问题后存款保险费率有所上升，并且定期存款的保险费远远高于活期存款保险费。

**关键词：**流动性风险，Gap 期权，清算折扣因子，活期存款

## 一、引言

　　存款保险制度是一种金融保障制度，是指符合条件的存款性金融机构集中起来建立一个保险机构，各存款机构作为投保人按一定比例向其缴纳保费，当成员机构发生经营危机或面临破产倒闭时，存款保险机构向其提供财务救助或直接向存款人支付部分或全部存款，以减少存款人的损失。存款保险制度起源于美国，美国在推进利率市场化进程中，曾发生过大量的商业银行倒闭事件，初期每年商业银行倒闭的数量为两位数，1987-1991 年平均每年则达到 20 家，最高一年达到 250 家。银行的大量倒闭使得人们的储蓄顷刻间灰飞烟灭，严重损害了存款人的利益和

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金融系统的稳定。而建立存款保险制度后，投保银行可以从保险机构取得资金救助或得到存款理赔，从而最大限度地保护了存款人的利益，也极大地减轻了银行的压力和风险。历史经验和实践表明，存款保险制度已成为抵御金融风险的一道重要防线。

我国金融市场发育尚不完善，金融监管理念、手段和方法尚不能与时俱进。同时，商业银行经营管理机制尚未完全转变，资产负债结构不合理现象较为突出，从而严重影响了风险抵御能力，因而建立存款保险制度不仅有利于革新传统观念，提高公众风险意识，还有利于加强中央银行的监管力度，减轻中央银行的负担。存款保险机构负责监督银行日常的经营活动，定期检查银行的财务状况，审查其上报的统计报表和账目。当银行经营管理不善或经营非法、风险较大的业务时，存款保险机构对其提出警告，勒令整改，帮助银行渡过难关，或促成其它银行的并购，从而实现中央银行的监管意图。1993年国务院首次提出了"要建立存款保险基金，保障社会公众利益"，历时21年，2014年11月30日，我国的存款保险制度终于诞生，并于2015年5月1日正式实施《存款保险条例》。这一标志性的事件，将政府长期以来承担的隐性担保责任通过制度而显性化。《条例》的颁布，亦使得我国以银行为主导的金融体系有了抵御风险的制度性保障，意义重大。虽然是首次颁布并实施的《条例》，但我国的存款保险制度更好的进行了国外成熟经验借鉴和充分考虑了国情，以风险为导向的差别化费率更是其突出特点。

存款保险制度的核心是存款保险的定价。判断一个存款保险制度建立的是否得当，能否最大限度的发挥其保障作用，关键在于存款保险的定价是否合理，合理的费率厘定及费率结构，能够起到维护金融体系稳定的关键作用。我国的《条例》中虽明确规定采用风险差别费率以区分不同银行的风险，但是如何衡量不同银行的风险，却未给出明确答案。同时，尽管学术界对存款保险公平定价问题进行了研究，但大部分模型都是建立在Merton（1977）[1]期权定价模型的经典范式之上，未能充分考虑活期存款和定期存款的差异及流动性风险的影响。有鉴于此，基于制度完善的理论与实践需要，文章建立了基于流动性风险的存款保险定价模型，运用Gap期权定价模型并纳入清算折扣因子，给出了不考虑和考虑技术性破产情况下定期存款和活期存款的存款保险费率进一步，文章运用我国25家上市银行的数据进行实证测算,对不同情况下的存款保险费率进行比较并最终给出政策建议。

## 二、国内外文献回顾

存款保险费率的公平定价是存款保险改革中争论最激烈的问题之一，与其他形式的保险一样，存款保险应根据风险定价。存款保险机构通过担保金融机构的存款，承担了潜在责任，投保机构的违约风险越大，存款保险机构潜在的责任就越大。因此，存款保险应根据投保机构对保险机构施加的潜在责任进行定价。

在存款保险定价的文献中，人们一直致力于公平定价问题的研究。存款保险费率定价的经典模型是 Merton（1977）[1]的存款保险期权定价模型，他将定期存款的保险视为存款保险机构发行的看跌期权，并运用 Black-Scholes（1973）[2]的期权定价模型推导出了存款保险费率的计算公式。后续的众多研究工作都是在这个模型基础之上进行拓展和完善的，如 Marcus 和 Shaked（1984）[3]发现银行的资产价值在获得存款保险前后是不同的，并在 Merton 模型中加入了股利发放率的情况；Ronn（1986）[4]在模型中加入监管宽容系数，并在 Merton 模型的基础上，利用股权与欧式看涨期权之间的关系，建立银行资产市场价值与银行权益市场价值，以及银行资产价值波动性与银行权益价值波动性之间的约束关系，求解银行资产价值和银行资产隐含波动率，从而克服了 Merton 存款保险定价模型无法估计银行资产价值及其波动率的问题；Duan（1994）[5]指出 Ronn（1986）模型中股权收益波动率为常数的假设，并不符合 Merton（1977）存款保险定价模型中股票波动率为随机波动的前提假设，因此 Duan(1994) 将极大似然估计法引入存款保险定价中，以估计银行资产价值及其波动性；Duan（1995）[6]在 Merton（1977）的模型中引入随机利率，明确解释了期限结构对存款保险定价的影响；之后 Duan（1999）[7]开发出一种全新的方法---GARCH 期权定价理论，应用于存款保险定价问题中。

为了解决实际估计中遇到的多种问题，创造更加符合现实条件的模型，后人基于上述存款保险定价的研究，通过改变存款保险合约的特征，如改变保险合约期限、加入制度因素等，使模型更加具有现实意义。Merton（1978）[8]将 Merton（1977）的标准单期模型扩展为多期，将存款保险视为一份无限到期的看跌期权，并在模型中加入了监管、审计成本等因素；Pennacchi (1987)[9-10]考虑了替代性监管政策和银行关闭政策，扩展了政府监管机构在存款保险定价中的有效性，并在 Merton（1978）的基础之上，加入资本比率因子更进一步考虑了无限期限的存款保险合约，将单期模型扩展为多期模型；Allen（1993）[11]通过将存款保险模型作为可赎回看跌期权来修正 Merton 的定价方法。也就是说，存款保险机构在没有到期日的情况下向银行提出一个认沽期权，但在任何时候都有

权终止看跌期权；Kerfriden（1993）[12]在有价付息债券期权定价公式中考虑了银行资产负债表中的期限结构、市场参数因素，进一步测算了存款保险的价格；Dermine（2001）[13]通过分析银行借贷资产的风险特征构建了基于银行借贷风险的费率模型；Kim（2004）[14]假设银行在保险期间内任何时候均可能发生破产，此时存款保险被视为美式期权合约，同时假设银行资产和负债价值的分布均为对数正态分布，在此假设下推出了封闭解。Camara（2020）[15]在具有代表性的三种存款保险制度下分别推导银行价值的封闭解，从而在考虑市场信息的情况下得到精算公平费率。

在国内，早期对存款保险定价的研究大多是对国外经验的介绍和比较分析。例如，林志琦（1985）[16]详细介绍了美国存款保险制度的发展过程。而关于存款保险定价的理论研究从 2002 年才逐渐开始出现，沈福喜（2002）[17]借鉴国外存款保险定价的经验，采用加权最小二乘法对我国的存款增长模型进行了模拟预测，分析认为将基准费率设定为 0.05%至 0.06%比较适合我国银行目前的状况。赖叔懿（2008）[18]建议我国采用根据风险定价的存款保险制度，认为我国在存款保险制度的建设中采取单一的存款保险费率，极易引发逆向选择和道德风险。李敏波（2015）[19]基于 Ronn（1986）的模型，采用数值方法对 16 家上市商业银行所获的政府隐性担保及其蕴含的存款保险基本费率进行了测算，测算结果表明，不同银行风险水平存在差异，所获政府隐性担保亦不同，有力支持了存款保险的差别费率设计思路。

监管宽容是存款保险定价时不可忽略的重要制度因素，故国内许多学者针对此问题亦进行了探究。张金宝（2006）[20]在 Merton（1977）的基础上，构建了加入监管宽容的存款保险价格的偏微分方程边界条件，分析监管宽容对存款保险费率的影响，得出适当的监管宽容能够降低存款保险费率的结论；孙晓琳等（2011）[21]基于 Ronn（1986）的模型，将资本展期与监管宽容合理结合，给出了监管宽容下的资本展期存款保险定价模型；明雷（2019）[22]将监管惩罚和监管宽容同时引入 Merton 基础定价模型，并以银行风险偏好为中介，探讨了监管惩罚、监管宽容程度对不同银行存款保险费率的影响。

国内许多学者也基于 Merton（1977）模型进行了拓展和改进。姜兴坤（2013）[23]在 Merton（1977）的定价模型中，考虑了所得税因素对投保银行的费率影响，以 10 家上市银行为研究对象计算了费率，发现引入所得税后的费率明显减低；罗宏锋（2014）[24]、程孝强（2019）[25]、袁金建（2019）[26]分别考虑了拨备充足率、银行监管资本与破产成本、

债务清偿顺序与时变波动率等因素影响下的存款保险费率。黄虹（2016）[27]通过质疑理性投资者假设，在 Knight 不确定性环境下考察存款保险定价模型，指出在市场不确定状态下，存款保险费率区间会收到不确定性参数的影响。进一步地，程志富等（2016）[28]在 RV 模型基础上，类比存款负债与其他普通债务，利用上市银行债券数据测算了银行存款保险费率和风险中性违约概率，发现银行的存款风险与债券收益波动率呈正相关，说明债券市值可以反映存款风险及其相对应的保费。

此外，除了通常的期权定价方法和期望损失方法之外，部分学者也提出了创新模型。刘海龙（2011）[29]使用银行破产时的期望损失作为存款保险定价的依据，运用极大似然估计方法进行测算，并通过实证分析检验了模型的风险敏感度、实践可行性，建立了基于银行监管资本的客观定价模型。刘鸿伟（2017）[30]则以金融机构系统性风险贡献度作为关键因素衡量存款风险，构建了宏观审慎监管视角下的存款保险定价模型。

尽管使用 Merton 期权定价模型分析存款保险费有其优点，但存在两个关键的缺陷。首先，Merton 的分析并没有区分定期和活期存款，相反，他假定活期存款的到期时间是到下一次银行监管审计的时间长度。这种方法忽略了在当前时刻和下一次监管审计之间随时撤回活期存款的可能性。在 Crouhy（1991）[31]中，活期存款被看作是一系列隔夜存款。例如，100 天的活期存款被视为 100 个隔夜存款。但是，100 天定期存款的违约风险与 100 个隔夜定期存款的风险不同。总之，以前的研究并没有注意到时间和活期存款对存款保险机构施加不同程度的风险这一事实，在实践中也没有注意到这种差异。例如，在美国，存款保险费是按照国内存款总额的固定比例计算的，而不区分时间和活期存款。通过对这两种形式存款的差异性风险的研究，可以更准确地确定存款保险费率，从而加深对存款保险定价的理解。

其次，该模型主要借助银行资产波动率进行反映银行风险，忽视了资本充足率、管理水平、盈利能力和流动性等指标，导致模型的实际应用有一定的片面性。银行关闭可能是由于"破产"或"技术性破产"，即银行虽然有清偿能力，但无法及时获得充足资金或无法以合理成本及时获得充足资金以应对资产增长或支付到期债务。在很多情况下，甚至建议在严重的技术性破产出现时进行早期监管干预和关闭银行（Lastra，1996）[32]。因此，根据这一观点，不仅在银行无力支付所有债务的情况下，而且在银行技术性破产的情况下，也可能触发银行的关闭。如果银行发生流动性问题或银行被证明无力偿债，存款保险机构向储户支付存款，在这两种情况下，银行都很快关闭，其资产被出售，而出售资产可

能会产生清算成本。Mullins（1994）[33]虽然考虑了清算成本来确定公平的保险费率，但未能将银行关闭程序的清算成本纳入存款保险定价模型，因此，从期权定价模型计算的保费将低于保险的公允价值。

通过国内外文献的梳理与回顾，特别是当前研究中存在的两个主要缺陷，文章将聚焦于此，抓住流动性风险这个存款保险定价的核心要素，运用 Gap 期权定价模型给出了存款保险的一般性定价公式。具体而言，文章的边际贡献与创新在于：1、运用 Ronn-Verma 存款保险定价模型，对定期和活期存款采取了不同的处理方式。即，假设两家银行 A 和 B 除负债外其他方面是相同的。银行 A 的负债仅包括到期时间为 T 的定期存款，而银行 B 的负债仅包括可在任何时间撤回的活期存款（0 <t≤T）。为了便于分析，假定这两家银行将在 T 年底解散。文章将定期存款保险作为看跌期权，把活期存款保险视为复合期权，推导出了存款保险的一般定价模型。2、采用 Gap 期权定价模型来处理第二个缺点。Boyle（1994）[34]介绍的 Gap 期权模型允许期权合约的触发价值与其执行价格不同。与普通期权不同，其执行价格既决定了期权的触发条件，也决定了到期时的赔付，而 Gap 期权的触发价值与其行权价格不同。由于忽略清算折扣因子会低估存款保险费，文章试图建立存款保险合同，考虑与随机存款相关的流动性问题，并纳入清算折扣因子。

## 三、模型构建与求解

（一）Merton 期权定价模型与 RV 改进模型

本文以 Merton(1977)的期权定价模型作为基础进行模型构建。该模型源于 Black-Scholes 的期权定价公式，其核心思想是将银行的存款保险看作是银行资产价值的一份看跌期权。Merton 认为，保险人通过担保银行的债务，本质上是对银行资产出售了一份看跌期权，这个看跌期权的执行价格是存款保险合同到期时银行的负债 $L$，标的资产价格是银行的资产价值 $A$。从而，存款保险合同到期后，保险公司的赔付为 $max(0, L - A)$，则存款保险的定价公式为：

$$G(T) = Le^{-rT} \cdot N(-d_2) - A_0 \cdot N(-d_1) \qquad （1）$$

其中，$d_1 = \dfrac{ln(\dfrac{A_0}{L}) + (r + \sigma^2 / 2)T}{\sigma\sqrt{T}}, d_2 = d_1 - \sigma\sqrt{T}, N(\cdot)$ 表示标准正态分布的累计分布函数，$r$ 为无风险利率，$A_0$ 为期初银行的资产价值，$T$ 为

存款保险的期限。

对于存款本金和利息都被担保的存款保险而言，被保险存款将是无风险的，且其现值为：

$$L_0 = Le^{-rT} \qquad （2）$$

记 $g = G(T) / L_0$，则 $g$ 为单位存款的保险价值。根据公式（1）和（2），$g$ 又可以写成：

$$g = N(-d_2) - \frac{1}{d} N(-d_1) \qquad （3）$$

其中，$d = \dfrac{L_0}{A_0}$。

在 Merton 模型基础上，Ronn（1986）通过两个联立方程求解银行资产的价值 $A$ 及其波动性 $\sigma$ 的估计值。第一个方程将可以直接观察到的银行股票市场价值 $E$ 作为对银行资产价值 $A$ 的看涨期权，执行价格等于保险合同到期时银行的债务，建立方程：

$$E = A \cdot N(x) - L \cdot N(x - \sigma\sqrt{T}) \qquad （4）$$

其中，$x = \dfrac{ln(\dfrac{A}{L}) + \sigma^2 \cdot T / 2}{\sigma\sqrt{T}}$。第二个方程将股权波动率 $\sigma_E$ 与银行资产收益率的波动率 $\sigma$ 联系起来：

$$E \cdot \sigma_E = A \cdot (\frac{\partial E}{\partial A}) = A \cdot \sigma \cdot N(x) \qquad （5）$$

联立（4）、（5）可求出银行的资产价值 $A$ 及银行资产收益率的波动率 $\sigma$。将其代入式（3）可求出每单位存款的保险费率。

RV 模型还有一个改进的地方是引入了监管者宽容（regulatory forbearance）。实际中，当银行资产价值 $A$ 小于负债 $L$ 时，监管者会通过直接注入资金或临时暂缓关闭的方式来挽救这一银行。因此，可以合理地假设银行的资产价值 $A$ 有一个限度，这一限度表示为银行负债的比例，即 $\rho L$，$\rho$ 为监管宽容参数，且 $\rho \le 1$。若超过这一限度，即银行的资产价值 $A$ 小于 $\rho L$，则清算银行资产。

在考虑监管者宽容后的定价模型为：

$$E = A \cdot N(x) - \rho L \cdot N(x - \sigma\sqrt{T}) \qquad （6）$$

其中，$x = \dfrac{ln(\dfrac{A}{\rho L}) + \sigma^2 \cdot T / 2}{\sigma\sqrt{T}}$。

同理，通过联立（5）、（6）可求出考虑监管宽容情形下的存款保险费率。

正如文献述评中指出的那样，Merton 模型存在两个关键的缺陷，一是未考虑定期存款与活期存款的区别，二是忽视了在银行资产波动率之外反映银行风险的流动性等指标。因此，文章下面将基于流动性风险扩展 Merton 模型，分定期存款与活期存款两钟情况，并考虑技术性破产因素。

（二）不考虑技术性破产的存款保险定价模型

1. 定期存款的存款保险定价

假设一家银行的负债只包括到期日相同的定期存款。如果存款保险到期后,银行的资产价值小于它所承担的债务,在这种情况下,银行将很快关闭且其资产将被出售,而出售资产可能产生清算成本。假设银行出售资产的清算折扣因子为 $\delta$，则银行资产的清算价值为 $\delta V(T)$，到期时，存款保险公司的赔付为：

$$f(V(T),T) = \begin{cases} L - \delta V(T)\,, & V(T) < L \\ 0\,, & V(T) \geq L \end{cases} \qquad （7）$$

利用 Gap Option 的定价公式，给出定期存款保险的定价公式：

$$\delta \cdot GP(A_0, \frac{1}{\delta}L, L, \sigma, 0, T) = \delta \cdot \left[ \frac{1}{\delta}L_0 \cdot N(-d + \sigma\sqrt{T}) - A_0 \cdot N(-d) \right]$$

$$（8）$$

其中，$d = \dfrac{ln(\dfrac{A_0}{L_0}) + \dfrac{\sigma^2 \cdot T}{2}}{\sigma\sqrt{T}}$，$L_0 = Le^{-rT}$，$N(\cdot)$ 表示标准正态分布的累计

分布函数，$GP(A_0,\frac{1}{\delta}L,L,\sigma,t_0,t_1)$ 表示执行价格为 $\frac{1}{\delta}L$ 且触发价格为 $L$ 的看跌期权的价值。

记 $g = G(T,m)/L_0$，$d = \frac{L_0}{A_0}$，则 $g$ 为单位存款的保险价值：

$$g = \delta\left[\frac{1}{\delta}N(-d+\sigma\sqrt{T}) - \frac{1}{d}N(-d)\right] \quad （9）$$

2. 活期存款的存款保险定价

由于 Merton 模型并没有区分定期存款和活期存款，相反，他假定活期存款的到期时间是到下一次银行监管审计的时间，这种方法忽略了在当前时刻和下一次监管审计之间随时撤回活期存款的可能性。

与定期存款不同，活期存款可在任何时间点提取。为了便于处理，假设一家银行的负债只包括活期存款，并且一旦需要资金，存款将被全部提取。此外，假定存在 $m$ 个离散的取款点，并且任意两点之间的间隔相同并等于 $T/m$，$\pi_i$ 表示在 $i$ 时刻取款的概率。这样，活期存款的保险成本可以被确定为所有可能的取款点预期现金流量折现的现值的总和：

$$G(T,m) = \sum_{i=1}^{m}\left\{e^{-rt_i}\hat{E}\left[\pi_i max(L_i - \delta A_i, 0_i)\right]\right\} = \sum_{i=1}^{m}\left\{\pi_i e^{-rt_i}\hat{E}\left[max(L_i - \delta A_i, 0_i)\right]\right\} \quad （10）$$

$$= \sum_{i=1}^{m}\left[\pi_i G(t_i)\right] = E\left[G(\tilde{t})\right]$$

$$G(t_i) = \delta \cdot GP(A_0, \frac{1}{\delta}L_i, L_i, \sigma, 0, t_i) = \delta \cdot \left[\frac{1}{\delta}L_0 \cdot N(-d_i + \sigma\sqrt{t_i}) - A_0 \cdot N(-d_i)\right]$$

$$（11）$$

其中，$d_i = \dfrac{ln(\dfrac{A_0}{L_0}) + \dfrac{\sigma^2 \cdot t_i}{2}}{\sigma\sqrt{t_i}}$，$t_i = \dfrac{iT}{m}$，$\hat{E}$ 表示风险中性定价，$r$ 为无风险利率，$\pi_i$ 表示在 $i$ 时刻取款的概率。记 $g = G(T,m)/L_0$，$d = \dfrac{L_0}{A_0}$，则 $g$

为单位存款的保险价值：

$$g = \sum_{i=1}^{m} \pi_i \cdot \delta \left[ \frac{1}{\delta} N(-d_i + \sigma\sqrt{t_i}) - \frac{1}{d} N(-d_i) \right] \qquad （12）$$

为了便于处理，假设取款点服从离散的均匀分布，即：

$$\pi_1 = \pi_2 = \cdots = \pi_m = \frac{1}{m} \qquad （13）$$

这样，（12）又可以写成：

$$g = \sum_{i=1}^{m} \frac{1}{m} \cdot \delta \left[ \frac{1}{\delta} N(-d_i + \sigma\sqrt{t_i}) - \frac{1}{d} N(-d_i) \right] \qquad （14）$$

（三）考虑技术性破产的存款保险定价模型

1. 定期存款的存款保险定价

假设一家银行的负债只包括到期日相同的定期存款，初始时刻的资产和负债分别为 $A_0$ 和 $L_0$。

首先，银行资产价值的变化 $\Delta A$ 受两个因素的影响：自主因素 $\Delta A_a$ 和诱导因素 $\Delta A_i$，且假设这两个因素的影响是相互独立的。其中，自主因素部分是银行资产价值变化的随机过程，服从几何布朗运动，即 $dA_a = \mu_a A dt + \sigma_A A dz_A$，$\mu_a$ 为银行资产的瞬时收益，$\sigma_A$ 为资产收益的波动率，$z$ 遵循标准的维纳过程。资产价值变动的诱导因素部分由银行负债端的变动引起，存款价值的变化会导致资产价值的均等变化，即 $\Delta A_i = \Delta L$。

其次，令 $W-1$ 表示存款的百分比变化，即 $\Delta L = (W-1)L_0$。如果 $W=0$，那么保险合同到期时，全部存款都被提取；如果 $W=1$，那么保险合同到期时，没有存款被提取；如果 $W>1$，那么保险合同到期时，存款有所增加。由于存款的变化有下界但没有上界，即 $-L_0 \leq \Delta D < \infty$，因此假设 $W$ 服从对数正态分布。

最后，假设银行净提款高于银行储备金加银行市场信贷额度时发生

流动性风险。银行储备金表示为初始时刻资产价值 $A_0$ 的比例，即 $\alpha A_0$，银行市场信贷额度表示为初始净值的比例，即 $\beta(A_0 - D_0)$。如果净提款超过银行的流动资产，即 $\Delta L < -\alpha A_0$，那么银行就需要向资金市场借款；如果净提款高于银行储备金加银行市场信贷额度，即 $\Delta L < -[\alpha A_0 + \beta(A_0 - L_0)]$，那么银行不能满足提款人的需要，需要被立即关闭。

当银行发生流动性问题或被证明无力偿债，在这两种情况下，银行都将很快关闭且其资产将被出售，而出售资产可能产生清算成本。假设银行出售资产的清算折扣因子为 $\delta$，由于资产价值变动的诱导因素部分不需要参与清算折扣，存款保险合同到期时银行的清算价值为 $\delta(A_0 + \Delta A_a) + \Delta A_i$，银行的负债为初始负债和应计利息之和，以及存款的变化，即 $e^{rt} L_0 + \Delta L$。保险人的赔付等于银行资产的清算价值和到期负债的差额部分，而 $\Delta A_i = \Delta L$，故保险人的赔付可表示为 $L - \delta A^*$，其中 $A^* = \rho(A_0 + \Delta A_a)$，$L = e^{rt} D_0$。

如上所述，有两种情况可能触发存款保险合同的执行。在第一种情况下，银行的资产价值小于它所承担的债务，即 $A^* < L$，银行资产的清算价值和到期负债的差额部分（即储户的损失）将由保险公司进行赔付，即 $L - \delta A^*$。在第二情况下，银行的资产价值小于它所承担的债务，但是发生流动性风险，即

$$A^* \geq L \ and \ \Delta L < -[\alpha A_0 + \beta(A_0 - L_0)]$$

由于银行具有偿付能力，其资产的清算价值可能足够支付它所承担的债务，因此在这种情况下，保险公司的赔付为 $max(L - \delta A^*, 0)$。

从而到期时，存款保险公司的赔付为：

$$f(V(T),\mathrm{T}) = \begin{cases} L - \delta A^*, & A^* < L \\ max(L - \delta A^*, 0), & A^* \geq L \ and \ \Delta L < -[\alpha A_0 + \beta(A_0 - L_0)] \end{cases}$$

（15）

即存款保险的定价为：

$$p = e^{-rt} \left\{ \begin{array}{l} prob\{A^* < L\} \cdot E\left[L - \delta A^* \mid A^* < L\right] \\ + prob\{A^* \geq L \ and \ \Delta L < -[\alpha A_0 + \beta(A_0 - L_0)]\} \\ \cdot E\left[Max\{L - \delta A^*, 0\} \mid A^* \geq L \ and \ \Delta L < -[\alpha A_0 + \beta(A_0 - L_0)]\right] \end{array} \right\}$$

（16）

上式第一部分可表示为：

$$\delta \cdot GP(A_0, \frac{1}{\delta}L, L, \sigma, 0, T) = \delta \cdot \left[\frac{1}{\delta}L_0 \cdot N(-d + \sigma\sqrt{T}) - A_0 \cdot N(-d)\right]$$

（17）

其中，$d = \dfrac{ln(\dfrac{A_0}{L_0}) + \dfrac{\sigma^2 \cdot T}{2}}{\sigma\sqrt{T}}$，$L_0 = Le^{-rT}$，$A^* = \rho(A_0 + \Delta A_a)$，$L = e^{rt}D_0$。

上式第二部分可表示为：

$$e^{-rt} \cdot prob\{A^* \geq L\} \cdot prob\{\Delta L < -[\alpha A_0 + \beta(A_0 - L_0)]\}$$
$$\cdot E\left[Max\{L - \delta A^*, 0\} \mid A^* \geq L \ and \ \Delta L < -[\alpha A_0 + \beta(A_0 - L_0)]\right]$$
$$= \delta \cdot \wedge(1 - (\alpha^* + \beta^*))$$
$$\cdot \left[BP\left(A_0, \frac{1}{\delta}L, \sigma, 0, T\right) - GP\left(A_0, \frac{1}{\delta}L, L, \sigma, 0, T\right)\right]$$

（18）

其中，$\alpha^* = \alpha\dfrac{A_0}{D_0}$，$\beta^* = \beta(\dfrac{A_0 - D_0}{D_0})$，$\wedge(\cdot)$ 表示存款和资产的累积对

数正态密度函数。从而，存款保险的定价公式为：

$$P = (1-\wedge) \cdot \delta \cdot GP(A_0, \frac{1}{\delta}L, L, \sigma, 0, T) + \wedge \cdot \delta \cdot BP\left(A_0, \frac{1}{\delta}L, \sigma, 0, T\right)$$

（19）

2. 活期存款的存款保险定价

同上，考虑一家银行的负债只包括活期存款。为了便于处理，假设一家银行的负债只包括活期存款，并且一旦需要资金，存款将被全部提取。此外，假定存在 $m$ 个离散的取款点，并且任意两点之间的间隔相同并等于 $T/m$，$\pi_i$ 表示在 $i$ 时刻取款的概率。这样，活期存款的保险成本可以被确定为所有可能的取款点预期现金流量折现的现值的总和：

$$G(T,m) = \sum_{i=1}^{m} \pi_i e^{-rt_i} \left\{ \begin{array}{l} prob\left\{A_i^* < L_i\right\} \cdot E\left[L_i - \delta A_i^* \mid A_i^* < L_i\right] \\ + prob\left\{A_i^* \geq L_i \text{ and } \Delta L < -\left[\alpha A_0 + \beta(A_0 - L_0)\right]\right\} \\ \cdot E\left[Max\left\{L_i - \delta A_i^*, 0\right\} \mid A_i^* \geq L_i \text{ and } \Delta L_i < -\left[\alpha A_0 + \beta(A_0 - L_0)\right]\right] \end{array} \right\}$$

$$= \sum_{i=1}^{m} \left[\pi_i G(t_i)\right]$$

$$= E\left[G(\tilde{t})\right]$$

（20）

$$G(t_i) = (1-\wedge) \cdot \delta \cdot GP(A_0, \frac{1}{\delta}L, L, \sigma, 0, t_i) + \wedge \cdot \delta \cdot BP\left(A_0, \frac{1}{\delta}L, \sigma, 0, t_i\right)$$

$$= (1-\wedge) \cdot \delta \cdot \left[\frac{1}{\delta}L_0 \cdot N(-d_i + \sigma\sqrt{T}) - A_0 \cdot N(-d_i)\right]$$

$$+ \wedge \cdot \delta \cdot \left[L_0 \cdot N(-d_i + \sigma\sqrt{T}) - A_0 \cdot N(-d_i)\right]$$

（21）

其中，$d_i = \dfrac{ln(\frac{A_0}{L_0}) + \frac{\sigma^2 \cdot t_i}{2}}{\sigma\sqrt{t_i}}$，$t_i = \dfrac{iT}{m}$，$\hat{E}$ 表示风险中性定价，$r$ 为无风险

利率，$\pi_i$ 表示在 $i$ 时刻取款的概率。记 $g = G(T,m) / L_0$，$d = \dfrac{L_0}{A_0}$，则 $g$

为单位存款的保险价值：

$$g = \sum_{i=1}^{m} \pi_i \cdot \left\{ \begin{array}{l} (1-\wedge) \cdot \delta \cdot \left[ \dfrac{1}{\delta} \cdot N(-d_i + \sigma\sqrt{t_i}) - \dfrac{1}{d} \cdot N(-d_i) \right] \\[3mm] \quad + \wedge \cdot \delta \cdot \left[ N(-d_i + \sigma\sqrt{t_i}) - \dfrac{1}{d} \cdot N(-d_i) \right] \end{array} \right\}$$

（22）

为了便于处理，假设取款点服从离散的均匀分布，即：

$$\pi_1 = \pi_2 = \cdots = \pi_m = \frac{1}{m} ,$$

（23）

这样，（22）又可以写成：

$$g = \sum_{i=1}^{m} \frac{1}{m} \cdot \left\{ \begin{array}{l} (1-\wedge) \cdot \delta \cdot \left[ \dfrac{1}{\delta} \cdot N(-d_i + \sigma\sqrt{t_i}) - \dfrac{1}{d} \cdot N(-d_i) \right] \\[3mm] \quad + \wedge \cdot \delta \cdot \left[ N(-d_i + \sigma\sqrt{t_i}) - \dfrac{1}{d} \cdot N(-d_i) \right] \end{array} \right\} .$$

（24）

## 四、实证测算

（一）RV 模型的实证测算与分析

选择我国 25 家上市银行 2017 年的相关数据来进行 RV 模型的实证测算，包括国有银行、股份制商业银行、城市商业银行以及农村商业银行 [1]。银行的股票市场价值 $E$，可通过银行股票每日的的收盘价与总股本乘积的算数平均求得；银行的总负债 $L$，可通过银行每年的年度报告得到；两期监管审计的时间间隔 $T$，假设为一年。测算过程如下：

首先，通过 Wind 数据库得到 2017 年 25 家上市银行的日收盘价和总股本，并计算银行的股票市场价值 $E$；其次，从年报中得到各银行的总负债 $L$，监管宽容参数 $\rho$ 假设为 0.95；最后，根据股票的日价格数据，

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[1] 我国的《存款保险条例》于 2015 年 5 月 1 日颁布，2016 年是制度实施的衔接期，因此选用 2017 年的数据进行实证测算才能更好地反映《条例》的实施情况。

求得股价收益率的波动率 $\sigma_E$：

$$\sigma_E = s\sqrt{t},\ s = \sqrt{\frac{1}{t-1}\sum_{i=1}^{n}(u_i - \overline{u})^2},\ u_i = ln(\frac{S_i}{S_{i-1}}) \qquad （25）$$

其中，$t$ 为我国股票市场每年交易天数；$S_i$ 为第 $i$ 个观测日的收盘价。

由上述步骤可得 25 家上市银行的股票市场价值 $E$、总负债 $L$ 和股价收益率的波动率 $\sigma_E$，见下表 1。

表 1 各银行股票市场价值、股价收益率波动率和负债

| 银行 | 2017 年 | | |
| | $E$（百万元） | $L$（百万元） | $\sigma_E$ |
| --- | --- | --- | --- |
| 平安银行 | 179792.5033 | 3026420.0000 | 0.2532864 |
| 宁波银行 | 76827.5577 | 974836.4700 | 0.3383263 |
| 浦发银行 | 360396.2362 | 5706255.0000 | 0.2260307 |
| 华夏银行 | 119371.5051 | 2339429.0000 | 0.2509158 |
| 民生银行 | 308201.0118 | 5512274.0000 | 0.1405802 |
| 招商银行 | 589422.9341 | 5814246.0000 | 0.2525883 |
| 南京银行 | 68163.8771 | 1072952.0000 | 0.4227466 |
| 兴业银行 | 343404.5409 | 5994090.0000 | 0.1406302 |
| 北京银行 | 140678.5352 | 2153091.0000 | 0.2463732 |
| 农业银行 | 1144513.237 | 19623985.0000 | 0.1436988 |
| 交通银行 | 460154.9853 | 7770759.0000 | 0.1399246 |
| 工商银行 | 1902464.465 | 23945987.0000 | 0.1662921 |
| 光大银行 | 190604.6429 | 3782807.0000 | 0.1312719 |
| 建设银行 | 1609312.465 | 20328556.0000 | 0.1724144 |
| 中国银行 | 1120012.83 | 17890745.0000 | 0.1387867 |
| 中信银行 | 315200.2557 | 5265258.0000 | 0.1750660 |
| 江阴银行 | 22958.51243 | 100048.8830 | 0.5608295 |
| 张家港行 | 27060.0076 | 94783.6898 | 0.6263355 |
| 无锡银行 | 21913.1125 | 127772.9600 | 0.5146919 |
| 江苏银行 | 104009.8169 | 1657723.2030 | 0.1848741 |

| | | | |
|---|---|---|---|
| 杭州银行 | 53113.90135 | 781507.7700 | 0.4442144 |
| 常熟银行 | 23296.73979 | 134715.9960 | 0.4189654 |
| 上海银行 | 139671.4063 | 1660325.5350 | 0.3554466 |
| 贵阳银行 | 35899.48366 | 438476.3450 | 0.2263343 |
| 吴江银行 | 16795.89732 | 86798.0810 | 0.5356732 |

联立方程式（5）和方程式（6），并代入 $E$ 、$\rho L$ 和 $\sigma_E$ 的值，求得商业银行资产价值 $A$ 和资产波动率 $\sigma$ ；将上一步求得的资产价值 $A$ 和资产波动率 $\sigma$ 带入式（3），便可得到考虑监管宽容的情况下各银行的存款保险费率 $g$ （见下表2）。

表 2 考虑监管宽容的 RV 模型求出的存款保险费率

| 银行 | $A$ （百万元） | $\sigma$ | $d = \dfrac{L}{A}$ | $g(\%)$ |
|---|---|---|---|---|
| 平安银行 | 3054891.2713 | 0.0149073 | 0.9906801 | 0.24127% |
| 宁波银行 | 1002914.3561 | 0.0259437 | 0.9720037 | 0.18253% |
| 浦发银行 | 5781338.4470 | 0.0140903 | 0.9870128 | 0.13534% |
| 华夏银行 | 2341828.9160 | 0.0127904 | 0.9989752 | 0.46087% |
| 民生银行 | 5544861.3118 | 0.0078139 | 0.9941230 | 0.10203% |
| 招商银行 | 6112956.1492 | 0.0243554 | 0.9511349 | 0.01817% |
| 南京银行 | 1087396.7521 | 0.0266986 | 0.9867162 | 0.53088% |
| 兴业银行 | 6037790.0409 | 0.0079985 | 0.9927622 | 0.07942% |
| 北京银行 | 2186114.8862 | 0.0158546 | 0.9848938 | 0.14353% |
| 农业银行 | 19787298.9870 | 0.0083117 | 0.9917465 | 0.06992% |
| 交通银行 | 7842376.0353 | 0.0082101 | 0.9908679 | 0.05468% |
| 工商银行 | 24651152.1150 | 0.0128337 | 0.9713942 | 0.00533% |
| 光大银行 | 3784271.2929 | 0.0066118 | 0.9996131 | 0.24492% |
| 建设银行 | 20921440.6650 | 0.0132624 | 0.9716614 | 0.00721% |
| 中国银行 | 18116220.5800 | 0.0085803 | 0.9875539 | 0.02773% |
| 中信银行 | 5317195.3557 | 0.0103778 | 0.9902322 | 0.09622% |
| 江阴银行 | 117882.2259 | 0.1118931 | 0.8487190 | 0.38409% |
| 张家港行 | 116844.7765 | 0.1503614 | 0.8111932 | 0.62227% |
| 无锡银行 | 143221.9105 | 0.0801677 | 0.8921328 | 0.29512% |
| 江苏银行 | 1678846.8597 | 0.0114535 | 0.9874178 | 0.07823% |
| 杭州银行 | 795462.5075 | 0.0299678 | 0.9824571 | 0.51778% |

| 常熟银行 | 151262.8115 | 0.0648195 | 0.8906088 | 0.10120% |
| 上海银行 | 1716957.3235 | 0.0289619 | 0.9670162 | 0.17994% |
| 贵阳银行 | 452452.0082 | 0.0179584 | 0.9691113 | 0.02971% |
| 吴江银行 | 99181.7768 | 0.0926290 | 0.8751414 | 0.33213% |

根据表 2 可以得到以下结论：

(1) 存款保险费率呈现出较为明显的分层现象，国有银行与其余上市银行之间费率差距较大。其中，国有银行、股份制商业银行、城市商业银行以及农村商业银行的平均费率分别为 0.03297%、0.17228%、0.23752% 以及 0.34696%。

(2) 在国有银行、股份制商业银行、城市商业银行以及农村商业银行各层之间也产生分层现象。其中，国有银行中农业银行和交通银行的费率较高，约为其它国有银行的 10 倍；股份制商业银行中招商银行和兴业银行的费率与国有银行相近，而其它股份制商业银行费率约高出 10 倍以上；城市商业银行中杭州银行和南京银行的费率较高，而其他银行的费率与股份制商业银行差不多；农村商业银行分层现象较弱，整体费率较其他银行高。

然而，由于 RV 模型和 Merton 模型一样，假设投保银行风险状况的信息完全由银行资产价格的标准差来度量，并不能准确度量投保银行的风险状况，而且也隐含了投保银行的风险是系统外生的假设。对于证券投资者而言，在不考虑红利的情况下，资产价格波动对其收益的影响是显著的，因此资产价格的标准差作为风险度量指标，能够反映出资产的风险状况。但对一家投保银行而言，其风险不仅反映在资产价格的波动中，银行的资本充足率、流动性比率、存贷款比率等指标更能全面地反映银行的风险水平。

（二）基于流动性风险的存款保险定价模型的实证测算

1. 不考虑技术性破产的存款保险定价模型

银行的股票市场价值 $E$、总负债 $L$、股价收益率的波动率 $\sigma_E$、银行资产价值 $A$ 和资产波动率 $\sigma$ 的计算方法同上节 RV 模型的实证测算。将求得的资产价值 $A$ 和资产波动率 $\sigma$ 带入式（9）和式（14），便可得到不考虑技术性破产情况下各银行的存款保险费率 $g$。

首先，考虑清算折扣因子 $\delta$ 对存款保险费率的影响。分别取 $\delta=1$（即 RV 模型求得的存款保险费率）、$\delta=0.99$、$\delta=0.95$、$\delta=0.9$，代入式（9）计算各银行的存款保险费率（见下表 3）。

表 3 不同清算折扣因子下的存款保费率

| 银行 | $\delta = 1$ | $\delta = 0.99$ | $\delta = 0.95$ | $\delta = 0.9$ |
|---|---|---|---|---|
| 平安银行 | 0.24127% | 0.50626% | 1.56624% | 2.89122% |
| 宁波银行 | 0.18253% | 0.32043% | 0.87204% | 1.56156% |
| 浦发银行 | 0.13534% | 0.31259% | 1.02158% | 1.90782% |
| 华夏银行 | 0.46087% | 0.92685% | 2.79081% | 5.12075% |
| 民生银行 | 0.10203% | 0.32750% | 1.22940% | 2.35678% |
| 招商银行 | 0.01817% | 0.03843% | 0.11944% | 0.22071% |
| 南京银行 | 0.53088% | 0.83851% | 2.06904% | 3.60720% |
| 兴业银行 | 0.07942% | 0.26157% | 0.99019% | 1.90097% |
| 北京银行 | 0.14353% | 0.31261% | 0.98892% | 1.83431% |
| 农业银行 | 0.06992% | 0.22958% | 0.86824% | 1.66656% |
| 交通银行 | 0.05468% | 0.18692% | 0.71590% | 1.37712% |
| 工商银行 | 0.00533% | 0.01735% | 0.06539% | 0.12545% |
| 光大银行 | 0.24492% | 0.72045% | 2.62257% | 5.00021% |
| 建设银行 | 0.00721% | 0.02248% | 0.08359% | 0.15997% |
| 中国银行 | 0.02773% | 0.10023% | 0.39027% | 0.75282% |
| 中信银行 | 0.09622% | 0.26870% | 0.95861% | 1.82100% |
| 江阴银行 | 0.38409% | 0.45953% | 0.76125% | 1.13841% |
| 张家港行 | 0.62227% | 0.71006% | 1.06121% | 1.50014% |
| 无锡银行 | 0.29512% | 0.37539% | 0.69650% | 1.09788% |
| 江苏银行 | 0.07823% | 0.21316% | 0.75288% | 1.42753% |
| 杭州银行 | 0.51778% | 0.79505% | 1.90410% | 3.29042% |
| 常熟银行 | 0.10120% | 0.13983% | 0.29435% | 0.48749% |
| 上海银行 | 0.17994% | 0.30454% | 0.80292% | 1.42591% |
| 贵阳银行 | 0.02971% | 0.07051% | 0.23368% | 0.43765% |
| 吴江银行 | 0.33213% | 0.41054% | 0.72418% | 1.11624% |

根据表 3，可以得到以下结论：

(1) 银行面临的流动性风险对其存款保险费率有显著影响。由于 Merton(1977)没有考虑到银行的流动性风险，而考虑银行出售资产的清算成本后，各银行的存款保险费率都有显著增加，说明 Merton(1977)低估了存款保险费率；

(2) 银行的存款保险费率与其面临的流动性风险成正比。随着清算折扣因子的减小，即银行资产的清算价值减小，存款保险费率呈上升趋

势，说明银行资产的流动性风险越大，其存款保险费率越高。结果表明，与不考虑银行资产的清算价值相比，$\delta = 0.99$ 时的存款保险费率约扩大了 2 倍，而 $\delta = 0.95$ 和 $\delta = 0.9$ 时分别扩大了 7 倍和 14 倍，清算折扣因子会显著影响保险费率，因而在对各上市银行的存款保险定价时，应该选取一个合适的清算折扣因子，采取最优的关闭政策，如果商业银行的流动性风险较小，即风险防范能力已经很强，还要求其与风险较高的商业银行一样缴纳较高的存款保险费率，则对该银行是不公平的。

以 $\delta = 0.95$ 为例，可以得到国有银行、股份制商业银行、城市商业银行以及农村商业银行的平均费率分别为 0.42468%、1.41236%、1.08908%以及 0.7075%，与 RV 模型的实证测算结果相比，国有银行、股份制商业银行和城市商业银行的存款保险费率扩大了 10 倍左右，而农村商业银行的平均费率仅扩大了 2 倍多，说明国有银行、股份制商业银行和城市商业银行面临的流动性风险较农村商业银行高。根据银监会2017 年第四季度的统计数据，国有银行、股份制商业银行和城市商业银行的不良贷款率分别为 1.53%、1.71%以及 1.52%，而农村商业银行仅为0.53%；农村商业银行的流动性比例为 98.17%，远远高于国有银行、股份制商业银行和城市商业银行的 48.1%、50.78%以及 51.48%。

其次，考虑活期存款的取款点个数对存款保险费率的影响。实际上，由于活期存款可以随时取回，$m$ 是无限的，$m$ 的值越高，则结果越接近于活期存款的实际保险成本。为了便于处理，取 $\delta = 0.95$，然后分别取 $m = 1$（即表 3 第 4 列）、$m = 5$、$m = 10$、$m = 20$、$m = 30$，代入式（14）计算各银行的存款保险费率（见表 4）。

表 4 不同取款点个数下求出的存款保费率表

| 银行 | $m = 1$ | $m = 5$ | $m = 10$ | $m = 30$ | $m = 50$ | $m = 100$ |
|---|---|---|---|---|---|---|
| 平安银行 | 1.56624% | 1.09104% | 1.00892% | 0.95564% | 0.94516% | 0.93731% |
| 宁波银行 | 0.87204% | 0.46086% | 0.41499% | 0.38527% | 0.37941% | 0.37503% |
| 浦发银行 | 1.02158% | 0.59983% | 0.54600% | 0.51141% | 0.50456% | 0.49943% |
| 华夏银行 | 2.79081% | 2.59622% | 2.55452% | 2.51569% | 2.50489% | 2.49497% |
| 民生银行 | 1.22940% | 0.80669% | 0.74144% | 0.69990% | 0.69167% | 0.68550% |
| 招商银行 | 0.11944% | 0.04164% | 0.03489% | 0.03067% | 0.02986% | 0.02925% |
| 南京银行 | 2.06904% | 1.51563% | 1.41242% | 1.34151% | 1.32768% | 1.31731% |
| 兴业银行 | 0.99019% | 0.59393% | 0.54174% | 0.50826% | 0.50162% | 0.49665% |
| 北京银行 | 0.98892% | 0.56881% | 0.51677% | 0.48325% | 0.47661% | 0.47164% |
| 农业银行 | 0.86824% | 0.49515% | 0.44959% | 0.42015% | 0.41432% | 0.40996% |
| 交通银行 | 0.71590% | 0.38260% | 0.34509% | 0.32073% | 0.31591% | 0.31232% |

| | | | | | | |
|---|---|---|---|---|---|---|
| 工商银行 | 0.06539% | 0.02137% | 0.01761% | 0.01528% | 0.01483% | 0.01450% |
| 光大银行 | 2.62257% | 2.50659% | 2.48007% | 2.45429% | 2.44679% | 2.43962% |
| 建设银行 | 0.08359% | 0.02819% | 0.02343% | 0.02047% | 0.01990% | 0.01947% |
| 中国银行 | 0.39027% | 0.17593% | 0.15505% | 0.14162% | 0.13899% | 0.13702% |
| 中信银行 | 0.95861% | 0.56058% | 0.51014% | 0.47768% | 0.47125% | 0.46644% |
| 江阴银行 | 0.76125% | 0.32135% | 0.27997% | 0.25359% | 0.24844% | 0.24460% |
| 张家港行 | 1.06121% | 0.45678% | 0.39927% | 0.36256% | 0.35537% | 0.35002% |
| 无锡银行 | 0.69650% | 0.30311% | 0.26549% | 0.24143% | 0.23672% | 0.23320% |
| 江苏银行 | 0.75288% | 0.40300% | 0.36352% | 0.33790% | 0.33284% | 0.32905% |
| 杭州银行 | 1.90410% | 1.31921% | 1.22012% | 1.15516% | 1.14241% | 1.13287% |
| 常熟银行 | 0.29435% | 0.11127% | 0.09492% | 0.08462% | 0.08262% | 0.08113% |
| 上海银行 | 0.80292% | 0.40953% | 0.36721% | 0.33979% | 0.33438% | 0.33035% |
| 贵阳银行 | 0.23368% | 0.09213% | 0.07928% | 0.07115% | 0.06957% | 0.06839% |
| 吴江银行 | 0.72418% | 0.31130% | 0.27208% | 0.24703% | 0.24213% | 0.23848% |

根据表 4，可以得到以下结论：

(1) 定期存款的保险费远远高于活期存款的保险费。表明活期存款的风险水平比定期存款的风险水平低，因此存款保险费率低于定期存款。这一发现与银行在借出活期存款收益方面更加保守的事实是一致的，因为他们必须随时准备提取活期存款。

(2) 取款点的个数 $m$ 与存款保险费率呈负相关。随着 $m$ 的增加，存款保险费率呈下降趋势，$m$ 的值越高，则模拟结果越接近于活期存款的实际存款保险费率。

综上，我国存保起步时的费率水平大概在万分之一到万分之二，远低于绝大多数国家存保起步时的水平以及现行水平，而上述实证测算结果也表明我国存保起步时的费率水平没有完全体现各银行的风险水平，因此我国急需一个合理的费率厘定及费率结构。

2. 考虑技术性破产的存款保险定价模型

银行的股票市场价值 $E$、总负债 $L$、股价收益率的波动率 $\sigma_E$、银行资产价值 $A$ 和资产波动率 $\sigma$ 的计算方法同上节 RV 模型的实证测算。假设存款的变化 $W$ 遵循对数正态分布，为了简单起见，假设 $\mu_W = 0$，

$\sigma_W = 0.5$。同时，假设 $\delta = 0.95$。

首先，考虑银行储备金系数 $\alpha$ 对存款保险费率的影响。为了便于处理，固定 $\beta = 0.8$，分别取 $\alpha = 0.02$、$\alpha = 0.05$、$\alpha = 0.1$，代入式（19）计算各银行的存款保险费率（见表 5）。

表 5 不同储备金系数下求出的存款保费率表

| 银行 | $\delta = 0.95$ | $\beta = 0.8$ | | |
|---|---|---|---|---|
| | | $\alpha = 0.02$ | $\alpha = 0.05$ | $\alpha = 0.1$ |
| 平安银行 | 1.56624% | 1.56701% | 1.56643% | 1.56625% |
| 宁波银行 | 0.87204% | 0.87255% | 0.87217% | 0.87205% |
| 浦发银行 | 1.02158% | 1.02241% | 1.02178% | 1.02159% |
| 华夏银行 | 2.79081% | 2.79145% | 2.79096% | 2.79081% |
| 民生银行 | 1.22940% | 1.23037% | 1.22964% | 1.22941% |
| 招商银行 | 0.11944% | 0.11972% | 0.11951% | 0.11945% |
| 南京银行 | 2.06904% | 2.06957% | 2.06917% | 2.06904% |
| 兴业银行 | 0.99019% | 0.99120% | 0.99044% | 0.99020% |
| 北京银行 | 0.98892% | 0.98969% | 0.98911% | 0.98892% |
| 农业银行 | 0.86824% | 0.86925% | 0.86849% | 0.86825% |
| 交通银行 | 0.71590% | 0.71693% | 0.71615% | 0.71590% |
| 工商银行 | 0.06539% | 0.06605% | 0.06555% | 0.06540% |
| 光大银行 | 2.62257% | 2.62328% | 2.62274% | 2.62257% |
| 建设银行 | 0.08359% | 0.08424% | 0.08375% | 0.08359% |
| 中国银行 | 0.39027% | 0.39130% | 0.39052% | 0.39028% |
| 中信银行 | 0.95861% | 0.95955% | 0.95884% | 0.95862% |
| 江阴银行 | 0.76125% | 0.76132% | 0.76127% | 0.76125% |
| 张家港行 | 1.06121% | 1.06126% | 1.06122% | 1.06121% |
| 无锡银行 | 0.69650% | 0.69660% | 0.69652% | 0.69650% |
| 江苏银行 | 0.75288% | 0.75380% | 0.75311% | 0.75289% |
| 杭州银行 | 1.90410% | 1.90458% | 1.90422% | 1.90410% |
| 常熟银行 | 0.29435% | 0.29443% | 0.29437% | 0.29435% |
| 上海银行 | 0.80292% | 0.80336% | 0.80303% | 0.80293% |
| 贵阳银行 | 0.23368% | 0.23424% | 0.23382% | 0.23369% |
| 吴江银行 | 0.72418% | 0.72426% | 0.72420% | 0.72418% |

注：第二列表示不考虑技术性破产的存款保险费率。

根据表 5，可以得到以下结论：

(1) 技术性破产对定期存款的保险费率有一定影响。考虑银行的技术性破产，即银行发生流动性问题时将关闭银行，存款保险费率有所上升，但上升幅度不大。

(2) 银行储备金系数与存款保险费率负相关。随着银行储备金系数的提高，存款保险费率呈下降趋势。这一趋势与拥有大量储备资产的银行出现流动性问题的概率较小的事实是一致的。

银行储备金即银行需要提留一定比例的存款，以保证储户提款，其余的存款才能用于放贷或投资。中央银行通过调整存款准备金率，可以影响金融机构的信贷扩张能力，从而间接调控货币供应量。表 5 结果表明，银行储备金系数越高，进一步增加银行储备金系数对存款保险费率的边际效应递减，因此中央银行应对各银行设置一个合理的存款准备金率。对于风险较高的银行，通过调高银行准备金，可以减少银行贷款，迫使各银行提高自身抗风险水平，促进商业银行提高资金管理水平，优化资产结构；而对于风险较低的银行，调低银行准备金利率，增加存款利息对银行本身的压力，可以迫使其中部分存款流出银行，刺激投资增长。

其次，考虑银行信贷额度系数 $\beta$ 对存款保险费率的影响。为了便于处理，固定 $\alpha = 0.05$，分别取 $\beta = 0.8$、$\beta = 0.9$、$\beta = 0.95$，代入式（19）计算各银行的存款保险费率（见表 6）。

<center>表 6 不同信贷额度系数下求出的存款保费率表</center>

| 银行 | $\delta = 0.95$ | $\alpha = 0.05$ | | |
| --- | --- | --- | --- | --- |
| | | $\beta = 0.8$ | $\beta = 0.9$ | $\beta = 0.95$ |
| 平安银行 | 1.56624% | 1.57331% | 1.56643% | 1.56624% |
| 宁波银行 | 0.87204% | 0.87665% | 0.87217% | 0.87204% |
| 浦发银行 | 1.02158% | 1.02917% | 1.02178% | 1.02158% |
| 华夏银行 | 2.79081% | 2.79668% | 2.79096% | 2.79081% |
| 民生银行 | 1.22940% | 1.23832% | 1.22964% | 1.22940% |
| 招商银行 | 0.11944% | 0.12198% | 0.11951% | 0.11944% |
| 南京银行 | 2.06904% | 2.07389% | 2.06917% | 2.06904% |
| 兴业银行 | 0.99019% | 0.99942% | 0.99044% | 0.99019% |
| 北京银行 | 0.98892% | 0.99604% | 0.98911% | 0.98892% |
| 农业银行 | 0.86824% | 0.87753% | 0.86849% | 0.86824% |

| 交通银行 | 0.71590% | 0.72537% | 0.71615% | 0.71590% |
| 工商银行 | 0.06539% | 0.07139% | 0.06555% | 0.06539% |
| 光大银行 | 2.62257% | 2.62908% | 2.62274% | 2.62257% |
| 建设银行 | 0.08359% | 0.08962% | 0.08375% | 0.08359% |
| 中国银行 | 0.39027% | 0.39976% | 0.39052% | 0.39027% |
| 中信银行 | 0.95861% | 0.96724% | 0.95884% | 0.95861% |
| 江阴银行 | 0.76125% | 0.76184% | 0.76127% | 0.76125% |
| 张家港行 | 1.06121% | 1.06167% | 1.06122% | 1.06121% |
| 无锡银行 | 0.69650% | 0.69740% | 0.69652% | 0.69650% |
| 江苏银行 | 0.75288% | 0.76132% | 0.75311% | 0.75288% |
| 杭州银行 | 1.90410% | 1.90853% | 1.90422% | 1.90410% |
| 常熟银行 | 0.29435% | 0.29509% | 0.29437% | 0.29435% |
| 上海银行 | 0.80292% | 0.80690% | 0.80303% | 0.80292% |
| 贵阳银行 | 0.23368% | 0.23885% | 0.23382% | 0.23368% |
| 吴江银行 | 0.72418% | 0.72493% | 0.72420% | 0.72418% |

结果表明，银行信贷额度系数与存款保险费率负相关。随着银行信贷额度系数的提高，存款保险费率呈下降趋势。这一趋势与信贷额度较高的银行出现流动性问题的可能性较小的事实是一致的。

信贷额度是银行可发放贷款的最高额度，是实行限制性货币政策阶段控制信贷量的一种方式。在信贷需求过大的情况下，货币当局为了控制信贷的投放量，通过有关限额或配额的法令和规则限制商业银行的信贷供给。对于不良贷款率较低、资产状况较好的银行，可适当提高信贷额度，提高其盈利能力；而对于不良贷款率较高、资产状况较差的银行，应降低银行的信贷额度，迫使其提高自身抗风险水平，有效的降低道德风险。

最后，考虑活期存款的存款保险费率。为了方便出来，取 $m = 10$，$\alpha = 0.05$，$\beta = 0.8$，代入式（24）计算各银行的存款保险费率（见表 7）。

表 7 考虑技术性破产的活期存款的保险费率表

| 银行 | $m = 10$ | $m = 1$ | $m = 10$ |
| --- | --- | --- | --- |
| 平安银行 | 1.00892% | 1.56643% | 1.00915% |
| 宁波银行 | 0.41499% | 0.87217% | 0.41514% |
| 浦发银行 | 0.54600% | 1.02178% | 0.54624% |
| 华夏银行 | 2.55452% | 2.79096% | 2.55469% |

| | | | |
|---|---|---|---|
| 民生银行 | 0.74144% | 1.22964% | 0.74172% |
| 招商银行 | 0.03489% | 0.11951% | 0.03494% |
| 南京银行 | 1.41242% | 2.06917% | 1.41259% |
| 兴业银行 | 0.54174% | 0.99044% | 0.54202% |
| 北京银行 | 0.51677% | 0.98911% | 0.96686% |
| 农业银行 | 0.44959% | 0.86849% | 0.44987% |
| 交通银行 | 0.34509% | 0.71615% | 0.34537% |
| 工商银行 | 0.01761% | 0.06555% | 0.01777% |
| 光大银行 | 2.48007% | 2.62274% | 2.48026% |
| 建设银行 | 0.02343% | 0.08375% | 0.02359% |
| 中国银行 | 0.15505% | 0.39052% | 0.15532% |
| 中信银行 | 0.51014% | 0.95884% | 0.51040% |
| 江阴银行 | 0.27997% | 0.76127% | 0.27998% |
| 张家港行 | 0.39927% | 1.06122% | 0.39928% |
| 无锡银行 | 0.26549% | 0.69652% | 0.26551% |
| 江苏银行 | 0.36352% | 0.75311% | 0.36378% |
| 杭州银行 | 1.22012% | 1.90422% | 1.22028% |
| 常熟银行 | 0.09492% | 0.29437% | 0.09493% |
| 上海银行 | 0.36721% | 0.80303% | 0.36733% |
| 贵阳银行 | 0.07928% | 0.23382% | 0.07943% |
| 吴江银行 | 0.27208% | 0.72420% | 0.27210% |

注：第二列表示不考虑技术性破产的活期存款的保险费率；第三列表示考虑技术性破产的定期存款的保险费率；第四列表示考虑技术性破产的活期存款的保险费率。

根据表 7，可以得到以下结论：

(1) 技术性破产对活期存款的保险费率有一定影响。考虑银行的技术性破产，即银行发生流动性问题时将关闭银行，存款保险费率有所上升，但上升幅度不大。

(2) 定期存款的保险费远远高于活期存款保险费。表明活期存款的风险水平比定期存款的风险水平低，因此存款保险费率低于定期存款。

综上，银行的储备金系数和信贷额度系数在一定程度上反映了银行发生流动性问题而关闭的概率，进而影响存款保险定价。因此选择合适的定价模型对存款保险进行定价，准确反映银行的各项风险，能保证存款保险作用的有效发挥，减少逆向选择和道德风险，从而起到降低金融风险、维护经济稳定的作用。

（三）因子分析与聚类分析

为了进一步研究国有银行、股份制商业银行、城市商业银行以及农村商业银行的存款保险费率是否存在显著差异，基于上述实证测算的结果对 25 家上市银行的资产价值、负债、资产波动率以及存款保费率进行因子分析[29]，得到两个公因子，第一个公因子在资产价值、负债、资产波动率上的载荷值很大，反映了银行的资产负债结构，第二个公因子在存款保险费率上的载荷值较大，反映了银行的风险情况。

根据得到的因子得分，进一步对 25 家上市银行进行聚类分析，结果如下图 1 所示：
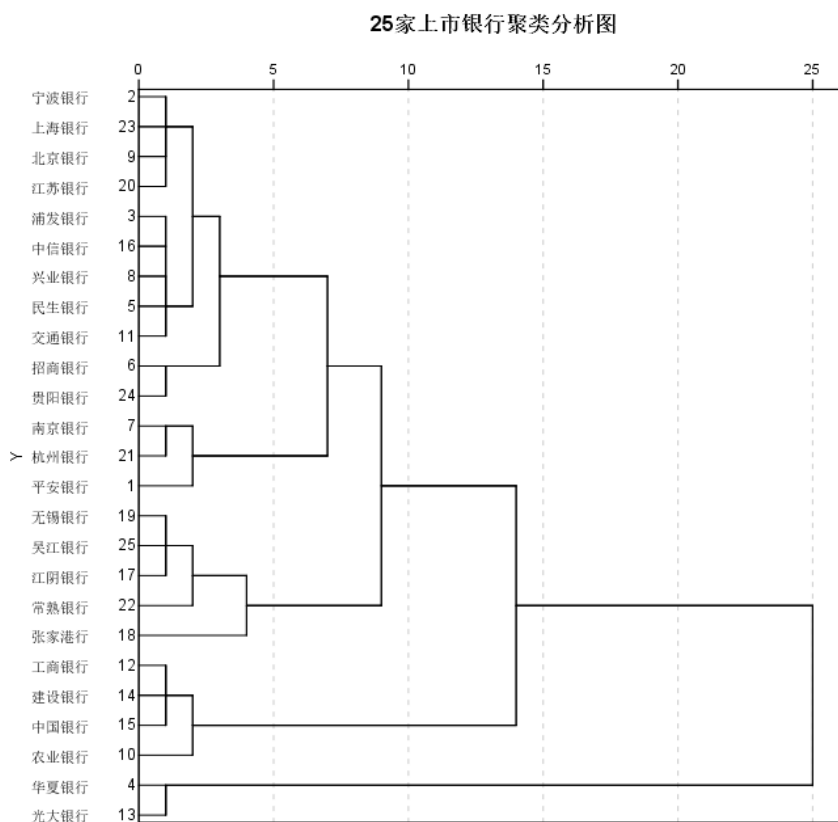


图 1  25 家上市银行聚类分析谱系图

根据图 1，可以将 25 家上市银行大致分为四个层次（见表 8）：

表 8  各银行资产负债结构及存款保险费率分类表

| 银行 | 分类 |
|---|---|
| 1——风险较小的国有制银行 | 工商银行、建设银行、农业银行、中国银行 |

| | |
|---|---|
| 2——风险较大的国有制银行、风险较小的股份制商业银行及城市商业银行 | 兴业银行、中信银行、浦发银行、民生银行、交通银行、宁波银行、上海银行、北京银行、江苏银行、平安银行、招商银行、贵阳银行、南京银行、杭州银行 |
| 3——风险较大的股份制商业银行 | 华夏银行、光大银行 |
| 4——农村商业银行 | 无锡银行、吴江银行、江阴银行、张家港行、常熟银行 |

第一层次是四大国有商业银行，包括工商银行、建设银行、农业银行、中国银行，平均费率为 0.26332%。其特点是：

(1) 资产规模大，资金实力雄厚。根据银监会 2017 年第四季度的统计数据，五大国有商业银行的资产规模占了银行总资产的 36.77%，股份制商业银行占 17.81%，城市商业银行占 12%。我国各类金融机构的资产状况见图 2。

2017年第四季度银行业金融机构资产情况



图 2 我国金融机构 2017 年度资产比例

(2) 在盈利能力方面，国有银行经营管理完善，市场化操作经验丰富，以及拥有幼稚的客户群，这些是其他金融机构所不具备的优势。根据银监会 2017 年的统计数据，国有银行 2017 年第四季度的资产利润率为 1.02%，而股份制商业银行、城市商业银行以及其他金融机构都在 1%以下。

(3) 国有商业银行兼具国有背景，在国家的扶持和政策倾斜下，使其在与其他金融机构竞争中更显优势十足。

第二层次是交通银行、风险水平较小的股份制商业银行及城市商业

银行，包括交通银行、兴业银行、中信银行、浦发银行、民生银行、宁波银行、上海银行、北京银行、江苏银行、平安银行、招商银行、贵阳银行、南京银行、杭州银行，其平均费率为0.81703%。这类金融机构经营效率高，盈利能力强，资产质量相对较高，不良资产比例低。根据银监会2017年的统计数据,股份制商业银行和城市商业银行的平均资本利润率为0.83%，且其不良资产比例一直比较低，据统计，截止到2017年底，股份制商业银行和城市商业银行的不良贷款总额为5674亿元，不良贷款率分别为 1.71%和 1.52%。但股份制商业银行和城市商业银行尚处于发展阶段，资产规模偏小，只占金融机构资产总额的18％左右，抵御风险的能力偏弱。
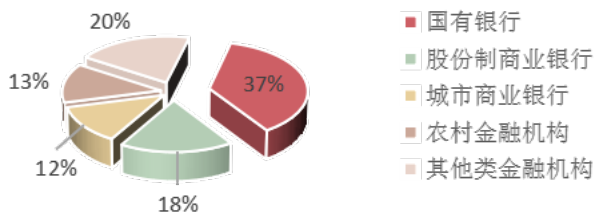
第三层次是风险水平较大的股份制商业银行，包括华夏银行和光大银行，其平均费率为2.51283%。

第四层级是农村商业银行，包括无锡银行、吴江银行、江阴银行、张家港行、常熟银行，平均费率为0.51958%。

综上所述，可以针对这四个风险层次的金融机构分别制定不同的费率水平。

## 五、结论与建议

基于风险的存款保险制度已被广泛提出，然而，大多数研究仅适用于定期存款，并且没有考虑到银行的流动性风险。本文从 Gap 期权定价模型出发，纳入清算折扣因子，推导了不考虑技术性破产和考虑技术性破产情况下定期存款和活期存款的存款保险费率,并基于 25 家上市商业银行的资产负债表数据和股价数据，利用上市银行的股票市场价值 $E$，

股价收益率的波动率 $\sigma_E$ 和银行的总负债 $L$，采用数值方法求解出银行的

资产价值以及隐含波动率，分别计算了不同情况下定期存款和活期存款的存款保险费率，最后根据实证测算结果对我国商业银行进行风险层次分类。具体的结论如下：

(1) 银行流动性风险与存款保险费率呈正相关关系，流动性风险越大，存款保险费率越高。银行的存款保险费率由两部分组成，一部分是银行的预计损失是无力偿还能力,另一部分是技术破产造成的预期损失。首先,较高的银行准备金率和信用额度降低银行发生流动性风险的概率，从而降低存款保险费率；其次，随着清算折扣因子的减小，即银行资产的清算价值减小，存款保险费率呈上升趋势。

(2) 定期存款的保险费远远高于活期存款保险费。当取款点的数量

足够大时，如 100，模型结果接近活期存款的保险成本。总的来说，这项研究表明，活期存款比定期存款的风险水平低，因此保险费率低于定期存款。这一发现与银行在借出活期存款收益方面更加保守的事实是一致的，因为他们必须随时准备提取活期存款。

然而本文还存在一定的局限性。首先，分析假设银行的负债完全由定期存款或活期存款构成，这是不切实际的。实际上，所有银行的负债都是由两种存款组合而成的。因此，该模型不容易适用于任何银行或任何特定类型的存款机构。第二，分析认为，与活期存款相关的随机性仅在于取款时间。在这种情况下，它假设，如果从活期存款中提款，则整个存款被撤回。显然，这是一个极端的假设，将使活期存款看起来比他们的保险成本更低。因此，这两种存款在公平保险费上的差异有些夸大。然而，尽管存在这些局限性，但这项研究的目的是证明这两种存款对保险机构造成的差别风险，并根据风险对更准确的存款保险定价做出贡献。

自 2015 年正式施行《存款保险条例》起，我国存款保险制度就在实践中不断地完善。从施行初期较低的单一基准费率（万分之一点六），到 2016 年 8 月，参考投保机构的监管评级试行差别化费率机制，纳入经营及风险情况等因素进行综合定量和定性测评。然而，相较国际平均万分之五的存保费率，我国的差异化费率机制不论在平均水平还是差异化程度上都有不小的距离。直至 2018 年底，存保基金余额仅为 821.2 亿元，相对高风险金融机构处置需求来说基金规模缺口很大，差异费率考察的因素也相对单一。针对存款保险制度的实施现状，并结合实证结果及结论，本文的政策建议如下：

(1) 健全相关法律法规

合理的存款保险制度需要完善的法律体系保障才能有效执行，我国虽然已经实施《存款保险条例》，但这只是基础，还需要一系列的法律法规来保证存款保险制度的有效运行。例如，可以出台相应的《金融机构破产法》，建立银行业良好的退出机制，迫使各银行提高自身经营水平，保护存款人利益。在健全法律体系的过程中，还应该大力宣传存款保险制度，加强公众的法律及风险意识，让其了解存款保险制度的运行机制，有效保护自身利益。

(2) 构建商业银行风险监测指标体系

合理的风险费率体系的建立需要对银行的风险水平进行综合评估，因此需要建立一个完善的商业银行风险监测指标体系，包括反映资本风险状况、利率风险、流动性风险和经营风险的指标体系等。根据商业银行风险监测预警指标的构建原则，可以选出一系列相关指标。然后根据

历史上发生金融危机事件的统计资料,对这些指标进行有效性统计检验,即可得出一组有效的金融监测指标。商业银行风险监测指标体系的构建不仅全面的表现商业银行面临的风险水平，使存款保险定价更加准确，还可以使银行了解自身风险水平，不断提升自身抗风险水平。

(3) 设计合理的差别费率体系

我国应该尽快建立起合理的风险费率体系，设定合理的风险评价标准及相应的费率水平。不仅要考虑银行的产权结构差异，还应该综合考虑银行面临的各种风险水平，针对不同的风险水平采用不同的费率，这样，成员机构才能按照评定的风险水平"对号入座"。但由于我国的金融体系整体水平较低、内部发展不平衡，缺乏统一、成熟的评价体系，不能对国内每一个金融机构的风险水平进行评估,因此在进行风险定价时，可以采用风险层次定价的方法，即通过聚类的方法，按照上述风险监测指标把我国的金融机构分为几个层次，对每个层次分别制定不同的保险费率，其中可以将国有商业银行作为风险最低的第一层次，其存款保险费牢可以作为基准费率，其他层次的费率在此基础上视风险状况调整。

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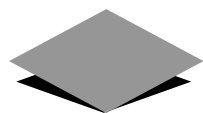
# 3

## 老 龄 化 与 保 险

# 养老保险对农村中老年居民心理健康的影响——基于 CHARLS 的实证研究

<div align="right">孙舒羽 [1]</div>

**摘要：**

  基于 CHARLS 数据库 2011、2013、2015 和 2018 年的平衡面板数据，采用工具变量-固定效应模型(IV-FE)来研究养老保险对农村中老年居民心理健康的影响，并进行异质性分析和中介效应分析。结果表明：参加养老保险可以显著提升农村中老年居民的心理健康水平，尤其是可以显著减少其抑郁情绪，但是政策效力有限；同时该影响具有异质性，男女居民、60 岁及以上和 60 岁以下老人、已婚或同居群体、中等或低收入群体和各地区居民参加养老保险都可以显著改善其心理健康水平，但是仅对 60 岁及以上老人、中低收入群体表现出更明显的积极影响；而从传导路径看，参加养老保险可以通过提高生活满意度和增加社会资本，进而降低农村中老年居民的抑郁程度，促进心理健康。

**关键词：** 养老保险，心理健康，IV-FE，异质性，中介效应

## 一、引言

  2001 年中国 65 周岁及以上人口占比突破 7%，中国自此进入"老龄化社会"；而到 2021 年，中国 65 周岁及以上人口超过 2 亿，占总人口的 14.2%，中国也正式进入"老龄社会"，比 2019 年版《世界人口展望》测算的 2025 年提前 4 年，老龄化进程也明显加快（如图 1.1 所示）。横向进行国际比较，中国老龄化的速度也明显快于其他发达国家，从老龄化社会到老龄社会，法国经历 115 年、美国经历 64 年、日本经历 24 年，预计全球平均需经历 35 年，但中国仅用 20 年就完成了这一转变。而在城乡二元结构的背景下，中国老龄化还出现明显的城乡差异，农村老龄化水平明显高于城镇，以 2020 年为例，农村地区 60 和 65 周岁及以上人口占比都高于城镇地区（如图 1.2 所示）。

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[1] 孙舒羽，北京大学经济学院保险硕士。
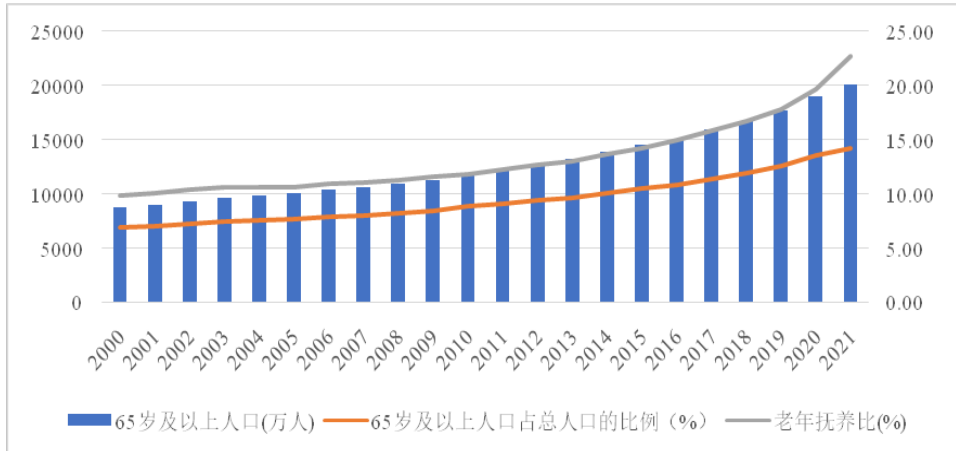
图 1.1：2000-2021 年中国人口年龄结构 [1]
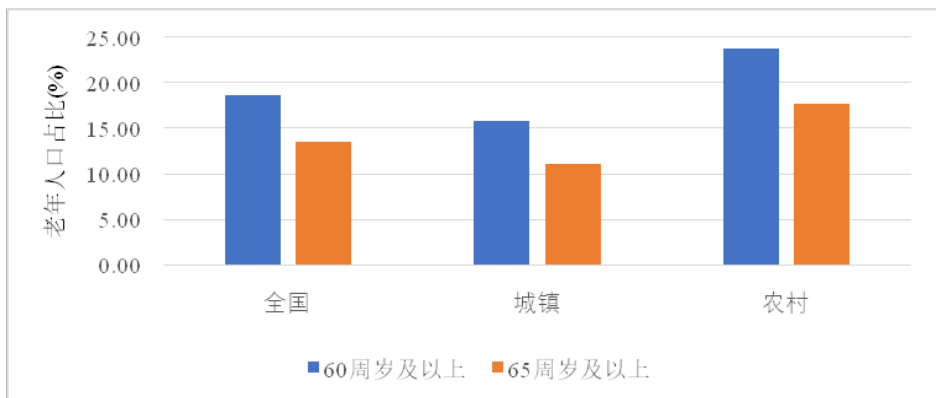


图 1.2：2020 年中国 60 和 65 周岁及以上人口占比（%）[2]

WHO(2017)指出，抑郁障碍是全球范围内常见的精神障碍之一。截至 2015 年，在全球范围内有 3.22 亿人患有抑郁症，相当于世界人口的 4.4%，且女性（5.1%）比男性（3.6%）更加常见；患病率也因年龄而异，在老年群体中水平最高。"中国精神障碍疾病负担及卫生服务利用的研究"的结果显示，我国成人抑郁障碍终生患病率为 6.8%；女性患病率(8.0%)高于男性 (5.7%)；分居、丧偶或离婚者的抑郁症终生患病率是已婚或同居者的 1.87 倍( Lu et al., 2021)。2015 年 CHARLS 数据显示，60 岁及以上的老年人共有 33.1%存在高抑郁风险；农业户口人群（38.3%）的高抑郁风险比例高于非农业户口人群（22.2%），且在各个年龄段都存在。总之，中国老年人的心理健康问题不容忽视，尤其是农村老人。

在如今全球老龄化背景下，逐步推行和落实针对老年人的福利政策是

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[1] 数据来源：国家统计局
[2] 数据来源：国家卫生健康委员会《2020 年度国家老龄事业发展公报》

各国政府的必然选择；而在人民物质生活水平日益提高的今天，对于心理健康的重视程度也逐步增加，精神障碍已成为全球疾病负担的一个重要问题，各国政府也在尽可能采取措施改善居民心理健康状况。而中国正面临老龄化和低心理健康水平的双重危机，二者叠加之下，中国农村老人的心理健康问题就非常值得关注。在中国农村，"养儿防老"是传统的养老观念，这代表着"家庭养老"；土地为农村居民提供了社会保障和失业保险（姚洋，2000），土地养老是其"自我养老"重要的组成部分；自 2009 年推行的"新农保"和 2014 年起实施的城乡居民养老保险制度则为农村居民提供"社会养老"保障；"家庭养老"、"自我养老"和"社会养老"逐渐成为农村中老年居民生活的三大支柱。其中，养老保险作为政府可以直接干预的政策手段，有必要确定其对于农村居民心理健康的影响。

本文选取 CHARLS 数据库 2011、2013、2015 和 2018 年的平衡面板数据，使用工具变量-固定效应模型(IV-FE)评估养老保险对农村中老年居民心理健康的影响，并进行异质性分析和中介效应分析。文章的创新之处在于：第一，对 CES-D 进行因素分析，将其分成"抑郁情绪"维度和"积极情绪"维度，从而能够更加直观地体现居民的心理健康结构，也使得养老保险对农村中老年居民心理健康的影响更加明确；第二，进行中介效应分析，选取生活满意度和社会资本两个中介变量，考察养老保险与农村中老年居民心理健康之间的传导机制。

## 二、文献综述

（一）老年心理健康的影响因素

对于老年人而言，健康状况的恶化是其面临的一大挑战，对其生活品质会产生巨大影响，也会影响其心理健康水平(Chen et al., 2012)。首先在身体疾病方面，患有慢性病或者残疾的老年人更容易出现心理健康问题，且随着疾病数量的增多抑郁风险也会增加(Chi et al., 2005; Lue et al., 2010)。其次当老年人出现功能障碍时也更容易患有抑郁症(Luppa et al., 2012)。认知能力也是抑郁症的重要影响因素，认知状态差会导致居民的应对能力差，进而导致抑郁，而抑郁也会进一步使其认知能力下降，陷入恶性循环(Liu et al., 1997)。自评健康状况也会影响老年人的心理健康，自评健康差的人往往更容易患有抑郁，这在没有身体疾病或功能障碍的老人中也同样存在(Chen et al., 2012; Luppa et al., 2012)。

社会资本指的是能够促进集体行动的社会结构、社会制度和信息渠道(温兴祥等，2017)，包括个体层面和社区层面的社会资本，二者都能以社会网络为媒介来促进居民的心理健康。一方面社会资本可以通过社会网络实现成员之间的物质支持、情感支持等，使居民可以从社会网络中获取一定资源从而促进其心理健康；另一方面社会资本可以使成员以社会网络为连

接来进行信息共享，从而降低信息获取成本，这对于医疗资源相对不充分的农村地区而言更为重要(孙博文等, 2016; 温兴祥等, 2017)。在对日本居民的实证研究中，Yamaguchi et al.(2019)研究发现，社区社会资本中的公民参与度与抑郁症发病率呈现负相关，而社会凝聚力和互惠性则与之无关。在对中国居民的实证研究中，Gu et al.(2021)发现社区社会资本中的公民参与和社会凝聚力提高都有助于城乡居民的心理健康；而个人层面的社会资本也是老年心理健康的重要保护因素(Lin et al., 2019; Norstrand & Xu, 2012; 温兴祥等, 2017)。

经济收入与健康之间存在正相关，这在身体和心理健康方面都存在(Wang , 2001; Wang & Zheng, 2021)。除了实际收入外，自我感知的财务状况也是影响心理健康的重要因素，如 Lue et al.(2010)对台湾老年人的研究发现，自评经济压力大的老人更容易患抑郁症，且在女性中表现更明显。但是收入和心理健康之间的因果关系很难确定，可能是较高的收入导致心理健康水平的改善，也可能是良好的心理健康带来更高的生产力和收入。

此外，一些人口统计学因素也是老年居民心理健康的重要影响因素。一般来说女性抑郁症患病率更高(Cheung & Mui, 2021; Li et al., 2014)。从婚姻状态来看，未婚、离异、丧偶或分居也是老年人心理健康的危险因素(Li et al., 2014; Luppa et al., 2012; Zhang & Li, 2011)；Förster et al.(2019)对德国老人的研究还发现，丧偶的男性比丧偶的女性更容易患上抑郁症。对于受教育程度，一般较低的受教育程度更容易导致抑郁，这可能是因为缺乏教育会限制人们解决问题、应对压力的能力(Chen et al., 2012; Hu et al., 2018)；但也有少数研究者持相反观点，Zhang et al.(2010)就认为受过良好教育的人会有更高的生活预期，那么当不能实现其预期的时候，就会导致更大的失望和焦虑。关于年龄和心理健康之间的关系，尚无统一的结论，有的学者认为年龄是老年人心理健康的重要影响因素(Chong et al., 2001; Kok & Reynolds, 2017)；但也有人认为二者无关，年龄的影响本质是其背后如身体健康状况等在发挥作用(Li et al., 2014; Yu et al., 2012)。

（二）养老保险与老年心理健康

养老金作为一项重要的社会福利政策，对于老年人尤其是农村老人的心理健康而言起到了重要作用。养老金被认为是老年人的重要收入来源，它可以保障其获取和利用资源的能力，保证其生活质量，同时养老金还可以提供稳定的收入预期，这都有助于保护老年人的心理健康(He et al., 2021)。

实证研究表明，养老保险对老年人心理健康的促进作用在全球范围内都存在。Golberstein(2015)利用美国政策变化造成的社会保障福利"缺口"，研究发现1000美元的额外社会保障收入能将女性抑郁症患病率显著降低8个百分点。Galiani et al.(2016)发现墨西哥的非缴费型养老金计划可以使老

年人在老年抑郁量表上的得分下降 12%；他们后续将研究拓展到秘鲁和巴拉圭，也发现类似的结论(Bando et al., 2020; Bando et al., 2022)。此外，在西欧、韩国、南非和巴西等国家也都发现了养老金与心理健康之间的正相关关系(Esser & Palme, 2010; Kim et al., 2018; Lloyd-Sherlock et al., 2012)。

在中国范围内，学者常常利用"新农保"这一政策来进行研究。张川川等(2015)使用断点回归和双重差分法处理CHARLS数据库 2011 年的数据，发现养老金收入可以显著提高老年人的主观福利；何泱泱和周钦(2016)发现，"新农保"不仅可以减轻农村中老年居民的抑郁症状，并且处于领取阶段的老年人比处于缴费阶段的人受益更多；Pan et al.(2021)也指出，养老金计划不仅减轻了农村中老年居民的抑郁症状，还降低了抑郁症的患病率。

学者还对养老保险对心理健康影响的异质性进行了研究。一般来说，女性、低收入群体、身体健康较差、受教育程度较低和独居老人更容易受到养老保险政策的影响，这些人也往往是原本心理健康水平较差的群体(Chen et al., 2019; Pan et al., 2021; Wang & Zheng, 2021; 何泱泱, 周钦, 2016; 张川川等, 2015; 周钦等, 2018)。地区方面，以中国为例，Wang 和 Zheng(2021)研究发现养老保险对中西部地区的老年人而言才可以改善其心理健康状况；但是 Pan et al.(2021)的研究则认为养老保险只在中部地区才有显著作用，作者指出与东部地区相比，中部地区的受访者平均收入较低，加上中部地区靠近相对发达的东部地区，因此该地区的居民可能会有更强的被剥夺感。

尽管既往关于养老保险与心理健康之间关系的研究众多，但对二者之间的传导机制尚不明晰，相关研究较少。Gao 和 Feng(2020)研究发现劳动参与度和养老保险收入之间存在负相关，但劳动参与度和抑郁症状之间无关；进一步做有调节的中介效应分析，结果显示性别起到调节作用，即对女性而言养老金收入通过劳动参与度对抑郁症状的间接影响是显著的，但在男性中没有影响；作者认为这可能是养老金收入使老年居民减少劳动参与时间从而有更多机会参与社会活动，而后者被认为与老年人抑郁症状减轻有关，尤其是对其更感兴趣的老年妇女(Kim et al., 2015; Zhang et al., 2017)。Pan et al.(2021)研究则发现养老金计划可以通过减少劳动力供应、改善家庭关系和增加消费支出来缓解老年人的抑郁症。He et al.(2021)考察代际支持是否起到中介作用，发现情感支持和工具支持并没有显著作用，仅经济支持才可以在养老金和抑郁症状之间发挥中介作用，这是因为目前农村地区养老金的数额有限，因此老人仍然难以负担生活或医疗费用，也就需要子女的经济支持，而后者与老人较低的抑郁水平相关(Chao, 2012)。

## 三、研究设计
（一）数据来源与变量定义

本文使用中国健康与养老追踪调查（China Health and Retirement Longitudinal Study, CHARLS）中 2011 年的基期数据和 2013、2015、2018 年的追踪数据构建平衡面板数据。由于本文意在通过构造四期平衡面板数据来追踪考察社会保险对农村中老年居民心理健康的影响，故删去不足四期数据的居民；另外考虑到研究对象是农村中老年居民，因此对于受访对象中的城镇居民和 45 岁以下的农村居民不予考虑。将这些样本剔除之后，满足条件的有 3619 人，四期有效样本共计 14476 个。

本文的因变量是居民心理健康状况，通过抑郁量表总得分（0-30 分）、各因子得分和是否有抑郁症状（1 表示有；0 表示无）来衡量。CHARLS 使用的是简化版的流调中心抑郁量表(the Center for Epidemiological Studies Depression Scale, CES-D)，共计 10 道问题(CESD-10)。该量表包含 8 个描述消极情绪的条目和 2 个描述积极情绪的条目，与过去一周的感觉和行为有关；每个条目都有四个选项，即"很少或者根本没有（<1 天）"、"不太多（1-2 天）"、"有时或者说有一半的时间（3-4 天）"和"大多数的时间（5-7 天）"，分别记为 0、1、2、3 分，在计分时将 2 个描述积极情绪的条目进行反向计分，然后再计算量表总分（如表 3.1 所示），总计 30 分，得分越高表示心理健康状况越差，得分超过 10 分被认为出现抑郁症状。对于 CESD-10 量表的因子结构，目前尚未得到统一结论，存在单因素说(Björgvinsson et al., 2013)、双因素说(Lee & Chokkanathan, 2008; Bradley et al., 2010; 黄庆波等, 2015)和三因素说(Cheng et al., 2006)，考虑到被试的文化背景和身份信息不足以解释 CESD-10 的因子结构差异，故本文会先进行因素分析，确定该样本更合适的因子模型，再计算各因子得分。

表 3.1：CESD-10 量表内容与计分方向

| 序号 | 条目内容 | 计分方向 |
|---|---|---|
| 1 | 我因一些小事而烦恼 | 正向 |
| 2 | 我在做事时很难集中精力 | 正向 |
| 3 | 我感到情绪低落 | 正向 |
| 4 | 我觉得做任何事都很费劲 | 正向 |
| 5 | 我对未来充满希望 | 反向 |
| 6 | 我感到害怕 | 正向 |
| 7 | 我的睡眠不好 | 正向 |
| 8 | 我很愉快 | 反向 |
| 9 | 我感到孤独 | 正向 |
| 10 | 我觉得我无法继续我的生活 | 正向 |

本文的自变量是居民是否参加养老保险的虚拟变量。具体而言，若居民加入了城乡居民养老保险或"新农保"，则认为当期参加了养老保险，记

为 1；反之则没有，记为 0。

本文的控制变量参考前人研究，具体包括个人、家庭和区域三个层面。个人层面包括性别（1 表示男性；0 表示女性）、年龄、受教育程度（1 表示高中及以上学历；0 表示高中以下学历）、婚姻状况（1 表示已婚或同居；0 表示未婚、离异、丧偶或分居）、是否有慢性病（1 表示有；0 表示）、是否参加新型农村合作医疗制度或城乡居民基本医疗保险（1 表示参加；0 表示未参加）。家庭层面包括家庭规模和人均收入两个变量。区域层面指样本所在地区，考虑不同区域的社会经济发展状况，将我国的经济区域划分为东部、中部、西部和东北四大地区（1 表示东部；2 表示中部；3 表示西部；4 表示东北）。

本文使用的工具变量为样本所在社区的参保率。社区参保率通常是政策执行、地方政府财政能力、地方政府声誉等的函数，通常与个人参加养老保险的可能性呈正相关，而与个人的心理健康状况无关，因此是一个合格的工具变量(Pan et al., 2021)。

本文使用的中介变量包括生活满意度和社会资本。CHARLS 问卷中有专门询问受访者生活满意度的问题，即"总体来看，您对自己的生活是否感到满意？是极其满意，非常满意，比较满意，不太满意还是一点也不满意？"，将结果分别记为 4、3、2、1、0 分，作为受访者生活满意度得分。考虑到 CHRALS 问卷中并没有涉及社区层面的社会资本的问题，因此本文仅考虑个人层面的社会资本，问卷中包含"您过去 1 个月是否参与了以下社交活动？"的问题，将个体参与 10 项活动的频次加总作为个人社会资本(孙博文等, 2016; 薛新东, 刘国恩, 2012)。

（二）模型设计

1.基准回归模型

本文使用固定效应模型(FE)来研究参加养老保险对农村中老年居民心理健康的影响，模型设定如下：

$$Y_{it} = \alpha + \beta D_{it} + \gamma X_{it} + \lambda_i + \varepsilon_{it} \qquad (1)$$

其中，$Y_{it}$ 表示个体 i 在第 t 年的心理健康状况；$D_{it}$ 表示个体 i 在第 t 年的参加养老保险的情况（$D_{it} = 1$ 表示个体 i 在第 t 年参保，$D_{it} = 0$ 表示个体 i 在第 t 年未参保）；$X_{it}$ 表示一组影响居民心理健康水平的特征变量，如性别、年龄、婚姻状况、受教育程度、是否有慢性病、是否参加医疗保险、家庭规模、家庭人均收入和地区等；$\lambda_i$ 表示个体固定效应。此外，当 $Y_{it}$

为是否有抑郁症状这个虚拟变量的时候，本文使用基于 Probit 的固定效应模型来进行分析。

为避免内生性对结果的影响，本文又使用两阶段最小二乘法(Two Stage Least Square, 2SLS)，工具变量-固定效应模型(IV-FE)设定如下：

$$\widehat{D}_{it} = \alpha_1 + \beta_1 IV_{it} + \gamma_1 X_{it} + \lambda_i + u_{it} \qquad (2)$$

$$Y_{it} = \alpha_2 + \beta_2 \widehat{D}_{it} + \gamma_2 X_{it} + \lambda_i + v_{it} \qquad (3)$$

式(2)为一阶段方程式，式(3)为二阶段方程式。其中，$IV_{it}$表示工具变量社区参保率；其他变量均不变。

2.中介效应模型

本文使用温忠麟和叶宝娟(2014)归纳总结出的中介效应检验流程（如图 3.1 所示）。模型设定如下：

$$MV_{it} = \alpha_3 + \beta_3 D_{it} + \gamma X_{it} + \lambda_i + \varepsilon_{it} \qquad (4)$$

$$Y_{it} = \alpha_4 + \beta_5 D_{it} + \beta_4 MV_{it} + \gamma X_{it} + \lambda_i + \varepsilon_{it} \qquad (5)$$

其中，$MV_{it}$表示中介变量；其他变量均不变。若方程(1)的系数$\beta$显著，则按中介效应理论，反之按遮掩效应立论。若$\beta_3$和$\beta_4$都显著则表示间接效应显著；若至少一个不显著则使用偏差校正的非参数百分位 Bootstrap 法检验$\beta_3\beta_4$，偏差校正的置信区间不包括 0 则说明间接效应显著，反之间接效应不显著。在存在间接效应的情况下，检验系数$\beta_5$，若不显著则说明直接效应不显著，只有中介效应；反之说明直接效应显著，此时比较$\beta_3\beta_4$和$\beta_5$的符号，若同号则说明是部分中介效应，报告的效应量为中介效应占总效应的比例$\beta_3\beta_4 / \beta$；若异号则说明是遮掩效应，报告的效应量为间接效应与直接效应比值的绝对值$|\beta_3\beta_4 / \beta_5|$。
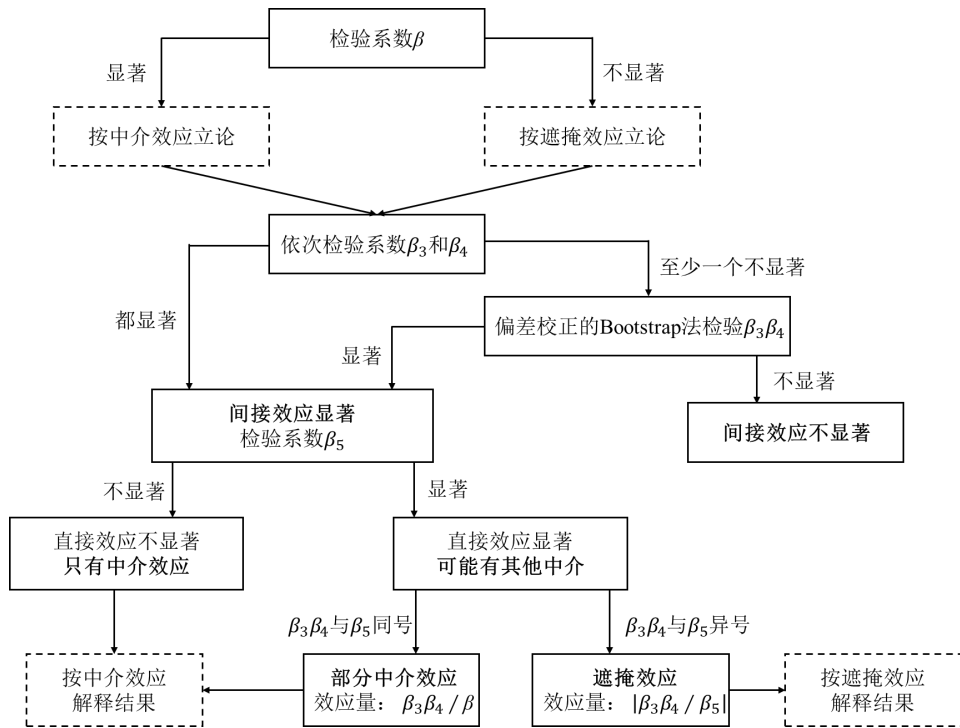
图 3.1：中介效应检验流程

（三）信效度检验与因素分析

1.效度检验与因素分析

对于 2011 年的数据，KMO=0.878 接近 1，且 Bartlett 检验结果显著（p<0.001），可见各条目之间有较强的相关性，效度很好，可以进行因素分析；碎石图结果显示二因子结构最合适；然后使用"方差最大正交旋转"，生成的旋转后成分矩阵如表 3.2 所示，可以看到条目 1、2、3、4、6、7、9、10 在成分 1 中得分较高，定义为"抑郁情绪"因子，条目 5 和 8 在成分 2 中得分较高，定义为"积极情绪"因子。对 2013、2015 和 2018 年的数据也进行因素分析，得到的结果类似。综上，CES-D 量表对于农村中老年居民而言具有很好的效度，且更符合二因子结构。

表 3.2：旋转后的成分矩阵

| 序号 | 2011 年 | | 2013 年 | | 2015 年 | | 2018 年 | |
|---|---|---|---|---|---|---|---|---|
| | 成分 1 | 成分 2 | 成分 1 | 成分 2 | 成分 1 | 成分 2 | 成分 1 | 成分 2 |
| 1 | 0.729 | 0.093 | 0.693 | 0.036 | 0.704 | 0.120 | 0.700 | 0.109 |
| 2 | 0.707 | -0.011 | 0.670 | -0.056 | 0.710 | -0.011 | 0.707 | 0.010 |
| 3 | 0.757 | 0.187 | 0.781 | 0.062 | 0.781 | 0.150 | 0.783 | 0.085 |

| | | | | | | | |
|---|---|---|---|---|---|---|---|
| 4 | 0.709 | 0.112 | 0.684 | 0.098 | 0.718 | 0.082 | 0.741 | 0.054 |
| 5 | 0.004 | 0.849 | -0.030 | 0.850 | -0.031 | 0.858 | -0.075 | 0.866 |
| 6 | 0.543 | 0.158 | 0.564 | 0.034 | 0.596 | 0.066 | 0.633 | 0.016 |
| 7 | 0.486 | 0.149 | 0.526 | 0.104 | 0.542 | 0.002 | 0.537 | 0.120 |
| 8 | 0.282 | 0.743 | 0.198 | 0.795 | 0.254 | 0.749 | 0.302 | 0.735 |
| 9 | 0.532 | 0.333 | 0.644 | 0.089 | 0.618 | 0.224 | 0.643 | 0.157 |
| 10 | 0.525 | 0.344 | 0.623 | 0.161 | 0.622 | 0.187 | 0.676 | 0.099 |

2.信度检验

信度检验结果如表 3.3 所示。对 2011 年的数据而言：在"抑郁情绪"维度上，基于标准化的 Cronbach's Alpha=0.805，可靠性很高，且 8 个条目任意一项删除后都会使得 Cronbach's Alpha 降低，因此不存在不合适的条目；在"积极情绪"维度上，Cronbach's Alpha=0.507，可靠性较好。对 2013、2015 和 2018 年的数据进行信度检验，结果类似。综上，CES-D 量表对于农村中老年居民而言具有很好的信度。

表 3.3：CES-D 量表的信度检验

| 年份 | 维度 | Cronbach's Alpha | 基于标准化项的 Cronbach's Alpha | 条目数 |
|---|---|---|---|---|
| 2011 | "抑郁情绪" | 0.801 | 0.805 | 8 |
| | "积极情绪" | 0.567 | 0.568 | 2 |
| 2013 | "抑郁情绪" | 0.801 | 0.808 | 8 |
| | "积极情绪" | 0.555 | 0.555 | 2 |
| 2015 | "抑郁情绪" | 0.822 | 0.826 | 8 |
| | "积极情绪" | 0.516 | 0.517 | 2 |
| 2018 | "抑郁情绪" | 0.835 | 0.838 | 8 |
| | "积极情绪" | 0.490 | 0.491 | 2 |

（四）描述性统计

对样本进行描述性统计，总样本共计 14476 个，CES-D 平均得分为 8.284，其中"抑郁情绪"维度的平均得分为 5.784，"积极情绪"维度的得分为 3.501，35.7%的人出现了抑郁症状。对样本进行分组描述性统计，参保样本共计 9075 个，CES-D 平均得分为 8.327，其中"抑郁情绪"维度的平均得分为 5.77，"积极情绪"维度的得分为 3.443，35.4%的人出现了抑郁症状；未参保样本共计 5401 个，CES-D 平均得分为 8.211，其中"抑郁情绪"维度的平均得分为 5.809，"积极情绪"维度的得分为 3.597，35.9%的人出现了抑

郁症状。其他变量的描述性统计如表 3.4 所示。

表 3.4：描述性统计

| 变量名 | 参保样本 | | 未参保样本 | | 总样本 | |
| --- | --- | --- | --- | --- | --- | --- |
| | 均值 | 标准差 | 均值 | 标准差 | 均值 | 标准差 |
| CES-D 得分 | 8.327 | 6.253 | 8.211 | 6.203 | 8.284 | 6.234 |
| 抑郁情绪 | 5.770 | 5.440 | 5.809 | 5.336 | 5.784 | 5.401 |
| 积极情绪 | 2.557 | 2.044 | 2.403 | 1.988 | 2.499 | 2.025 |
| 抑郁症状 | 0.359 | 0.480 | 0.354 | 0.478 | 0.357 | 0.479 |
| 性别 | 0.497 | 0.500 | 0.518 | 0.500 | 0.505 | 0.500 |
| 年龄 | 60.269 | 7.857 | 58.619 | 8.217 | 59.653 | 8.033 |
| 受教育程度 | 0.077 | 0.267 | 0.085 | 0.279 | 0.080 | 0.272 |
| 婚姻状况 | 0.901 | 0.298 | 0.91 | 0.286 | 0.905 | 0.294 |
| 是否有慢性病 | 0.791 | 0.407 | 0.726 | 0.446 | 0.767 | 0.423 |
| 是否参加医疗保险 | 0.985 | 0.122 | 0.911 | 0.284 | 0.957 | 0.202 |
| 家庭规模 | 2.774 | 1.573 | 3.223 | 1.760 | 2.941 | 1.660 |
| 人均收入 | 13420.697 | 40708.683 | 14315.074 | 41030.608 | 13754.389 | 40829.966 |
| 地区 | 2.147 | 0.889 | 2.132 | 0.943 | 2.141 | 0.910 |
| 生活满意度 | 2.261 | 0.747 | 2.149 | .726 | 2.219 | 0.742 |
| 社会资本 | 0.758 | 1.205 | 0.707 | 1.152 | 0.739 | 1.186 |

## 四、实证结果与分析

1.基准分析

为增加本文选用固定效应模型而非随机效应模型的可信度，首先使用 Hausman 检验来进行比较，结果显示 p<0.001，因此使用固定效应模型更

为合适；其次内生性检验结果显示 p<0.001，即存在内生性，有必要使用工具变量法。分别以 CES-D 总得分及各因子得分、是否有抑郁症状为因变量建立方程，结果如表 4.1 和表 4.2 所示。

模型(1)~(3)以 CES-D 量表总得分为因变量，IV-FE 结果显示参加养老保险的系数为-1.063，且在 1%的水平上显著，表示农村居民参加养老保险可以使其 CES-D 得分降低 1.063，心理健康水平得到显著改善，此外，年龄增加、婚姻状况不佳、有慢性病和家庭规模较小也都会使其 CES-D 得分增加；使用 OLS 模型和 FE 模型也可以得到类似的结论。模型(4)~(6)以是否出现抑郁症状为因变量，IV-FE 结果显示参加养老保险的平均边际效应为-0.009，表示与未参保群体相比，参保群体出现抑郁症状的概率降低0.9%，但是效果并不显著；使用 OLS 模型和 FE 模型得到的结果也类似。模型(7)~(9)以"抑郁情绪"维度得分为因变量，IV-FE 结果显示参加养老保险的系数为-1.687，且在 1%的水平上显著，表示农村居民参加养老保险可以使其在"抑郁情绪"维度上的得分降低 1.687，抑郁情绪显著减弱，此外，年龄增加、婚姻状况不佳、有慢性病和家庭规模较小也都会使其 CES-D得分增加；使用 OLS 模型和 FE 模型也可以得到类似的结论。模型(10)~(12)以反向计分的"积极情绪"维度得分为因变量，IV-FE 结果显示参加养老保险的系数为 0.624，且在 1%的水平上显著，表示农村居民参加养老保险可以使其在反向计分的"积极情绪"维度的得分增加 0.624，积极情绪也显著减弱；使用 OLS 模型和 FE 模型得到的结果也类似。

综上所述，参加养老保险有助于通过减少抑郁情绪来改善农村居民心理健康水平，但是并不足以使其出现抑郁症状的概率显著降低，这可能与农村养老保险保障力度不足有关(Wang & Zheng, 2021)。

表 4.1：养老保险对农村居民 CES-D 总得分和是否出现抑郁症状的影响

| | CES-D 总得分 | | | 是否有抑郁症状 | | |
|---|---|---|---|---|---|---|
| | (1)<br>OLS | (2)<br>FE | (3)<br>IV-FE | (4)<br>OLS | (5)<br>FE | (6)<br>IV-FE |
| 是否参保 | -0.226** | -0.316*** | -1.063*** | -0.016* | -0.003 | -0.009 |
| | (-2.16) | (-3.14) | (-6.16) | [-1.92] | [-0.39] | [-0.65] |
| 性别 | -1.888*** | 0.000 | 0.000 | -0.125*** | 0.000 | -0.125*** |
| | (-18.84) | (.) | (.) | [-16.41] | [.] | [-16.39] |
| 年龄 | 0.003 | 0.067*** | 0.105*** | -0.000 | 0.002 | 0.000 |
| | (0.51) | (3.71) | (5.43) | [-0.05] | [0.58] | [0.08] |
| 受教育程度 | -0.961*** | 0.000 | 0.000 | -0.057*** | 0.000 | -0.054*** |
| | (-5.19) | (.) | (.) | [-3.76] | [.] | [-3.57] |
| 婚姻状况 | -1.757*** | -1.449*** | -1.447*** | -0.100*** | -0.131*** | -0.099*** |

|  | (-10.16) | (-4.52) | (-4.50) | [-7.66] | [-5.92] | [-7.55] |
|---|---|---|---|---|---|---|
| 是否有慢性病 | 2.708*** | 0.795*** | 0.880*** | 0.192*** | 0.074*** | 0.191*** |
|  | (22.91) | (4.81) | (5.28) | [20.30] | [5.81] | [20.13] |
| 是否参加医疗保险 | -0.227 | -0.121 | 0.003 | -0.023 | -0.028 | -0.022 |
|  | (-0.91) | (-0.51) | (0.01) | [-1.21] | [-1.57] | [-1.10] |
| 家庭规模 | -0.146*** | -0.117*** | -0.113*** | -0.007*** | -0.002 | -0.008*** |
|  | (-4.66) | (-3.44) | (-3.33) | [-2.94] | [-0.59] | [-3.18] |
| 人均收入 | -0.434*** | -0.009 | -0.015 | -0.026*** | -0.003 | -0.026*** |
|  | (-13.40) | (-0.29) | (-0.49) | [-10.65] | [-0.90] | [-10.68] |
| 地区 | 0.634*** | 0.000 | 0.000 | 0.044*** | 0.000 | 0.041*** |
|  | (11.60) | (.) | (.) | [10.46] | [.] | [9.32] |
| 个体固定效应 | 否 | 是 | 是 | 否 | 是 | 是 |
| 样本量 | 14476 | 14476 | 14476 | 14476 | 14476 | 14476 |
| $R^2$ | 0.102 | 0.503 | / | 0.059 | 0.139 | / |

注：当因变量为 CES-D 总得分时，使用系数估计值；当因变量为是否有抑郁症状时，使用平均边际效应；***、**、*分别表示在 1%、5%和 10%的水平上显著；( )中为 t 值，[ ]中为 z 值。

表 4.2：养老保险对农村居民 CES-D 量表各维度得分的影响

|  | "抑郁情绪"维度得分 | | | "积极情绪"维度得分 | | |
|---|---|---|---|---|---|---|
|  | (7) | (8) | (9) | (10) | (11) | (12) |
|  | OLS | FE | IV-FE | OLS | FE | IV-FE |
| 是否参保 | -0.328*** | -0.522*** | -1.687*** | 0.102** | 0.206*** | 0.624*** |
|  | (-3.61) | (-5.79) | (-10.86) | (2.90) | (5.19) | (9.16) |
| 性别 | -1.754*** | 0.000 | 0.000 | -0.134*** | 0.000 | 0.000 |
|  | (-20.13) | (.) | (.) | (-3.97) | (.) | (.) |
| 年龄 | -0.008 | 0.058*** | 0.118*** | 0.011*** | 0.009 | -0.013* |
|  | (-1.29) | (3.59) | (6.76) | (4.83) | (1.27) | (-1.65) |
| 受教育程度 | -0.706*** | 0.000 | 0.000 | -0.255*** | 0.000 | 0.000 |
|  | (-4.38) | (.) | (.) | (-4.08) | (.) | (.) |
| 婚姻状况 | -1.491*** | -1.423*** | -1.419*** | -0.267*** | -0.027 | -0.028 |
|  | (-9.91) | (-4.95) | (-4.90) | (-4.57) | (-0.21) | (-0.22) |
| 是否有慢性病 | 2.271*** | 0.551*** | 0.683*** | 0.437*** | 0.244*** | 0.196*** |
|  | (22.10) | (3.72) | (4.56) | (10.97) | (3.75) | (2.99) |
| 是否参加医 | -0.036 | -0.021 | 0.171 | -0.191** | -0.100 | -0.169* |

| 疗保险 | (-0.17) | (-0.10) | (0.80) | (-2.28) | (-1.07) | (-1.79) |
|---|---|---|---|---|---|---|
| 家庭规模 | -0.145*** | -0.146*** | -0.141*** | -0.001 | 0.029** | 0.028** |
| | (-5.35) | (-4.81) | (-4.59) | (-0.05) | (2.20) | (2.05) |
| 人均收入 | -0.316*** | 0.012 | 0.002 | -0.118*** | -0.021* | -0.017 |
| | (-11.23) | (0.44) | (0.07) | (-10.81) | (-1.73) | (-1.42) |
| 地区 | 0.527*** | 0.000 | 0.000 | 0.107*** | 0.000 | 0.000 |
| | (11.09) | (.) | (.) | (5.83) | (.) | (.) |
| 个体固定效应 | 否 | 是 | 是 | 否 | 是 | 是 |
| 样本量 | 14476 | 14476 | 14476 | 14476 | 14476 | 14476 |
| $R^2$ | 0.096 | 0.469 | / | 0.033 | 0.269 | / |

注：当因变量为 CES-D 量表各因子得分时，使用系数估计值；***、**、* 分别表示在 1%、5%、10% 的水平上显著；( )中为 t 分数。

2.稳健性检验

本文在进行实证分析的时候，是以个体为研究单位的，但考虑到参加养老保险后得到的经济及其他保障可以作用于整个家庭，比如领取的养老金或获得的未来收入预期实际上对夫妻双方均有益，因此也可以使用家户为研究单位，并将其作为稳健性检验，结果如表 4.3 所示：无论是以 CES-D 量表总得分或各因子得分为因变量，还是以是否出现抑郁症状为因变量，以个体为单位和以家户为单位的结果都是一致的，因此本文研究结论具有较好的稳健性。

表 4.3：稳健性检验

| | CES-D 总得分 | | "抑郁情绪" | | "积极情绪" | | 抑郁症状 | |
|---|---|---|---|---|---|---|---|---|
| | 个体 | 家户 | 个体 | 家户 | 个体 | 家户 | 个体 | 家户 |
| 是否参保 | -1.063*** | -1.123*** | -1.687*** | -1.663*** | 0.624*** | 0.540*** | -0.009 | -0.018 |
| | (-6.16) | (-6.07) | (-10.86) | (-9.94) | (9.16) | (7.43) | [-0.65] | [-1.17] |
| 性别 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | -0.125*** | -0.124*** |
| | (.) | (.) | (.) | (.) | (.) | (.) | [-16.39] | [-11.73] |
| 年龄 | 0.105*** | 0.113*** | 0.118*** | 0.121*** | -0.013* | -0.008 | 0.000 | -0.000 |
| | (5.43) | (5.39) | (6.76) | (6.38) | (-1.65) | (-0.97) | [0.08] | [-0.82] |
| 受教育程度 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | -0.054*** | -0.084*** |
| | (.) | (.) | (.) | (.) | (.) | (.) | [-3.57] | [-5.33] |
| 婚姻状况 | -1.447*** | -1.457*** | -1.419*** | -1.422*** | -0.028 | -0.035 | -0.099*** | -0.109*** |
| | (-4.50) | (-4.75) | (-4.90) | (-5.13) | (-0.22) | (-0.29) | [-7.55] | [-8.06] |

| 是否有慢性病 | 0.880*** | 0.871*** | 0.683*** | 0.696*** | 0.196*** | 0.175** | 0.191*** | 0.199*** |
|---|---|---|---|---|---|---|---|---|
| | (5.28) | (4.45) | (4.56) | (3.94) | (2.99) | (2.28) | [20.13] | [16.16] |
| 是否参加医疗保险 | 0.003 | 0.077 | 0.171 | 0.213 | -0.169* | -0.136 | -0.022 | -0.036 |
| | (0.01) | (0.28) | (0.80) | (0.85) | (-1.79) | (-1.25) | [-1.10] | [-1.46] |
| 家庭规模 | -0.113*** | -0.098*** | -0.141*** | -0.129*** | 0.028** | 0.031** | -0.008*** | -0.011*** |
| | (-3.33) | (-2.65) | (-4.59) | (-3.87) | (2.05) | (2.16) | [-3.18] | [-3.86] |
| 人均收入 | -0.015 | -0.018 | 0.002 | 0.002 | -0.017 | -0.020 | -0.026*** | -0.028*** |
| | (-0.49) | (-0.53) | (0.07) | (0.07) | (-1.42) | (-1.53) | [-10.68] | [-9.78] |
| 地区 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.041*** | 0.042*** |
| | (.) | (.) | (.) | (.) | (.) | (.) | [9.32] | [8.50] |
| 个体固定效应 | 是 | 是 | 是 | 是 | 是 | 是 | 是 | 是 |
| 样本量 | 14476 | 14476 | 14476 | 14476 | 14476 | 14476 | 14476 | 14476 |

注：当因变量为 CES-D 量表总得分或各因子得分时，使用系数估计值；当因变量为是否有抑郁症状时，使用平均边际效应；***、**、*分别表示在 1%、5%和 10%的水平上显著；( )中为 t 值，[ ]中为 z 值。

3.异质性分析

（1）性别

为了进一步探究养老保险对农村居民心理健康的影响，对性别分样本进行异质性分析，结果如表 4.4 所示。无论是男性居民还是女性居民，参加养老保险都可以使其 CES-D 得分显著降低（男性：-1.095；女性：-1.022），且主要体现在"抑郁维度"的得分降低上，而在反向计分的"积极维度"的得分都略有升高；虽然出现抑郁症状的概率有所降低，但是并不显著。对比男性和女性的 IV-FE 结果来看，参加养老保险对于男性农村居民心理健康的影响虽然大于对于女性农村居民心理健康的影响，但是差异并不显著，可能的原因是养老保险对农村劳动力供给有负面影响，男性作为劳动力的主力较之女性会有更明显的负担减轻，但女性可以在减轻劳动负担后有更多的时间参加社交活动并能从中受益更多，因此总体来看养老保险在男女之间的政策效力并无显著差异(Kim et al., 2015；吴海青等，2020)。

表 4.4：养老保险对不同性别农村居民心理健康的影响

| | CES-D 总得 | "抑郁情绪" | "积极情绪" | 抑郁症状 |
|---|---|---|---|---|

| | 分 | | | |
|---|---|---|---|---|
| 男性 | -1.095*** | -1.689*** | 0.594*** | -0.007 |
| | (-4.87) | (-8.44) | (6.33) | [-0.39] |
| 女性 | -1.022*** | -1.676*** | 0.653*** | -0.010 |
| | (-3.88) | (-7.00) | (6.61) | [-0.48] |
| 男性 vs.女性 | -0.072 | -0.014 | -0.059 | / |

注：当因变量为 CES-D 量表总得分或各因子得分时，使用系数估计值；当因变量为是否有抑郁症状时，使用平均边际效应；***、**、*分别表示在 1%、5%和 10%的水平上显著；( )中为 t 值，[ ]中为 z 值。本表回归控制了个体固定效应和除性别外所有的控制变量，限于篇幅未予列示。

（2）年龄

考虑到 60 岁是可以开始领取养老金的年龄，因此以 60 岁为界限对年龄分样本进行异质性分析，结果如表 4.5 所示。结果显示：参加养老保险可以使 60 岁以上农村居民的 CES-D 得分显著降低 1.057，且在 1%的水平上显著；使 60 岁以下农村居民的 CES-D 得分显著降低 0.374，且在 10%的水平上显著；此外，政策效力在两个年龄群体之间存在显著差异(p = 0.055)。具体而言，对于 60 岁及以上农村居民，参加养老保险虽然使其在反向计分的"积极情绪"维度上的得分略有增加，但能显著降低其在"抑郁维度"上的得分，因此总的心理健康水平得到显著改善；但是对 60 岁以下的农村居民来说，参加养老保险的确可以降低其在"抑郁情绪"维度上的得分，但是也会显著增加其在反向计分的"积极情绪"维度上的得分，使得总的心理健康水平改善程度并不高；而且养老保险对 60 岁以下居民在"积极情绪"方面的影响显著高于对 60 岁及以上居民的影响，而在"抑郁情绪"方面并无显著差异。此外，参加养老保险对居民出现抑郁症状的比例的影响在两个年龄群体中都不显著。出现上述情况可能的原因是，对于 60 岁以下的居民而言仍然有缴费压力存在，这对农村居民而言是一笔不小的负担，从而减弱了养老保险对心理健康的积极影响；而对于 60 岁及以上的居民，可以直接享受领取养老金的好处，因此对心理健康的改善更显著(周钦等，2018)。

表 4.5：养老保险对不同年龄农村居民心理健康的影响

| | CES-D 总得分 | "抑郁情绪" | "积极情绪" | 抑郁症状 |
|---|---|---|---|---|
| 60 岁以下 | -0.374* | -1.122*** | 0.749*** | 0.009 |
| | (-1.74) | (-5.86) | (8.70) | [0.51] |
| 60 岁及以上 | -1.057*** | -1.449*** | 0.392*** | -0.025 |

| | | | | |
|---|---|---|---|---|
| | (-4.20) | (-6.36) | (3.95) | [-1.19] |
| 60 岁及以上 vs.60 岁以下 | -0.683* | -0.327 | -0.356** | / |

注：当因变量为 CES-D 量表总得分或各因子得分时，使用系数估计值；当因变量为是否有抑郁症状时，使用平均边际效应；***、**、*分别表示在 1%、5%和 10%的水平上显著；( )中为 t 值，[ ]中为 z 值。本表回归控制了个体固定效应和除年龄外所有的控制变量，限于篇幅未予列示。

（3）婚姻

为了进一步探究养老保险对农村居民心理健康的影响，对婚姻状况分样本进行异质性分析，结果如表 4.6 所示。结果显示：对于不同婚姻状况的农村居民而言，参加养老保险对其出现抑郁症状的概率并无显著影响。但是参加养老保险可以使已婚或同居的农村居民的 CES-D 得分降低 1.054 分，且在 1%的水平上显著；具体而言，参加养老保险可以显著减少其在"抑郁维度"上的得分，同时也会增加其在反向计分的"积极情绪"维度上的得分。而参加养老保险对于处在未婚、离异、丧偶或分居状态的农村居民来说并不能显著降低其 CES-D 的得分；具体来说，参加养老保险的确可以降低其在"抑郁情绪"维度上的得分，但是也会使其在反向计分的"积极情绪"维度上的得分略有增加，使得总的心理健康水平并没有显著改善。尽管如此，参加养老保险的政策效力在两种婚姻状态的群体中并无显著差异，可能的原因是以家庭为单位来看，对于已婚或同居的人而言往往意味着双倍的社会保障，参保的受益可以与其伴侣分享，因此表现出显著的心理健康水平的改善；但同时，对于未婚、离异、丧偶或分居的居民而言，养老保险为其独立生活提供了必要的支持，因此也可以适当保护其心理健康(Engelhardt et al., 2005)；这就使得二者之间的差异并不显著。

表 4.6：养老保险对不同婚姻状况的农村居民心理健康的影响

| | CES-D 总得分 | "抑郁情绪" | "积极情绪" | 抑郁症状 |
|---|---|---|---|---|
| 已婚或同居 | -1.054*** | -1.738*** | 0.684*** | -0.013 |
| | (-5.86) | (-10.76) | (9.50) | [-0.89] |
| 未婚、离异、丧偶或分居 | -0.852 | -1.265** | 0.413* | 0.028 |
| | (-1.25) | (-2.03) | (1.71) | [0.58] |
| 已婚或同居 vs. 未婚、离异、丧偶或分居 | -0.202 | -0.473 | 0.271 | / |

注：当因变量为 CES-D 量表总得分或各因子得分时，使用系数估计值；当因变量为是否有抑郁症状时，使用平均边际效应；***、**、*分别表示在 1%、

5%和10%的水平上显著；( )中为t值，[ ]中为z值。本表回归控制了个体固定效应和除婚姻状况外所有的控制变量，限于篇幅未予列示。

（4）家庭经济情况

为了进一步探究养老保险对农村居民心理健康的影响，对家庭经济情况分样本进行异质性分析，结果如表 4.7 所示。结果显示：对于不同收入群体的农村居民而言，参加养老保险对其出现抑郁症状的概率并无显著影响。但是参加养老保险可以显著降低中低收入群体的农村居民的 CES-D 得分（中等收入：-1.281；低收入：-1.485），具体而言，参加养老保险可以显著减少其在"抑郁维度"上的得分，同时也会增加其在反向计分的"积极情绪"维度上的得分；但是对于高收入群体的农村居民而言，参加养老保险虽然同样可以减少"抑郁情绪"维度得分、增加"积极情绪"维度得分，但是整体而言，对于 CES-D 得分并无显著影响。对比不同收入群体农村居民参加养老保险的影响，低收入群体从中受益最多，其次是中等收入群体，而高收入群体并无显著变化，但是只有在中、高收入群体和低、高收入群体之间才存在显著差异，低、中收入群体之间并无显著差异。可能的原因是养老金收入在更为贫困的群体收入中占比更高，因此养老保险可以为贫困群体放宽财务约束，从而对其心理健康状况的改善起到更大的作用(Chen et al., 2019)；此外，截至 2017 年底，城乡居民养老保险年人均待遇 1500 元 1，占中、低收入群体年均收入的比例较大（中：16.26%；低：42.60%），占高收入群体年均收入的比例较小（5.26%），因此前两者从养老保险中受益的差异并不显著，而高收入群体与他们相比则有显著差异。

表 4.7：养老保险对不同收入群体的农村居民心理健康的影响

| | CES-D 总得分 | "抑郁情绪" | "积极情绪" | 抑郁症状 |
|---|---|---|---|---|
| 高收入群体 | -0.167 | -0.957*** | 0.789*** | 0.005 |
| | (-0.51) | (-3.32) | (5.69) | [0.22] |
| 中等收入群体 | -1.281*** | -1.876*** | 0.595*** | -0.005 |
| | (-4.47) | (-7.29) | (5.14) | [-0.19] |
| 低收入群体 | -1.485*** | -2.064*** | 0.579*** | -0.025 |
| | (-5.02) | (-7.63) | (5.37) | [-1.08] |
| 低 vs.中 | -0.204 | -0.188 | -0.016 | / |
| 中 vs.高 | -1.113** | -0.919** | -0.194 | / |

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1 数据来源：国家人力资源和社会保障局

| 低 vs.高 | -1.317*** | -1.108** | -0.21 | / |

注：当因变量为 CES-D 量表总得分或各因子得分时，使用系数估计值；当因变量为是否有抑郁症状时，使用平均边际效应；***、**、*分别表示在 1%、5%和 10%的水平上显著；( )中为 t 值，[ ]中为 z 值。本表回归控制了个体固定效应和除家庭经济情况外所有的控制变量，限于篇幅未予列示。

（5）地区

对地区分样本进行异质性分析，结果如表 4.8 所示。结果显示：无论是东部、中部、西部还是东北地区，参加养老保险虽然对于降低出现抑郁症状的概率都没有显著作用，但是都可以显著降低该地区农村居民的 CES-D 得分（东部：-1.258；中部：-1.315；西部：-0.713；东北：1.640），具体来说会降低其在"抑郁情绪"维度上的得分，增加其在反向计分的"积极情绪"维度上的得分。对比四个地区 IV-FE 结果的差异，东北地区享受参加养老保险对心理健康的积极影响最多，其次是中部地区和东部地区，最后是西部地区；但是只有东西地区在"抑郁情绪"维度上存在显著差异，其他差异均不显著。可能的原因是地方人民政府可以根据实际情况适当提高养老金标准，因此东部地区往往会享受更为丰厚的养老金待遇，也拥有更为完善的养老保障体系，从而使东部地区老年人相对于最为贫困的西部地区老年人而言，对未来的担忧明显降低更多。

表 4.8：养老保险对不同地区农村居民心理健康的影响

| | CES-D 总得分 | "抑郁情绪" | "积极情绪" | 抑郁症状 |
| --- | --- | --- | --- | --- |
| 东部地区 | -1.258*** | -2.020*** | 0.762*** | -0.008 |
| | (-3.89) | (-6.97) | (5.48) | [-0.31] |
| 中部地区 | -1.315*** | -1.854*** | 0.539*** | -0.016 |
| | (-4.22) | (-6.70) | (4.46) | [-0.63] |
| 西部地区 | -0.713** | -1.362*** | 0.649*** | -0.010 |
| | (-2.44) | (-5.10) | (5.90) | [-0.45] |
| 东北地区 | -1.640** | -1.898*** | 0.258 | -0.001 |
| | (-2.19) | (-2.89) | (0.89) | [-0.02] |
| 东北 vs.中部 | -0.325 | -0.044 | -0.281 | / |
| 中部 vs.东部 | -0.057 | 0.166 | -0.222 | / |
| 东部 vs.西部 | -0.545 | -0.658* | 0.112 | / |

| | | | | |
|---|---|---|---|---|
| 东北 vs.东部 | -0.382 | 0.121 | -0.504 | / |
| 东北 vs.西部 | -0.927 | -0.536 | -0.391 | / |
| 中部 vs.西部 | -0.602 | -0.492 | -0.110 | / |

注：当因变量为 CES-D 量表总得分或各因子得分时，使用系数估计值；当因变量为是否有抑郁症状时，使用平均边际效应；***、**、*分别表示在 1%、5%和 10%的水平上显著；( )中为 t 值，[ ]中为 z 值。本表回归控制了个体固定效应和除地区外所有的控制变量，限于篇幅未予列示。

4.中介效应

考虑到对于含有工具变量的模型，无法使用 Bootstrap 法进行中介效应检验；前文基准分析中 FE 模型与 IV-FE 模型下结果一致，因此中介效应以 FE 模型为基础来进行。

（1）生活满意度

以"生活满意度"为中介变量进行中介效应检验，结果如表 4.9 所示。

以 CES-D 总得分为因变量，系数$\beta$, $\beta_3$, $\beta_4$, $\beta_5$显著，且$\beta_3\beta_4$和$\beta_5$同号，表示存在部分中介效应，总效应为-0.316，直接效应为-0.262，间接效应为-0.054，效应量为 17.20%；以"抑郁情绪"维度得分为因变量，系数$\beta$, $\beta_3$, $\beta_4$, $\beta_5$显著，且$\beta_3\beta_4$和$\beta_5$同号，表示存在部分中介效应，总效应为-0.522，直接效应为-0.482，间接效应为-0.040，效应量为 7.63%；以"积极情绪"维度得分为因变量，系数$\beta$, $\beta_3$, $\beta_4$, $\beta_5$显著，且$\beta_3\beta_4$和$\beta_5$异号，表示存在遮掩效应，总效应为 0.206，直接效应为 0.220，间接效应为-0.015，效应量为 6.60%。综上，"生活满意度"是养老保险对农村居民心理健康影响的中介变量；$\beta_3$为正$\beta_4$为负，表示参加养老保险可以通过提高农村居民的生活满意度，进而促进其心理健康。事实上，养老保险对于农村居民而言，是使其"老有所养"的重要举措，是除了"家庭养老"和"土地养老"之外重要的养老方式，可以为其提供稳定的收入预期，因此对于提高其生活满意度非常重要，而对生活满意正是好的心理健康状况的体现。

表 4.9：中介效应检验之"生活满意度"

|  | CES-D 总得分 | "抑郁情绪" | "积极情绪" |
| --- | --- | --- | --- |
| $\boldsymbol{\beta}$ | -0.316$^{***}$ | -0.522$^{***}$ | 0.206$^{***}$ |
| （总效应） | (-3.14) | (-5.79) | (5.19) |
| $\boldsymbol{\beta_3}$ | 0.037$^{***}$ | 0.037$^{***}$ | 0.037$^{***}$ |
|  | (2.63) | (2.63) | (2.63) |
| $\boldsymbol{\beta_4}$ | -1.451$^{***}$ | -1.063$^{***}$ | -0.388$^{***}$ |
|  | (-21.87) | (-17.76) | (-14.69) |
| $\boldsymbol{\beta_5}$ | -0.262$^{***}$ | -0.482$^{***}$ | 0.220$^{***}$ |
| （直接效应） | (-2.65) | (-5.42) | (5.61) |
| $\boldsymbol{\beta_3\beta_4}$ | -0.054 | -0.040 | -0.015 |
| （间接效应） | | | |
| 效应量 | 17.20% | 7.63% | 6.60% |

注：$^{***}$、$^{**}$、$^{*}$分别表示在 1%、5%和 10%的水平上显著；( )中为 t 值。本表回归控制了个体固定效应和所有的控制变量，限于篇幅未予列示。

（2）社会资本

以"社会资本"为中介变量进行中介效应检验，结果如表 4.10 所示。以 CES-D 总得分为因变量，系数$\boldsymbol{\beta}$，$\boldsymbol{\beta_3}$，$\boldsymbol{\beta_4}$，$\boldsymbol{\beta_5}$显著，且$\boldsymbol{\beta_3\beta_4}$和$\boldsymbol{\beta_5}$同号，表示存在部分中介效应，总效应为-0.316，直接效应为-0.296，间接效应为-0.020，效应量为 6.36%；以"抑郁情绪"维度得分为因变量，系数$\boldsymbol{\beta}$，$\boldsymbol{\beta_3}$，$\boldsymbol{\beta_4}$，$\boldsymbol{\beta_5}$显著，

且$\boldsymbol{\beta_3\beta_4}$和$\boldsymbol{\beta_5}$同号，表示存在部分中介效应，总效应为-0.522，直接效应为-0.507，间接效应为-0.015，效应量为 2.80%；以"积极情绪"维度得分为因变量，系数$\boldsymbol{\beta}$，$\boldsymbol{\beta_3}$，$\boldsymbol{\beta_4}$，$\boldsymbol{\beta_5}$显著，且$\boldsymbol{\beta_3\beta_4}$和$\boldsymbol{\beta_5}$异号，表示存在遮掩效应，总效应为 0.206，直接效应为 0.211，间接效应为-0.005，效应量为 2.60%。综上，"社会资本"是养老保险对农村居民心理健康影响的中介变量；$\boldsymbol{\beta_3}$为正$\boldsymbol{\beta_4}$为负，表示参加养老保险可以通过增加农村居民的社会资本，进而促进其心理健康。养老金计划使得农村居民可以减少劳动时间从而有更多的机会

参与社会活动(Kim et al., 2015)，也就有利于其个人社会资本的增加，而个人层面的社会资本是老年心理健康的重要保护因素(Lin et al., 2019; Norstrand & Xu, 2012; 温兴祥等, 2017)。

表 4.10：中介效应检验之"社会资本"

| | CES-D 总得分 | "抑郁情绪" | "积极情绪" |
|---|---|---|---|
| $\beta$（总效应） | -0.316*** | -0.522*** | 0.206*** |
| | (-3.14) | (-5.79) | (5.19) |
| $\beta_3$ | 0.112*** | 0.112*** | 0.112*** |
| | (4.91) | (4.91) | (4.91) |
| $\beta_4$ | -0.179*** | -0.130*** | -0.049*** |
| | (-4.24) | (-3.44) | (-2.94) |
| $\beta_5$（直接效应） | -0.296*** | -0.507*** | 0.211*** |
| | (-2.94) | (-5.62) | (5.33) |
| $\beta_3\beta_4$（间接效应） | -0.020 | -0.015 | -0.005 |
| 效应量 | 6.36% | 2.80% | 2.60% |

注： ***、**、*分别表示在 1%、5%和 10%的水平上显著；( )中为 t 值。本表回归控制了个体固定效应和所有的控制变量，限于篇幅未予列示。

## 五、结论与政策建议

本文基于中国健康与养老追踪调查（CHARLS）2011、2013、2015 和 2018 年的数据，采用工具变量-固定效应模型(IV-FE)评估了养老保险对农村中老年居民心理健康的影响，并进行了异质性分析和中介效应分析。

研究表明：首先，参加养老保险可以显著提升农村中老年居民的心理健康水平，尤其是可以显著减少其抑郁情绪，但是政策效力有限，不能显著降低出现抑郁症状的概率；其次，该影响具有异质性，男女居民、60 岁及以上和 60 岁以下老人、已婚或同居群体、中等或低收入群体和各地区居民参加养老保险都可以显著改善其心理健康水平，但是在不同性别之间、不同婚姻状况之间、不同地区之间并无显著差异，而对于 60 岁及以上老人、中低收入群体则表现出更明显的积极影响；最后，从传导路径看，参加养老保险可以通过提高生活满意度和增加社会资本，进而降低农村中老年居民的抑郁程度，促进心理健康。

从本文的实证研究结果来看，养老保险对于农村中老年居民心理健康水平的提高具有重要影响，因此未来有必要进一步完善养老保险制度从而更好地发挥其作用。本文提出以下政策建议：

第一，逐步提高基本养老金待遇。本文研究发现，养老保险会农村中老年居民心理健康的影响有限，不能显著减少其出现抑郁症状的概率。事实上，虽然农村保险制度实施以来，养老金水平在稳步提高，但仍存在增长幅度不及农村地区物价增长率的情况；2020年，城乡居民基本养老保险月人均待遇水平约170元，仅仅相当于农村低保平均保障标准的34.9%，农村居民人均消费支出的14.9%，难以对农村居民的日常生活产生实质影响。未来还需要结合当前经济发展水平与物价变动情况，全面落实城乡居民基本养老保险待遇确定和基础养老金正常调整机制，逐步提高城乡居民基础养老金标准，为农村居民提供更好的经济保障。

第二，强化缴费激励机制，提高参保补贴标准。本文研究发现，养老保险仅对60岁及以上居民有积极作用，且政策效力在两个年龄群体之间的差异边际显著，这背后恰恰反映了缴费与领取阶段的农村居民的差异。对于农村居民而言，养老保险的缴费仍是一笔不小的经济压力，这也在一定程度上导致其参保积极性不高，即便参保也普遍选择按照较低档次缴费。因此未来有必要切实根据农村居民生活现状，建立农村养老保险缴费标准的动态调整机制，如针对困难群体或在特殊时期适当降低缴费金额，并提高农村居民的参保补贴标准，使其可以在获取未来养老保障的前提下又不会过分加重当前经济负担，增加制度吸引力。

第三，逐步实施城乡居民养老保险省级统筹、全国统筹。2022年1月1日起企业职工基本养老保险全国统筹开始实施，资金在全国范围内互济余缺。但城乡居民养老保险至今尚未实现省级统筹，统筹层次低，加上各地经济发展水平、财政支持力度的差异，使得不同地区之间城乡居民养老保险待遇出现明显差异，且对于东北等老龄化严重的地区，养老保险基金的支出压力较大。不同地区之间养老保险基金结构性矛盾日益突出，因此应尽快学习企业职工基本养老保险全国统筹的经验，建立城乡居民养老保险省级统筹机制，并逐步实现全国统筹，在更大范围内调剂使用基金，确保养老金按时足额发放，且适当缩小养老金待遇的地区差异。

第四，建立农村老年居民活动中心，鼓励参与社交活动。本文研究指出，参加养老保险可以通过增加农村中老年居民的社会资本，进而促进其心理健康。对于农村居民而言，养老金计划在一定程度上将其从农业耕作等劳动中解放出来，为其老年生活提供了经济支持，这就使其有机会更多地参与社交活动。但是目前农村地区老年活动中心建设不完善，使得老年居民缺乏足够的活动场所和设施，不能完全释放养老保险政策以社会资本为媒介改善心理健康状况的效力。因此未来有必要在广大农村地区建立老

年居民活动中心，为其提供娱乐、互助、继续教育等场所，增进邻里关系，构建社会网络，增加社会资本，从而更好地促进养老保险对农村居民心理健康的积极影响。

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# 老农保参保经历与养老金预期

吕有吉、郑伟、谢志伟 [1]

**摘要：**

　　基于 2011-2018 年中国健康与养老追踪调查（CHARLS）数据，本文构建了养老金预期偏差和养老金准确预期者两个变量衡量个体的养老金预期，并采用 PSM 方法检验老农保参保经历对个体养老金预期的影响。研究发现，老农保参保经历使得个体养老金预期偏差程度降低 0.191，其成为准确预期者的概率会提高 3.4 个百分点。同时，机制分析表明，老农保参保经历降低了个体的极端乐观预期，并提升了个体对养老金的认知能力，进而改善个体对未来养老金的准确预期程度。

**关键词：** 养老金预期，个人经历，农村社会养老保险

## 一、引言

　　近年来，伴随人口年龄结构的加速转型，养老议题日益成为社会关注的焦点，公共养老金制度，特别是居民基本养老保险制度历经多次改革，已经进入提质增效的改革深水区。公共养老金制度的完善不能就制度论制度，而应在系统观念指导下准确把握制度设计与居民决策之间的互动机制，深刻理解制度运行的底层逻辑。已有研究表明，养老金预期是连接公共养老金制度与个人决策的关键路径，并通过储蓄（Chetty et al.，2014）和劳动供给（Chan and Stevens，2004）等宏微观经济变量直接影响个人福利。错误预期未来养老金会导致个人决策偏离最优路径，从而无法实现一生福利最大化，并削弱公共养老金制度改革的福利效应（Bottazzi et al.，2006；Baldini et al.，2019）。因此，有必要分析影响养老金预期的各类因素，为进一步完善公共养老金制度提供决策参考。

　　已有研究指出，个人的信息掌握程度、处理信息的能力和主观态度均会影响其养老金预期。一方面，如果个人无法掌握与未来养老金相关的政

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策变动、宏观趋势等信息，就无法在此基础上对未来养老金做出理性预期，从而导致养老金预期偏差（Mastrobuoni，2011；Dolls et al.，2018；Bai et al.，2021）。另一方面，养老金作为一类典型的金融资产，需要个人掌握必要的金融知识及与保险相关的特定知识，才能够理解相关制度安排，否则即使个人掌握了必要的信息，也无法对这些信息进行有效的加工处理，从而导致养老金预期偏差（Gustman and Steinmeier，2005；Alessie et al.，2011；Barrett et al.，2015）。此外，个人对养老金制度的信任程度（Van Duijn et al.，2013；Okumura and Usui，2014）和个人对未来的乐观程度（Bissonnette and Van Soest，2015）等主观态度因素也会影响个人的养老金预期。特别地，当养老金制度频繁变动时，更可能出现信息更新不及时和处理信息能力滞后等问题，从而影响个人养老金预期的准确度（Delavande and Rohwedder，2011；Guiso et al.，2013）。由此可以预期，历经多次改革的居民基本养老保险的参保者更可能经历由养老金预期偏差导致的福利损失。

经济学和社会学的研究表明，个人成长经历和社会履历会对个人决策产生长期影响（以下本文称之为经历学习效应）。在庞杂的信息面前，有限理性的个人总是在自身认知能力和价值观的影响下，选择性地过滤掉部分信息并做出各项决策，而成长经历和社会履历会塑造个人的认知能力和价值观（蔡庆丰等，2016；Ehling et al.，2018；Malmendier et al.，2020）。就成长经历而言，个人成长过程中经历的宏观经济波动，比如通货膨胀（Malmendier and Nagel，2016）、经济衰退（Knüpfer et al.，2017）、大规模失业（Kuchler and Zafar，2019），自然灾害（Gao et al.，2020；Chen et al.，2021）等均会影响个人成年后在相似情形下所做的决策。就社会履历而言，个人成年后的参军（赖黎等，2016）、投资（Choi et al.，2009；Seru et al.，2010；Chiang et al.，2011）、履职（姜付秀等，2016；戴维奇等，2016；周楷唐等，2017；何瑛等，2019）等经历均会显著影响个人之后的投资、经营、管理等决策。农村社会养老保险（以下简称"老农保"）是我国居民基本养老保险制度的"先行者"，自1991年试点至1999年叫停，九年间历经探索发展和整顿调整，顶峰时期曾覆盖8000万人，无疑对我国居民和后续的居民基本养老保险制度产生了深远影响。然而，鲜有人关注老农保参保经历如何影响个人的养老金预期，更缺乏对上述作用的实证刻画。

具体而言，老农保参保经历可能会从正反两个方面影响个人的养老金预期。一方面，1999年劳社部宣布将中国人民银行公布的一年期定期存款利率作为退保计息标准，这远低于老农保之前承诺的个人账户计息利率；部分地区甚至只清退本金，不付利息，使得老农保参保者遭受到了实际上的储蓄收益损失（郑伟，2007），这损害了政府的公信力。而居民基本养老保险的政策实施效果很大程度上依赖于居民对政府和相关政策的信任程度（Giné and Yang，2009；高明等，2021），由此可以预见，遭受损失的

老农保参保者更可能会低估居民基本养老保险制度的保障程度。另一方面，老农保将养老保险的概念引入广大居民，让居民对养老保险的运作机制有了浅显的认知，为后续居民基本养老保险的实施打下了认知基础（李轩红，2011）。从这一角度出发，老农保参保者可能会因为养老金认知水平的提升及对制度相关信息的额外关注而作为较为准确的养老金预期。综上所述，我们难以直观判别老农保参保经历对个人养老金预期的影响，也无法确定上述影响发挥作用的渠道，需要通过实证研究加以分析。

本文使用中国健康与养老追踪调查（CHARLS）2011-2018 年的数据，利用倾向得分匹配（Propensity Score Matching，PSM）方法，检验了老农保参保经历对个体养老金预期的影响。结果显示，如果个人曾经参加过老农保，其养老金预期偏差程度会降低 0.191，其成为准确预期者的概率会提高 3.4 个百分点。在更换不同匹配方法、不同样本窗口期以及将样本缩小至农村户口群体时，上述结论依旧稳健。同时，机制分析的结果表明，老农保参保经历主要通过"降低极端乐观预期"和"提升认知能力"两条路径改善个体的养老金预期偏差。

相较于现有研究，本文的贡献主要体现在以下三个方面：第一，在研究范畴上，本文将个人老农保参保经历纳入养老金预期的影响因素，为养老金预期偏差提供了新的解读视角，为经历学习效应发掘了新的应用场景。第二，在研究结论上，本文证明老农保参保经历会显著改善居民基本养老保险参保者的养老金预期准确程度，证实了老农保在提升居民养老金认知水平方面的积极作用。第三，在机理分析上，本文提出了老农保参保经历影响养老金预期的"降低极端乐观预期"和"提升认知能力"两条路径，并通过实证方式予以检验，证实了两条老农保参保经历的影响机制。

本文剩余内容安排如下：第二部分梳理政策背景并提出理论假说；第三部分为研究设计，介绍数据、变量构造和模型设定；第四部分汇报实证检验的回归结果；第五部分总结本文的研究结论并提出相应政策建议。

## 二、政策背景

居民基本养老保险的发展历程可以分为两个阶段，第一个阶段为老农保制度的建设与整顿，第二个阶段是新农保和城居保制度的兴起及整合。

1991 年，民政部在山东等地进行了首批农村社会养老保险试点。1992 年 1 月，民政部正式出台《县级农村社会养老保险基本方案（试行）》，开始在全国范围内推广老农保。截至 1997 年底，老农保覆盖的农村居民已达 7452 万人，[1]发展势头迅猛。另一方面，由于各地的老农保制度资金管理层级低，部分地区有强制摊派或高息揽储以提高参保率的现象，老农保制度存在一定的债务风险（胡晓义，2019）。伴随着 1997 年亚洲金融危机爆

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发，1999 年 7 月国务院发布《国务院批转整顿保险业工作小组保险业整顿与改革方案的通知》，明确指出"目前我国农村尚不具备实行社会保险的条件"，要求对老农保"进行清理整顿，停止接受新业务"。自此，老农保陷入停滞和整顿阶段。

随着养老金制度改革的逐步深化，对覆盖城乡居民的基本养老保险制度的探索被再次提上日程。2009 年国务院发布《关于开展新型农村社会养老保险试点的指导意见》，为农村居民提供基本的养老保障，到 2012 年，全国参加新农保（不含城居保）的人数达到 4.63 亿，覆盖面显著扩大。2011 年国务院发布《关于开展城镇居民社会养老保险试点的指导意见》，将不符合职工基本养老保险参保条件的城镇非从业居民首次纳入到基本养老保险的覆盖范围中来。由于新农保和城居保在缴费方式和待遇计发方法上均较为相似，2014 年国务院发布《关于建立统一的城乡居民基本养老保险制度的意见》，将两种制度进行合并为统一的"城乡居保"。至此，居民基本养老保险制度基本建立。

老农保的失败直接导致了参保人的利益受损，在清退老农保时计息标准从承诺的个人账户利率改为一年期定期存款利率，部分地区甚至只退本金而不付利息（郑伟，2007）。由此可以预期，个体在参与老农保制度后对未来养老金的预期会更加悲观，即降低对未来养老金的预期。但需要说明的是，个体降低对未来养老金的预期并不一定导致养老金预期偏差的上升。为便于讨论，本文将预期养老金高于依据政策计算的真实养老金的个体，称作乐观预期的个体，反之则为悲观预期的个体。养老金预期偏差的变化，与个体自身的预期乐观与否以及由老农保制度导致的预期变动程度有关。首先，若个体自身为悲观参保者，则老农保参保经历会导致个体的悲观预期进一步加深，从而造成养老金预期偏差上升。其次，若个体最初为乐观参保者，在经历了老农保失败后，预期养老金的降低可能导致个体乐观程度下降，也可能导致个体由乐观预期转为悲观预期。这两种情形下，养老金预期偏差的变化方向并不确定。基于上述逻辑推演，本文提出假说 1a 和假说 1b。

**假说 1a**：相比于未参加老农保的个体，参加过老农保的个体会增加养老金预期偏差；

**假说 1b**：相比于未参加老农保的个体，参加过老农保的个体会缩小养老金预期偏差。

但不可否认的是，老农保制度的失败意味着政府公信力的下降，导致居民降低对新农保制度的信任程度（高明等，2021），而居民基本养老保险的政策实施效果很大程度上依赖于居民对政府和相关政策的信任程度（Giné and Yang，2009）。因此，老农保失败会降低参保者对未来养老金的极端乐观预期，并进而影响其养老金预期偏差程度。据此本文提出假说

2。

**假说 2**：老农保参与经历会通过降低参保者的极端乐观预期进而影响参保者的养老金预期偏差程度。

另外，老农保参保经历作为个体成长经历之一，往往会对个体决策产生长期影响，形成经历学习效应。具体而言，老农保是首个覆盖农村居民的养老保险制度，增加了农村居民对养老保险运作机制的认知（李轩红，2011），有利于农村居民提升养老金预期能力。同时，个体对外部信息的处理和学习能力是影响养老金预期偏差的重要因素（Mitchell，1988；Alessie et al.，2011；Barrett et al.，2015），个体的学习能力越强，对未来养老金预期越准确。而个人的教育背景和人力资本水平与个体的信息处理能力和学习能力呈正相关。由此本文推测，老农保参保经历会通过影响个体的养老金认知能力进而影响养老金预期偏差，表现为教育背景越高的个体，老农保造成的影响越大。综上，本文提出假说 3。

**假说 3**：老农保参与经历会通过提升参保者的养老金认知能力进而影响参保者的养老金预期偏差程度。

## 三、研究设计

（一）数据来源

本文实证分析所用数据来源于 CHARLS。CHARLS 于 2011 年开展首次基线调查，随后在 2013 年、2015 年和 2018 年进行了三次追访。CHARLS 在个人、家庭和社区层面搜集了丰富的信息，包括受访者养老保险参保和缴费情况、家庭内部转移支付、工作时间及收入、主要的人口特征等。此外，CHARLS 在 2014 年进行了"中国居民生命历程调查"，对 2011 年和 2013 年的所有在世受访者的参保历史、工作历史、教育历史等生命历程信息进行了追访，为研究受访者过往经历的长期影响提供了良好的数据基础。

本文分以下四个步骤筛选研究样本。第一步，鉴于现行制度下居民基本养老保险最低领取年龄为 60 岁，为确保所有样本均处于缴费期，选择年龄在 45 岁到 59 岁之间的居民基本养老保险参保者。第二步，为尽量提高本文对消费者价格指数、基础养老金最低标准等宏观数据进行预测所用假设的置信度，选择预期在 2033 年前领取养老金的样本。第三步，在清理相关变量数据后，剔除那些关键变量缺失的样本。第四步，对养老金预期偏差程度、缴费水平和当期工作收入 1% 和 99% 分位数以外的非零观测值进行缩尾处理。以上四个步骤生成本文实证分析使用的样本，共有 22296 个有效观测值，表 1 报告了该样本的描述性统计，呈现了每个变量的均值、中位数和标准差。[1]

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[1] 考虑到近十年来我国快速城镇化会导致部分居民由农村户口转向非农户口，部分非农户口的样本汇报曾参与过老农保是合理的，本文生成研究样本的过程并未将持有非

（二）变量构造

1. 养老金预期偏差

为了衡量城乡居民养老保险参保人对未来养老金的预期偏差，本文首先需要明确两个概念：预期养老金和政策养老金。其中，预期养老金是指居民基本养老保险参保者主观预期其到达领取年龄时可以从居民基本养老保险领取的养老金，CHARLS 数据库询问受访者"未来预期能领多少养老金"，本文以此作为参保者的预期养老金。政策养老金是指居民基本养老保险参保者在 60 岁时点依据现行相关政策能够从居民基本养老保险领取的养老金。考虑到老农保与其余三种制度在个人缴费、集体补助、政府补贴等制度设计方面存在较大差异，本文依据居民基本养老保险参保者是否曾经参与老农保将其分为第一类参保者（未参与老农保）及第二类参保者（曾参与老农保），并分别计算两类人群的政策养老金（具体计算过程见附录A）。

基于个人的预期养老金和政策养老金，本文构建被解释变量——养老金预期偏差程度。参考 Bottazzi et al.（2006）和 Baldini et al.（2019）的做法，本文将养老金预期偏差程度定义为：

$$养老金预期偏差程度 = \left| \frac{预期养老金 - 政策养老金}{政策养老金} \right| \tag{1}$$

在养老金预期偏差程度变量的基础上，本文进一步定义了另一个被解释变量——准确预期者，取值为 1 表示个人的养老金预期偏差程度小于等于 0.25，否则为 0。表 1 显示，养老金预期偏差程度逐年缓慢降低，由 2011 年的 0.76 下降到 2018 年的 0.74，准确预期者的比例逐年上升，由 2011 年的 5% 提高到 2018 年的 11%。由此可知，一方面，居民基本养老保险参保者的养老金预期能力普遍较弱，难以形成准确预期；另一方面，随着时间的推移，居民基本养老保险参保者的养老金预期能力得到了一定程度的增强。

2. 参与老农保

基于个人自报的老农保参保状态，本文构建核心解释变量——曾参与老农保，取值为 1 表示个人曾参与过老农保，否则为 0。由表 1 可知，在本文所用样本中，老农保参保者的比例逐年下降，由 2011 年的 18% 降低至 2018 年的 11%，这是因为部分老农保参保者在 CHARLS 的追访过程中到达了 60 岁，从而退出了本文所关注的研究样本。平均而言，每 8 位居民基本养老保险参保者中，就有 1 位曾经参与过老农保，这也侧面反映了老农保失败可能造成的广泛影响。

3. 控制变量

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农户口的观测值剔除。

如前文所述，个体主观态度也会影响养老金预期偏差。本章使用乐观参保者变量（参保者是否为乐观预期）控制个体主观态度的影响。为准确刻画老农保参保经历对养老金预期能力的作用，参考 Gustman and Steinmeier（2005）的做法，本文使用受教育年限、使用网络、子女数量等变量控制影响个人搜集信息和处理信息能力的因素。此外，本文同样将缴费水平、年龄、性别等个体和家户特征变量纳入控制变量，以控制这些变量对个人养老金预期能力的潜在影响。特别地，本文将家户成员的老农保参保状态纳入控制变量，以排除老农保参保经历在家户成员间的溢出效应。

（三）模型设定

本文首先使用式（2）估计老农保参保经历对养老金预期的影响：

$$Y_{i,t} = \alpha + \beta ORBP\ Participation_i + \gamma X_i + \delta Z_{i,t} + \chi Community_c + \phi Year_t + \varepsilon_{i,t}. \quad (2)$$

其中 $Y_{i,t}$ 表示第 $t$ 期的第 $i$ 个人的养老金预期，由养老金预期偏差程度和准确预期者两个变量衡量，$ORBP\ Participation_i$ 代表第 $i$ 个人的老农保参保经历，控制变量包括不随时变的控制变量 $X_i$ 和随时变的控制变量 $Z_{i,t}$，$Community_c$ 和 $Year_t$ 分别代表社区（村庄）固定效应和年份固定效应。由于核心解释变量为过去是否曾参加老农保，模型不存在反向因果和估计偏差问题，但遗漏变量问题仍然有可能导致模型的内生性。

本文使用 PSM 方法解决上述遗漏变量问题。以参与过老农保作为处理组，利用式（3）中的控制变量，本文使用一对四匹配方法，对处理组的个体进行一一匹配，倾向得分回归方程如下所示：

$$ORBP\ Participation_i = \alpha + \gamma X_i + \delta Z_{i,t} + \chi Community_c + \phi Year_t + \varepsilon_{i,t}. \quad (3)$$

其中变量的定义与式（2）中一致。在经过倾向得分匹配后，本文得到平均处理效应（ATT），即处理组（参与过老农保）与处理组（未参与老农保）在养老金预期偏差上的差异。同时，在稳健性检验中，本文使用了 k 近邻匹配、卡尺匹配、卡尺内最近邻匹配和核匹配以验证结果的稳健性。

## 四、实证结果

（一）主要结论

表 2 呈现了老农保参保经历对养老金预期能力的提升效应。列（1）的结果表明，不考虑其他因素对个人养老金预期偏差程度的影响，如果个人曾经参与过老农保，其养老金预期偏差程度会降低 0.070。列（2）控制了个体层面和家户层面的特征变量，上述效应增强至 0.143，表明一些个人特征和家户特征可能会作用于个人的老农保参保决策，从而影响其对养老金预期偏差程度的影响。列（3）在列（2）的基础上额外控制了年份固定效应和社区（村庄）固定效应，进一步解决了回归中可能存在的遗漏变量问题，使得上述效应下降至 0.121。列（4）至列（6）将被解释变量更换为准确预期者，得到了与列（1）至列（3）基本一致的结论。由列（6）可知，

如果个人曾经参与过老农保，其是准确预期者的概率会提高 2.0 个百分点，考虑到本文研究样本中的平均养老金预期偏差程度为 0.75，且仅 8.3% 的样本是准确预期者，老农保参保经历对养老金预期能力的提升效应在经济学意义上无疑也是显著的。

表 3 使用 PSM 方法重新检验了老农保参保经历对养老金预期能力的影响。本文首先使用一对一匹配方法进行估计，结果表明，如果居民基本养老保险参保者参与过老农保，其养老金预期偏差会显著减少 0.166，其为准确预期者的概率会显著提高 4.5 个百分点。Abadie et al.（2004）建议使用一对四匹配，以最小化均方误差。本文使用一对四匹配方法的重新进行估计得到了与一对一匹配方法下类似的结论。具体而言，如果居民基本养老保险参保者参与过老农保，其养老金预期偏差会显著减少 0.191，其为准确预期者的概率会显著提高 3.4 个百分点。为进一步检验匹配方法对本文主要结论的影响，本文同样汇报了使用卡尺内匹配方法、卡尺内最近邻匹配方法、核匹配方法估计的结果，在所有匹配方法下，本文均发现老农保参保经历可以显著降低居民基本养老保险参保者的养老金预期偏差程度。在各类核函数中，Epanechnikov 核函数能够使得积分均方误差最小化，从而得到较为可信的匹配结果。因此，在后续分析中，本文主要使用一对四匹配方法和基于 Epanechnikov 核函数的核匹配方法进行分析。

附图 1 汇报了 PSM 平衡性检验的结果。无论是使用一对四匹配方法，还是基于 Epanechnikov 核函数的核匹配方法，匹配后的老农保参保者群体和非老农保参保者群体之间在各个可观测变量上均不存在显著差异，即平稳性要求得到满足。

（二）稳健性检验

表 4 针对各种可能影响实证结果的干扰因素进行稳健性检验。首先，本文针对极端值的处理方法进行稳健性检验。为排除极端值对本文实证结果的干扰，在构建研究样本时，本文对养老金预期偏差程度、缴费水平和当期工作收入 1% 和 99% 分位数以外的非零观测值进行缩尾处理。为检验本文结论的稳健性，本文使用 PSM 方法，在未经缩尾处理的研究样本内重新进行了分析。结果表明，在一对四匹配方法下，参加过老农保的居民基本养老参保者的养老金预期偏差程度会低 0.256，且他们是准确预期者的概率会高 3.0 个百分点；在基于 Epanechnikov 核函数的核匹配方法下，上述两种效应分别变化至 0.230 和 3.5 个百分点。因此，考虑变量极端值的影响后，本文的主要结论并不改变。

其次，本文针对研究样本窗口期进行稳健性检验。居民基本养老保险在 2010 年至 2013 年期间迅速扩面，其覆盖人群从 2010 年的 1.03 亿人迅速增长至 2013 年的 4.98 亿人。本文研究样本的窗口期为 2011 年至 2018 年，与上述扩面期存在重叠。为检验本文研究结论在扩面期结束后是否仍

然稳健，本文使用 2013 年至 2018 年的研究样本重新检验了老农保参保经历对个人养老金预期能力的影响。结果表明，在上述样本内，若使用一对四匹配方法，参加过老农保的居民基本养老参保者的养老金预期偏差程度会低 0.184；若使用基于 Epanechnikov 核函数的核匹配方法，上述效应为 0.182。综上，排除居民基本养老保险扩面期影响后，主要结论依旧稳健。最后，本文针对农村户口持有者进行稳健性检验。2014 年前，居民基本养老保险以新农保和城居保两种形态分别覆盖农村户口居民和非农户口居民。虽然二者的制度设计类似，且最终于 2014 年合并为城乡居保，但仍有必要关注新农保和城居保覆盖人群的差异，特别是考虑到我国近年来城乡差异不断扩大的宏观背景和本文研究样本中非农户口持有者仅占 5%的数据特征。为此，本文使用农村户口子样本对本文研究问题重新进行了检验。结果表明，在农村户口子样本内，老农保参保经历使得个人的养老金预期偏差程度显著降低 0.208，且他们是准确预期者的概率会高 3.0 个百分点，结果较为稳健。

表 1：描述性统计

| | 2011 年 (N=4876) | | | 2013 年 (N=6011) | | | 2015 年 (N=6273) | | | 2018 年 (N=5136) | | |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
| | 均值 | 中位数 | 标准差 | 均值 | 中位数 | 标准差 | 均值 | 中位数 | 标准差 | 均值 | 中位数 | 标准差 |
| 预期养老金 | 69.57 | 38.41 | 134.60 | 95.90 | 52.21 | 190.60 | 130.40 | 64.23 | 259.90 | 162.80 | 82.42 | 299.10 |
| 政策养老金 | 191.90 | 127.00 | 228.10 | 213.60 | 135.00 | 259.80 | 246.90 | 152.10 | 304.20 | 286.40 | 171.70 | 347.60 |
| 养老金预期偏差程度 | 0.76 | 0.71 | 0.50 | 0.73 | 0.62 | 0.72 | 0.76 | 0.60 | 0.89 | 0.74 | 0.58 | 0.97 |
| 准确预期 | 0.05 | 0.00 | 0.21 | 0.08 | 0.00 | 0.27 | 0.09 | 0.00 | 0.29 | 0.11 | 0.00 | 0.32 |
| 曾参与老农保 | 0.18 | 0.00 | 0.38 | 0.14 | 0.00 | 0.35 | 0.12 | 0.00 | 0.32 | 0.11 | 0.00 | 0.31 |
| 乐观参保者 | 0.06 | 0.00 | 0.24 | 0.08 | 0.00 | 0.27 | 0.11 | 0.00 | 0.32 | 0.13 | 0.00 | 0.34 |
| 缴费水平 | 160.30 | 100.00 | 154.50 | 154.50 | 94.93 | 155.20 | 183.80 | 91.76 | 229.60 | 190.80 | 86.75 | 236.80 |
| 使用网络 | 0.01 | 0.00 | 0.08 | 0.02 | 0.00 | 0.15 | 0.05 | 0.00 | 0.23 | 0.16 | 0.00 | 0.37 |
| 受教育年限 | 5.79 | 6.00 | 4.13 | 5.98 | 6.00 | 3.99 | 6.22 | 6.00 | 3.80 | 6.56 | 6.00 | 3.58 |
| 子女数量 | 2.26 | 2.00 | 0.95 | 2.34 | 2.00 | 1.01 | 2.26 | 2.00 | 0.94 | 2.19 | 2.00 | 0.93 |
| 家户成员曾参与老农保 | 0.24 | 0.00 | 0.43 | 0.20 | 0.00 | 0.40 | 0.17 | 0.00 | 0.37 | 0.15 | 0.00 | 0.36 |
| 年龄 | 51.65 | 52.00 | 4.18 | 52.14 | 51.00 | 4.17 | 51.91 | 52.00 | 4.08 | 52.92 | 53.00 | 3.59 |
| 男性 | 0.48 | 0.00 | 0.50 | 0.48 | 0.00 | 0.50 | 0.47 | 0.00 | 0.50 | 0.46 | 0.00 | 0.50 |
| 已婚 | 0.95 | 1.00 | 0.21 | 0.95 | 1.00 | 0.22 | 0.96 | 1.00 | 0.20 | 0.95 | 1.00 | 0.22 |
| 城镇居民 | 0.26 | 0.00 | 0.44 | 0.25 | 0.00 | 0.43 | 0.27 | 0.00 | 0.44 | 0.28 | 0.00 | 0.45 |
| 非农户口 | 0.04 | 0.00 | 0.19 | 0.05 | 0.00 | 0.22 | 0.05 | 0.00 | 0.22 | 0.05 | 0.00 | 0.22 |
| 低抑郁风险 | 0.59 | 1.00 | 0.49 | 0.63 | 1.00 | 0.48 | 0.66 | 1.00 | 0.48 | 0.61 | 1.00 | 0.49 |
| 当期工作收入 | 12931.00 | 5000.00 | 20051.00 | 13823.00 | 5063.00 | 21839.00 | 16606.00 | 6652.00 | 26127.00 | 19536.00 | 9044.00 | 26962.00 |
| 拥有房产 | 0.96 | 1.00 | 0.20 | 0.96 | 1.00 | 0.21 | 0.92 | 1.00 | 0.28 | 0.92 | 1.00 | 0.27 |

注：为更好地体现样本特征，此处分别构建了受教育水平为没有接受过正式教育、可以读写、小学毕业和中学及以上的 0-1 变量，并在表中汇报其均值。

表 2：老农保参保经历与养老金预期能力

| 被解释变量 | (1) | (2) | (3) | (4) | (5) | (6) |
| --- | --- | --- | --- | --- | --- | --- |
| | 养老金预期偏差程度 | | | 准确预期者 | | |
| 回归方法 | OLS | OLS | OLS | Probit | Probit | Probit |
| 曾参与老农保 | -0.070*** | -0.143*** | -0.121*** | 0.045*** | 0.023** | 0.020** |
| | (0.020) | (0.029) | (0.031) | (0.006) | (0.009) | (0.010) |
| 乐观参保者 | | 1.224*** | 1.162*** | | 0.113*** | 0.093*** |
| | | (0.055) | (0.054) | | (0.006) | (0.007) |

| | | | | | | |
|---|---|---|---|---|---|---|
| Ln(缴费水平) | | 0.028*** | 0.026** | | -0.002 | 0.000 |
| | | (0.011) | (0.012) | | (0.004) | (0.004) |
| 使用网络 | | 0.054* | 0.075*** | | 0.013 | -0.001 |
| | | (0.030) | (0.028) | | (0.008) | (0.009) |
| 受教育年限 | | -0.003** | -0.000 | | 0.002*** | 0.000 |
| | | (0.001) | (0.001) | | (0.001) | (0.001) |
| 子女数量 | | -0.021*** | -0.008 | | -0.004** | -0.004* |
| | | (0.005) | (0.005) | | (0.002) | (0.002) |
| 家户成员曾参与老农保 | | 0.032 | 0.009 | | 0.007 | 0.001 |
| | | (0.025) | (0.026) | | (0.008) | (0.009) |
| 年龄 | | -0.014*** | -0.014*** | | 0.009*** | 0.010*** |
| | | (0.001) | (0.001) | | (0.001) | (0.001) |
| 男性 | | 0.061*** | 0.040*** | | -0.013*** | -0.008* |
| | | (0.011) | (0.010) | | (0.004) | (0.004) |
| 已婚 | | 0.006 | -0.006 | | -0.012 | -0.013 |
| | | (0.024) | (0.023) | | (0.008) | (0.009) |
| 城镇居民 | | 0.010 | -1.875*** | | -0.002 | 0.275** |
| | | (0.015) | (0.079) | | (0.006) | (0.134) |
| 非农户口 | | 0.348*** | 0.260*** | | -0.020** | -0.013 |
| | | (0.054) | (0.055) | | (0.009) | (0.013) |
| 低抑郁风险 | | -0.030*** | -0.031*** | | 0.018*** | 0.020*** |
| | | (0.010) | (0.010) | | (0.004) | (0.004) |
| ln(当期工作收入+1) | | -0.001 | -0.000 | | 0.001 | 0.001** |
| | | (0.002) | (0.002) | | (0.001) | (0.001) |
| 拥有房产 | | -0.015 | 0.009 | | -0.002 | -0.000 |
| | | (0.021) | (0.024) | | (0.008) | (0.009) |
| 样本量 | 22296 | 22296 | 22296 | 22296 | 22296 | 22296 |
| R-squared | 0.001 | 0.251 | 0.355 | | | |
| Wald p-value | | | | 0.000 | 0.000 | 0.000 |
| 年份固定效应 | NO | NO | YES | NO | NO | YES |
| 社区（村庄）固定效应 | NO | NO | YES | NO | NO | YES |

注：*、**、***分别代表在 10%、5%、1%的程度上显著不等于 0。表格汇报了对应指标的边际效应，括号里的数字为社区层面的聚类标准差。回归包含常数但未报告。表中预期养老金偏差程度是指受访者的预期养老金偏离政策养老金的程度，准确预期是指受访者的预期养老金偏差程度小于 25%。

表 3：倾向得分匹配（PSM）结果

| *k 近邻匹配（k=1）* | | | | | | |
|---|---|---|---|---|---|
| | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |

| | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差程度 | 0.686 | 0.853 | **-0.166** | 0.049 | -3.390*** | 22296 |
| 准确预期者 | 0.129 | 0.084 | **0.045** | 0.015 | 3.040*** | 22296 |

*k 近邻匹配（k = 4 ）*

| | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差程度 | 0.686 | 0.877 | **-0.191** | 0.039 | -4.890*** | 22296 |
| 准确预期者 | 0.129 | 0.094 | **0.034** | 0.013 | 2.720*** | 22296 |

*卡尺内匹配（ε = 0.05 ）*

| | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差程度 | 0.686 | 0.853 | **-0.166** | 0.049 | -3.390*** | 22296 |
| 准确预期者 | 0.129 | 0.084 | **0.045** | 0.015 | 3.040*** | 22296 |

*卡尺内匹配（ε = 0.01 ）*

| | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差程度 | 0.686 | 0.853 | **-0.167** | 0.049 | -3.440*** | 22296 |
| 准确预期者 | 0.129 | 0.084 | **0.045** | 0.015 | 3.050*** | 22296 |

*卡尺内最近邻匹配（k = 4 ， ε = 0.01 ）*

| | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差程度 | 0.686 | 0.880 | **-0.194** | 0.039 | -5.080*** | 22296 |
| 准确预期者 | 0.129 | 0.095 | **0.034** | 0.012 | 2.730*** | 22296 |

*核匹配（Epanechnikov 核函数）*

| | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差程度 | 0.686 | 0.853 | **-0.166** | 0.049 | -3.390*** | 22296 |
| 准确预期者 | 0.129 | 0.084 | **0.045** | 0.015 | 3.040*** | 22296 |

*核匹配（Gaussian 核函数）*

| | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差程度 | 0.686 | 0.853 | **-0.166** | 0.049 | -3.390*** | 22296 |
| 准确预期者 | 0.129 | 0.084 | **0.045** | 0.015 | 3.040*** | 22296 |

*核匹配（Uniform 核函数）*

| | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差程度 | 0.686 | 0.853 | **-0.166** | 0.049 | -3.390*** | 22296 |
| 准确预期者 | 0.129 | 0.084 | **0.045** | 0.015 | 3.040*** | 22296 |

*核匹配（Tricube 核函数）*

| | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差程度 | 0.686 | 0.853 | **-0.166** | 0.049 | -3.390*** | 22296 |
| 准确预期者 | 0.129 | 0.084 | **0.045** | 0.015 | 3.040*** | 22296 |

注：*、**、***分别代表在 10%、5%、1%的程度上显著不等于 0。表中处理组是指老农保参保者群体，对照组是指经过匹配后得到的非老农保参保者群体。

表 4：稳健性检验

| $k$ 近邻匹配（$k=4$） | | | | | | |
|---|---|---|---|---|---|---|
| *不对极端值进行缩尾处理* | | | | | | |
| | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |
| 养老金预期偏差程度 | 0.732 | 0.988 | **-0.256** | 0.0563 | -4.090*** | 22296 |
| 准确预期者 | 0.129 | 0.099 | **0.030** | 0.013 | 2.360*** | 22296 |
| *去掉 2011 年数据* | | | | | | |
| | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |
| 养老金预期偏差程度 | 0.693 | 0.877 | **-0.184** | 0.047 | -3.920*** | 17420 |
| 准确预期者 | 0.129 | 0.107 | **0.022** | 0.015 | 1.420 | 17420 |
| *仅考虑农村户口样本* | | | | | | |
| | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |
| 养老金预期偏差程度 | 0.675 | 0.883 | **-0.208** | 0.039 | -5.350*** | 21198 |
| 准确预期者 | 0.128 | 0.099 | **0.030** | 0.013 | 2.260** | 21198 |
| 核匹配（Epanechnikov 核函数） | | | | | | |
| *不对极端值进行缩尾处理* | | | | | | |
| | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |
| 养老金预期偏差程度 | 0.732 | 0.962 | **-0.230** | 0.043 | -5.290*** | 22296 |
| 准确预期者 | 0.129 | 0.094 | **0.035** | 0.011 | 3.140*** | 22296 |
| *去掉 2011 年数据* | | | | | | |
| | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |
| 养老金预期偏差程度 | 0.692 | 0.875 | **-0.182** | 0.038 | -4.740*** | 17420 |
| 准确预期者 | 0.128 | 0.111 | **0.017** | 0.013 | 1.280 | 17420 |
| *仅考虑农村户口样本* | | | | | | |
| | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |
| 养老金预期偏差程度 | 0.675 | 0.850 | **-0.175** | 0.029 | -6.060*** | 21198 |
| 准确预期者 | 0.128 | 0.094 | **0.034** | 0.011 | 2.92*** | 21198 |

注：*、**、***分别代表在 10%、5%、1%的程度上显著不等于 0。表中处理组是指老农保参保者群体，对照组是指经过匹配后得到的非老农保参保者群体。加入的额外的控制变量包括自报健康较差、过去一个月曾经和朋友交往及家户规模。

（三）机制分析

表 5 报告了乐观预期和认知能力两个影响机制的检验结果。在乐观预期方面，如前所述，若老农保参保经历会通过影响参保者的极端乐观预期，从而改变个体的养老金预期偏差，那么整体而言乐观参保者受到老农保参保经历的影响会更强。结果显示，在悲观参保者子样本内，如果个人有过老农保参保经历，其养老金预期偏差程度会显著降低 0.067，其为准确预期者的概率会提高 2.1 个百分点。而在乐观参保者子样本内，老农保参保经历使得个人的养老金预期偏差程度显著降低 1.059，使得个人是准确预期者的概率提高 12.6 个百分点。组间系数差异检验显示，两组 ATT 数值在统计意义上呈现显著差异，即老农保参保经历对乐观参保者的影响更强，表明老农保主要降低了乐观参保者对未来养老金的极端乐观预期，进而导致乐观参保者的养老金预期偏差程度下降，成为准确预期者的概率上升。

在认知能力方面，本文将是否接受 9 年以上的教育年限作为标准将总样本划分为两个群体。若教育年限越长的个体受到老农保参保经历的影响越大，由于教育年限与认知能力正相关，则可以说明老农保参保经历提升了个体的养老金认知能力，进而改变了养老金预期偏差。表 6 的结果显示，在受教育年限小于等于 9 年的子样本内，老农保参保经历能够使得个人的养老金预期偏差程度显著降低 0.183。而在受教育年限大于 9 年的子样本内，上述影响提高至 0.384，且老农保参保经历能够显著提升个人是准确预期者的概率。进一步的系数差异检验表明，老农保参保经历对个人养老金预期能力的影响确实在受教育年限较长的受访者内更强。这表明，老农保经历对于教育年限较长的个体的养老金预期能力改善作用更强，侧面证实了老农保对养老金认知能力的提升作用。

最后，本文从性别维度验证认知能力这一机制。由于女性平均寿命更长，面临更大的长寿风险，因此从理性的角度更愿意去主动搜集养老金信息，因此相比与男性，会更有可能从老农保参保经历中学到更多的有用信息。因此，若实证结果显示女性受到的影响更大，则可以侧面检验老农保参保经历对认知能力的影响。结果表明，在女性子样本内，如果个人有过老农保参保经历，其养老金预期偏差程度会显著降低 0.113，其是准确预期者的概率会提高 4.3 个百分点。而在男性子样本内，上述两种影响分别变为 0.241 和 1.6 个百分点。系数差异检验显示，女性的养老金预期能力更可能得到老农保参保经历的锻炼，从而降低其养老金预期偏差程度。

表 5：机制检验结果

| 分组变量：乐观参保者 | | | | | | |
|---|---|---|---|---|---|---|
| 分组变量取值 | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |

| 分组变量取值 | 处理组 | 对照组 | ATT | S.E. | t 检验量 | 样本量 |
|---|---|---|---|---|---|---|
| 养老金预期偏差 | 0.570 | 0.637 | **-0.067** | 0.011 | -6.210*** | 20158 |
| 准确预期者 | 0.101 | 0.080 | **0.021** | 0.012 | 1.780** | 20158 |
| 养老金预期偏差 | 1.420 | 2.480 | **-1.059** | 0.255 | -4.160*** | 2138 |
| 准确预期者 | 0.293 | 0.167 | **0.126** | 0.050 | 2.490*** | 2138 |
| 养老金预期偏差（1-0） | | | **-0.992** | 0.182 | -5.451*** | |
| 准确预期者（1-0） | | | **0.105** | 0.035 | 3.001*** | |
| 分组变量：受教育年限 | | | | | | |
| 分组变量取值 | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |
| 养老金预期偏差 | 0.626 | 0.740 | **-0.113** | 0.038 | -3.00*** | 11799 |
| 准确预期者 | 0.123 | 0.080 | **0.043** | 0.015 | 2.800*** | 11799 |
| 养老金预期偏差 | 0.741 | 0.982 | **-0.241** | 0.075 | -3.220*** | 10497 |
| 准确预期者 | 0.131 | 0.115 | **0.016** | 0.020 | 0.890 | 10497 |
| 养老金预期偏差（1-0） | | | **-0.201** | 0.084 | -2.393** | |
| 准确预期者（1-0） | | | **0.035** | 0.022 | -1.591 | |
| 分组变量：男性 | | | | | | |
| 分组变量取值 | 处理组 | 对照组 | **ATT** | S.E. | t 检验量 | 样本量 |
| 养老金预期偏差 | 0.626 | 0.740 | **-0.113** | 0.038 | -3.00*** | 11799 |
| 准确预期者 | 0.123 | 0.080 | **0.043** | 0.015 | 2.800*** | 11799 |
| 养老金预期偏差 | 0.741 | 0.982 | **-0.241** | 0.075 | -3.220*** | 10497 |
| 准确预期者 | 0.131 | 0.115 | **0.016** | 0.020 | 0.890 | 10497 |
| 养老金预期偏差（1-0） | | | **-0.128** | 0.056 | -2.56** | |
| 准确预期者（1-0） | | | **-0.027** | 0.017 | 2.704*** | |

注：*、**、***分别代表在 10%、5%、1%的程度上显著不等于 0。表中处理

组是指老农保参保者群体，对照组是指经过匹配后得到的非老农保参保者群体。

## 五、结论与建议

养老金预期是连接养老金制度改革和个人经济决策的关键路径，探究养老金预期偏差的影响因素，对进一步完善养老金制度存在重要意义。基于 CHARLS 2011-2018 年的数据，本文采用 PSM 方法检验了老农保参保经历对个体养老金预期的影响。

本文发现，老农保参与经历会显著降低个人的养老金预期偏差。具体而言，如果个人曾经参加过老农保，其养老金预期偏差程度会降低 0.191，其成为准确预期者的概率会提高 3.4 个百分点。在更换不同匹配方法、不同样本窗口期以及将样本缩小至农村户口群体时，上述结论依旧稳健。同时，机制分析的结果表明，老农保参保经历降低了个体的极端乐观预期，并提升了个体对养老金的认知能力，使得个体养老金预期准确程度得到改善。

本文的研究结论对于完善公共养老金制度具有一定的政策启示。首先，公共养老金制度改革不能就制度论制度，不仅需要着眼于民众收入水平、年龄结构、利率水平等客观因素，还应当将公众对制度的理解程度等主观因素考虑在内，从全局出发，对公共养老金制度做出适当调整，确保公共养老金发挥其应有的保障职能。

其次，政府应当做好公共养老金的配套制度设计，尤其是公共养老金制度的教育普及工作和养老金公开信息披露工作。一方面，公共养老金的待遇回报计算需要一定的金融专业能力，而普通民众的金融知识有限，对公共养老金的预期难免存在偏差，这需要政府对民众做好教育普及工作。另一方面，个人收集养老金信息的渠道有限，民众往往无法准确评估公共养老金账户的资金，这同样影响个人养老金预期。因此，政府及时公开披露养老金账户的相关信息尤为重要。

最后，政府应当从长远出发，合理设计公共养老金制度待遇，同时也应积极兑现承诺的养老金待遇，不能因外部因素变动为由而随意变更承诺。老农保失败的教训仍历历在目，个体因老农保失败降低了对未来养老金的乐观预期，期间造成的对政府公共政策的误解和怀疑仍然会影响公共政策的实施效果。因此，政府在设计公共养老金待遇时，不仅需要关注当下，还应当放眼未来，合理适当调整养老金待遇，既尽力而为也量力而行。同时，政府对已经发布的政策承诺不应当随意违约，保证公共养老金的公信力。

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# "舐犊"抑或"反哺"？

## ——长期护理保险的代际效应

王煜正 [1]

**摘要：**

　　随着家庭内老年人照护需求日益增长，长期护理保险能否改善家庭转移支付行为受到关注。本文基于 2011 年-2018 年中国健康与养老追踪调查（CHARLS）数据和中国长期护理保险试点政策数据库（CLIP），使用渐进双重差分法和事件研究法分析长期护理保险对家庭代际向下的转移支付行为（"舐犊"）及家庭代际向上的转移支付行为（"反哺"）的影响。研究发现，长期护理保险增加了家庭代际向下的转移支付行为，而对家庭代际向上的转移支付行为没有影响，存在"舐犊"现象而非"反哺"现象。发现长期护理保险通过提高家庭整体效用水平和老年人利他动机进而促进"舐犊"行为，而没有发现交易动机的证据。长期护理保险增加家庭代际向下转移支付行为揭示出通过社会保险制度助推家庭代际经济资源合理配置的必要性。

**关键词：**长期护理保险，代际效应，转移支付，家庭经济

## 一、引言

　　随着我国人口老龄化水平日益攀升，我国政府面临的一个主要挑战是提供可持续的社会保障项目，以确保老年人家庭福利。为了应对这一挑战，我们考虑引入长期护理保险这一新型社会保险的试点计划，探索通过建立长期护理保险制度对老年人家庭的保障作用。然而，为了实现改善老年人福利的目标，必须了解家庭内部在老年人口方面的经济资源分配和交换活动。长期护理保险制度的引入可能会导致老年人口在家庭内部中经济条件的改变。例如，长期护理保险制度可能会增加老年人的经济资源，对其子女产生溢出效应；长期护理保险制度可能会改变子女对老年人赡养行为的激励，扭曲家庭内部经济交换行为。因此，一个重要但尚未研究的问题是长期护理保险制度的引入是否会影响家庭内部的代际经济资源分配和交换行为。

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有关社会保险制度代际效应影响评估的一项挑战在于是否参保社会保险和家庭内部的转移支付本身存在相互作用，我国的长期护理保险这一外生制度冲击为本文识别社会保险的家庭代际效应提供了机会。在长期护理保险制度的试点过程中，我国采取了先试点后推广的政策实施模式，由中央政府制定统一的制度原则，各城市的地方政府进行具体的政策设计。山东省青岛市于 2012 年最先开展长期护理保险制度试点，随后，各试点城市从参保对象、筹资机制和待遇给付等多个维度完善了长期护理保险制度。未来我国的长期护理保险制度有望成为世界上参保人数最多、覆盖范围最大的长期护理保险制度。我国的长期护理保险制度实行市级统筹，且不同地级市试点实施长期护理保险的时间不一致，为本文提供了估计条件。

从理论上讲，在实施长期护理保险时，家庭代际转移支付可能会增加也可能会减少，即可能出现"舐犊"现象，也可能出现"反哺"现象。一方面，长期护理保险的引入可能会减轻家庭老年人护理负担，增加家庭总体经济资源进而增加家庭内部的转移支付行为。另一方面，负担的减轻可能会引起更多的护理需求或增加预防性储蓄动机，减少家庭总体经济资源进而减少家庭内部的转移支付行为。随着长期护理保险试点实施，家庭内部成员可能会重新调整家庭内部经济资源的交换和分配行为。因此，长期护理保险是否会增加或减少代际转移支付是一个经验性问题。

本文使用中国健康与养老追踪调查（CHARLS）数据和中国长期护理保险试点政策数据库（CLIP）分析长期护理保险的代际效应，得到长期护理保险增加了家庭代际向下的转移支付行为，而对家庭代际向上的转移支付行为没有影响，存在"舐犊"现象而非"反哺"现象。

本文主要包含以下三点贡献。首先，本文从家庭内部福利的视角丰富了长期护理保险的影响评估文献。当前研究仍聚焦于长期护理保险对参保老人本身的医疗费用、健康水平和经济行为影响（马超等，2019；于新亮等，2019；Kim, and Lim，2015）及对医疗机构服务的影响（王贞和封进，2021；Lu et al.，2020），鲜有长期护理保险对参保家庭成员的代际效应进行评估。本文补充了长期护理保险对影响家庭内部的代际经济资源分配和交换行为的影响评估。

其次，本文致力于识别长期护理保险产生代际效应的动机。理论层面认为出于交易动机和利他动机，家庭内部存在代际转移支付。交易动机发现代际转移支付存在利己心理，父代对子代的代际转移支付是出于在年老时获得子代转移支付和时间照料的考量，子代对父代的代际转移支付是希望父代帮助自己处理家务或照料孙代（Bernheimetal，1985；Perozek，1998）。利他动机表明父代出于利他心理对子代进行经济供养，而子代也会对父代进行经济支持（Barro，1974；Abel，1987）。部分文献验证了长期护理保险中存在的利他动机（Klimaviciute et al.，2020）和养老保障中存在的交易动机（Wu and Li，2014）。在长期护理保险的代际效应中分别识别交易动机和利他动机有利于完善保险机制和家庭资源配置。

最后，本文通过估计长期护理保险的代际挤出效应或挤入效应确定家庭内部福利分配机制，更好地助推长期护理保险制度发展和完善。部分文

献认为社会保险对家庭代际转移支付具有挤出效应。发现养老保险挤出了家庭代际经济支持，增加了参保老人的经济独立性，养老金收入替代了子女的转移支付（张川川和陈斌开，2014；Zhang，2019），减轻了子女的养老负担（陈华帅和曾毅，2013）。居民医疗保险减少了子女向参保老人的转移支付（白重恩等，2012），减少了参保老人获得经济供养的规模和概率（刘西国，2015）。另一部分文献则认为社会保险对家庭代际转移支付具有挤入效应。得到养老保险增加子女对老人的经济供养（王翌秋和陈青霞，2017）。居民医疗保险促进了家庭代际经济交换，增加了对子女的经济资助（于大川等，2020）。长期护理保险和养老保险及医疗保险同属社会保险制度，制度设计和运行模式有较多相似之处（Feng et al.，2020），可能存在类似于养老保险及医疗保险的代际效应。

## 二、长期护理保险制度背景

长期护理保险是以社会互助共济方式筹集资金，为长期失能人员的基本生活照料和与基本生活密切相关的医疗护理提供资金或服务保障的社会保险制度，是解决医养问题的关键所在（顾海和吴迪，2021）。长期护理保险制度试点工作，是应对人口老龄化、促进社会经济发展的战略举措，也是发展社会保险的"第六险"，健全社会保障体系的重要制度探索。

国家层面的长期护理保险制度于 2016 年开始试点，此前已有部分城市先行探索。2016 年 7 月，《人力资源社会保障部办公厅关于开展长期护理保险制度试点的指导意见》（人社厅发〔2016〕80 号）宣布在河北省承德市、吉林省长春市等 15 个城市开展长期护理保险（长护险）试点，吉林和山东两省为国家试点的重点联系省份。2020 年 9 月，为进一步推进试点工作，《国家医保局、财政部关于扩大长期护理保险制度试点的指导意见》（医保发〔2020〕37 号）宣布新增内蒙古自治区呼和浩特市、天津市等 14 个城市为长期护理保险试点城市。

各城市实际试点的时间与国家层面不一致，山东省青岛市于 2012 年最先开展长期护理保险制度试点，《转发市人力资源社会保障局等部门关于建立长期医疗护理保险制度的意见（试行）的通知》（青政办字〔2012〕91 号）是地方政府首次制定的长期护理保险相关条例。随后，各试点城市从参保对象、筹资机制和待遇给付等多个维度完善了长期护理保险制度。

在长期护理保险制度的推进过程中，中国采取了先试点后推广的政策实施模式，由中央政府制定统一的制度原则，各城市的地方政府进行具体的政策设计。从中央政府的统一制度原则来看，长期护理保险制度是一项重要的社会保险制度，其参保对象是城镇职工，也可以扩大到其他群体；保障范围为符合条件的失能老年人和重度残疾人，也会逐步扩大到其他群体；筹资机制为医保统筹基金转移，不需要公共财政支持；待遇给付方面资金支付水平总体控制在 70% 左右，采取现收现付制，由地方政府向商业保险公司支付资金运行长期护理保险。
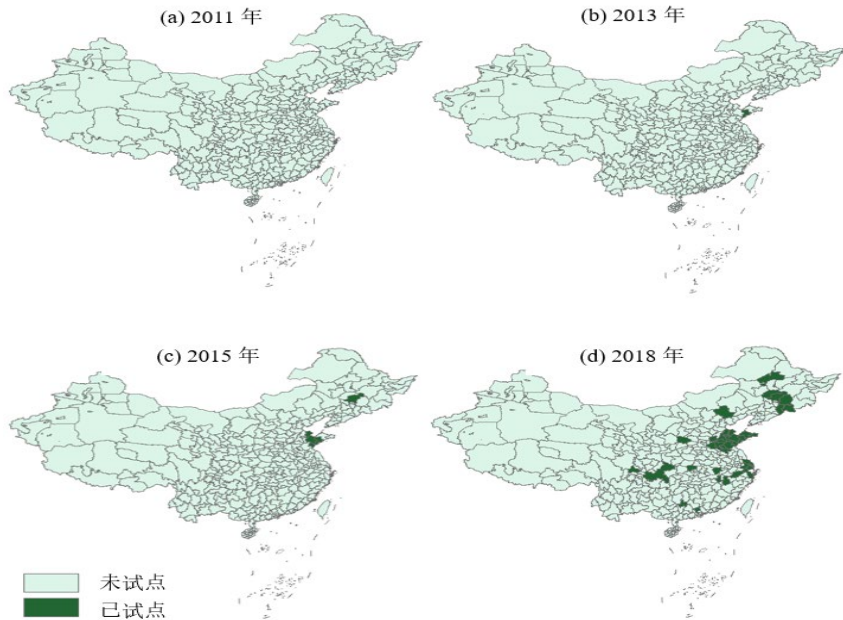
图 1：长期护理保险试点城市图

从长期护理保险试点城市的分布来看，同时涉及到东部、中部、西部和东北四大区域，在同一省份内部试点城市选择上，也体现了对不同经济发展地区的兼顾，这说明实现对不同社会经济发展水平的兼顾是试点地区选择的一个重要特征，也表明对本文所要研究的每个微观家庭的代际转移支付行为而言，长期护理保险试点城市的选择具有较好的外生性。

从地方政府的具体政策设计来看，不同城市长期护理保险的参保对象、待遇保障和失能认定存在差异。参保对象方面，在 29 个试点城市中，10 个试点城市实现了城乡居民全覆盖，而其余城市的参保对象仅为城镇职工或所有城镇居民；筹资机制方面，9 个试点城市在个人缴费和单位缴费基础上有政府财政补贴，而其余城市无政府财政补贴；失能认定方面，13 个试点城市按照 ADL 标准认定失能，而其余城市为自行制定失能标准。

长期护理保险制度的覆盖范围大。根据国家医保局统计数据，截止2021 年 8 月，全国 49 个长护险试点城市参保人数达 1.34 亿人，累计享受待遇人数 152 万人。当前中国已有 29 个省份至少有一个城市试点实施长期护理保险制度，这一制度预期可成为世界上覆盖地区最广，参保人数最多的长期护理保险制度。讨论这一制度有利于和我国养老保险制度和医疗保险制度互补，构建完善的"医+养+护"保险体系；也可为其他国家的长期护理保险制度提供保险模式借鉴。

## 三、数据来源与研究设计

（一）数据来源

本文使用中国健康与养老追踪调查（CHARLS）数据和中国长期护理

保险试点政策数据库（CLIP）分析长期护理保险的代际效应。本文根据CHARLS 调查得到代际转移支付数据。CHARLS 是由北京大学国家发展研究院组织，北京大学中国社会科学调查中心执行的全国性社会调查，调查群体为 45 周岁以上中老年人。本文合并使用 2011 年、2013 年、2015 年和2018 年的 CHARLS 数据，形成一个包含长期护理保险状况和代际转移支付的四期面板数据。

本文还根据 CLIP 试点政策数据库得到全国各个地级市的长期护理保险实施数据。CLIP 由中国人民大学公共管理学院长护险研究小组提供，包含详细的各城市长期护理保险试点文件。本文从中得到长期护理保险试点年份信息、参保对象信息、保障范围信息、资金筹集信息和待遇支付信息，并与 CHARLS 数据合并使用。

本文还从《中国城市统计年鉴》中获取了城市层面的数据。《中国城市统计年鉴》是国家统计局编撰的全国城市经济与社会发展的年度总结，可以有效反映长期护理保险试点城市和非试点城市的经济与社会差异。由于长期护理保险是基于城市层面试点的，本研究将《中国城市统计年鉴》数据与 CHARLS 数据及 CLIP 数据合并使用，能够将城市经济与社会差异纳入考虑范围内，有效控制非长期护理保险试点引发的城市差异。

使用上述数据研究长期护理保险的代际效应具有三方面明显优势。第一，CHARLS 调查样本覆盖了所有长期护理保险试点城市所在的省份，样本具有全国范围代表性。第二，《中国城市统计年鉴》数据详细记录了长期护理保险试点城市的经济水平和社会保障水平信息，CLIP 全面地提供了各城市长期护理保险的参保对象信息、保障范围信息、待遇支付信息和失能标准信息，有助于讨论不同城市实施效果的差异并提出可能的改进方案。第三，CHARLS 调查时间涵盖 CLIP 政策文件发布时间，包含长期护理保险的试点期和推广期，能够有效捕捉长期护理保险的代际效应。

（二）识别策略

本文使用渐进双重差分法和事件分析法估计长期护理保险的代际效应。基于长期护理保险在不同地级市试点时间的差异，本文使用渐进双重差分法估计随着长期护理保险试点年份推进中代际转移支付发生的变化。第一重差分为年份层面长期护理保险实施年份的前后差异，第二重差分为参保老人和子女所在的地级市层面已实施长期护理保险市与尚未实施长期护理保险市的差异。具体估计模型为：

$$IntergenerationalSupport_{ict}^{s}$$

$$= \beta_0^s + \beta_1^s LTCI_{ct}^s + \sum_{j=1}^{J} \omega_j^s X_{ict}^s + \delta_c^s + \delta_t^s + \varepsilon_{ict}^s \quad (1)$$

(1)式中被解释变量为 t 年份在地级市 c 的参保老人 i 的代际转移支付情况 $IntergenerationalSupport_{ict}^{s}$，上标 s 表示特定的子样本；核心解释变量为地级市 c 在 t 年的长期护理保险试点实施情况 $LTCI_{ct}^{s}$，当地级市 c 在 t 年已经实施长期护理保险，则 $LTCI_{ct}^{s}=1$，而若尚未实施长期护理保险，则 $LTCI_{ct}^{s}=0$；$\sum_{j=1}^{J}\omega_{j}^{s}X_{ict}^{s}$ 为所有控制变量的加总，包括家户人口指标和城市经济指标等；$\delta_{c}^{s}$ 为城市固定效应；$\delta_{t}^{s}$ 为年份固定效应；$\varepsilon_{ict}^{s}$ 为随机扰动项。

本文重点关注核心解释变量 $LTCI_{ct}^{s}$ 的系数 $\beta_{1}^{s}$ 的方向，若系数 $\beta_{1}^{s}$ 显著为正，则说明长期护理保险增加了代际转移支付；若系数 $\beta_{1}^{s}$ 显著为负，则说明长期护理保险减少了代际转移支付；若系数 $\beta_{1}^{s}$ 不显著，则说明长期护理保险对代际转移支付没有影响。

为了检验长期护理保险试点实施前不同地级市的代际转移支付趋势是否存在显著差异，在模型（1）的基础上，使用事件分析法设定检验渐进 DID 平行趋势的模型（2）：

$$IntergenerationalSupport_{ict}^{s}$$
$$=\alpha_{0}^{s}+\sum_{i=-3}^{i=3}\alpha_{\tau}^{s}LTCI_{ic,t-\tau}+\sum_{j=1}^{J}\omega_{j}^{s}X_{ict}^{s}+\delta_{c}^{s}+\delta_{t}^{s}+\varepsilon_{ict}^{s} \quad (2)$$

模型（2）的被解释变量与模型（1）一致，区别在于解释变量 $LTCI_{ic,t-\tau}$ 是长期护理保险的政策虚拟变量，若参保老人 i 所在的地级市 $i$ 在 $t-\tau$ 期实施了长期护理保险，则赋值为 1，否则为 0。数据年份为 2011 年、2013 年、2015 年和 2018 年，长期护理保险试点政策为 2016 年发布，分别取长期护理保险试点前三期和后三期，将最早一期作为基准年份，估计系数 $\alpha_{\tau}^{s}$ 的变动。

（三）样本限定

本文依据长期护理保险的政策文件规定对样本范围进行限定。长期护理保险的基本覆盖范围为城镇职工医疗保险参保群体，仅有部分城市额外将城乡居民医疗保险参保群体纳入覆盖范围内。为了排除城乡居民医疗保险代际效应对本文估计结果的干扰，本文将样本范围限定为至少有一名家庭成员参保城镇职工医疗保险的家庭。同时，本文基于地级市政府文件公布的长期护理保险制度试点实施时间确定该城市实施长期护理保险的年份，而不是基于全国统一的长期护理保险试点实施时间，亦非地级市政府文件发布的时间。这是为了利用准确的长期护理保险试点实施时间年份差异和地级市城市差异创造的估计条件检验长期护理保险的代际效应。此外，由于本文讨论的是长期护理保险对家庭代际转移支付的影响，本文通过取均值或至少家中有一名家庭成员存在观测行为的方式将所有变量汇总到家庭层面，最终得到 498 个家庭的 4 期平衡面板数据。

（四）变量选取

本文选取的被解释变量为家庭代际转移支付，具体包含家庭代际是否发生向下的转移支付、家庭代际是否发生向上的转移支付、家庭代际向下转移支付额和家庭代际向上转移支付额四项指标 [1]。基于长期护理保险的年份差异和城市差异，本文定义的解释变量为家庭所在的城市在调查年份是否已试点实施长期护理保险制度。本文选取的控制变量包含家户人口指标和城市经济指标。家户人口指标包含家庭平均年龄及其平方、家庭是否至少有一位家庭成员完成义务教育、家庭是否至少有一位家庭成员在婚和家庭人均收入的对数值。城市经济指标包含城市生产总值对数值、城市财政规模对数值、城市工业水平对数值和城市基础设施对数值。

表 1 报告了本文的数据来源和变量描述性统计。可以发现，家庭内发生向下代际转移支付的概率达 64%，发生向上代际转移支付的概率达 74%，不同年份间向下代际转移支付额和向下代际转移支付额差异较大。

<div align="center">表 1 描述性统计</div>

| 变量名称 | 全样本 | | | | 分年份均值 | | | |
|---|---|---|---|---|---|---|---|---|
| | 数据来源 | 观测值 | 均值 | 标准差 | 2011 | 2013 | 2015 | 2018 |
| **被解释变量** | | | | | | | | |
| 是否下移 | A | 1992 | 0.64 | 0.48 | 0.53 | 0.69 | 0.62 | 0.70 |
| 是否上移 | A | 1992 | 0.74 | 0.44 | 0.68 | 0.76 | 0.70 | 0.81 |

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[1] 家庭代际向下转移支付额和家庭代际向上转移支付额对原始转移支付额进行加 1 后取对数的处理。

| 变量 | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| 下移支付额 | A | 1992 | 3.69 | 4.28 | 1.68 | 3.42 | 5.16 | 4.49 |
| 上移支付额 | A | 1992 | 4.02 | 4.15 | 1.42 | 3.95 | 5.61 | 5.09 |
| **解释变量** | | | | | | | | |
| 长护险试点后 | B | 1992 | 0.06 | 0.23 | 0.00 | 0.02 | 0.03 | 0.15 |
| **控制变量：家户人口指标** | | | | | | | | |
| 平均年龄 | A | 1992 | 61.56 | 9.17 | 58.42 | 60.31 | 62.31 | 65.20 |
| 平均年龄平方 | A | 1992 | 38.74 | 11.70 | 34.91 | 37.15 | 39.61 | 43.29 |
| 完成义务教育 | A | 1992 | 0.81 | 0.39 | 0.83 | 0.84 | 0.78 | 0.80 |
| 是否在婚家庭 | A | 1992 | 0.85 | 0.36 | 0.88 | 0.86 | 0.86 | 0.81 |
| 家庭人均收入 | A | 1992 | 10.98 | 1.33 | 10.72 | 11.08 | 10.70 | 11.38 |
| **控制变量：城市经济指标** | | | | | | | | |
| 城市生产总值 | C | 1992 | 7.85 | 0.87 | 7.60 | 7.80 | 7.92 | 8.09 |
| 城市财政支出 | C | 1992 | 7.34 | 0.34 | 7.21 | 7.29 | 7.39 | 7.46 |
| 城市工业水平 | C | 1992 | 3.83 | 0.23 | 3.92 | 3.89 | 3.82 | 3.70 |
| 城市基础设施 | C | 1992 | 9.12 | 0.54 | 8.91 | 9.01 | 9.17 | 9.39 |

数据来源：A：中国健康与养老追踪调查（CHARLS）数据；B：中国长期护理保险试点

政策数据库（CLIP）数据；C：《中国城市统计年鉴》数据。

## 四、实证结果

（一）基准回归结果

表 2 报告了长期护理保险对家庭代际向下转移支付和向上转移支付的基准回归结果，得到长期护理保险对家庭代际向下转移支付存在显著的正向影响，而对家庭代际向上转移支付没有影响。其中，第（1）列和第（2）列是面板 LogitFE 模型的回归结果，第（3）列和第（4）列使用 LMP 模型进行检验，第（5）列和第（6）列是固定效应 FE 回归结果。第（1）列报告了长期护理保险对家庭是否发生代际向下转移支付的边际效应，得到长期护理保险试点后参保家庭发生代际向下转移支付的概率显著增加18.7%。第（2）列报告了长期护理保险对家庭是否发生代际向上转移支付的边际效应，得到长期护理保险对家庭是否发生代际向上转移支付不存在显著影响。第（3）列和第（4）列使用 LMP 模型进行检验，结果仍然稳健，得到长期护理保险使得家庭发生代际向下转移支付的概率增加了22.0%，而对家庭代际向上转移支付没有影响。

本文进一步使用具体向下转移支付额和向上转移支付额更加准确地度量家庭代际向下转移支付和向上转移支付。第（5）列报告了长期护理保险对家庭代际向下转移支付额的影响，得到长期护理保险试点使得家庭代际向下转移支付额显著增加 1.206,在经济意义上表明长期护理保险试点使得家庭代际向下转移支付的概率增加 32.7%。第（6）列报告了长期护理保险对家庭代际向上转移支付额的影响，得到长期护理保险对向上转移支付不存在显著影响。总之，长期护理保险增加了家庭代际向下转移支付，而对家庭代际向上转移支付没有影响。这一结果意味着长期护理保险挤入了家庭代际向下转移支付，长期护理保险带来了"舐犊"现象，但不存在"反哺"现象。

可能的原因是，长期护理保险对符合条件的失能参保老人提供复合式待遇支付，显著降低了失能家庭长期照护费用，且保障了家庭内老年人的身心健康，在家庭内对老年人子女产生了溢出效应，带来了"舐犊"现象。子女作为家庭内老年人的赡养者，家庭代际向上转移支付是子女"反哺"的主要手段而长期护理保险报销带来的照护费用减少抵消了子女所应承担的家庭代际向上代际转移支付额，长期护理保险与家庭代际向上转移支付互为替代品，不存在"反哺"现象。这体现了长期护理保险制度试点实施带来家庭福利的溢出效应的政策效果，具体表现为挤入家庭代际向下转移支付。

表 2 长期护理保险的代际效应

| 变量名称 | LogitFE 模型 | | LMP 模型 | | FE 模型 | |
|---|---|---|---|---|---|---|
| | 是否下移 | 是否上移 | 是否下移 | 是否上移 | 向下转移 | 向上转移 |

|  |  |  |  |  | 支付额 | 支付额 |
| --- | --- | --- | --- | --- | --- | --- |
|  | (1) | (2) | (3) | (4) | (5) | (6) |
| 长护险试点后 | 0.187** | 0.058 | 0.220*** | 0.093 | 1.206** | 0.166 |
|  | (0.047) | (0.053) | (0.053) | (0.069) | (0.550) | (0.483) |
| 平均年龄 | -0.130** | -0168** | -0.137*** | -0.160*** | 0.885*** | 0.411 |
|  | (0.053) | (0.072) | (0.034) | (0.031) | (0.244) | (0.283) |
| 平均年龄平方 | 0.089** | 0.143** | 0.099*** | 0.130*** | -0.776*** | -0.226 |
|  | (0.036) | (0.060) | (0.022) | (0.021) | (0.187) | (0.182) |
| 完成义务教育 | -0.001 | 0.035 | 0.014 | 0.053 | -0.057 | 0.061 |
|  | (0036) | (0.038) | (0.046) | (0.042) | (0.389) | (0.355) |
| 是否在婚家庭 | -0.029 | -0.012 | -0.028 | -0.009 | 0.582 | 0.602 |
|  | (0.050) | (0.049) | (0.054) | (0.050) | (0.516) | (0.380) |
| 家庭人均收入 | 0.007 | 0.013 | 0.010 | 0.018 | 0.059 | 0.283*** |
|  | (0.008) | (0.009) | (0.009) | (0.012) | (0.088) | (0.095) |
| 城市生产总值 | 0.289* | 0.382* | 0.250 | 0.309* | 0.128 | 0.112 |
|  | (0.168) | (0.196) | (0.200) | (0.157) | (1.461) | (1.707) |
| 城市财政支出 | -0.051 | 0.142 | -0.078 | 0.150 | -1.749 | 0.131 |
|  | (0.118) | (0.128) | (0.137) | (0.110) | (1.179) | (1.007) |
| 城市工业水平 | -0.050 | -0.232 | 0.055 | -0.152 | 0.683 | -0.102 |
|  | (0.182) | (0.203) | (0.224) | (0.197) | (1.765) | (2.012) |
| 城市基础设施 | 0.268 | 0.181 | 0.308** | 0.164 | 0.200 | -0.071 |
|  | (0.130) | (0.117) | (0.147) | (0.117) | (1.051) | (0.913) |
| 家户固定效应 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 |
| 年份固定效应 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 |
| 观测值 | 1,298 | 1,235 | 1,921 | 1,921 | 1,921 | 1,921 |
| $R^2$/Pseduo $R^2$ | 0.177 | 0.161 | 0.432 | 0.343 | 0.503 | 0.580 |

注：括号内为聚类稳健标准误，聚类在地级市层面上。*、**、***分别表示在10%、5%和1%水平上显著。

（二）事件分析结果

长期护理保险的代际效应成立的一个重要前提是在长期护理保险试点前，试点实施长期护理保险城市的家庭代际转移支付和不试点实施长期护理保险城市的家庭代际转移支付间没有显著差异。为了检验长期护理保险代际效应的平行趋势，本文使用事件分析法重新估计长期护理保险对代际转移支付的影响。图2报告了长期护理保险代际效应事件分析的估计结果，结果发现，长期护理保险城市的家庭代际转移支付和不试点实施长期护理保险城市的家庭代际转移支付间不存在显著差异，满足估计长期护理保险

代际效应的前提条件。

从动态角度考查长期护理保险的代际效应，得到长期护理保险制度试点实施第 2 期和第 3 期存在对家庭代际向下转移支付的正向影响。在长期护理保险制度试点实施当年，长期护理保险并不会对代际转移支付产生影响，在长期护理保险制度试点实施后的第 2 期和第 3 期，长期护理保险增加家庭代际向下转移支付，且在实施第 3 期长期护理保险对家庭向下转移支付的代际挤入效应更大。这是由于长期护理保险制度试点对家庭代际转移支付的影响存在一个滞后的传导过程，从城市层面的长期护理保险制度试点到家庭层面"舐犊"现象的影响机制有待深入探讨。
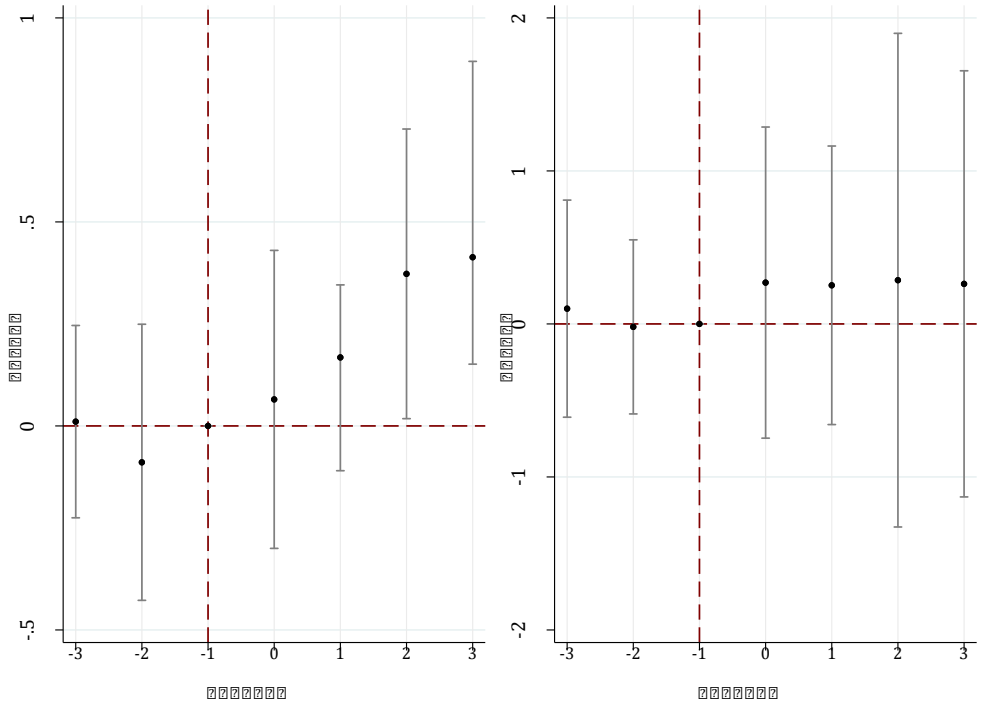


注：选取长期护理保险制度试点实施前一期作为基准，用事件分析法估计长期护理保险的代际效应。

图 2 长期护理保险代际效应的事件分析

（三）安慰剂检验

在长期护理保险制度试点实施的同时，其他可能影响家庭代际转移支付的政策或不可观测的特征会对干扰长期护理保险代际效应的估计结果，需要对长期护理保险制度试点进行安慰剂检验。本文使用随机抽取长期护

理保险制度试点实施时间的方法，首先将数据按照地级市分组，然后在每个地级市组内的年份变量中随机抽取一个年份作为长期护理保险制度的伪试点时间，检验估计系数是否显著。

在 2011-2018 年间对长期护理保险制度试点实施时间进行重复 500 次的随机抽样。图 3 为长期护理保险代际效应安慰剂检验估计系数的核密度分布图，对家庭代际转移支付进行估计，得到 500 次的抽样估计结果均小于真实估计系数，且大部分估计系数在 10%显著性水平线之上，且服从正态分布，说明长期护理保险代际效应的估计结果不太可能是偶然得到的，在 2011-2018 年间的其他政策对家庭代际转移支付不存在显著影响。
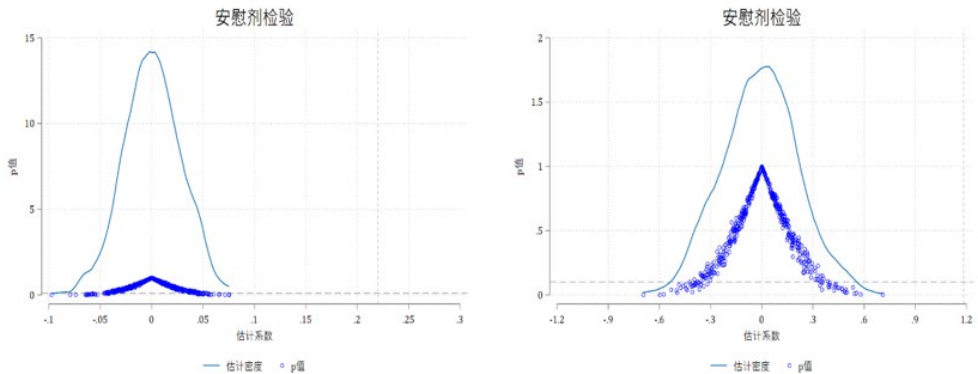


注：图3为随机抽取长期护理保险试点实施年份500次的安慰剂检验结果，垂直虚线是渐进双重差分模型真实估计值0.220和1.206，水平虚线是显著性水平0.1。

图 3 长期护理保险试点安慰剂检验

（四）其他稳健性检验

我们从时间和空间两个维度检验了基准回归结果的稳健性。在时间上，干净地识别出长期护理保险代际效应的前提是排除同期医疗保险和养老保险政策和家庭共同财富的干扰。为了解决该问题，一方面我们引入医疗保险和养老保险的虚拟变量作为控制变量，以此控制医疗保险和养老保险政策对基准回归的干扰。我们发现长期护理保险代际效应的系数出现了小幅度的下降，但仍然是正向显著的，这意味着基准回归的结果是稳健的。另一方面，由于家庭共同财富不计入家庭代际效应中，即使我们在基准回归中纳入了共同财富较多的家庭，其结果也只可能是低估了长期护理保险的代际效应。为了进一步验证家庭共同财富的影响，我们在实证分析中剔除了父代和子代共同居住的样本，这是因为父代和子代共同居住的家庭往往存在较多的家庭共同财富。回归分析显示目标系数没有明显变化，基准回归的结果是稳健的。除此之外，我们还采用基于插补的估计法纠正了渐进双重差分的偏误。

在空间上，由于这次改革实施的地区选取可能存在非随机性，这会导致一些不易观察到的因素同时影响长期护理保险试点和家庭代际转移支付进而干扰本文对长期护理保险代际效应的识别。为了解决该问题，我们首先控制了城市特征趋势项，将观测初期 2011 年的各城市特征变量与年份线性趋势项的交互项加入回归方程，从而控制原有差异在时间趋势层面对估计结果的影响，得到稳健结果。其次，我们使用双重差分的倾向评分加权（Propensity Score Weighting-DID）方法对该问题进行解决。具体而言，我们首先使用 Probit 模型来估计各个城市实施长期护理保险的概率，随后我们根据倾向得分对样本进行剔除和加权，最后我们基于加权后的样本进行估计。我们发现长期护理保险代际转移支付的系数没有出现明显的变化，这意味着基准回归的结果是稳健的。最后，我们采用空间配比双重差分的方法（BD-DID）进行处理。具体来说，我们将长期护理保险试点城市视作实验组，将那些和长期护理保险试点城市相邻但未实施长期护理保险的城市作为对照组 [1]。这种研究设计可以使得我们在很大程度上克服试点城市和非试点城市之间可能存在的系统性差异。那么，长期护理保险实施之后在这两组样本之间呈现的差异就可以归咎于长期护理保险试点的效应。结果显示，长期护理保险的代际效应在空间配比的研究设计中依然存在并且系数没有明显变化，这意味着基准回归的结果并不是由样本选取非随机性所驱动的 [2]。

## 五、拓展分析

（一）异质性分析

为了识别受到长期护理保险代际效应影响的群体，本文检验了长期护理保险对不同健康水平、受教育水平和参保比例家庭的影响。图 4 报告了长期护理保险代际效应的异质性分析结果，研究发现，长期护理保险向下转移支付的代际挤入效应主要集中于无日常活动能力障碍的家庭、受教育水平较低的家庭和所有参保家庭。

第一，家庭成员的健康水平长期受到文献关注，长期护理保险并不会影响有日常活动能力障碍家庭的代际向下转移支付，但会增加无日常活动能力障碍的家庭的代际向下转移支付。有日常活动能力障碍家庭的老年人护理成本较高，家庭经济资源较多地分配给老年人，而无日常活动能力障碍的家庭存在向下的代际挤入效应，存在"舐犊"现象。第二，长期护理保险不影响受教育水平较高的家庭的代际向下转移支付，而增加受教育水平较低家庭的代际向下转移支付，这可能是由于较高受教育水平的家庭每位家庭成员有较强的经济独立意识，而受教育水平较低的家庭存在较多的"赡

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[1] 在思想上，这种处理方法借鉴了断点回归的逻辑；在具体操作上，我们提取城市质心并计算长期护理保险试点城市和非试点城市的质心距离，分别选取质心距离 100km、200km 和 300km 范围内的城市进行估计，均得到稳健结果。

[2] 由于篇幅限制，其他稳健性检验结果在本文附录中呈现。

养"现象。第三，无论家庭参保比例高低，长期护理保险都会对家庭代际向下转移支付产生正向影响，这进一步说明了长期护理保险对参保家庭经济资源分配和交换的重要挤入作用。
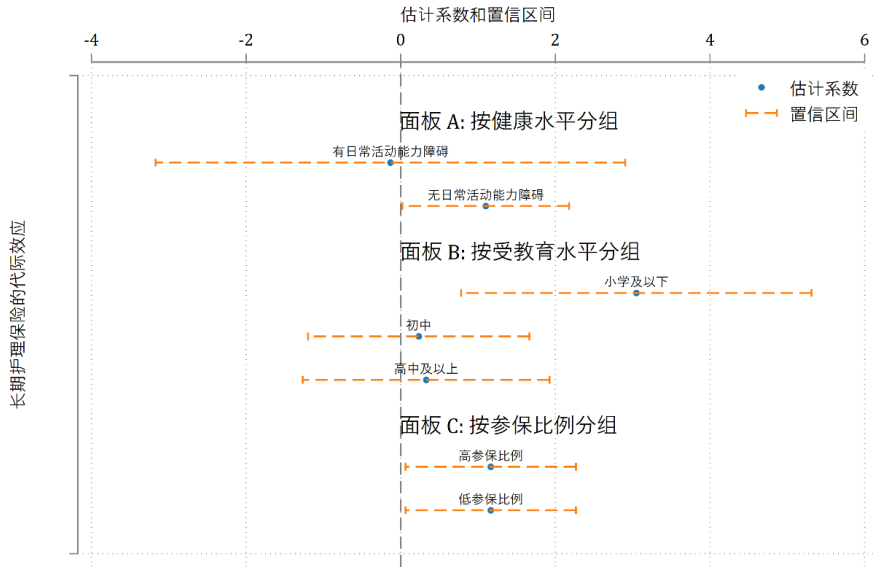


注：图4中的点表示长期护理保险对相应子样本影响的估计系数，置信区间计算聚类于地级市层面。

图 4 长期护理保险代际效应在不同类型家庭中的差异

不同城市长期护理保险的政策设计存在差异。为了比较不同类型的政策设计在家庭代际效应中的适用性，本文区分了不同城市长期护理保险的参保范围、筹资机制和失能评定标准，检验了长期护理保险政策设计对家庭代际转移支付的异质性影响。具体而言，我们在基准回归的基础上加入了长期护理保险实施后和长期护理保险是否覆盖所有城乡居民，长期护理保险是否有政府财政补贴和失能评定是否依据ADL的交互项。结果发现，长期护理保险覆盖所有城乡居民，筹资有政府财政补贴和失能评定依据ADL能够显著增加家庭代际向下的转移支付。

表 3 政策文本异质性分析

| 变量名称 | LogitFE 模型 | | | LMP 模型 | | | FE 模型 | | |
|---|---|---|---|---|---|---|---|---|---|
| | 是否下移 | | | 是否下移 | | | 向下转移支付额 | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| 全覆盖 ×长护险 | 0.193*** | | | 0.205*** | | | 1.378* | | |
| | (0.055) | | | (0.071) | | | (0.786) | | |

| | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| 政府补贴 × 长护险 | 0.210*** | | | 0.259*** | | | 1.358** | | |
| | (0.050) | | | (0.048) | | | (0.541) | | |
| 评估依据 × 长护险 | | 0.192*** | | | 0.205*** | | | 1.378* | |
| | | (0.055) | | | (0.071) | | | (0.786) | |
| 控制变量 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 |
| 家户固定 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 |
| 年份固定 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 |
| 观测值 | 1,303 | 1,303 | 1,303 | 1,926 | 1,926 | 1,926 | 1,926 | 1,926 | 1,926 |
| $R^2$/Pseduo $R^2$ | 0.176 | 0.179 | 0.176 | 0.430 | 0.432 | 0.430 | 0.503 | 0.503 | 0.503 |

注：括号内为聚类稳健标准误，聚类在地级市层面上。*、**、***分别表示在10%、5%和1%水平上显著。

### （二）机制分析

为了探究长期护理保险代际向下挤入效应的作用机制，我们综合考虑了家庭效用和转移支付的动机。表4报告了长期护理保险的代际效应机制检验及动机识别的结果。我们希望获取长期护理保险实施后家庭得到的护理水平提高的证据证明家庭整体效用水平得到提高。然而，我们的数据中缺乏直接的护理水平变量，我们使用能够衡量护理结果的健康水平作为护理水平的代理变量。第(1)列显示长期护理保险试点实施使得家庭心理健康均值显著增加了 0.044，第(2)列显示长期护理保险试点实施使得家庭身体健康均值显著增加了 0.078，发现长期护理保险实施后家庭健康水平提高，而家庭健康水平的提高有助于改善家庭内部经济资源的交换和分配行为。进一步，我们加入了消费水平的交互项，第(3)列和第(4)列得到消费水平交互项对家庭代际向下转移支付的显著正向影响，发现家庭效用提高确实可以增加代际向下转移支付。

我们进一步识别了长期护理保险试点后家庭产生代际效用效用的动机。第(3)列和第(4)列的结果也意味着父代对子代存在利他动机。而第(5)列和第(6)列发现长期护理保险实施后对家庭代际向上转移支付的影响不显著，因而没有直接证据显示子代对父代存在利他动机。同时，第(7)列结果显示长期护理保险实施后父代得到的时间照料并没有发生显著变化，可见交易动机的假设在本文中不成立。

结合前人的研究结论，家庭健康水平提高促进了老年人对子女的经济支持（于大川，2020），而长期护理保险替代了子女对老年人的转移支付（张川川和陈斌开，2014；Zhang，2019）。由此，长期护理保险代际向下挤

入效应的作用机制得到验证。长期护理保险对家庭健康水平和效用水平起到提高作用，并且通过该影响进一步促进了农村家庭内部代际经济资源的交换和分配，促进家庭代际向下的转移支付，而用长期护理保险福利替代向上的转移支付，即通过健康机制和家庭效用提高机制产生了"舐犊"现象，而不存在"反哺"现象。且上述证据表明，这一"舐犊"现象是由家庭中父代对子代的利他动机驱动的。

<div align="center">表 4 长期护理保险的代际效应机制检验及动机识别</div>

| 变量名称 | 心理健康 (1) | 身体健康 (2) | 是否下移 (3) | 向下转移支付额 (4) | 是否上移 (5) | 向上转移支付额 (6) | 父代得到时间照料 (7) |
|---|---|---|---|---|---|---|---|
| 长护险试点 | 0.044* | 0.078* | 0.233*** | 1.124** | 0.093 | 0.184 | -0.104 |
| | (0.026) | (0.043) | (0.050) | (0.557) | (0.069) | (0.484) | (0.177) |
| 消费水平×长护险试点 | | | 0.254*** | 2.773*** | | | |
| | | | （0.069） | （0.587） | | | |
| 控制变量 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 |
| 家户固定效应 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 |
| 年份固定效应 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 | 控制 |
| 观测值 | 1,870 | 1,919 | 1,607 | 1,607 | 1,928 | 1,928 | 557 |
| $R^2$/Pseduo $R^2$ | 0.501 | 0.536 | 0.459 | 0.531 | 0.344 | 0.578 | 0.624 |

注：括号内为聚类稳健标准误，聚类在地级市层面上。*、**、***分别表示在10%、5%和1%水平上显著。

## 六、研究结论与政策建议

（一）研究结论

本文使用中国健康与养老追踪调查数据和长期护理保险政策实施数据，利用渐进双重差分方法评估长期护理保险试点实施对家庭代际转移支付行为的影响，得到以下五点结论。第一，长期护理保险增加了家庭代际向下的转移支付行为，而对家庭代际向上的转移支付行为没有影响，存在"舐犊"现象而非"反哺"现象。第二，本文验证了长期护理保险的代际向下挤入效应，且该效应动态扩大。第三，长期护理保险向下转移支付的代际挤入效应主要集中于无日常活动能力障碍的家庭、受教育水平较低的家庭

和所有参保家庭。第四，长期护理保险提高了护理水平，带来了家庭整体效用水平的提高从而增加了家庭福利。第五，长期护理保险代际向下挤入效应存在健康机制和经济机制，长期护理保险通过提高家庭健康水平和家庭效用水平进而促进家庭代际向下转移支付，且主要受到父代对子代的利他动机驱动。

（二）政策建议

上述发现为长期护理保险的制度设计提供了一定的经验借鉴。第一，扩大长期护理保险的试点范围，加快推进长期护理保险制度实施。本文发现长期护理保险使得家庭健康水平提高和家庭经济资源增加，对参保家庭产生了良好的健康绩效和经济绩效，提高了参保家庭的福利水平。因而应当总结长期护理保险试点经验，推进长期护理保险在全国范围内的有效实施。第二，逐步扩大长期护理保险的参保群体。当前长期护理保险的参保对象为参保城镇职工医疗保险的群体，家庭代际经济资源的交换和分配行为的改善不应仅限于城镇职工医疗保险参保者，应当逐步将城乡居民医疗保险参保者列为参保对象，使得长期护理保险对家庭福利的溢出效应同时覆盖城镇和农村群体。第三，通过为长期护理保险提供政府财政补贴的模式在长期护理保险试点时期是有效的，这种方式可以增加家庭福利。第四，重点关注长期护理保险给老年人带来的福利。长期护理保险带来了"舐犊"现象的增加，而对"反哺"现象没有影响，可见长期护理保险的福利较多地传导到了家庭子代，而长期护理保险制度设立的初衷是为了保障失能老人自身的健康和经济福利，需要在确保家庭内部老年人充分享受到长期护理保险带来福利的前提下提供对家庭的福利溢出效应。第五，要以家庭为整体评估长期护理保险带来的福利效用，发挥代际利他动机促进家庭和谐。

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# 健康保险
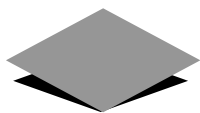
# The Impact of Urban-rural Health Insurance Integration on Labor Supply——Evidence from Rural China

曹园、陈华、王沁雨、丁宇刚 [1]

**Abstract：**

This paper examines the impact of China's urban-rural health insurance integration policy on labor supply of rural residents using data from the China Health and Retirement Longitudinal Study 2011-2018. Taking advantage of the city-by-city rollout of the integration policy, we employ a staggered difference-in-differences model for estimation. The results show that the implementation of the integration policy significantly increases the probability of rural residents having more than one job by 0.62 percentage points and significantly decreases the likelihood of missing work due to illness by 3.73 percentage points. This effect is heterogeneous. The impact of increasing the probability of having more than one job occurs mainly in the low-income and younger groups, while the impact of reducing the likelihood of missing work due to illness occurs mainly in the high-income and younger groups. In addition, we find a significant increase in the probability of non-agricultural employment and a significant decrease in the probability of quitting the labor market due to illness as a result of integration policies. Overall, we find that urban-rural health insurance integration policy is important for promoting labor mobility, narrowing the urban-rural gap, reducing health insurance inequality, and improving the well-being of rural residents.

**Keywords：** public health insurance, labor supply, integration policy

## 1 Introduction

In recent years, many developing countries have introduced or expanded public health insurance schemes to cover more people and improve resilience to health shocks and disease risks. Some developing countries, such as China, Vietnam, Ghana, Mexico, and India, have established universal health coverage (1). However, unequal benefit effects have accompanied the expansion of health insurance coverage. Therefore, adjusting health insurance schemes and

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social protection programs to reach all population groups, especially excluded groups, and alleviating health insurance inequalities will be the focus of universal health insurance reform (2).

In China, the inequality in public health insurance (i.e., basic medical insurance) has long existed. China's basic medical insurance system consists of three main pillars: the New Rural Cooperative Medical Scheme (NRCMS) for rural residents, the Urban Resident Basic Medical Insurance (URBMI) for unemployed and informally employed urban residents, and the Urban Employee Basic Medical Insurance (UEBMI) for formal urban employees. By the end of 2021, the basic medical insurance scheme in China had covered around 1.36 billion people, with the coverage rate exceeded coverage exceeding 95%[1]. Despite the high coverage rate, the system still has some limitations, and there is a significant disparity among the three pillars in terms of fund financing, contribution levels, and benefits. Furthermore, China's long-standing urban-rural hukou system has led to negative impacts of the fragmentation of the health insurance system and the dualistic division between urban-rural areas. Compared to urban formal employees, rural residents are always at a disadvantaged position in terms of the level of health insurance benefits (3). Consequently, in 2016, the State Council decided to end the urban-rural hukou restriction in the public health insurance systems and gradually integrate the NRCMS and URBMI nationwide, establishing the Urban and Rural Resident Basic Medical Insurance (URRBMI) to cover both urban and rural residents. The URRBMI has broken the urban-rural dual structure of the past medical insurance system, unified the reimbursement drug catalog and designated medical institutions, and eliminated the gap between urban and rural medical insurance benefits.By the end of 2019, most cities have completed the integration.

Scrutinizing and evaluating the effects of integration policy is crucial to finding healthcare solutions that promote social equity and benefit all people. The labor supply, as a direct source of wealth creation, is of paramount importance to individuals' survival and development.For this purpose, we investigatethe impact of integration policy on labor supplywhichis a non-negligible economic variable. Throughout the relevant literature, there are two opposite paths of the impact of health insurance on labor supply. On the one hand, health insurance can compensate partial healthcare expenses of the insured, which is actually equivalent to increasing the household income. At this point, workers will choose to replace labor with more leisure so that health insurance will reduce labor supply(4, 5). On the other hand, getting access to health insurance can promote workers to increase the utilization of healthcare services, thus promoting health improvement, reducing the time spent sick, and ultimately increasing the probability that the insured choose to participate in work and increase working hours (6-9). Measures of the urban-rural health insurance integration include increasing reimbursement rates, expanding the list of reimbursed drugs, allowing non-local reimbursement, and reducing deductibles, etc. Theoretically both income effectand health effect exist, and the magnitude of each effect is ambiguous. The extent of the integration policy's

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impacts on labor supply and how they are channeled calls for rigorous empirical analysis.

Taking advantage of the gradual rollout of the integration policy at the prefecture level, we use a staggered difference-in-differences (DID) model to estimate the treatment effects. We exploit data from four waves of a large Chinese household survey, China Health and Retirement Longitudinal Study(CHARLS), to match individual-level information with the integration timeof cities. From baseline results, we find that the implementation of integration policy significantly increases the probability of rural residents having more than one job by 0.618% and significantly decreases the likelihood of rural residents missing work due to illness by 3.73%. This result suggests that integration policy increases the employment flexibility and provides protection for rural residents' work,which has brought significant welfare effects to rural residents in China.

To test the parallel trend hypothesis and observe dynamic treatment effects, we use event study estimation. Moreover,considering the heterogeneity treatment effects, we applySun and Abraham (10)'s approach. The event study estimation results show that the coefficients on interaction terms are small and indistinguishable from zero before the reform occurs. This attests the parallel trend hypothesis to ensure that the target variables in the treatment and control groups share a common pre-trend. The estimation results of the event study also verify the robustness of the staggered DID results. After implementation of the policy, the absolute values of estimated coefficients become larger and significantly distinguishable from zero. This implies that the labor supply in the treatment group is significantly different from that in the control group at post-integration period.

Next, we further scrutinize the effects of integration policy on eightjob statuses and five reasons for leaving the labor market. The findings reveal that integration policy increases the probability of non-agricultural employment and reduces the likelihood of quitting the labor market for health reasons. It demonstrates additional evidence of the welfare effects of integration policy, suggesting a role in promoting rural-urban labor mobility, improving the health status of participants, and delaying retirement. Moreover, the impact of increasing the probability of having more than one job occurs mainly in the low-income and younger groups, and the impact of reducing the likelihood of working due to illness occurs mainly in the high-income and younger groups.

We furtherverify the robustness and validity of our model. We restrict the samples to the cities that implemented integration policy over 2011-2018, and the results remained consistent with the baseline. We then conduct two placebo tests, including randomly setting policy time and treatment group. The true policy effects are significantly different from the placebo test results, suggesting that the changes in labor supply are not caused by other random factors. The robustness of our results is demonstrated.

Our study contributes in the following two aspects. First, we address the endogeneity problem using a quasi-natural experiment and a causal inference approach. There have been many researches combining the quasi-experiment with DID methods (e.g., 11-15). To alleviate the endogeneity between health insurance participation and outcome variables, these literature use DID method.

We complement this type of literature by studying the impact of a quasi-natural experiment, urban-rural health insurance integration policy, on labor supply using a staggered approach which differs from traditional DID. In addition, we keep up with the latest research and take heterogeneous treatment effects into consideration, making our arguments more rigorous and credible.

Second, our study contributes to the literature on the evaluation of the effects of urban-rural health insurance integration policies in China. Most scholars have focused on the effects of URBMI or NRCMS on labor supply (16-19), while relatively little literature has previously focused on the implementation effects of the urban-rural health insurance integration policy, especially on labor supply. Chang, Su, and Zhou (20) use Chinese Labor-force Dynamics Survey (CLDS) data, find that integration policy benefits residents' health and mitigates the negative effects of health shock on work, and this effect is more pronounced among rural residents. Huang and Wu (21) use CHARLS data of year 2011-2015 and find that integration significantly increases the utilization of inpatient care among middle-aged and older rural residents, and this positive impact is particularly pronounced in poorer areas due to the increase in health insurance benefits. This paper fills a gap in the literature by exploring the impact of urban-rural health insurance integration from the labor supply perspective.

The remainder of this paper is organized as follows. Section 2 introduces the data sources, variable selection and empirical strategy. Section 3 discusses the empirical results, including the baseline results, heterogeneity analysis, robustness test and placebo test. Section 4 concludes.

## 2 Materials and methods

### 2.1 Data

The data used in this paper are from the China Health and Retirement Longitudinal Study (CHARLS), a survey conducted by the National School of Development of Peking University. The baseline survey of CHARLS started in 2011, targeting residents aged 45 years and older in China. These samples are followed up every two or three years, with a follow-up rate of about 80%. The CHARLS questionnaire contains a large amount of information about individuals and households, including demographic backgrounds, health status, cognitive status, labor supply, household income and expenditure, etc. It can provide strong support for our empirical study. In this paper, we use data from four waves of the CHARLS in year 2011, 2013, 2015, and 2018, which basically cover the pre- and post-integration periods of urban and rural health insurance integration in most cities.

To meet the needs of our research, first, we restrict the sample to rural hukou individuals, which is because the integration policy mainly raises the benefits of rural residents. Second, we only retain the sample aged 45-75 years old, which is because the elderly people have lost their labor capacity, and including the elderly in the regression will cause the estimation errors. Third, we also exclude individuals who have never participated in any social health insurance program in the four waves of CHARLS for the reason that they are not affected by the integration policy. Finally, we obtain 50,012 observations

from 15,888 individuals, which are from 124 cities in 28 provinces in China.

2.2 Variables

The key explanatory variable $Integration_{ct}$ is a dummy variable that indicates whether a prefecture-level city $c$ implemented the urban-rural basic medical insurance integration policy in year $t$. Data on policy implementation in prefecture-level cities are mainly obtained from two sources: the Peking University (PKU) Law database, and the official websites of each city's government and human resources and social security bureau. PKU Law database includes national laws, administrative regulations, local regulations, etc. By searching local regulations and documents, we get the accurate time when some prefecture-level cities implemented the integration policy. For those cities not included in the database, we choose to search documents and news on the official websites of the government. Ultimately, we obtain data on the timing of integration policy integration for 124 prefecture-level cities, all of which hand completed integration between 2004 and 2020. The policy implementation dates are precise to the date. By comparing the survey date and the implementation date of the integration policy, we obtain the values of the key explanatory variable $Integration_{ct}$.

The explanatory variable in this paper is the status of individuals' labor supply. We choose whether the individual is currently working now ($Work1$), whether the individual has more than one job ($Work2$), working hours per week ($Work\_time$), whether the individual missed work because of sickness in the past month ($Sick$). In addition to the variables in the main regression, variables such as types of work and reasons for leaving the labor market are selected to represent the labor force status of older adults in further analysis.

To mitigate endogeneity and reduce the estimation error, we also select individual-level and city-level control variables for regression. Individual-level control variables include marital status, age, gender, education level, and number of children. City-level control variables include public expenditure, GDP per capita, GDP growth rate, and unemployment rate, which represent the level of cities' economic and social development. The city-level data are obtained from the China City Statistical Yearbook sponsored by National Bureau of Statistics. Considering currency inflation, we take the 2011 CPI as the baseline and calculate the real GDP and real government expenditure excluding the inflation rate. Moreover, we take the natural logarithm of the variables age, government expenditure, and GDP per capita to alleviate data volatility and solve the heteroskedasticity problem. Detailed definitions of the variables are displayed in Table 1.

2.3 Summary statistics

Table 2 shows the summary statistics for each variable. In our sample, 72.1% of the rural elderly are still currently working and 2.2% have more than one job, which is consistent with the actual situation and indicates that "ceaseless toil" still exists in the rural elderly (22). On average, rural residents work 44.02 hours per week, and the probability of missing work due to illness is 34%. As for the control variables at city-level, we find that over the four

waves, the average city public expenditure is 2.93 million ($e^{14.89}$) yuan, the average real GDP per capita is 43,045 ($e^{10.76}$) yuan, the average GDP growth rate is 10.4%, and the average unemployment rate is 2.9%. At the individual level, the average age of the sample is 58 ($e^{4.06}$) years old, 82.4% of the sample are married, and the average education background is elementary school and below. 46.9% of the sample are male, and there are 2-3 living children in the household on average.

2.4 Empirical strategy

As a common method to assess the effects of policies, DID has been widely used in empirical economics research. The principle of DID is to construct a treatment group with policy treatment and a control groupwithout policy treatment, and to explain the policy effect by comparing the differences between the treatment group and the control group before and after the policy, while controlling for other factors. Traditional DID assumes that all individuals in the treatment group start to experience the policy shock at exactly the same point in time. However, in some cases, individuals in the treatment group do not receive the treatment at the same time, such as when the policy starts as a pilot and is gradually rolled out, which constitutes a staggered DID model (23-25). Since the integration policy is rolled out on a city-by-city basis, we use this feature to estimate by a staggered DID approach. We use a two-way fixed effects (TWFE) model for estimation. The regression equation is as follows:

$$Y_{ict} = \alpha + \theta Integration_{ct} + \beta X_{ict} + \mu_c + \lambda_t + \varepsilon_{ict} \#(1)$$

Where $Y_{ict}$is an outcome variable indicating the labor supply of individual $i$at city $c$ in year $t$, including work status, work hours, and whether missed work due to illness. $Integration_{ct}$ is the treatment state variable, i.e., whether city $c$ has implementedthe integration policy in year$t$. $X_{ict}$ is the control variable at individual level and city level (see subsection 2.2). $\mu_c$ is the time-invariant unobservable city fixed effect. $\lambda_t$ is the year fixed effect. $\varepsilon_{ict}$ is the error term. Standard errors are interactively clustered at the level of city and year.$\theta$is our coefficient of interest, i.e., the DID estimator. $\theta$ equals the average treatment effect (ATE), as shown in the following equation:

$$ATE = \{E[y_1|D_{ct} = 1] - E[y_1|D_{ct} = 0]\}\#$$
$$- \{E[y_0|D_{ct} = 1] - E[y_0|D_{ct} = 0]\}$$
$$= (Y_{after} - Y_{before}) - (C_{after} - C_{before}) = (\theta + \lambda_t) - \lambda_t$$
$$= (Y_{after} - C_{after}) - (Y_{before} - C_{before}) = (\theta + \mu_c) - \mu_c = \theta\#(2)$$

The staggered DID has been widely used in the field of policy evaluation, however, some theoretical econometricians have recently found that the staggered DID method has the estimation bias (26-28). While using TWFE for staggered DID estimates, the heterogeneous treatment effects are often overlooked. The same policy treatment produces different effects for different individuals, and such differences may be manifested in two dimensions: the length of time since receiving the treatment or the group receiving the treatment at different points in time. When policy effects change over time, the result

estimated by the staggered DID, i.e., $\boldsymbol{\theta}$ of Eq. (1), is not well-defined average treatment effects, but rather a weighted average of ATE estimated by multiple standard DID methods, and the weights may be negative (27). This implies that the single coefficient estimates of the staggered DID are no longer credible in the presence of heterogeneous treatment effects. Instead, the dynamic effects test of the staggered DID, i.e., the event study method is an analytical tool that may cope with this situation, so we employ it as an alternative empirical strategy. We use the method proposed by Sun and Abraham (10) to obtain the "heterogeneity-robust" estimator. The principle of this method is to select "not yet treated" and "never treated" samples as control groups, calculate the cohort-specific average treatment effects on the treated (CATT), and obtain the dynamic treatment effects by weighting the sum on the basis of CATT[1]. The specification of the event study model can be defined as:

$$Y_{ict} = \alpha + \sum_{j=2}^{4} \eta_j (Lagj)_{ct} + \sum_{k=0}^{4} \gamma_k (Leadk)_{ct} + \beta X_{ict} + \mu_c + \lambda_t + \varepsilon_{ict} \#(3)$$

Where $Lagj$ and $Leadk$ are binary variables indicating the $j$ years before and $k$ years after the integration policy respectively, and the other variables are the same as those in Eq. (1). To avoid collinearity, the variable $Lag1$ (one year before the event) is omitted from the estimation. In general, a wider event window will cover a larger sample size and the estimation is more efficient but there may be a larger bias. Therefore, the width of the event window should not be too large, so we choose four years before/after the policy as the event window period[2]. It is a remarkable fact that Eq. (3) can not only display the dynamic change of treatment effects, but also test the parallel trend hypothesis. The underlying assumption of the DID design is that the treatment and control groups do not differ systematically prior to the event, namely, all individuals have similar time trends in labor force outcomes. We can examine whether the parallel pre-trend holds by observing whether the coefficient of the lag term $\eta_j$ is significantly different from zero. The result we expect to obtain is that the coefficients are not significantly different from zero before the policy, while the coefficients become distinguishable from zero after the policy.

## 3 Empirical results

In this section we report the empirical results. First, we report the regression results of the main model staggered DID, i.e., the effect of the integration policy on labor supply. Next, we present event study estimates to manifest the dynamic treatment effects and validate theparallel trend hypothesis.Then, we examine the effect of the integration policy on the types of jobs and reasons for quitting the labor market to further disentangle the effect

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[1] Intuitively, since treatment effects may be heterogeneous due to the different time points of first treatment and the different duration of treatment, treatment effects are the same for samples treated at the same time point and for the same duration after treatment.
[2] We also chose 5, 6, and 7 years before/after the policy as window widths, and the results remain robust.

and mechanism. We also group the sample by income and age to explore the heterogeneity of the effects. Finally, we test the robustness of the model using reduced sample size and placebo tests.

3.1 Baseline results

The baseline results are shown in Table 3. Table 3 presents the effects of the integration policy on the labor force variables of currently working, having more than one job, work time, and missing work due to illness, respectively. Columns (1), (3), (5), and (7) demonstrate the results without individual-level and city-level covariates; columns (2), (4), (6), and (8) demonstrate the results with covariates. Columns (1)-(8) all include city fixed effects and year fixed effects. The results show that the implementation of the health insurance integration policy significantly increases the probability of having more than one job by 0.618% and significantly decreases the likelihood of missing work due to illness by 3.73% for older adults. After controlling individual and city-level variables, the integration policy significantly increases the probability of having more than one job by 0.524% and significantly decreases the likelihood of missing work due to illness by 3.26%, which is consistent with the regression results without covariates, indicating the robustness of the results. In addition, the integration policy increases the probability of currently working and decreases work hours, which it is not statistically significant. In summary, it can be concluded that the integration policy has a welfare effect, increasing the flexibility of work and improving the health of the insured, reducing the likelihood of missing work due to health shocks such as illness.

3.2 The dynamic impacts of the policy reform on labor supply

We plot the results of the event study estimates in Figures 1 and 2, showing the dynamic impacts of the integration policy on the variables $Work2$ and $Sick$, respectively. For both variable $Work2$ and $Sick$, we observe small and insignificant estimates in the years before the integration policy implementation, suggesting that there are no pre-trends which might confound our findings. In other words, the trends of individual labor supply in the treatment and control groups before the policy are largely consistent. The parallel trend hypothesis can be validated. For the variable $Work2$, although the pre-trend is not 100% satisfactory, i.e., the impact at year -3 is small but statistically different from zero, the impact becomes positive and distinguishable from zero in the years after experiencing the policy reform. The magnitude of impact is around 0.8% in the year t=1 (one year after integration) and then steadily goes up to 1.6% in the year t=3 (three years after integration). The estimated impact of integration on $Sick$ undergoes a similar change. Before the reform occurs, the coefficients are not significantly different from zero and are all positive. However, from year 0 (the policy year),we observe a conspicuous sharp jump which indicates that the integration policy has a significant impact on the likelihood of missing work due to illness. The impact is around -5% to -7%, which is approximately the same as the staggered DID estimates.

According to the graph, we can find that for $Work2$, the impact does not take place immediately, but it starts to change at one year after reform. For

*Sick*, the impact begins to occur in the year when the policy was launched. We can attribute this phenomenon to the time and labor costs involved in finding a job. People will decide whether to find a part-time job after deliberation, and it is not easy for rural residents to find a non-agricultural job, which means that changes in labor supply require reaction time. On the contrary, the increased availability of health care resources allows rural residents to seek immediate medical care and reduces the likelihood of missing work due to illness. As a result, the impact has occurred since the reform year and gradually increased.

3.3 Decomposition of work status

First, we split the work status into eight types to further investigate the impact of the integration policy on work status. The eight work statuses include: agricultural employed, agricultural self-employed, non-agricultural employed, non-agricultural self-employed, non-agricultural family business, unemployed, retired, and never work. The results are shown in Table 4. The integration policy significantly increases the probability of non-agricultural employed by 2.79%, and has no significant impact on other work types. Combined with the findings of the baseline regressions, the insured's probability of having more than one job increases, who choose to engage in non-agricultural employment as a side job. The increase in non-agricultural employment alleviates the problem of rural residents' employment being restricted to agriculture for a long time, promotes rural residents' job-hunting by self-determination, and improves their employment flexibility. This shows that the policy of urban-rural health insurance integration has achieved its original purpose, which is conducive to narrowing the urban-rural gap, coordinating urban-rural development, and promoting the orderly flow and competition of labor force between urban and rural areas.

Second, we analyze the effect on the reasons for rural residents to quit the labor market. We classify the reasons into five categories according to the questionnaire content: retired, fired, laid off, health reason, family reason, and find new job. The results are shown in Table 5. The integration policy significantly reduces the likelihood of quitting the labor market for health reasons by 7.62%, and has no significant impact on other quitting reasons. Combined with the findings in subsection 3.1, the integration policy improves the health of the insured and enables them to continue working in the labor market in a healthy way. At the same time, we can attribute the increase in the number of jobs held to the improvement in health. The integration policy has increased the level of benefits for rural residents, reduced their burden of access to health care, improved their access to health care behavior, and thus improved their health. Rural residents have more time and energy to engage in agricultural and non-agricultural work. Therefore, the probability of having more than one job has increased.

3.4 Heterogeneity analysis

We have already obtained the conclusion that integration policy increases the probability of having more than one job and decreases the likelihood of missing work due to illness. To further investigate the heterogeneity impact on different groups, we classify the sample by income level and age, and the results are shown in Tables 6 and 7.

First, we divide the sample into three groups according to individual income: low income, middle income, and high income[1]. Individual income includes individual wage income, pensions, government subsidies, social income transfer sources, and other transfer income. The results in Table 6 show that the integration policy significantly increases the probability of having more than one job for the low-income group by 0.97%, significantly decreases the likelihood of missing work due to illness for the middle-income group by 5.11%, and significantly decreases the likelihood of missing work due to illness for the high-income group by 4.31%. The impact of increasing work occurs mainly in the low-income group and the impact of being less affected by health shocks occurs mainly in the middle and high-income groups. This is because the low-income group is more focused on improving their income, and with better health care coverage, the low-income group are able to find side jobs to increase their income. The middle and high-income groups, on the other hand, have more wealth and are more resilient to risk, so they are less likely to look for new jobs and are healthier to cope with their original jobs. It is worth noting that the fact that low-income people are not less likely to miss work due to illness does not indicate that their health has not improved. However, low-income people choose to keep working even when their health is poor, so as to maintain the general level of household income (29).

Second, we regress the sample into three different age groups using a group of 10 years. The results are shown in Table 7. For individuals aged 45-55, the likelihood of missing work due to illness decreases significantly by 4.69%. For individuals aged 55-65, the probability of having more than one job increases significantly by 0.712%, and the likelihood of missing work due to illness decreases significantly by 5.36%. For individuals aged 65-75, the integration policy has no significant impact. This is also in line with the reality. The younger groups are more likely to change their labor supply behavior and obtain welfare due to the improvement of health care. Older people, on the other hand, are less likely to participate in the labor market themselves, so the policy has less impact on their labor force performance.

3.5 Robustness test

In addition to changing the covariates and applying the event study method for robustness test, we use reduced sample for robustness test in this section. We restrict the sample to cities that have implemented integration policies in 2011 to 2018, leaving 41,335 observations from 101 cities. The results are shown in Table 8 and are consistent with the baseline results in Table 3, indicating that our estimation results are robust.

3.6 Placebo test

Torule out that our results are not caused by accidental causes, we also conduct a placebo test. The placebo test is conducted by randomly setting the implementation time of the integration policy in each city and randomly selecting the treatment group. Since the pseudo policy times and pseudo

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[1] Similar conclusions are obtained when the samples are grouped using per capita household income, which includes: personal wage income and transfer income of household members, income from household agricultural activities, income from self-employment or starting a business, and income from public transfers received by the household.

treatment groups are randomly generated, the policy variables should not have a significant effect on the labor outcome variables, i.e., the regression coefficients of the pseudo treatment variables should be around zero. Otherwise, the model setting is biased. The model is estimated by repeating above stochastic process 500 times and plotting the kernel density of the estimated coefficients of the pseudo policy variables. The results are displayed in Figure 3 and Figure 4. The estimated coefficients under the two stochastic processes are distributed around zero and obey a normal distribution, as expected. Meanwhile, the actual estimated coefficients lie within the range of small probability events in the kernel density plots. In other words, the treatment effect of the integration policy on labor supply is not a chance event. Our findings are reliable and robust.

## 4 Discussion and conclusion

This paper estimates the impact of China's urban-rural health insurance integration policy on the labor supply of rural residents using CHARLS 2011-2018 data. Taking advantage of the gradual implementation of the integration policy, we utilize the staggered DID approach for estimation. We also use an event study approach to examine parallel trends and dynamic treatment effects. We conclude that the implementation of the integration policy significantly increases the probability of rural residents having more than one job by 0.618% and significantly decreases the likelihood of rural residents missing work due to illness by 3.73%. This suggests that the integration policy increases the availability of medical resources by improving rural residents' insurance benefits, improves their health, increases their employment flexibility, and provides protection for their work in a healthy way.

Our findings have important policy implications. First, we provide evidence that expanding public health insurance in developing countries will lead to significant welfare effects. Many previous studies have found that public insurance programs will discourage work motivation and increase employment lock-in. Policymakers may refrain from further expanding public health insurance for fear of creating adverse labor market incentives (4, 30, 31). The findings in this paper eliminate this concern by showing that public health insurance expansion does not result in efficiency loss. Second, we find new evidence on the role of the health insurance integration policy in mitigating rural-urban development inequality. The integration policy increases the probability of non-agricultural employment of rural residents, which suggests that integration policy liberates farmers from the land, increases urban-rural labor mobility, and helps break the urban-rural dual structure. Third, after conducting heterogeneity analysis, we find that low-income groups and older groups do not receive more significant welfare than higher-income groups and younger groups. In the future, policymakers should focus more on protecting vulnerable groups. A small portion of resources may generate substantial welfare improvements when these resources are properly allocated to vulnerable groups, such as the rural low-income residents and elderly, to ultimately enhance the well-being of the entire population.

There are some limitations in our study. Due to data limitations, the survey data we use does not include individuals' choices of work locations, and we are

not able to study the effect of integration policy on labor migration and mobility. Moreover, the reform was initiated only in recent years and the limited welfare effects may be attributed to the short duration of our study. Future research should pay more attention on the long-term impacts of integration policy and other welfare effects.

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## 5 Conflict of Interest

The authors declare that the research was conducted in the absence of any commercial or financial relationships that could be construed as a potential conflict of interest.

## 6 Funding

## 7 Data Availability Statement

Publicly available datasets were analyzed in this study. This datacan be found here: http://charls.pku.edu.cn/

Table 1 Variable definitions

| Variables | Definition |
|---|---|
| *I. Independent variables* | |
| Integration | =1 if the city has implemented the integration policy; =0 if not |
| *II. Dependent variables* | |
| Work | =1 if the individual is currently working; =0 if not |
| Work2 | =1 if the individual has more than one job; =0 if not |
| Work_time | The hours of work per week at current job, including primary work and other work |
| Sick | =1 if missed work because of sickness in the past month ; =0 if not |
| *III. Control variables: city level* | |
| Ln (Expenditure) | The logarithm of the public expenditure of the city (yuan) |
| Ln (GDP) | The logarithm of the GDP per capita of the city (yuan) |
| GDP_growth | Real GDP growth |
| Unemp_rate | The unemployment rate of the city |
| *IV. Control variables: individual level* | |
| Marriage | =1 if married; =2 if partnered; =3 if separated; =4 if divorced; =5 if widowed; =6 if never married |
| Ln (Age) | The logarithm of the individual's real age |
| Gender | =1 if male; =0 if female |
| Edu | =1 if no formal education (illiterate); =2 if did not finish primary school; =3 if sishu (an old-style private school in China)/home school; =4 if elementary school; =5 if middle school; =6 if high school; =7 if vocational school; =8 if two/three-year college/associate degree; =9 if four-year college/bachelor's degree; =10 if post-graduated (master/PhD) |
| Child | The number of living children |

Table 2 Summary Statistics

| Variables | Mean | Std. Dev. | Minimum | Maximum | Median |
|---|---|---|---|---|---|
| Work | 0.721 | 0.449 | 0 | 1 | 1 |
| Work2 | 0.0220 | 0.147 | 0 | 1 | 0 |
| Work_time | 44.02 | 25.50 | 0 | 312 | 42 |
| Sick | 0.340 | 0.474 | 0 | 1 | 0 |
| Ln (Expenditure) | 14.89 | 0.716 | 13.07 | 18.14 | 14.84 |
| Ln (GDP) | 10.76 | 0.656 | 8.736 | 15.68 | 10.76 |
| GDP_growth | 0.104 | 0.0390 | -0.0290 | 0.212 | 0.102 |
| Unemp_rate | 0.0290 | 0.0170 | 0.00600 | 0.0760 | 0.0260 |
| Marriage | 1.484 | 1.222 | 1 | 6 | 1 |
| Ln (Age) | 4.061 | 0.135 | 3.807 | 4.317 | 4.060 |

| | | | | | |
|---|---|---|---|---|---|
| Gender | 0.469 | 0.499 | 0 | 1 | 0 |
| Edu | 3.167 | 1.752 | 1 | 10 | 4 |
| Child | 2.641 | 1.280 | 0 | 10 | 2 |

Table 3 The impact of the integration policy on labor supply

| Variables | (1) work1 | (2) work1 | (3) work2 | (4) work2 | (5) work_time | (6) work_time | (7) sick | (8) sick |
|---|---|---|---|---|---|---|---|---|
| Integration | 0.00526 | 0.00506 | 0.00618*** | 0.00524** | -0.889 | -0.459 | -0.0373** | -0.0326* |
| | (0.0165) | (0.0170) | (0.00237) | (0.00249) | (0.985) | (1.026) | (0.0181) | (0.0176) |
| Ln (Expenditure) | | -0.0425*** | | 0.00265* | | 0.416 | | -0.0223* |
| | | (0.0122) | | (0.00150) | | (0.644) | | (0.0117) |
| Ln (GDP) | | -0.0307* | | -0.00247* | | -0.730 | | -0.00592 |
| | | (0.0165) | | (0.00145) | | (0.540) | | (0.0154) |
| GDP_growth | | -0.0577 | | -0.00604 | | -3.172 | | 0.243 |
| | | (0.192) | | (0.0286) | | (11.33) | | (0.191) |
| Unemp_rate | | 0.0431 | | -0.128** | | 9.106 | | -0.332 |
| | | (0.340) | | (0.0504) | | (19.83) | | (0.321) |
| Marriage | | -0.0165*** | | -2.33e-05 | | -0.338** | | 0.00148 |
| | | (0.00185) | | (0.000492) | | (0.136) | | (0.00184) |
| Ln (Age) | | -0.735*** | | -0.0705*** | | -35.59*** | | 0.0273 |
| | | (0.0367) | | (0.00716) | | (1.839) | | (0.0259) |
| Gender | | -0.143*** | | -0.0152*** | | -6.436*** | | 0.0142** |
| | | (0.00766) | | (0.00173) | | (0.456) | | (0.00691) |
| Edu | | -0.00532** | | 0.000837 | | -0.281** | | -0.0131*** |
| | | (0.00207) | | (0.000516) | | (0.129) | | (0.00228) |
| Child | | -0.00869** | | 0.000855 | | 0.0843 | | 0.00850** |
| | | (0.00381) | | (0.000730) | | (0.188) | | (0.00356) |
| City FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Year FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Covariates | No | Yes | No | Yes | No | Yes | No | Yes |
| Observations | 49,652 | 46,796 | 35,828 | 34,085 | 34,726 | 33,063 | 36,109 | 34,368 |
| R-squared | 0.028 | 0.139 | 0.003 | 0.012 | 0.029 | 0.076 | 0.001 | 0.108 |

*Note:* (1) Robust standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1. (2) Standard errors are interactively clustered at the level of city and year. (3) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level.

Table 4 The impact of the integration policy on types of work

| Variables | (1) agri_emp | (2) agri_sfemp | (3) nonagri_emp | (4) nonagri_sfemp | (5) nonagri_fb | (6) unemp | (7) retired | (8) neverwork |
|---|---|---|---|---|---|---|---|---|
| Integration | -0.00343 | -0.0217 | 0.0279* | -0.00108 | 0.00348 | 0.00542 | -0.00700 | -0.00363 |
| | (0.00515) | (0.0214) | (0.0147) | (0.0106) | (0.00519) | (0.00339) | (0.0157) | (0.00407) |
| Ln (Expenditure) | -0.000222 | -0.0579*** | 0.0264*** | -0.00528 | 0.000335 | 0.00261 | 0.0353*** | -0.00127 |
| | (0.00324) | (0.0155) | (0.00834) | (0.00583) | (0.00254) | (0.00191) | (0.0109) | (0.00292) |
| Ln (GDP) | 0.00262 | -0.0497** | 0.00740 | 0.0119** | -0.00190 | 0.000505 | 0.0276* | 0.00155 |
| | (0.00214) | (0.0214) | (0.00495) | (0.00601) | (0.00169) | (0.000477) | (0.0149) | (0.00265) |
| GDP_growth | 0.105* | 0.158 | -0.143 | -0.164* | 0.0409 | -0.00895 | -0.0253 | 0.0372 |
| | (0.0563) | (0.217) | (0.126) | (0.0969) | (0.0419) | (0.0140) | (0.181) | (0.0478) |
| Unemp_rate | 0.322*** | -0.161 | 0.198 | -0.217 | -0.0502 | 0.0149 | -0.183 | 0.0753 |
| | (0.112) | (0.373) | (0.229) | (0.145) | (0.0763) | (0.0232) | (0.318) | (0.0939) |
| Marriage | -0.00220*** | -0.0217*** | 0.0102*** | -0.00226*** | -0.00113*** | -1.54e-05 | 0.0157*** | 0.00146** |
| | (0.000493) | (0.00159) | (0.00154) | (0.000628) | (0.000371) | (0.000117) | (0.00184) | (0.000579) |

| | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| Ln (Age) | -0.0528*** | 0.188*** | -0.610*** | -0.302*** | -0.0208** | -0.0192** | 0.799*** | 0.0173* |
| | (0.0110) | (0.0339) | (0.0265) | (0.0221) | (0.00850) | (0.00924) | (0.0344) | (0.00919) |
| Gender | 4.45e-05 | 0.0332*** | -0.150*** | -0.0432*** | 0.00541** | -0.00250* | 0.143*** | 0.0145*** |
| | (0.00268) | (0.00708) | (0.00706) | (0.00455) | (0.00241) | (0.00135) | (0.00737) | (0.00212) |
| Edu | -0.00349*** | -0.0232*** | 0.0162*** | 0.00558*** | 0.00189** | 3.56e-05 | 0.00384* | -0.000920** |
| | (0.000735) | (0.00224) | (0.00188) | (0.00126) | (0.000775) | (0.000188) | (0.00207) | (0.000421) |
| Child | -0.00483*** | 0.00462 | -0.0135*** | 0.00469*** | 0.000824 | 0.00102 | 0.00673* | 0.000406 |
| | (0.00122) | (0.00367) | (0.00235) | (0.00154) | (0.000854) | (0.000957) | (0.00368) | (0.000943) |
| City FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Year FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Covariates | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | 48,441 | 48,441 | 48,441 | 48,441 | 48,441 | 48,441 | 48,441 | 48,441 |
| R-squared | 0.012 | 0.071 | 0.143 | 0.048 | 0.004 | 0.008 | 0.115 | 0.050 |

*Note:* (1) Robust standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1. (2) Standard errors are interactively clustered at the level of city and year. (3) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level.

Table 5 The impact of the integration policy on reasons for quitting the labor market

| Variables | (1) retired | (2) fired | (3) health | (4) family | (5) newjob |
|---|---|---|---|---|---|
| Integration | -0.0210 | 0.0199 | -0.0762** | 0.0201 | 0.00132 |
| | (0.0214) | (0.0172) | (0.0329) | (0.0187) | (0.0108) |
| Ln (Expenditure) | 0.0319** | 0.00281 | -0.0207 | -0.00583 | 0.00977 |
| | (0.0130) | (0.0118) | (0.0202) | (0.0128) | (0.00802) |
| Ln (GDP) | -0.00723 | 0.0266* | -0.0428* | -0.00254 | 0.00200 |
| | (0.00757) | (0.0136) | (0.0252) | (0.0117) | (0.00484) |
| GDP_growth | 0.0757 | -0.0155 | 0.0307 | -0.00266 | -0.224 |
| | (0.186) | (0.183) | (0.354) | (0.180) | (0.136) |
| Unemp_rate | 0.0530 | 0.292 | 0.416 | -0.509 | 0.497 |
| | (0.348) | (0.349) | (0.595) | (0.392) | (0.352) |
| Marriage | -0.00448* | 0.00303 | -0.00577 | 0.00635** | -0.00411*** |
| | (0.00229) | (0.00242) | (0.00384) | (0.00285) | (0.00139) |
| Ln (Age) | 0.287*** | -0.121** | 0.271*** | -0.412*** | 0.0234 |
| | (0.0599) | (0.0473) | (0.0773) | (0.0616) | (0.0696) |
| Gender | -0.00592 | -0.0266** | -0.0723*** | 0.0941*** | -0.00403 |
| | (0.0121) | (0.0118) | (0.0168) | (0.0125) | (0.00789) |
| Edu | 0.00957*** | 0.00506 | -0.0268*** | 0.00751* | 0.0117*** |
| | (0.00352) | (0.00321) | (0.00555) | (0.00421) | (0.00266) |
| Child | -0.0129*** | 0.00439 | 0.00936 | 0.00468 | -0.0133 |
| | (0.00422) | (0.00419) | (0.00669) | (0.00429) | (0.00815) |
| City FE | Yes | Yes | Yes | Yes | Yes |
| Year FE | Yes | Yes | Yes | Yes | Yes |
| Covariates | Yes | Yes | Yes | Yes | Yes |
| Observations | 9,652 | 9,652 | 9,652 | 9,652 | 9,652 |
| R-squared | 0.092 | 0.046 | 0.171 | 0.124 | 0.066 |

*Note:* (1) Robust standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1. (2) Standard errors are interactively clustered at the level of city and year. (3) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level.

Table 6 Heterogeneous effects of the integration policy: grouped by income

| (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|

| Variables | Low income | | Middle income | | High income | |
|---|---|---|---|---|---|---|
| | work2 | sick | work2 | sick | work2 | sick |
| Integration | 0.00970** | 0.00314 | 0.00259 | -0.0511** | 0.00409 | -0.0431* |
| | (0.00392) | (0.0236) | (0.00450) | (0.0256) | (0.00445) | (0.0240) |
| Ln (Expenditure) | 0.000271 | -0.0401** | 0.00347 | -0.0279* | 0.00305 | -0.00271 |
| | (0.00225) | (0.0163) | (0.00267) | (0.0143) | (0.00286) | (0.0155) |
| Ln (GDP) | -0.00353* | -0.00546 | -0.00328 | -0.00820 | -0.000622 | 0.00153 |
| | (0.00201) | (0.0140) | (0.00243) | (0.0151) | (0.00326) | (0.0262) |
| GDP_growth | 0.0852** | 0.466* | -0.0515 | -0.0347 | -0.0511 | 0.180 |
| | (0.0393) | (0.276) | (0.0527) | (0.222) | (0.0570) | (0.274) |
| Unemp_rate | -0.124* | -1.109** | -0.134 | -0.0576 | -0.106 | 0.252 |
| | (0.0700) | (0.474) | (0.0926) | (0.414) | (0.0991) | (0.440) |
| Marriage | -0.000480 | 0.00409 | -0.00121 | 0.00501 | 0.00144 | -0.00345 |
| | (0.000613) | (0.00281) | (0.00114) | (0.00346) | (0.000917) | (0.00297) |
| Ln (Age) | -0.0629*** | 0.0196 | -0.0734*** | 0.0792 | -0.0669*** | 0.00694 |
| | (0.0102) | (0.0459) | (0.0136) | (0.0503) | (0.0138) | (0.0450) |
| Gender | 0.0108*** | 0.00155 | 0.0179*** | -0.0365*** | 0.0170*** | -0.00668 |
| | (0.00247) | (0.0106) | (0.00320) | (0.0116) | (0.00330) | (0.0124) |
| Edu | -6.96e-06 | -0.0103*** | 0.000379 | -0.0135*** | 0.00231** | -0.0174*** |
| | (0.000734) | (0.00331) | (0.000941) | (0.00385) | (0.000997) | (0.00376) |
| Child | 0.000119 | 0.00871 | 0.000621 | 0.00314 | 0.00155 | 0.0143*** |
| | (0.000988) | (0.00558) | (0.00141) | (0.00601) | (0.00141) | (0.00549) |
| City FE | Yes | Yes | Yes | Yes | Yes | Yes |
| Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| Covariates | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | 11,620 | 11,932 | 11,819 | 11,957 | 10,646 | 10,479 |
| R-squared | 0.012 | 0.109 | 0.013 | 0.122 | 0.012 | 0.108 |

*Note:* (1) Robust standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1. (2) Standard errors are interactively clustered at the level of city and year. (3) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level.

Table 7 Heterogeneous effects of the integration policy: grouped by age

| Variables | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Age$\in$ [45, 55) | | Age$\in$ [55, 65) | | Age$\in$ [65,75] | |
| | work2 | sick | work2 | sick | work2 | sick |
| Integration | 0.00458 | -0.0469** | 0.00712* | -0.0536** | 0.00113 | 0.0190 |
| | (0.00468) | (0.0228) | (0.00371) | (0.0232) | (0.00371) | (0.0312) |
| Ln (Expenditure) | 0.00713** | -0.0212 | -9.34e-05 | -0.0197 | 0.000371 | -0.00850 |
| | (0.00283) | (0.0144) | (0.00222) | (0.0131) | (0.00229) | (0.0163) |
| Ln (GDP) | -0.00689** | -0.00302 | -0.000150 | -0.0312** | 0.000575 | 0.0166 |
| | (0.00298) | (0.0122) | (0.00227) | (0.0144) | (0.00178) | (0.0132) |
| GDP_growth | 0.0266 | 0.302 | -0.0447 | 0.151 | 0.00481 | 0.397 |
| | (0.0524) | (0.220) | (0.0429) | (0.226) | (0.0444) | (0.306) |
| Unemp_rate | -0.0684 | -0.241 | -0.208*** | -0.454 | -0.0866 | -0.313 |
| | (0.0911) | (0.399) | (0.0760) | (0.398) | (0.0799) | (0.523) |
| Marriage | -3.81e-05 | 0.00778** | -0.000674 | -0.00106 | 0.000152 | -0.000370 |
| | (0.00120) | (0.00374) | (0.000739) | (0.00259) | (0.000558) | (0.00304) |
| Ln (Age) | -0.0690** | 0.0873 | -0.0736*** | 0.161* | -0.0191 | -0.118 |
| | (0.0283) | (0.0901) | (0.0259) | (0.0878) | (0.0295) | (0.157) |
| Gender | 0.0248*** | -0.0164 | 0.0117*** | -0.00529 | 0.00119 | -0.0261 |
| | (0.00313) | (0.0105) | (0.00265) | (0.0103) | (0.00270) | (0.0165) |

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| Edu | 0.000588 | -0.0134*** | 0.00169** | -0.0146*** | -0.000124 | -0.00917* |
| | (0.000962) | (0.00340) | (0.000767) | (0.00320) | (0.000844) | (0.00519) |
| Child | 0.00298* | 0.0148** | 1.08e-05 | -0.000961 | -0.000285 | 0.0132** |
| | (0.00168) | (0.00593) | (0.00112) | (0.00503) | (0.000864) | (0.00584) |
| City FE | Yes | Yes | Yes | Yes | Yes | Yes |
| Year FE | Yes | Yes | Yes | Yes | Yes | Yes |
| Covariates | Yes | Yes | Yes | Yes | Yes | Yes |
| Observations | 13,436 | 13,590 | 13,994 | 14,101 | 6,655 | 6,677 |
| R-squared | 0.012 | 0.114 | 0.013 | 0.108 | 0.004 | 0.110 |

*Note:* (1) Robust standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1. (2) Standard errors are interactively clustered at the level of city and year. (3) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level.

Table 8 Robustness test: keep samples whichimplemented the policy in 2011-2018

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| Variables | work1 | work1 | work2 | work2 | work time | work time | sick | sick |
| Integration | -0.00499 | -0.00628 | 0.00871** | 0.00791** | -0.988 | -0.686 | -0.0988*** | -0.0399* |
| | (0.0219) | (0.0230) | (0.00342) | (0.00352) | (1.191) | (1.217) | (0.0248) | (0.0217) |
| Ln (Expenditure) | | -0.0398*** | | 0.00398** | | 0.613 | | -0.0223* |
| | | (0.0129) | | (0.00184) | | (0.665) | | (0.0120) |
| Ln (GDP) | | -0.0292 | | -0.00218 | | -1.124** | | 0.000237 |
| | | (0.0189) | | (0.00168) | | (0.474) | | (0.0154) |
| GDP_growth | | -0.0551 | | -0.0291 | | 0.784 | | 0.232 |
| | | (0.220) | | (0.0319) | | (13.08) | | (0.203) |
| Unemp_rate | | 0.433 | | -0.0879 | | -5.532 | | -0.154 |
| | | (0.400) | | (0.0567) | | (23.26) | | (0.365) |
| Marriage | | -0.0160*** | | -3.71e-05 | | -0.293** | | 0.00101 |
| | | (0.00196) | | (0.000544) | | (0.121) | | (0.00201) |
| Ln (Age) | | -0.732*** | | -0.0634*** | | -33.82*** | | 0.0295 |
| | | (0.0408) | | (0.00795) | | (1.949) | | (0.0280) |
| Gender | | 0.148*** | | 0.0161*** | | 6.091*** | | -0.00968 |
| | | (0.00856) | | (0.00193) | | (0.466) | | (0.00728) |
| Edu | | -0.00648*** | | 0.000566 | | -0.256* | | -0.0153*** |
| | | (0.00227) | | (0.000575) | | (0.138) | | (0.00241) |
| Child | | -0.0110** | | 0.000379 | | -0.0893 | | 0.00971** |
| | | (0.00440) | | (0.000804) | | (0.210) | | (0.00390) |
| City FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Year FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

| Covariates | No | Yes | No | Yes | No | Yes | No | Yes |
|---|---|---|---|---|---|---|---|---|
| Observations | 41,027 | 38,795 | 29,536 | 28,203 | 28,611 | 27,349 | 29,757 | 28,421 |
| R-squared | 0.024 | 0.133 | 0.003 | 0.012 | 0.030 | 0.076 | 0.003 | 0.110 |

*Note:* (1) Robust standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1. (2) Standard errors are interactively clustered at the level of city and year. (3) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level.
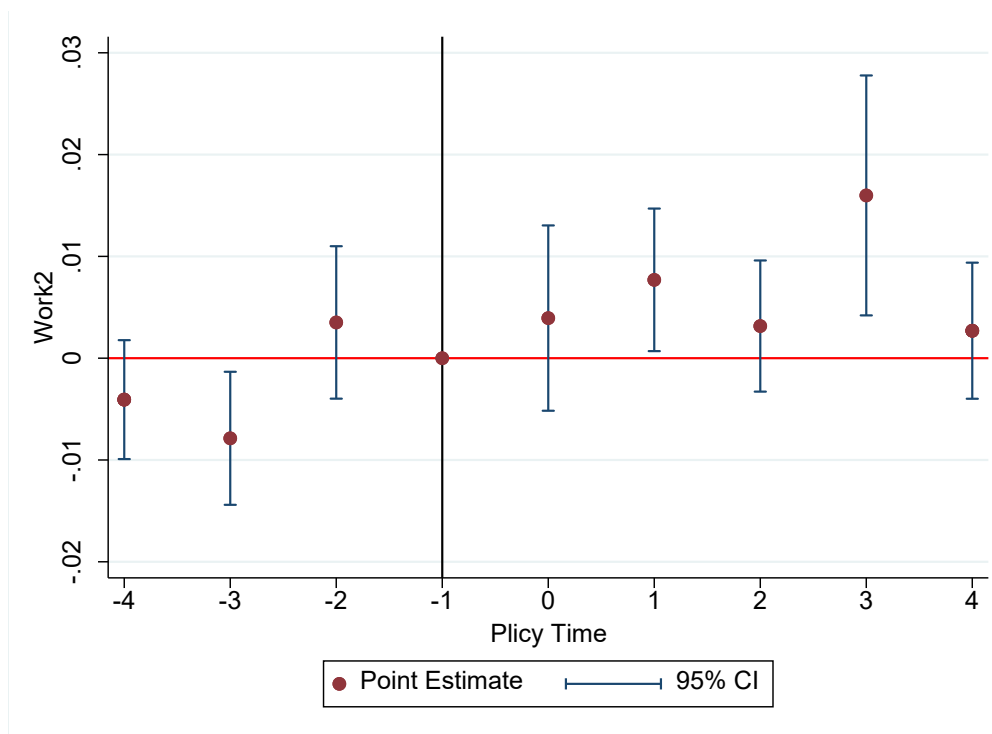


Figure 1 The dynamic impact of integration policy on the probability of having more than one job

*Note:* (1)This figure plots the estimated impact of the integration policy on the probability of having more than one job in a dynamic pattern based on Eq. (3). (2) Standard errors are interactively clustered at the level of city and year. (3) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level. (3) 95% confidence intervals displayed.
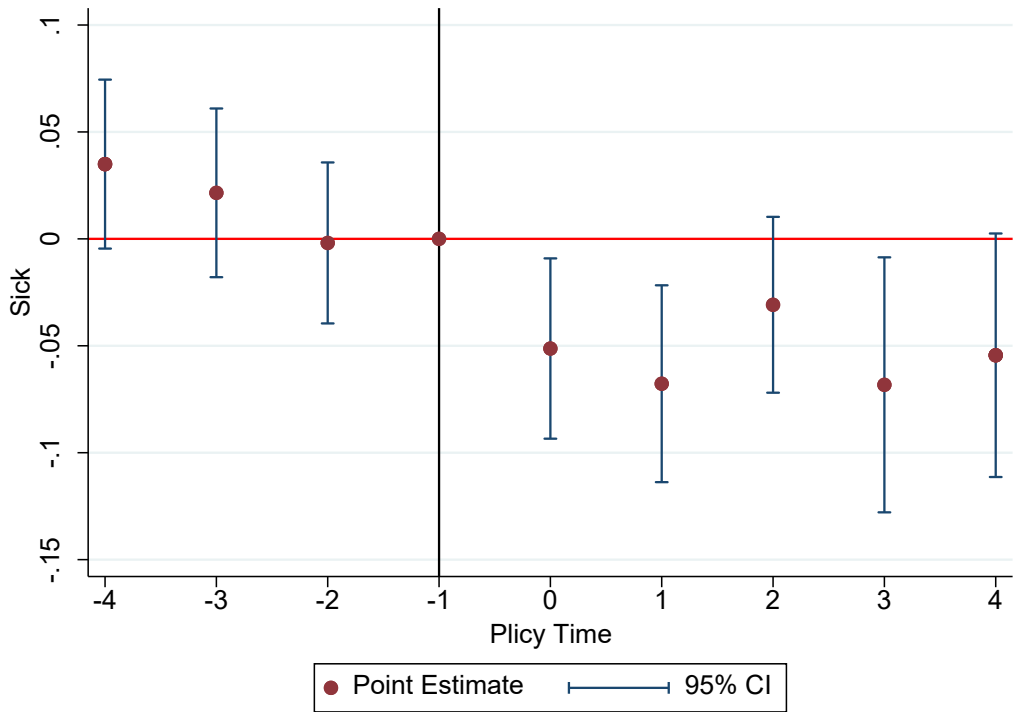
Figure 2The dynamic impact of integration policy on the likelihood of missing work due to illness

*Note:* (1)This figure plots the estimated impact of the integration policy on the likelihood of missing work due to illness in a dynamic pattern based on Eq. (3). (2) Standard errors are interactively clustered at the level of city and year. (3) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level. (3) 95% confidence intervals displayed.
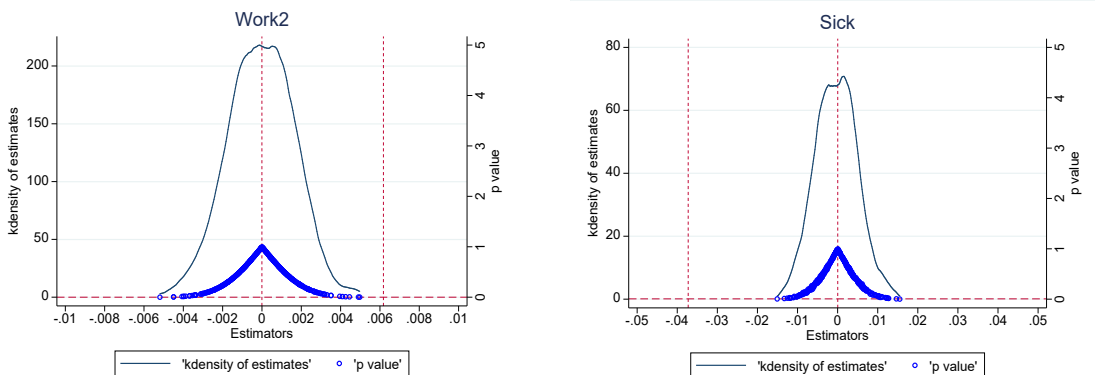


Figure 3 Placebo test: Construct pseudo treatment groups and pseudo control groups

*Note:* (1)This figure plots the kernel density and p values of the estimated coefficients when we construct pseudo treatment groups and pseudo control groups. (2) We repeat the

stochastic process 500 times. (3) Standard errors are interactively clustered at the level of city and year. (4) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level.
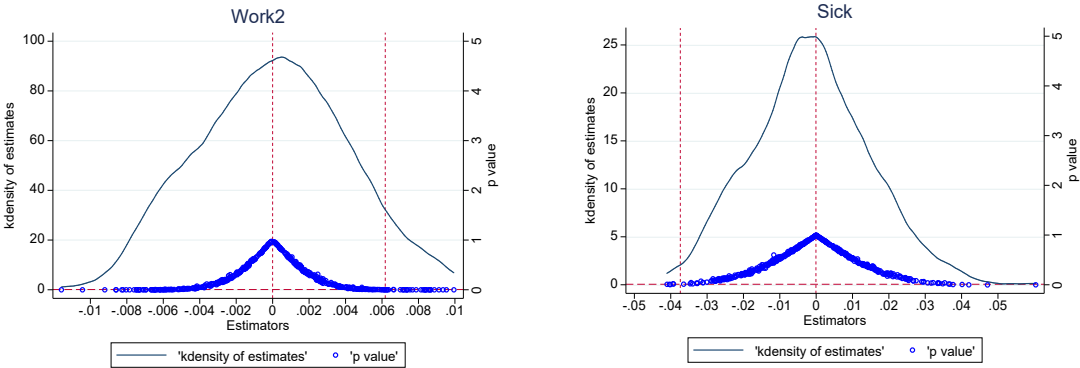


Figure 4 Placebo test: Construct pseudo policy times

*Note:* (1)This figure plots the kernel density and p values of the estimated coefficients when we construct pseudo policy times. (2) We repeat the stochastic process 500 times. (3) Standard errors are interactively clustered at the level of city and year. (4) Covariates include marital status, age, gender, education level, and number of children at individual-level; government expenditure, GDP per capita, GDP growth rate, and unemployment rate at city-level.

# The Impact of Medical Insurance on Family Support——Evidence from CHARLS in China

Yuan Cao, Hua Chen, and Xiaoxu Yang [1]

**Abstract:**

In this study, we investigate the effect of Urban and Rural Resident Basic Medical Insurance (URRBMI) on family supportin China.55,062 individual samples were obtained from four waves of the China Health and Retirement Longitudinal Study (CHARLS): 2011, 2013, 2015 and 2018. Linear-regression model and propensity score matching are used to determine the relationship between medical insurance and family support. Then, mediation model is introduced to identify the mediation mechanisms. Also, moderation model is used to estimate the moderation effect of parental education and health. We find that medical insurance has significantly increased family support between the insured parents and their children. Moreover, this positive effect is heterogeneous since only families living in rural areas were affected, and the direction of family support changed with the aging of the parents. The welfare of medical insurance on financial status have also been proven in this paper. The results indicate that medical insurance reduces the out-of-pocket ratio of medical expenses and increases health investment, which can perform as as two mediation mechanisms to affect family support. Besides, the education and health status of the insured parents play a role in moderating the effect of medical insurance.Our study demonstrates the importance of medical insurance to release the burden of medical expenses, improve health investments, and promote family support.

**Key words:** Medical insurance, Family support, Out-of-pocket ratio of medical expenses, Health expense

## Introduction

The public medical insurance system for urban workers was established in 1998, and since then, China's medical insurance coverage has been expanded to include both urban and rural residents. By the end of May 2022, the number

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of residents enrolled in Urban and Rural Resident Basic Medical Insurance (URRBMI) had reached 9.8 billion, and the national coverage rate had increased rapidly from 64.6% in 2009 to more than 95%[1]. A growing body of studies has shown that, as an important policy to transfer health risks and medical expenses, public medical insurance provides important health benefits and financial benefits to residents. Specifically, evidence from Critical Illness Insurance (CII) in rural China indicates that medical insurance has a positive effect on eliminating health inequalities and promoting the daily consumption of rural residents [1]. Evidence from New Rural Cooperative Medical Scheme (NCMS) shows that the enrollment in NCMS is associated with a lower incidence of catastrophic health expenditure [2] and better performance in activities of daily living and cognitive function [3]. Similarly, a study of expansions in public health insurance in the United States also indicates that an exposure to public health insurance in one's childhood has positive effects on both economic and health outcomes in adulthood, such as reducing out-of-pocket medical spending, increasing financial stability, and decreasing the probability and mortality of chronic disease [4]. In addition, a comparative analysis of different medical insurances (i.e., NCMS, Urban Employee Medical Insurance (UEMI), and the Urban Resident Medical Insurance (URMI) of China) finds that different reimbursement benefits may induce inequity in health service utilization among middle-aged and older adults [5].

While most of the existing research on medical insurance has focused on its direct effect on health outcomes, medical utilization, and financial protection, the essential benefit of medical insurance to regulate family support and intergenerational relations has received little attention. Admittedly, the effect of medical insurance on private financial decision, such as family support within a family, is not as noticeable as its direct effect on improving health status and reducing medical expenses.However, it still deserves closer attention, considering that Chinese attach great importance to mutual support and material exchanges between their families. It is highly possible that the expansion of medical insurance in China can work as a public channel, which firstly influence the financial status of the insured and theninfluence theirsupport between the children.As a result, the medical insurance can affect the net welfare of the family members.

Based on the preliminary assumption that the benefits from the public channel will affect the support from the private channel, this study empirically examines the impact of medical insurance on family support, based on panel data from the China Health and Retirement Longitudinal Study (CHARLS).China is a good case for this study because it has experienced rapid expansion in its medical insurance for urban and rural residents. What's more, the Chinese attach great importance to family support between parents and children. Frequent and intensive family support between parents and children is considered an essential symbol of family harmony and family happiness.

This study contributes to the literature on medical insurance and family welfare in three ways. First, most previous studies on the effects of insurance

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[1] Data Sources: Statistics from the National Healthcare Security Administration. [http://www.nhsa.gov.cn]

programs on family support have focused on public pension programs, while the effects of medical insurance are less discussed. As a part of social security systems, URRBMI may also crowd out or crowd in family support, resulting in a change in the welfare of the families. Therefore, exploring the effect of medical insurance on family support is essential for evaluating the net welfare of medical insurance. Second, we consider the heterogeneous effects of medical insurance, which has received scant attention in previous studies about public support and private support. Third, we introduce two mechanisms by which medical insurance affects family support.Thus, we are able to provide more comprehensive explanations for the inconclusive answers regarding whether and how medical insurance affects family support.

The remainder of this study is organized as follows. In Section 2, we introduce the background of the basic medical insurance of China. In Section 3, we review the relevant literature. Section 5 describes the dataset and descriptive statistics. Section 5 presents the empirical framework. Section 6 shows the empirical results. Section 7 provide a discussion and Section 8 gives a conclusion of this paper.

## Background

In 1998, China established the UEMI as a supporting program for the reform of state-owned enterprises. The UEMIwas the first formal medical insurance program in the country. Since then, a series of programs have been launched to provide medical services to both rural and urban residents. In 2003 and 2007, China established the NCMS and the URMI, respectively. The former aims to cover the vast majority of rural residents and the latter is designed to cover unemployed urban residents. Thus far, China has essentially achieved the goal of providing universal coverage of basic medical insurance for its citizens.

Since 2016, many regions have merged NCMS with URMI and implemented a unified URRBMI. Therefore,the coverage, financing mechanism, treatment, catalog, and management of NCMS and URMI have achieved equality between urban and rural residents.URRBMI employs a combination of individual contributions and financial subsidies financing mechanisms, residents can enjoy the benefits from URRBMI after paying the premiums for that year. What's more, the financing standard and government subsidies of URRBMI have been continuously improved since its establishment. According to the requirement of the National Medical Security Bureau, the individual payment standard of URRBMI in 2022 is 350 yuan per capita, and the financial subsidy standard is no less than 610 yuan per capita. After enrollment, residents can enjoy the benefits from URRBMI for inpatient and outpatient medical expenses. For the inpatient medical expenses incurred in the first, second and third level hospitals, the payment line, i.e., the deductible of URRBMI is 200 yuan, 400 yuan, and 600 yuan respectively.The payment proportions for the expenses between the payment line and maximum paymentline are 65%, 60%, and 50% respectively. As for the outpatientexpenses,the payment line is 600 yuan. The payment proportion for the expenses between the payment lineand the maximum payment line is 50%. The maximum annual payment is 3000 yuan per capita.

So far, a large number of literatures have studied the remarkable policy effects of URRBMI, including reducing medical burden [6] and the intensity of catastrophic health expenditure, ameliorating the impoverishment of the poorest residents [7], improving the benefit for outpatient care [8], and improving the health status of preschool children [9]. Several bodies of literature also estimate the impact of URRBMI on individual and family decisions, such as job selection [10] and stock market participation [11].

## Literature review

Many scholars have studied the impact of public policies on family support from both theoretical and empirical views, but most of them focused on public pension programs and ignored the potential effect of medical insurance. In the theoretical field, scholars generally believe that if a public program achieves its expected effect and improves the economic and health status of the insured, then both the insured parents and their children will change their support to each other. Whether a public program will crowd out or crowd in family support depends on the people's motivation. For example, Becker and Barro point out that under altruistic motivation, the improvement of parents' economic or health status will crowd out family support from their children [12,13]. However, Cox believes that due to exchange motivation, children expect their parents to offer more help or inheritance to them after insurance enrollment.Thus, they increase the support to their insured parents in order to obtain these benefits in the future [14]. As a result, the improvement of parents' economic or health status via insurance will crowd in family support from children. As for family support from parents to children, studies have shown that after economic or health status improves due to insurance coverage, parents tend to increase their support to children because of the behavioral habit of taking care of their children or in exchange for companionship and spiritual support [15,16].

In the empirical studies, Deindl and Brandt study the social insurance systems of 14 European countries and find that social insurance will crowd out family support received by parents and the intensity of the crowd-out effect increases with the amount of social insurance benefits [17]. Lin et al. find that formal insurance has a crowd-out effect on informal private support through a theoretical model, and further verify this theory through experimental economics [18]. More specifically, scholars have proved that the Long-term Care Insurance (LTCI) in the United States, the public old-age support system in South Korea, and the New Rural Resident Pension policy in China have negative effects on family support from children to parents [19–21]. However, little is known about the influence of medical insurance programs on family support, especially for those support from parents to children. Deng et al. uses qualitative methods and finds that urban families, who tend to have more access to public welfare due to China's dualistic welfare system, are less eager to invest in intergenerational support or expect less reciprocity. Thus, more empirical evidence from the medical insurance program needs to be provided [22].

Although medical insurance is less discussed as a public channel which

may affect the private channel within the family, many studies have sufficiently confirmedits effect on reducing the medical expense, improving the health status and financial expectations, and stimulating the consumption demand of the insured [3,23,24]. In addition, studies that concentrate on the determinant of individual financial behavior have shown that a household's financial literacy [25], financial education [26], financial status [27], number of children and the gender pattern of children [28] can affect the family supportsuch as financial transfers. Therefore, the change in out-of-pocket medical expense and consumption demand may act as mediators to regulate the effect of medical insurance on family support. For example, when faced with less pressure from medical bills, insured parents may increase financial support and labor support to their children. Furthermore, children may increase the support to their insured parents to meet their increased demand for consumption.

Based on the above literature, we can conclude that the existing studies have conducted in-depth research on the effect of benefits from public channel on private support. However, these studies mainly focus on public pension programs and have yielded inconclusive results. Moreover, existing studies generally use cross-sectional data, and thus it is difficult to provide a convincing examination of the causal relationship between insurance and family support. In addition, the heterogeneous effect and mediation mechanism are seldomly addressed in the existing literature. To address the aforementioned problems, this study uses the CHARLS panel database to analyze the impact of URRBMI on family support and investigate the heterogeneity and mediation mechanism of its effect comprehensively.

## Data and descriptive statistics

*Data and variables*

Our data comes from the CHARLS database in 2011, 2013, 2015, and 2018. CHARLS uses multistage stratified probability proportional to size sampling to conduct a nationwide survey of respondents aged 45 years or older. The survey follows detailed protocols for sampling, field surveying, and data quality verification. The baseline CHARLS was conducted in 2011. The second-, third-, and fourth-wave national surveys conducted in 2013, 2015, and 2018 aimed to revisit the same respondents sampled in the first wave. The information in CHARLS includes individual demographic backgrounds, family contacts and transfers, health care, and insurance. Thus, this dataset has been widely used in international health economics research, aging research, and family welfare analysis [29–31].

By the time the national follow-up was completed in 2018, the survey had covered 150 county-level units and 450 village-level units in 28 provinces, with 19,000 yearly respondents in approximately 12,400 families. We generate a series of variables required in our study based on the original database and questionnaires, such as URRBMI, family support from parents to children, and family support from children to parents. To identify the impact of URRBMI on family support, we exclude respondents who were enrolled in other social medical insurance, such as UEMI and Supplementary medical insurance.The participants from commercial medical insurance are also excluded from the full

sample. Since NCMS and URMI in some areas were merged into URRBMI after 2016, both those enrolled in NCMS and URMI are regarded as being enrolled in URRBMI in this study. Specifically, those enrolled in NCMS, URMI, and URRBMI constitute the treatment group. After excluding respondents with missing information, we finally obtain unbalanced cross-sectional data on 55,062 observations. We define the treatment group as "the parents enrolled in URRBMI" and the control group as "the parents without any medical insurance." There are 52,734 samples in the treatment group and 2,328 samples in the control group.

The independent variable in our study is URRBMI, the value of insured is 1, and the value of non-insured is 0. The dependent variables are $Support_{p-c}$ and $Support_{c-p}$, which capture family support from parents to children and family support from children to parents, respectively. We aggregate themoney support in total (includes providing living expenses, paying for water, electricity or telephone bill, paying for mortgage/ rent or other forms of regular expensesin the past year) and the total value of in-kind payment support (such as buying food, clothes or other itemsin the past year) as the total family support a parent has received from his/her children. The same strategy is also used to measure the family support from parents to children.

We are also interested in two factors that may mediate the effect of URRBMI on family support: the out-of-pocket ratio of medical expenses and health expense. The out-of-pocket ratio of medical expenses is equal to out-of-pocket inpatient medical expenses in the past year divided by the total inpatient medical expenses in the last year. Health expense is equal to the expense of health investment and health care in the last year. The control variables in our analysis include both individual-level and household-level characteristics. The individual-level variables include age, gender, urban residence, marriage status, education, self-rated health status, and public pensions. The household-level variables include the number of children, living arrangements, and family financial assets.

*Descriptive statistics*

Table 1 shows the definition and descriptive statistics of the variables. We compare the differences between the treatment group and the control group with the t-test and find that the two groups show significant differences in family support. The average of family support from URRBMI-insured parents to children is 4,196 yuan per mouth, which is significantly higher than that of uninsured parents (2,925 yuan per month). This significant difference in family support from parents to children initially indicates that URRBMI may have a positive effect on parental support to children. Similarly, the average of family support from children to URRBMI-insured parents (3,882 yuan per month) is significantly higher than that of those uninsured by URRBMI (2,421 yuan per month), which may be due to exchange motivation from children. In addition, compared with uninsured respondents, insured respondents are younger, better educated, more likely to be married, and have more family financial assets. On the other hand, uninsured respondents are healthier, have lower public pension enrollment rates, and are more likely to have more children and live with children. The characteristics of the uninsured group suggest that they may be

more dependent on family and children for support.

Table 1 Descriptive Statistics

| Variables | Definition | Treatment group | Control group | Differences |
|---|---|---|---|---|
| **Independent variable** | | | | |
| URRBMI | 1 if the respondent participates in URRBMI; 0 if not | | | |
| **Dependent variables** | | | | |
| Support $_{p\text{-}c}$ | Total family support from the respondent to his/her children | 4196.393 | 2925.427 | 1270.996*** |
| Support $_{c\text{-}p}$ | Total family support from children to the respondent | 3882.124 | 2420.817 | 1461.307*** |
| **Control variables** | | | | |
| Age | Age of the respondent | 60.664 | 61.509 | -0.845*** |
| Male | 1 if the respondent is male; 0 if female | 0.48 | 0.428 | 0.052*** |
| Urban residence | 1 if the respondent is urban resident; 0 if rural resident | 0.203 | 0.312 | -0.108*** |
| Marriage status | 1 if the respondent is married or cohabitating; 0 if separated, divorced or never married | 0.883 | 0.806 | 0.077*** |
| Education: | | | | |
| No formal education | 1 if the respondent has no formal education; 0 if not | 0.245 | 0.302 | -0.057*** |
| Can read and write | 1 if the respondent's highest level of education is "can read and write"; 0 if not | 0.197 | 0.222 | -0.024*** |
| Primary school | 1 if the respondent's highest level of education is primary school; 0 if not | 0.226 | 0.200 | 0.027*** |
| Secondary school or above | 1 if the respondent's highest level of education is secondary school or above; 0 if not | 0.332 | 0.277 | 0.055*** |
| Self-rated health status | 1 if very poor; 2 if poor; 3 if fair; 4 if good; 5 if very good | 3.042 | 3.083 | -0.042** |
| Public pensions | 1 if the respondent participates in a public pension program; 0 if not | 0.602 | 0.294 | 0.308*** |

| | | | | |
|---|---|---|---|---|
| Number of children | Total number of living and dead children, including biological children, stepchildren, and adopted children | 3.029 | 3.162 | -0.132*** |
| Living arrangement | 1 if the respondent living with his/her children; 0 if not | 0.279 | 0.349 | -0.069*** |
| Family financial assets | Total net family financial assets | 4708.932 | 3091.608 | 1617.324*** |

Note: *, **, and *** indicate that the coefficients significantly differ from 0 at the 10%, 5%, and 1% levels, respectively.

## Empirical framework

*Basic empirical model*

In empirical analysis, the dependent variables $Support_{p-c}$ and $Support_{c-p}$ are individually scaled by their logarithmic forms. Specifically, we estimate the following linear regression model:

$$Ln(Support_{it}) = \beta_0 + \beta_1 URRBMI_{it} + \beta_X X_{it} + \beta_2 Year_t + \beta_3 Province_i + u_{it}$$
(1)

where $URRBMI$ is a binary variable, indicating whether the respondent is enrolled in URRBMI. $\beta_1$ is the estimator that we are most interested in since it captures the effect of URRBMI on family support. $X_{it}$ is a vector of individual-level and family-level control variables. $Year_t$ is a series of year dummy variables with the coefficient $\beta_2$ representing the year fixed effects. $\beta_3$ represents the province fixed effects and $u_{it}$ is the error term.

The linear regression model in Equation (1) examines the effect of URRBMI by directly comparing the family support between the treatment and the control groups. However, it should be noted that enrolling in URRBMI is voluntary. Residents with higher income or greater risk awareness tend to have higher enrollment rates in insurance programs, and these wealthier respondents may also offer more support to their children. The unobserved heterogeneity between insured and uninsured individuals may lead to selection bias. The key point to solve this problem is to identify whether the family support for residents with URRBMI are higher than the potential results if they werenot enrolled in URRBMI. In other words, it is necessary to determine whether there is a significant average treatment effect on the treated (ATT). In the robustness test, we use the propensity score matching (PSM) method to eliminate the influence of selection bias and estimate the ATT. Besides, another problem that may interfere with our estimation results is reverse causation. Since the data used in this study come from a microsurvey, the timing of URRBMI enrollment cannot be precisely identified. The enrollment time may not necessarily precede

the occurrence of the family support, which may result in the endogenous problem of reverse causation. Therefore, we construct a new independent variable $URRBMI_{coverage}$ which captures the percentage of time that an individual was covered by URRBMI from 2011 to 2018. Using $URRBMI_{coverage}$ as the independent variable and family support in 2018 as the dependent variable, we can rule out potential reverse causality issues.

*Propensity score matching (PSM)*

Since enrollment in URRBMI is not randomly assigned, the decision to enroll in URRBMI is likely to be affected by individuals' personal characteristics. For example, people with higher education and stronger risk awareness are more inclined to enroll in medical insurance. Therefore, the observed differences in family support between the treatment group and the control group may be caused by two things: one is the effect of enrollment in URRBMI, and the other is the natural differences of the two groups. To eliminate the endogeneity problem caused by sample selection bias, we use the PSM method to check the robustness of the main results presented in the previous section. For individual $i$, family support may have two states, depending on whether or not $i$ is enrolled in URRBMI:

$$Support_{it} \begin{cases} Support_{1it} & if\ URRBMI_{it} = 1 \\ Support_{0it} & if\ URRBMI_{it} = 0 \end{cases} \tag{2}$$

where $Support_{0it}$ represents the family support when individual $i$ is not insured, and $Support_{1it}$ represents the family support when individual $i$ is enrolled in URRBMI. Thus, Equation (2) can be rewritten as:

$$Support_{it} = (1 - URRBMI_{it})Support_{0it} + URRBMI_{it}Support_{1it} = Support_{0it} + (Support_{1it} - Support_{0it})URRBMI_{it}$$

(3)

where $(Support_{1it} - Support_{0it})$ is the treatment effect of the URRBMI on family support. Since $(Support_{1it} - Support_{0it})$ is a random variable, we only focus on the expected value of the URRBMI treatment group, i.e., the ATT:

$$ATT \equiv E(Support_{1it} - Support_{0it}|URRBMI_{it} = 1) \tag{4}$$

For individual $i$ of the treatment group, if individual $j$ can be found in the control so that the distance between $X_i$ and $X_j$ is as small as possible, i.e., $X_i \approx X_j$. Based on the ignorability assumption, the probability of individual $i$ and individual $j$ being enrolled in URRBMI is similar. Thus, $Support_{jt}$ can be taken as an estimator of $Support_{0it}$, i.e., $\widehat{Support_{0it}} = Support_{jt}$. We can use $(Support_{it} - \widehat{Support_{0it}}) = (Support_{it} - Support_{jt})$ as a measurement of the treatment effect on individual $i$.

We use the "propensity score" proposed by Rosenbaum and Rubin to measure the distance between $X_i$ and $X_j$ [32]. The propensity score of individual $i$ is the conditional probability of individual $i$ being enrolled in URRBMI given $X_i$, i.e., $P(X_i) \equiv P(URRBMI_{it} = 1|X = X_i)$. We use the logit regression to estimate the propensity score $P(X_i)$, and use the propensity score as a distance function for matching. The ATT can be estimated from

$$ATT = \frac{1}{N_t}\sum_{i:URRBMI_i=1}(Support_{it} - \widehat{Support_{0it}})\ (5)$$

where $N_t = \sum_i I_i$ represents the number of individuals in the treatment group, and $\sum_{i:URRBMI_i=1}$ represents aggregating the treatment effect of individuals in the treatment group.

*Mediation model*

If the independent variable $URRBMI_{it}$ influences the dependent variable *Support* by affecting some mediator variables, then we can use the following regression equation to describe the causal relationship between the variables. This framework for mediation analysis was first proposed by Baron and Kenny, based on the linear structural equation model [33]. It was further developed by Imai et al. and has been used by many social science practitioners to examine the mediation mechanism [34]. The significant effect of medical insurance on reducing medical expenses and releasing health investment has been well documented in previous research. These two factors are important components of family finance, and may further influence other family decisions, such as family support within the family. Therefore, the mediator mechanisms we focus on are the out-of-pocket ratio of medical expenses and health expense. First, medical insurance may reduce the out-of-pocket ratio of medical expenses, thus promoting family support from insured parents to children. Second, medical insurance may stimulate the health investment of insured parents, thus increasing the family support from children to insured parents.

$$Mediator_{it} =$$
$$\beta_{10} + a \times URRBMI_{it} + \beta_{1X}X_{it} + \beta_{12}Year_t + \beta_{13}Province_i + u_{it1}$$
$$(6)$$
$$Ln(Support_{it}) = \beta_{20} + c \times URRBMI_{it} + \beta_{2X}X_{it} + \beta_{22}Year_t +$$
$$\beta_{23}Province_i + u_{it2}$$
$$(7)$$
$$Ln(Support_{it}) = \beta_{30} + c' \times URRBMI_{it} + b \times Mediator_{it} + \beta_{3X}X_{it} +$$
$$\beta_{32}Year_t + \beta_{33}Province_i + u_{it3}$$
$$(8)$$

Specifically, the mediator variables in this study are *Medical expenses*$_{Out-of-pocket\,ratio}$ and *Health expense*. The coefficient

$a$ in Equation (6) represents the effect of $URRBMI_{it}$ on $Mediator_{it}$. The coefficient $b$ in Equation (8) represents the effect of $Mediator_{it}$ on $Support_{it}$. The coefficient $c$ in Equation (7) represents the total effect of $URRBMI_{it}$ on $Support_{it}$, and the coefficient $c'$ represents the direct effect of $URRBMI_{it}$ on $Support_{it}$ after controlling the effect of $Mediator_{it}$.

Baron and Kenny suggest that mediation effects can be tested under the following conditions [35]: First, the variation in $URRBMI_{it}$ is a significant predictor of the variation in $Mediator_{it}$ in Equation (6); second, the variation $URRBMI_{it}$ is a significant predictor of the variation in $Support_{it}$ in Equation (7); and third, the variation in $Mediator_{it}$ is a significant predictor of the variation in $Support_{it}$ in Equation (8). The tested mediator is a valid mediator when all of these conditions are satisfied in the predicted direction, in which case the effect of $URRBMI_{it}$ on $Support_{it}$ must be smaller in Equation (8) than in Equation (7).

We also utilize Sobel–Goodman mediation tests to verify the robustness of our mediation analysis and estimate the ratios of the total effect that is mediated [36]. The Sobel–Goodman method directly tests the significance of the product of coefficient $a$ and coefficient $b$ in Equations (6) to (8), thus its testing power is superior to the sequential test, i.e., Baron and Kenny's step-by-step method.

*Moderation model*

The causal effect of URRBMI on family support might be affected by personal backgrounds, such as education level and health status. With higher education and better health status, parents may engage in more activities and generate more entertainment expenses, which may have a moderate effect on the supportbetween themselves and their children. In addition, healthier parents may also have less medical and health care consumption, thus moderating the support between themselves and their children. Therefore, we use hierarchical regression to perform moderation analysis in this section. First, we repeat the regression between URRBMI and family support in Equation (1). Other covariate variables, time fixed effect, and province fixed effect are also included, as described in section 4.1. Second, education and self-rated health status are entered into Equation (1) as independent variables. Third, the interaction terms of education and URRBMI, and self-rated health status and URRBMI are entered into Equation (1) to examine the moderation effects.

# Empirical Results

*The linear regression results*

Firstly, we use ordinary least squares (OLS) regression to analyze the effect of URRBMI on family support. Table 2 reports the OLS regression results. The estimators in Column (1) and Column (2) present the effect of URRBMI on family support from parents to children without and with covariates, respectively. The results indicate that URRBMI has a positive and significant effect on family support from parents to children. Specifically, the

total amount of support from parents to children increases by 21.9% for URRBMI enrollees, as shown in Column (1). This positive effect of URRBMI on support from parents to children remains significant after the addition of control variables, as shown in Column (2), with an increase of 18.7% for URRBMI enrollees. Since the average family support from parents to children of all samples is 4,143 yuan, this increase is equal to 775 yuan per year. The estimators in Column (3) and Column (4) present the effect of URRBMI on family support from children to parents without and with covariates, respectively. These results also indicate that URRBMI has a positive and significant effect on family support from children to parents. As shown in Column (4), the total amount of support from children to parents increases by 51.9% for URRBMI enrollees. Since the average family support from children to parents of all samples is 3,820 yuan, this increase is equal to 1,983 yuan per year.

Table 2：The effects of URRBMI on family support: OLS estimates

| | Ln (Support p-c) (1) | Ln (Support p-c) (2) | Ln (Support c-p) (3) | Ln (Support c-p) (4) |
|---|---|---|---|---|
| URRBMI | 0.219*** | 0.187*** | 0.704*** | 0.519*** |
| | (0.071) | (0.070) | (0.079) | (0.074) |
| Age | | -0.068*** | | 0.066*** |
| | | (0.002) | | (0.002) |
| Male | | -0.013 | | -0.231*** |
| | | (0.033) | | (0.032) |
| Urban residence | | 0.820*** | | -0.547*** |
| | | (0.049) | | (0.045) |
| marriage status | | 0.424*** | | 0.350*** |
| | | (0.044) | | (0.045) |
| Can read and write | | 0.197*** | | 0.143*** |
| | | (0.044) | | (0.043) |
| Primary school | | 0.364*** | | 0.148*** |
| | | (0.045) | | (0.044) |
| Secondary school or above | | 0.673*** | | 0.123*** |
| | | (0.048) | | (0.046) |
| Self-rated health status | | 0.127*** | | 0.033** |
| | | (0.016) | | (0.015) |
| Public pensions | | -0.161*** | | 0.289*** |
| | | (0.036) | | (0.035) |
| Number of children | | 0.125*** | | 0.279*** |
| | | (0.011) | | (0.011) |
| Living arrangement | | -0.416*** | | -1.676*** |
| | | (0.035) | | (0.035) |
| Family financial assets | | 0.577*** | | 0.148** |
| | | (0.070) | | (0.059) |
| Constant | 0.780** | 2.996*** | 0.213 | -3.682*** |
| | (0.342) | (0.373) | (0.324) | (0.346) |
| Observations | 55,062 | 55,062 | 55,062 | 55,062 |
| R-squared | 0.102 | 0.151 | 0.095 | 0.216 |
| Province fixed effects | Yes | Yes | Yes | Yes |
| Year fixed effects | Yes | Yes | Yes | Yes |

*Note:* *, **, and *** indicate that the coefficients significantly differ from 0 at the 10%, 5%, and 1% levels. The dependent variables are scaled by it's logarithmic form. Family

financial assets are scaled by 1/100,000.

Additionally, the coefficient of parents' age on family support from children to parents is significantly positive, while the coefficient of parents' age on family support from parents to children is significantly negative. This indicates that the older an individual is, the more support he/she receives from his/her children, and the less support he/she gives to his/her children. The coefficient of urban residence indicates that rural residents tend to receive more support from their children, while urban residents tend to provide more support to their children. The above results preliminarily show that the elderly and rural parents are more dependent on family support, which may lead to the heterogeneity of the effect of URRBMI.

*Heterogeneity analysis*

The results of heterogeneity analysis are shown in Table 3. The results in Columns (1) and (2) in Table 3 show that the positive effect of URRBMI on family support from parents to children and from children to parents for rural residents is significantly greater than that for urban residents. In Panel A, the total amount of family support from parents to children increases significantly by 14.4% for rural residents enrolled in URRBMI. In comparison, the total amount of family support from parents to children for urban residents increases by only 10.4%, and this effect is nonsignificant. Similarly, the results in Columns (1) and (2) of Panel B in Table 3 indicate the total amount of family support from children to parents increases significantly by 67.8% for rural residents enrolled in URRBMI, while this effect is not nonsignificant for urban residents.

Table 3：Heterogeneous effects of URRBMI on family support

| | Urban and rural residence | | Age | |
|---|---|---|---|---|
| | **Rural** | **Urban** | **Middle-aged** | **Elderly** |
| | (1) | (2) | (3) | (4) |
| **Dependent variables:** Ln (Support $_{p-c}$) | | | | |
| URRBMI | 0.144* | 0.104 | 0.522*** | 0.482*** |
| | (0.079) | (0.144) | (0.115) | (0.096) |
| Observations | 43,611 | 11,451 | 26,361 | 28,701 |
| R-squared | 0.147 | 0.145 | 0.203 | 0.166 |
| Covariates | Yes | Yes | Yes | Yes |
| Province fixed effects | Yes | Yes | Yes | Yes |
| Year fixed effects | Yes | Yes | Yes | Yes |
| **Dependent variables:** Ln (Support $_{c-p}$) | | | | |
| URRBMI | 0.678*** | 0.016 | 0.093 | 0.253*** |

| | | (0.086) | (0.144) | (0.109) | (0.087) |
|---|---|---|---|---|---|
| Observations | | 43,611 | 11,451 | 26,361 | 28,701 |
| R-squared | | 0.204 | 0.244 | 0.166 | 0.125 |
| Covariates | | Yes | Yes | Yes | Yes |
| Province fixed effects | | Yes | Yes | Yes | Yes |
| Year fixed effects | | Yes | Yes | Yes | Yes |

*Note:* *, **, and *** indicate that the coefficients significantly differ from 0 at the 10%, 5%, and 1% levels. The dependent variables are scaled by it's logarithmic form.

The results in Columns (3) and (4) of Table 3 show that family support from parents to children increases significantly by 48.2% for elderly URRBMI enrollees, while this amount increases by 52.2% for middle-aged enrollees. What's more, the effect of URRBMI on family support from children to parents is significant among elderly enrollees and nonsignificant among middle-aged enrollees.

*Robustness Tests*

In the first robustness test, we use three matching methods: nearest neighbor matching (k=4), radius matching, and kernel matching. The balance check of the distribution of the covariates between the treatment group and control group is shown in Table 4. The matching strategy is effective since the results indicate that all the covariates in the post-matching subsample pass the balancing test since the absolute standardized bias of the covariates is less than 10%[37]. Table 5 reports the ATT of URRBMI. The PSM estimates based on three different matching methods show that the ATT of URRBMI on family support from parents to children is 0.641, and the ATT of URRBMI on family support from children to parents is 0.959. Our results remain robust while the OLS estimator underestimates the effect of URRBMI. The heterogeneous results show that URRBMI has a positive and significant effect on rural, urban, middle-aged, and elderly enrollees.However, this effect on rural enrollees' support, middle-aged enrollees' support to children, and the elderly's support received from children is more remarkable.

Table 4：Covariates balancing test of PSM: Mean differences before and after matching

| Variable | | Mean | | | % reduction |
|---|---|---|---|---|---|
| | | Treated | Control | % bias | \|bias\| |
| Age | U | 60.664 | 61.509 | -8.4 | 46.0 |
| | M | 60.664 | 61.12 | -4.5 | |
| Male | U | 0.480 | 0.428 | 10.5 | 89.6 |
| | M | 0.480 | 0.475 | 1.1 | |

| | | | | | |
|---|---|---|---|---|---|
| Urban resident | U | 0.203 | 0.312 | -25 | 95.5 |
| | M | 0.203 | 0.208 | -1.1 | |
| Marriage status | U | 0.883 | 0.806 | 21.3 | 82.5 |
| | M | 0.883 | 0.869 | 3.7 | |
| No formal education | U | 0.245 | 0.302 | -12.9 | 95.1 |
| | M | 0.244 | 0.242 | 0.6 | |
| Can read and write | U | 0.197 | 0.221 | -6 | 96.9 |
| | M | 0.197 | 0.197 | 0.2 | |
| Primary school | U | 0.226 | 0.200 | 6.5 | 72 |
| | M | 0.226 | 0.219 | 1.8 | |
| Secondary school or above | U | 0.332 | 0.277 | 12 | 80 |
| | M | 0.332 | 0.343 | -2.4 | |
| Public pensions | U | 0.602 | 0.294 | 65.1 | 98.5 |
| | M | 0.602 | 0.606 | -1 | |
| Self-rated health status | U | 3.042 | 3.083 | -4.2 | 46.8 |
| | M | 3.042 | 3.020 | 2.3 | |
| Number of children | U | 3.029 | 3.162 | -7.7 | 74 |
| | M | 3.029 | 3.063 | -2 | |
| Living arrangement | U | 0.280 | 0.349 | -15 | 80.8 |
| | M | 0.279 | 0.293 | -2.9 | |
| Family financial assets | U | 0.047 | 0.031 | 6.6 | 89.2 |
| | M | 0.047 | 0.045 | 0.7 | |

*Notes:* % bias denotes mean standardized difference in percentage. "U" represent "the unmatched subsample", and "M" represent "the matched subsample". The balancing test results here are from nearest neighbor matching (k=4) of the whole sample. The matching test results of radius matching, kernel matching and other sub-groups are also effective, which are available upon request.

Table 5：The ATT of URRBMI on family support: PSM estimates

| Dependent variables | | Nearest neighbor matching (k=4) | | Radius matching | | Kernel matching | | The average of ATT |
|---|---|---|---|---|---|---|---|---|
| | | ATT | t-statistic | ATT | t-statistic | ATT | t-statistic | |
| Ln (Support p-c) | Full sample | 0.639*** | 6.04 | 0.685*** | 7.21 | 0.600*** | 7.23 | 0.641 |
| | Rural | 0.719*** | 3.61 | 0.793*** | 4.39 | 1.019*** | 6.27 | 0.844 |
| | Urban | 0.664*** | 5.62 | 0.626*** | 5.82 | 0.588*** | 6.26 | 0.626 |
| | Middle-aged | 0.753*** | 4.06 | 0.697*** | 4.19 | 0.610*** | 4.33 | 0.687 |
| | Elderly | 0.496*** | 4.06 | 0.556*** | 5.10 | 0.552*** | 5.63 | 0.535 |

| | | 0.879*** | 7.53 | 0.943*** | 8.99 | 1.056*** | 11.55 | 0.959 |
|---|---|---|---|---|---|---|---|---|
| Ln (Support c-p) | Full sample | 0.879*** | 7.53 | 0.943*** | 8.99 | 1.056*** | 11.55 | 0.959 |
| | Rural | 0.974*** | 7.13 | 0.922*** | 7.40 | 1.018*** | 9.38 | 0.971 |
| | Urban | 0.728*** | 3.60 | 0.786*** | 4.29 | 0.893*** | 5.42 | 0.802 |
| | Middle-aged | 0.681*** | 4.66 | 0.838*** | 6.43 | 0.957*** | 8.18 | 0.825 |
| | Elderly | 1.128*** | 6.16 | 1.149*** | 6.98 | 1.184*** | 8.50 | 1.154 |

*Note:* *, **, and *** indicate that the coefficients significantly differ from 0 at the 10%, 5%, and 1% levels, respectively. The dependent variables are scaled by it's logarithmic form. ATT means the Average Treatment Effect on the Treated.

In the second robustness test, we use $URRBMI_{coverage}$ as the independent variable to rule out reverse causality. The results in Table 6 report the effect of $URRBMI_{coverage}$ on family support and the results remain robust when compared with the results from Table 2 and Table 3.

Table 6：The robustness test: The effects of URRBMI Coverage on family support

| | Ln (Support p-c) | Ln (Support p-c) | Ln (Support c-p) | Ln (Support c-p) |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| URRBMI coverage | 0.993*** | 0.680*** | 0.674*** | 0.535** |
| | (0.219) | (0.200) | (0.233) | (0.226) |
| Observations | 15,349 | 15,349 | 15,349 | 15,349 |
| R-squared | 0.022 | 0.187 | 0.022 | 0.097 |
| Covariates | Yes | Yes | Yes | Yes |
| Province fixed effects | Yes | Yes | Yes | Yes |
| Year fixed effects | Yes | Yes | Yes | Yes |

| | Urban and rural residence | | Age | |
|---|---|---|---|---|
| | **Rural** | **Urban** | **Middle-aged** | **Elderly** |
| **Dependent variables:** Ln (Support p-c) | | | | |
| URRBMI coverage | 0.535** | 0.220 | 0.758*** | 0.119 |
| | (0.257) | (0.477) | (0.276) | (0.361) |
| Observations | 12,181 | 3,168 | 6,542 | 8,807 |
| R-squared | 0.091 | 0.102 | 0.053 | 0.108 |
| Covariates | Yes | Yes | Yes | Yes |
| Province fixed effects | Yes | Yes | Yes | Yes |
| Year fixed effects | Yes | Yes | Yes | Yes |

**Dependent variables:** Ln (Support c-p)

| | URRBMI coverage | 0.865*** | 0.248 | 0.297 | 0.896*** |
|---|---|---|---|---|---|
| | | (0.230) | (0.407) | (0.314) | (0.255) |
| | Observations | 12,181 | 3,168 | 6,542 | 8,807 |
| | R-squared | 0.165 | 0.250 | 0.192 | 0.125 |
| | Covariates | Yes | Yes | Yes | Yes |
| | Province fixed effects | Yes | Yes | Yes | Yes |
| | Year fixed effects | Yes | Yes | Yes | Yes |

*Note:* \*, \*\*, and \*\*\* indicate that the coefficients significantly differ from 0 at the 10%, 5%, and 1% levels. The dependent variables are scaled by it's logarithmic form.

*The mediation analysis*

In the previous section, we find that enrollment in URRBMI has a significant effect on family support between the insured parents and their children. We examine the indirect effects of URRBMI through the out-of-pocket ratio of medical expenses and health expense in this section. First, enrollment in URRBMI could reduce the out-of-pocket ratio of medical expenses and alleviate the burden of medical expenses. As a result, parents who are enrolled in URRBMI can provide more support to their children. Second, URRBMI may raise individuals' health awareness and stimulate health consumption, thus crowding in children's support to parents.

The results in Columns (1) to (3) of Table 7 report the regression results of the out-of-pocket ratio of inpatient medical expenses as a mediator. Column (1) shows the effect of URRBMI on the out-of-pocket ratio of inpatient medical expenses. The estimated coefficient on URRBMI is −0.03, which indicates that the out-of-pocket ratio decreases significantly if the respondent is enrolled in URRBMI. Column (2) shows the total effect of URRBMI on family support from parents to children. As explained above, the total amount of family support increases by 18.7% after enrolling in URRBMI. The last step of the mediation analysis is to test the impact of URRBMI on family support by including the out-of-pocket ratio as a covariate. The results in Column (3) show that after controlling for the impact of the out-of-pocket ratio, the effect of URRBMI on family support is significant.However, it drops from 0.187, the result in Column (2), to 0.177. Therefore, we conclude that URRBMI has a crowd-in effect on family support from parents to children, and part of this crowd-in effect is mediated by a reduction in the out-of-pocket ratio of inpatient medical expenses.

Table 7：The mediation analysis of URRBMI on family support

| | Medical expenses Out-of-pocket ratio | Ln (Support p-c) | Ln (Support p-c) | Ln (Health expenses) | Ln (Support c-p) | Ln (Support c-p) |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| URRBMI | -0.030*** | 0.187*** | 0.177** | 0.306*** | 0.518*** | 0.504*** |
| | (0.003) | (0.070) | (0.070) | (0.053) | (0.074) | (0.074) |
| Medical | | | -0.335*** | | | |

expenses

Out-of-pocket ratio

|  | | | | | | |
|---|---|---|---|---|---|---|
| | | | (0.095) | | | |

| | | | | | | |
|---|---|---|---|---|---|---|
| Ln (Health expenses) | | | | | | 0.048*** |
| | | | | | | (0.006) |
| Observations | 55,062 | 55,062 | 55,062 | 55,062 | 55,062 | 55,062 |
| R-squared | 0.053 | 0.151 | 0.151 | 0.103 | 0.216 | 0.217 |
| Covariates | Yes | Yes | Yes | Yes | Yes | Yes |
| Province fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Year fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |

*Note:* *, **, and *** indicate that the coefficients significantly differ from 0 at the 10%, 5%, and 1% levels. The dependent variable is scaled by it's logarithmic form. Family financial assets are scaled by 1/100,000.

Next, we examine whether URRBMI has a crowd-in effect on family support from children to parents by stimulating health consumption. The results presented in Columns (4) to (6) of Table 7 indicate that health consumption increases by 30.6% after enrolling in URRBMI, which corroborates earlier findings that medical insurance has a positive effect on consumption related to health investment and health care. The results in Column (6) show that after controlling for the impact of health consumption, the effect of URRBMI on family support is significant, and its coefficient decreases from 0.518 to 0.504. Therefore, URRBMI has a crowd-in effect on family support from children to parents by increasing the health consumption of insured parents.

Table 8：The Sobel-Goodman Mediation Tests of URRBMI on family support

| Mediator: **Medical expenses** Out-of-pocket ratio | Coefficient | Std Err | Z | P>Z |
|---|---|---|---|---|
| a coefficient | -0.030 | 0.004 | -8.452 | 0.000 |
| b coefficient | -0.335 | 0.091 | -3.680 | 0.000 |
| Indirect effect | 0.010 | 0.003 | 3.374 | 0.001 |
| Direct effect | 0.177 | 0.076 | 2.319 | 0.020 |
| Total effect | 0.187 | 0.076 | 2.453 | 0.014 |

Proportion of total effect that is mediated:0.054

Ratio of indirect to direct effect: 0.057

Ratio of total to direct effect: 1.057

| Mediator: **Ln (Health expenses)** | Coefficient | Std Err | Z | P>Z |
|---|---|---|---|---|
| a coefficient | 0.306 | 0.053 | 5.725 | 0.007 |
| b coefficient | 0.048 | 0.006 | 8.292 | 0.000 |
| Indirect effect | 0.015 | 0.003 | 4.711 | 0.000 |
| Direct effect | 0.504 | 0.073 | 6.955 | 0.000 |
| Total effect | 0.519 | 0.073 | 7.155 | 0.000 |

Proportion of total effect that is mediated: 0.028

| | | | | | Ratio of indirect to direct effect: | 0.029 |
|---|---|---|---|---|---|---|

Ratio of indirect to direct effect: 0.029

Ratio of total to direct effect: 1.029

The results of the Sobel−Goodman mediation tests are given in Table 8. The results shown in the upper part of Table 8 suggest that the out-of-pocket ratio of inpatient medical expenses is an effective mediator. The negative coefficient *a* indicates that enrollment in URRBMI significantly reduces the out-of-pocket ratio of inpatient medical expenses by 3.0 %, which is consistent with the results in Table 7. The negative coefficient *b* indicates that a 10% decrease in the out-of-pocket ratio significantly increases family support from parents to children by 3.35 %. Therefore,URRBMI promotes family support from parents to children by reducing the out-of-pocket ratio of inpatient medical expenses. This positive mediating role accounts for 5.4% of the total effect. The results of health consumption shown in the second part of Table 8 indicate that URRBMI has a positive effect on health consumption and the increase of health consumption could further increase family support from children to parents. In addition, the indirect effect, i.e.,the mediation effect of health consumption, is statistically significant, with approximately 2.8% of the total effect being mediated.

*The moderation analysis*

The results in Table 9 indicate the coefficient of the interaction term of education and URRBMI is significantly positive, which means the education of parents has a moderating effect on the relationship between URRBMI and family support from parents to children. A high level of parental education is conducive to increasing the support from parents to children in the insured group. The health status of parents also has a moderating effect on the relationship between URRBMI and family support from children to parents. Insured parents with a better health status receive significantly less support from their children than those whose health is worse.

Table 9：The hierarchical multiple regression results for the moderation analysis

| | Ln (Support p-c) (1) | Ln (Support p-c) (2) | Ln (Support p-c) (3) | Ln (Support c-p) (4) | Ln (Support c-p) (5) | Ln (Support c-p) (6) |
|---|---|---|---|---|---|---|
| URRBMI | 0.229*** | 0.187*** | -0.017 | 0.524*** | 0.518*** | 0.972*** |
| | (0.070) | (0.070) | (0.220) | (0.074) | (0.074) | (0.252) |
| Education | | 0.221*** | 0.091 | | 0.035** | 0.042 |
| | | (0.016) | (0.059) | | (0.015) | (0.062) |
| Self-rated health status | | 0.127*** | 0.126* | | 0.032** | 0.169** |
| | | (0.016) | (0.067) | | (0.015) | (0.074) |
| URRBMI*Education | | | 0.136** | | | -0.007 |
| | | | (0.059) | | | (0.062) |
| URRBM*Self-rated health status | | | 0.002 | | | -0.144* |
| | | | (0.069) | | | (0.075) |
| Observations | 55,062 | 55,062 | 55,062 | 55,062 | 55,062 | 55,062 |
| R-squared | 0.146 | 0.150 | 0.151 | 0.216 | 0.216 | 0.216 |
| Covariates | Yes | Yes | Yes | Yes | Yes | Yes |

| | | | | | | |
|---|---|---|---|---|---|---|
| Province fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Year fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |

*Note:* *, **, and *** indicate that the coefficients significantly differ from 0 at the 10%, 5%, and 1% levels. The dependent variable is scaled by it's logarithmic form. Family financial assets are scaled by 1/100,000.

## Discussion

The empirical findings in this paper make several contributions to the understanding of public program, family decisions, and family welfare. Before this study, empirical evidence on the positive effect of medical insurance on intergenerational support within Chinese family was limited. Our study is one of the first attempts to thoroughly examine the impact of medical insurance on family support. Through our evidence, we confirm the positive effect of URRBMI on personal financial status and family welfare. Our results indicate that URRBMI could significantly reduce the out-of-pocket ratio of inpatient medical expenses, thus reducing the burden of medical expenses. We also demonstrate that URRBMI has a positive effect on increasing the health investment and health consumption, which may further improve people's health. Different from the previous literature, which mainly focuses on the preliminary effect of URRBMI on health and personal finance[1,2,7,9,23], we further examine the positive effect of URRBMI on intergenerational support within the family.Since the family support between parents and childrenis an important symbol of family happiness and harmony in China, we believe that the positive effect of medical insurance on family welfare may be underestimated in previous studies.Our analysis also contributes to the literature on the evaluation of the spillover effect of medical insurance policies on non-target populations, such as the study of the indirect effect of medical insurance on increasing labor supply and releasing physical stress for the family members of the insured [38].

In addition, the results from the heterogeneity effect are also striking. The positive effect of URRBMI on family support is more significant among rural residents. This may be related to the fact that rural residents in China place more importance on family solidarity and support. The previous study has indicates thatthe core functions of the family as the major welfare provider in rural China have remained [39]. Therefore, after enrolling in URRBMI, rural parents are more willing to increase financial and material assistance to their families. This finding is also consistent with the study of attitudes toward family obligation among urban and rural adolescents, which has shown that urban male adolescents have a weaker sense of family obligation than do rural male adolescents [40]. Fueled by a stronger sense of family obligation, children in rural area also increase more support to their parents. Besides, the positive effect of URRBMI is also more significant in middle-aged parents' support to their children, and children's support to their elderly parents. And this heterogeneity by age can be explained by the difference in health and economic status of people at different ages. Generally, younger parents tend to have better physical and economic conditions than the elderly. Thus, after the enrollment in URRBMI, middle-aged parents have an advantage over elderly ones to provide more support to their children.

It is important to note some limitations in our study. First, although one of the strengths of our data is that CHARLS data are nationally representative, this data may not objectively reflect some characteristics of respondents. Slight errors may have occurred when a respondent was asked to recall the money and in-kind support between his/her children. In addition, the use of self-rated health status as a representation of physical health status may bring some deviation. Second, we consider that medical insurance has a positive effect on family harmony since it can increase support between family members. However, we did not measure this welfare improvement effect by an individual or family utility function. Thus, there are some difficulties in comparing the effects of different medical insurance policies. Future studies might introduce an expected utility model and measure the utility improvement of this effect.

## Conclusion

This article uses data from CHARLS for the years 2011-2018 to estimate the effect of URRBMI on family support. The results show that URRBMI significantly improves family support between insured parents and their children. The out-of-pocket ratio of inpatient medical expenses and the health expense of insured parents are important mechanisms that mediate the effect of URRBMI on family support. In addition, both the heterogeneity effect and moderation effect are discussed in our study, providing a comprehensive understanding of the effect of URRBMI on different family backgrounds. Overall, URRBMI could significantly promote support between family members, which we believe is conducive to family harmony and intergenerational mutual assistance.

The results from our study have several implications for the literature on medical insurance and other social welfare programs, and for literature focused on family economics and labor economics. Our results suggest that it is important to recognize the roles of both preliminary and spillover results in the evaluation of public policy. The medical program may not only influence the benefits of beneficiaries, but also affect their cognition and decision-making through a direct or mediation mechanism. In addition, the government and policymakers should pay attention to the fact that the benefits of a public program can be moderated by other factors. Therefore, it is important to consider the needs of different groups when developing and evaluating public policies.

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# Does Public Health Insurance Crowd Out Private Health Insurance——New Evidence from a Quasi-Experiment in China

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**Abstract:**

This paper provides new evidence of the effect of public health insurance expansion on private health insurance market by exploiting a quasi-experiment of Urban Resident Basic Medical Insurance (URBMI) in China. We find that the implementation of URBMI has no significant effect on the aggregated private health insurance market. When the entire market is divided into medical and illness insurance market, the former one is significantly crowded out by URBMI but the latter one not. This may help to explain the mixed findings in the literature, which typically lumps all segments of private health insurance markets together.In terms of the mechanisms,the crowding-out effect of public health insurance on private health insurance is through lowering the cost of health services, raising the level of accessibility to health services or reducing people's health risks.

**Key Words:** Health Insurance, Crowding-out Effect, Quasi Experiment, China

## 1. Introduction

Many countries are actively engaged in the development of public health insurance. Low-cost or even free public health insurance can significantly improve people' welfare. However, the development of the public health insurance may cause people to switch from private to the public health insurance, hindering the development of private market, reducing people' insurance coverage, and increasinggovernment's financial burden. As a result, the appropriate level of public health insurance is the focus of theoretical and practical discussions. One of the most importance aspectsof this issue is accurately identifying and estimating the crowding-out effect of public health insuranceon private health insurance.

Cutler and Gruber (1996) argued thirty years ago that the crowding-out effect of public health insurance on private health insurance could be significant. Many subsequent studies found consistent results (Gruber and

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Simon, 2008; Brown and Finkelstein, 2008; Hamersma et al., 2018), but their conclusion was challengedothers (Ham and Shore-Sheppard, 2005; Shore-Sheppard, 2008). Hence, whether and how much public health insurance crowds out private health insurance remains unclear and needs more empirical research, especially in developing countries where public health insurance has been growing in recent years, such as China.

We provide a reasonably precise estimate on the impact of expanding public health insurance on private health insurance based on a quasi-experiment in China. In July 2007, the launch of the Urban Resident Basic Medical Insurance (URBMI) pilot program marked a substantial improvement in China public health insurance system. Since URBMI gradually enrolled all cities between 2007 and 2009, we estimate itseffect on the private health insurance market using a staggered difference-in-differences (DD) method.

We show that the implementation of URBMI has a negative but statistically insignificant effect on the development of the entire private health insurance market. More interestingly, when splitting the entire private health insurance market into two segments, i.e., medical insurance and illness insurance markets, we find that the URBMI has a significant crowding-out effect on the former but an insignificant influence on the latter.

We then consolidate this causal relationship from several aspects. We show that the difference in private health insurance between treatment and control groups is insignificant before the launch of URBMI. This verifies the common trends assumption of DD estimation. To address the possible bias caused by heterogeneous treatment effects in staggered DD specification, we employ the difference-in-differences multiple method proposed by de Chaisemartin and d'Haultfoeuille (2020) and the event study model by Sun and Abraham (2021). We demonstrate that the bias from heterogeneous treatment effects is small and has no bearing on our main findings. Moreover, the placebo test shows that the implementation of URBMI has no significant effect on life insurance or property insurance. Therefore, bias due to other treatments or unobservable factors is unlikely.

We further provide some evidence of possible mechanism of the crowding-out effect from both supply and demand sides. On the supply side, we focus on the mechanisms of medical service price and medical service accessibility. The implementation of URBMI may control the price of medical services, lowering the cost of medical services and decreasing the value of private health insurance. In other words, if the cost of medical services is high, the value of private health insurance will be high and it is hard to be crowded out by URBMI. This mechanism is supported by our regression results. In the same spirit, if access to health services is widespread, the relative advantage of private health insurance on treatment dispatch and quality will be diminished, thus increasing medical accessibility contributes to the value of URBMI while adding little value to private health insurance.

On the demand side, we examine the mechanism of health risks. Lower health risks as a results of public health insurance may decrease people's demand for private health insurance. When there is a higher level of health risk,it is more difficult for the URBMI to cover the health risk and the value of private health insurance will be higher. Our findings indicate that URBMI is

less likely to crowd out the private health insurance market in cities with higher averaged health risks, which lending credence to the mechanism of health risks.

At the same time, there are some competing stories that support the null hypothesis that the URBMI has no significant crowding-out effect (or crowding-in effect) on the private medical insurance market. To make our results more compelling, we consider two of these stories that URBMI may increase people's consumption budget and financial literacy, potentiallyincreasing people's purchase of private medical insurance and causing a crowding-in effect on the private market. Yet, our results suggest that these two stories are unlikely to be true.

Our work contributes to the literature on the relationship between public and private health insurance. There is no unified conclusion about this question. In the case of China, some studies document that public health insurance has significant crowding-out effect private health insurance (Li and Tian, 2020), but others show that the expansion of public health insurance has no significantcrowding-out effect on or even promotes the private sector (Hou and Zhang, 2017). Our paper is distinguished by the following features.

First, previous research has focused onentire private health insurance market that aggregates the medical and illness insurance, which may make it difficult to determine the specific impact of social health insurance on different types of commercial health insurance. In contrast, we divide the entire private health insurance market into two segments, medical insurance and illness insurance, and estimate the effects of URMBI implementation on each separately. Our findings show that URMBI implementation has a significant negative impact on private medical insurance but not private illness insurance. This may help to explain the contradictory findings in the literature, which typically lumps all segments of private health insurance markets together.

Second, we employ a quasi-experiment in China and use the staggered DD method to obtain a convincing causal reference. Some studies also use the DD method based on policy shocks to identify the causal relationship, but they ignore the underlying assumption of homogeneous treatment effect, violation of which may lead to biased estimation (de Chaisemartin and d'Haultfoeuille, 2020; Sun and Abraham, 2021). We employ newly developed methods proposed in recent literature to address these concerns and make the results more convincing.

Third, we conduct extensive empirical testing of potential mechanisms of this crowding-out effect from both the supply and demand sides. Although some studies discuss the possible mechanisms, they do not conduct empirical tests (e.g., Hou and Zhang, 2017). Instead, we present plentiful empirical evidence of these mechanisms, which has meaningful

policy implications for the coordinated development of public and private health insurance.

## 2. Theory and Hypothesis

*2.1. The crowding-out effect*

According to Cutler and Gruber (1996), under the circumstance that everyone has the willingness to buy private health insurance initially, those who are eligible for free or very cheap public health insurance will be more likely to choose their public insurance option if

$$Value_{public} - Cost_{public} > Value_{private} - Cost_{private}$$

where $Value_{public}$ ($Value_{private}$) is the value of public (private) health insurance, and $Cost_{public}$ ($Cost_{private}$) denotes the cost of buying public (private) health insurance.

To facilitate comparisons, we distinguish between private medical insurance and private illness insurance, two components of the entire private health insurance market. The coverage of private medical insurance with a low insured amount (accounting for over 80% market share of total private medical insurance) is nearly the same as that of public medical insurance (Yu and Jia, 2020). Its annual premium fee is around 200 to 300 RMB, compared to an average of 180 RMB for insured residents under URBMI for the same period. Hence, the URBMI implementation is expected to have a significant crowding-out effect on private medical insurance due to the similar coverage and lower cost than URBMI.

(Critical) Illness insurance is designed to cover medical expenses for these specific serious illnesses or deaths, and it provides larger compensation through a refund of premiums paid or lump sum payment. The premium paid on illness insurance is much higher than public health insurance, and the private illness insurance becomes an essential complement to the public health insurance or private medical insurance (Feng, 2014). Even ifa person has had URBMI, private illness insurance can serve as a supplement by providing more extensive risk coverage. Therefore, it is possible that the URBMI will not have a significant impact on private illness insurance.

The realized impact of URBMI on the entire private health insurance market is determined by the proportion of private medical insurance and illness insurance. According to the China Insurance Regulatory Commission (CIRC), premium revenue for private illness insurance in China was RMB 668.1 billion from 2002 to 2013, and premium revenue for private medical insurance was RMB 402.51 billion. Because the proportion of illness insurance is noticeably higher than medical insurance, the impact of URBMI implementation on the entire private health insurance market is likely to be minor and insignificant. Consequently, we propose the following hypotheses.

**Hypothesis IA**: *The implementation of URBMI will have no significant impact on the whole private health insurance.*

**Hypothesis IB**: *The implementation of URBMI will exhibit a significant crowding-out effect on private medical insurance but not on private illness insurance.*

*2.2. The mechanisms of crowding-out effect*

We then analyze three possible mechanism of the crowding-out effect of URBMI on private medical insurance from both supply and demand sides. Generally, URBMI crowds out the private medical insurance market through lowering the medical cost, increasing the accessibility of medical service, and decreasing people's health risk.

Firstly, we focus on the mechanism of medical service price on the supply side. The implementation of URBMI directly reduces the out-of-pocket cost of medical services (Cardon and Hendel, 2001; Card et al., 2008; Anderson et al., 2012; Ying and Chang, 2019). Further, in the long run, the implementation of URBMI may aid in the control of medical price, which influences patients' expenditure on medical goods and services (Duggan and Morton, 2010; Lakdawalla and Yin, 2015; Ellis and Esson, 2021). International experience such as studies on Medicare confirm the effect of controlling costs with the implementation of public insurance.

The control of medical serviceprice by URBMI lower the cost of medical service, which reduces the value of private health insurance. As a result, the URBMI may crowd out the private medical insurance market. In other words, if the price of medical services is high (low), the value of private health insurance will be high (low) and it is hard (easy) to be crowded out by URBMI. Therefore, we propose:

**Hypothesis II**: *The lower (higher) the price of health services, the stronger (weaker) the crowding-out effect of URBMI on private medical insurance.*

Secondly, we turn to the accessibility of medical services in the supply-side mechanism. From 2009 to 2019, China made substantial progress in improving equal access to care and enhancing financial protection, especially for people with low income, though there remains a gap in the quality of healthcare (Yip et al., 2019; Gao and Wang, 2021; Milcent, 2021). Many studies document a strong positive relationship between public health insurance and access to health care (Eggleston et al., 2008; Hoffman and Paradise, 2008). Because private health insurance is advantageous in situations where access to healthcare is limited as it enables insured individuals to receive care more rapidly (Lakdawalla, 2015; Sohn and Jung, 2016), increasing access to medical services will increase the value of public health insurance but decrease the value of private health insurance.In other words, if the access to health services is high (low), the relative advantage of private health insurance on the dispatch and quality of receiving treatment will be weakened (strengthened), thusincreasing medical accessibility contributes to the value of URBMI while adding little value to private health insurance. We therefore propose:

**Hypothesis III**: *The higher (lower) the accessibility of health services, the stronger (weaker) the crowding-out effect of URBMI on private health insurance.*

Lastly, we examine the demand-side mechanism of health risk. Many empirical studies found that public health insurance has significant health benefits to individuals (Levy and Meltzer, 2008; Pan et al., 2016). Lower health risks as a results of public health insurance may decrease people's demand for private health insurance. In other words, if there is a higher (lower) level of health risk, it is more difficult for the URBMI to cover the health risk and the

value of private health insurance will be higher, thus URBMI is less likely to crowd out the private health insurance market. Therefore, we propose the following hypothesis.

*Hypothesis IV*: *The higher (lower) level of health risk, the weaker (stronger) the crowding-out effect of URBMI on private health insurance.*

## 3. Data, Variables, and Summary Statistics

This paper uses annual panel data of Chinese cities from 2002 to 2013 to estimate the impacts of URBMI on private health insurance market.[1] Chinese cities are defined as cities at the prefecture level and above, including 4 municipalities, 15 sub-provincial cities, and 297 prefecture-level cities.

*3.1. Explained variables*

Our explained variables are the premium income of private health insurance, private medical insurance, and private illness insurance.[2] They are expressed as a per capita value, i.e., premium income divided by the population at the city-year level. These data are from the non-public database of the CIRC. In addition to the overall premium income of private health insurance market, it also provides the premium income of two segments of health insurance market, i.e., the medical insurance market and illness insurance market. Hence, it allows us to examine the impacts of URBMI on the entire private health insurance market and its segments, thus providing more meaningful results of the crowding-out effect of public health policy. As shown in Table 1, in the sample period, the average per capita premium income of private health insurance, private medical insurance, and private illness insurance is 35.23 RMB/person, 12.61 RMB/person, and 22.62 RMB/person, respectively. In that time period, China's private health insurance market was quite small in comparison to developed countries. In the placebo tests, we use premium income of life insurance and property insurance as the explained variables. This data is from the public available *China Insurance Yearbook*.

*3.2. Treatment variable*

In July 2007, the State Council issued the guiding opinions on Piloting Urban Resident Basic Medical Insurance (URBMI), which was significant in closing the final gap in China's basic medical insurance system and achieving full coverage.Our treatment variable (*Treat*) is a dummy that equals one if the URBMI is implemented in a given city and year. The data is from

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[1] The sample begins in 2002 for two reasons: (1) The pilot of urban residents' basic medical insurance started in 2007, which requires data from several years earlier in the study; (2) China formally joined the WTO on December 11, 2001, and the framework of the insurance market system was basically established. The sample ends in 2013 because the critical illness insurance program was included in the basic medical insurance after 2013, which may confound the treatment effect of implementation of URBMI.

[2] The premium includes: private health insurance (basically consisting of private medical insurance and private illness insurance), private medical insurance, and private illness insurance; life insurance, divided into term life insurance (underwriting period of one year or less) and long-term life insurance (underwriting period of more than one year); accident insurance; property insurance, divided into motor vehicle insurance mainly. The property insurance sector is divided into motor vehicle insurance (mainly serving the lives of residents) and other property insurance (mainly serving corporate production activities in mainland China).

relevant policy documents issued by the Chinese government. The list of first 79 pilot cities comes from *the Approval of the list of pilot cities for urban residents' basic medical insurance in 2007 (Letter from the Ministry of Labor and Social Affairs [2007] No. 174)*.

Other 229 cities were added to *the Approval and Reply on the List of Expanded Pilot Cities for Basic Medical Insurance for Urban Residents in 2008 (Letter from the Ministry of Labor and Social Security [2008] No. 24)* issued on February 15, 2008. According to *the Notice on the Comprehensive Launch of Basic Medical Insurance for Urban Residents (Ministry of Human Resources and Social Security [2009] No. 35)* issued on April 8, 2009, more cities enact the policy between 2008 February 15 and 2009 April 8 and then the URBMI program becomes nationwide. See Figure 1 for the distribution of cities with different policy year.
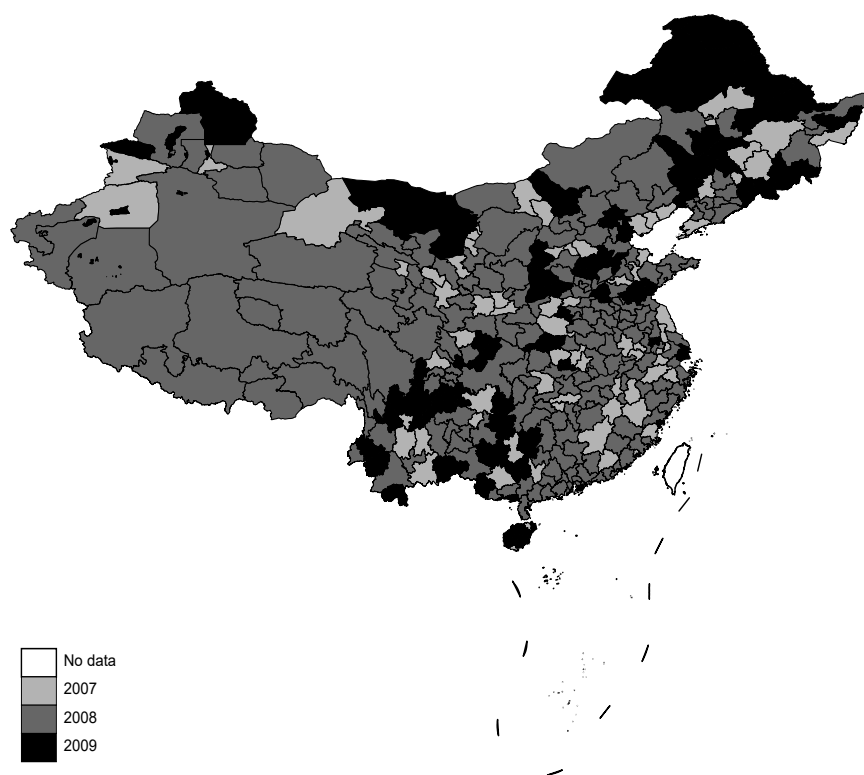


1. **Figure 1**. Distribution of cities with different policy year

*Note*: This figure shows the policy year when each city is included in the

URBMI.

*3.3. Control variables*

We also include other insurance market variables and economics variables to control the possible confounding factors. The insurance market variables

include the number of branches of insurance companies operating in each line of business in each city each year, which controls insurance supply factors. The source of the insurance market variables is the CIRC non-public database.

Other city-year economic and social variables include demographic structure, economic development level, education level, fiscal expenditure structure, and medical service conditions. Firstly, the demographic structure variables include female ratio. Secondly, the variable reflecting the level of economic development is GDP per capita and urbanization rate. Thirdly, the education variable is the number of students enrolled in higher education schools as a proportion of the resident population. Fourthly, variables measuring the structure of fiscal expenditures include the share of fiscal budget expenditures on health care (supporting the supply side of health care services) and the share of fiscal budget expenditures on social security and employment (supporting the demand side of health care services). Lastly, the medical variable is the number of hospital beds per 10,000 people. Data for these variables were obtained from the *China Statistical Yearbook for Regional Economy* and *China Statistical Yearbook of City*. Table 1 gives the details of descriptive statistics of variables.

**Table 1.** Summary statistics

| Variables | Definition | Mean | S.D. |
|---|---|---|---|
| *MedInc* | Premium of private medical insurance, yuan/person | 12.61 | 26.14 |
| *IllInc* | Premium of private illness insurance, yuan/person | 22.62 | 31.26 |
| *HealthInc* | Premium of private health insurance, yuan/person | 35.23 | 54.24 |
| *Treat* | =1 if the city has URBMI in a given year | 0.42 | 0.49 |
| *EntpMed* | Number of private medical insurer branches | 10.42 | 8.11 |
| *EntpIll* | Number of private illness insurer branches | 8.80 | 6.41 |
| *Gender* | Gender ratio (female=100) | 105.77 | 3.88 |
| *GDP* | log of GDP per capita | 9.74 | 0.84 |
| *Urban* | Urbanization rate | 0.38 | 0.17 |
| *Edu* | Proportion of students enrolled in higher education | 0.01 | 0.02 |
| *MedShare* | Health care share on government expenditure | 0.03 | 0.23 |
| *SocShare* | Social security share on government expenditure | 0.15 | 0.15 |
| *Bed* | Number of beds in hospitals per 10,000 people | 33.33 | 12.97 |

*Notes:* The money value is constant at 2002 price.

## 4. Impacts of URBMI on private health insurance consumption
*4.1 Estimation using staggered difference-in-differences*

Because the various treatment timing in different cities, we use the staggered difference-in-differences (DD) specification to estimate the impacts

of URBMI implementation on private health insurance market. The regression specification is:

$$y_{i,t} = \alpha + \theta Treat_{i,t} + \beta x_{i,t} + \mu_i + \lambda_t + \epsilon_{i,t} \tag{1}$$

where $y_{i,t}$ denotes the log of premium income per capita for a particular type of insurance in city $i$ and year $t$. The core explanatory variable is $Treat_{i,t}$ which equals one if the city $i$ has URBMI in year $t$ and zero otherwise. $x_{i,t}$ denotes a set of control variables as stated in last section. $\mu_i$ is the city fixed effects, $\lambda_t$ is the year fixed effects, and $\epsilon_{i,t}$ is the error term.

First two columns of Table 2 present the results for (entire) private health insurance. In Column 1, the estimatedcoefficient of *Treat* is -0.037 with a standard error of 0.064, which is statistically insignificant. After adding additional controls, the estimated coefficient of *Treat* is -0.023 that remains statistically insignificant. This means that the implementation of URBMI has no significant impact on the entireprivate health insurance.

When the explained variable is the log premium income per capita of private medical insurance, the coefficient on*Treat* (as shown in Column 3) is -0.259 with a standard error of 0.037. It is statistically significant at 1% level.After adding additional controls, the estimated coefficient remains significant. This means that the introduction of URBMI has a significant crowding-out effect on the private medical insurance market. Specifically, the implementation of URBMI causes the premium income of private medical insurance in treated cities to decline more by 27.7% compared to other cities, which is economically significant.

As for the private illness insurance market, the regression results in the last two columns in Table 2 show that the coefficient of *Treat* is statistically insignificant.Hence, the implementation of URBMI has no significant impact on private illness insurance. In short, these results above are consistent with our hypothesis I.

**Table2.** Impacts of URBMI on private health (medical and illness) insurance market

|  | (1) health | (2) health | (3) medical | (4) medical | (5) illness | (6) illness |
|---|---|---|---|---|---|---|
| *Treat* | -0.037 | -0.023 | -0.259*** | -0.277*** | 0.064 | 0.097 |
|  | (0.064) | (0.056) | (0.037) | (0.035) | (0.087) | (0.077) |
| *EntpMed* |  | -0.005 |  | -0.002 |  |  |
|  |  | (0.016) |  | (0.005) |  |  |
| *EntpIll* |  | -0.002 |  |  |  | -0.012** |
|  |  | (0.018) |  |  |  | (0.005) |
| *Gender* |  | 0.001 |  | -0.002 |  | 0.003 |
|  |  | (0.007) |  | (0.009) |  | (0.007) |
| *GDP* |  | 0.580*** |  | 0.522*** |  | 0.625*** |
|  |  | (0.092) |  | (0.102) |  | (0.104) |
| *Urban* |  | 0.475*** |  | 0.213* |  | 0.657*** |
|  |  | (0.133) |  | (0.114) |  | (0.203) |

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| *Edu* | | -1.278 | | 3.477 | | -5.313 |
| | | (3.268) | | (3.235) | | (3.712) |
| *MedShare* | | -0.457*** | | -0.380*** | | -0.560*** |
| | | (0.103) | | (0.103) | | (0.132) |
| *SocShare* | | 0.500* | | -0.033 | | 0.823** |
| | | (0.268) | | (0.285) | | (0.336) |
| *Bed* | | 0.008*** | | 0.007* | | 0.010*** |
| | | (0.003) | | (0.003) | | (0.003) |
| *Constants* | 2.899*** | -3.209** | 2.050*** | -3.110** | 2.273*** | -4.603*** |
| | (0.026) | (1.131) | (0.014) | (1.312) | (0.037) | (1.328) |
| *City FE* | Y | Y | Y | Y | Y | Y |
| *Year FE* | Y | Y | Y | Y | Y | Y |
| $R^2$ | 0.89 | 0.913 | 0.807 | 0.837 | 0.909 | 0.928 |
| *N* | 3812 | 3500 | 3812 | 3500 | 3812 | 3500 |

*Notes:*This table reports the regression results from estimating Equation (1). The regression sample size after adding the control variables is smaller than 3812 because of missing data for some control variables.Standard errors in parentheses are double-clustered at the city and year level. $^*$ $p < 0.10$, $^{**}$ $p < 0.05$, $^{***}$ $p < 0.01$.

*4.2 On the common trend assumption*

The validity of DD strategy relies on the common trend assumption thatoutcomes for the treatment and control groups should display parallel trend without the URBMI implementation. To test this hypothesis, werun the dynamic DD regression shown in Equation (2):

$$y_{i,t} = \sum_{k=-5}^{-2} \eta_k d_k + \sum_{k=0}^{4} \gamma_k d_k + \alpha + \beta x_{i,t} + \mu_i + \lambda_t + \epsilon_{i,t} \qquad (2)$$

where $d_k = 1(t - t_s^* = k)$ measures the time relative to policy year (i.e., the year launch the URBMI); $t_s^*$ means the policy year. We estimate the impact of URBMI implementation using a window of [−5, 4] around the event (i.e., five periods before the treatment and five periods after (including) the treatment), with ''binned'' indicators for five (four) or more years before (after) the policy event. Following the literature, we exclude the indicator of the year before the event, i.e. $d_{-1}$, using it as the reference year.Other variables and settings are the same as that in Equation (1).

The coefficients of interest, $\eta_k$ and $\gamma_k$, measure the difference in $y_{i,t}$ between treatment group and control group in five years leading up to URBMI's implementation and five years after relative to the reference year. The $\eta_k$ are falsification tests that capture the change in outcomes before URBMIimplementation. They are also used to test the common trends assumption. The $\gamma_k$ are intention-to-treat (ITT) effects of an additional percentage point change on outcomes after URBMI implementation. This specification identifies heterogeneity in URBMI's effect overtime. The estimates will equal zero if URBMI's implementation does not affect the private health insurance market across cities.

The results in Figure 2 give strong support to the common-trend assumption. The coefficients fluctuate around zero and are statistically insignificant before the URBIM implementation, suggesting that there were no heterogeneous trends

in the outcomes between treatment and control groups. This holds for all three outcomes.

Another important takeaway from Figure 2 is that the coefficients gradually decrease and become statistically significant after the implementation of URBMI when the outcome variable is the premium income of private medical insurance. This shows a crowding-out effect of URBMI on private medical insurance. Yet, when the explained variable is the premium income ofentire private health insurance or private illness insurance,  as shown in subfigure A and C, the coefficients are small and statistically insignificant after the implementation of URBMI. This means that the URBMI does not significantly affect the entire private health insurance or private illness insurance market. In short, from Figure 2, we verify the parallel-trend assumption and get dynamic estimation of crowding-out effects that are consistent with the baseline results.
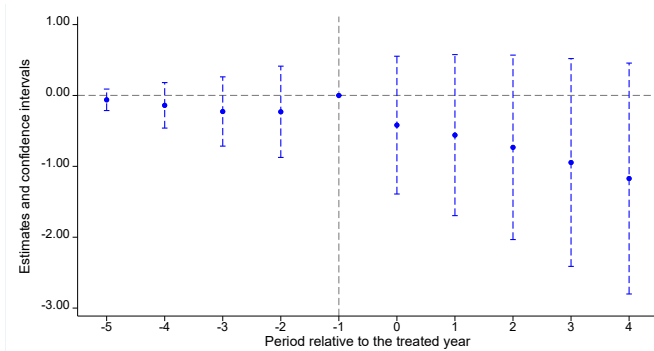


A. Health insurance
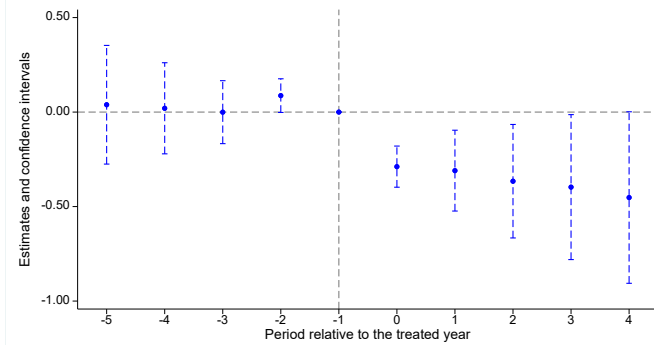


B. Medical insurance
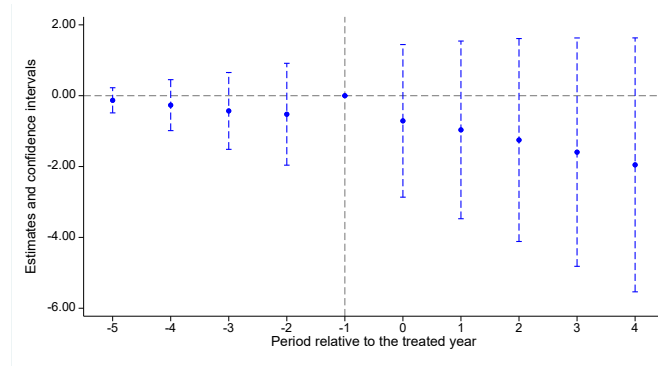
C. Illness insurance

**Figure 2.** Multiple-periods DD estimates of policy effects

*Notes:* The figure shows the regression results (point estimates and 95% confidence intervals) from estimating Equation (2) when the explained variable is log of premium income per capita of private health insurance (subfigure A), medical insurance (subfigure B), or illness insurance (subfigure C). Standard errors are double-clustered at the city and year level.

*4.3 Addressing the concern on heterogeneity treatment effect*

The recent literature in applied econometrics shows that two-way fixed-effects estimations of staggered difference-in-differences can lead to substantial biases when there are heterogeneous treatment effects (Baker et al., 2022; De Chaisemartin and d'Haultfoeuille, 2020). Following the literature, we address the concern on heterogeneous treatment effects using two methods.

First, referring to De Chaisemartin and d'Haultfoeuille (2020), we employ the difference-in-differences multiple (DDM) model to address the concern of heterogeneous treatment effect. The results, as shown in Table3, are similar to those of the canonical TWFE estimators presented in Table 2. The coefficient of *Treat* in the regression of private medical insurance is -0.199 with a standard error of 0.081 which is statistically significant.The coefficient remains statistically significant after additional controls are added.This means that the launch of URBMI causes the private medical insurance in treated cities to decline more by 23.5% than the control group. However, when the explained variable is the premium income of private health insurance or illness insurance, the coefficient of *Treat* is statistically insignificant, which is consistent with the results in Table 2.

**Table 3.** Estimations using interactive DDM model

|  | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
|  | health | health | medical | medical | illness | illness |
| *Treat* | -0.091 | -0.125 | -0.199*** | -0.235*** | -0.049 | -0.076 |
|  | (0.068) | (0.067) | (0.081) | (0.106) | (0.060) | (0.070) |
| *Controls* | N | Y | N | Y | N | Y |
| *City FE* | Y | Y | Y | Y | Y | Y |
| *Year FE* | Y | Y | Y | Y | Y | Y |
| *Switchers* | 264 | 236 | 264 | 236 | 264 | 236 |

| N | 3812 | 3500 | 3812 | 3500 | 3812 | 3500 |
|---|---|---|---|---|---|---|

*Notes:* Standard errors in parentheses are double-clustered at city and year level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Second, we follow Sun and Abraham (2021) to conduct the event study using the last-treated cities as the control group. It works by first estimating individual cohort-time-specific treatment effects, then aggregating them to produce measures of overall treatment effects while accounting for treatment effect heterogeneity. As shownin Figure 3, the treatment effects are small and statistically insignificant before thepolicy. They become larger and statistically significant after the URBMI introduction when the premium income of private medical insurance is the explained variable.Yet, the estimated coefficients are insignificant for the entire private health insurance market or private illness insurance market. Hence, the results are similar to those in Figure 2 for all outcome variables. As a result, the two bias-adjusted estimations abovealso confirm our main conclusion that the URBMI has a significant crowding out effect on the private medical insurance market but not on the private illness insurance market. In short, the bias from heterogeneous treatment effects is small in our setting and has no bearing on our main findings.
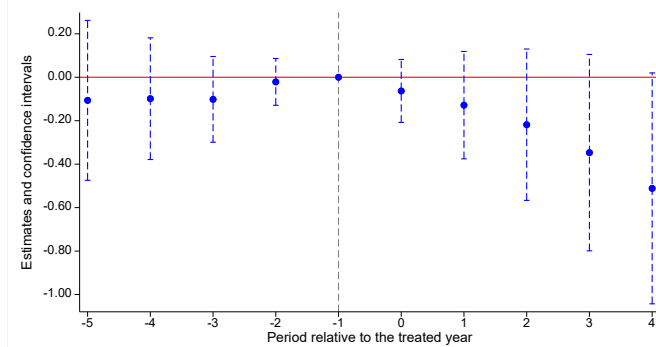


A. Health insurance
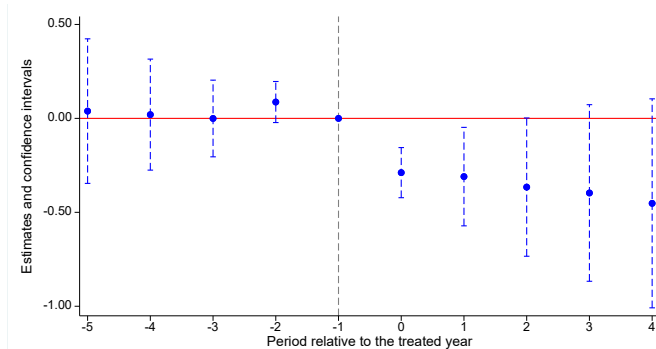


B. Medical insurance
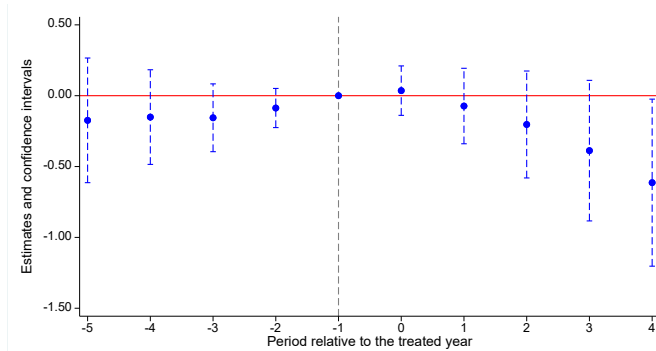
## C. Illness insurance
**Figure 3.** Event study estimates

*Notes:* The figure shows the regression results (point estimates and 95% confidence intervals) from estimating the event study model proposed by Sun and Abraham (2021). The explained variable is log of premium income per capita of private health insurance (subfigure A), medical insurance (subfigure B), or illness insurance (subfigure C). Standard errors are double-clustered at the city and year level.

### 4.4 Placebo test

In order to address the concern on other possible shocks or policy events that could influence the insurance market and bias the treatment effects of URBMI, we further conduct a placebo test in which we use the premium income per capita of life and property insurance as the explained variables and rerun the regression in Equation (1).If the URBMI had a significant crowding out effect on life or property insurance markets, there might be other treatments or unobservable factors that affect entire insurance market and drive our aforementioned results.

As shown in the first two columns in Table 4, when the log of premium income per capita of life insurance is the explained variable, the estimated coefficient of *Treat* is small and statistically insignificant. The results for property insurance are shown in the last two columns, and the coefficient on *Treat* is also statistically insignificant. These findings indicate that the URBMI has no significant effect on life or property insurance market. Therefore, the bias due to other treatments or unobservable factors is unlikely.

**Table 4.** Effects of URBMI on life and property insurance market

|  | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
|  | A. Life Insurance | | B. Property Insurance | |
| *Treat* | -0.022 | -0.007 | 0.016 | 0.048 |
|  | (0.026) | (0.020) | (0.040) | (0.028) |
| *Control Variables* |  | Y |  | Y |
| *City FE* | Y | Y | Y | Y |
| *Year FE* | Y | Y | Y | Y |
| $R^2$ | 0.913 | 0.934 | 0.896 | 0.919 |
| *N* | 3673 | 3397 | 3808 | 3498 |

## 5. Possible mechanisms

In this section, we empirically test some possible mechanisms of the crowding-out effect of URBMI on the private medical insurance market.

### 5.1 Medical cost

First, we test the mechanism of medical price (Hypothesis II) using the following regression:

$$y_{i,t} = \theta Treat_{i,t} + \varphi HighCost_i \times Treat_{i,t} + \beta x_{i,t} + \mu_i + \lambda_t + \epsilon_{i,t} \quad (3)$$

where $HighCost_i$ refers to a dummy variable that equals to one when the city's averaged medical price is larger than the median value before the launch of URBMI (before 2007) and zero otherwise. $HighCost_i$ is constant overtime for a given city, so itsestimated coefficient will be absorbed by the city fixed effects and omitted in the regression. The interaction term, $MedCPI_i \times Treat_{i,t}$, captures the heterogenous responses of outcomes to URBMI across cities with different medical costs and also shows the mechanism of health service price in the impact of URBMI on the outcomes. Other variables are the same as that in the Equation (1).

**Table 4.** Mechanism of medical cost

|  | (1) | (2) |
| --- | --- | --- |
| *Treat* | -0.279[***] | -0.301[***] |
|  | (0.059) | (0.060) |
| *HighCost×Treat* | 0.049[*] | 0.051[*] |
|  | (0.029) | (0.030) |
| *Controls* | N | Y |
| *City FE* | Y | Y |
| *Year FE* | Y | Y |
| $R^2$ | 0.807 | 0.837 |
| *N* | 3812 | 3500 |

*Notes:* The table reports the regression results from estimating Equation (3) when the explained variable is the log per capita premium income of private medical insurance. Standard errors in parentheses are double-clustered at city and year level. [*] $p < 0.10$, [**] $p < 0.05$, [***] $p < 0.01$.

Table 4 shows the results from estimating Equation (3)when the explained variable is log of per capita premium income of private medical insurance. The coefficient of *Treat* is negative and statistically significant, and the coefficient of interaction term is positive and statistically

significant.This means that the crowding out effect of URBMI on private medical insurance is more (less) significant in regions where medical service priceare lower (higher). The coefficient of interaction term remains positive and statistically significant after additional controls are added. Overall, the regression results are consistent with the Hypothesis II and the mechanism of medical costholds.

### 5.2 Medical service accessibility

To examine the mechanism of medical service accessibility (Hypothesis III), we run the following regression in Equation (4):

$$y_{i,t} = \theta Treat_{i,t} + \varphi HighAcc_i \times Treat_{i,t} + \beta x_{i,t} + \mu_i + \lambda_t + \epsilon_{i,t} \qquad (4)$$

where we use the number of health facility staff per capita as a proxy variable for medical service accessibility. $HighAcc_{i,t}$ refers to a dummy variable which turns one for cities whose medical service accessibility is larger than the median value before the launch of URBMI and zero otherwise.Other variables are the same as that in the Equation (1).

**Table 5.** Mechanism of medical service accessibility

|  | (1) | (2) |
|---|---|---|
| *Treat* | -0.147** | -0.180*** |
|  | (0.058) | (0.051) |
| *HighAcc×Treat* | -0.198** | -0.185*** |
|  | (0.065) | (0.057) |
| *Controls* | N | Y |
| *City FE* | Y | Y |
| *Year FE* | Y | Y |
| $R^2$ | 0.810 | 0.839 |
| *N* | 3812 | 3500 |

*Notes:* The table reports the regression results from estimating Equation (4) when the explained variable is the log per capita premium income of private medical insurance. Standard errors in parentheses are double-clustered at city and year level.* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 5 displays the regression results from estimating Equation (4). The estimated coefficient of *Treat* is -0.147 with the standard error of 0.058, and the coefficient on the interaction term is -0.198 with a standard error of 0.065, both of which are statistically significant.Similarly, after adding additional controls, the coefficient on the interaction term remains positive and significant. This means that the higher (lower) the accessibility of health services, the stronger (weaker) the crowding-out effect of URBMI on private medical insurance. Specifically, the crowding out effect of URBMI on private medical insurance is 18%-20% greater in regions with higher medical service accessibility. Therefore, the results provide strong support to the mechanism of medical service accessibility (Hypothesis III).

### 5.3 Health risks

We then investigate the mechanism of health risk in the demand side. We measure the health risk using the incidence of infectious disease and run the

regression with interaction term as before to test the hypothesis:

$$y_{i,t} = \alpha + \theta Treat_{i,t} + \varphi HighRisk_i \times Treat_{i,t} + \beta x_{i,t} + \mu_i + \lambda_t + \epsilon_{i,t} \quad (5)$$

where $HighRisk_i$ refers to a dummy variable which equals to one for cities whose incidence of infectious disease is above the median value and zero otherwise during the pre-policy period. $HighRisk_i$ is constant overtime for a given city, so it is absorbed by the city fixed effects. The interaction term shows the mechanism of health service price in the impact of URBMI on the outcomes. Other variables are the same as that in the Equation (1).

Table 6 shows the results from estimating Equation (5). In column (1), the estimated coefficient of the *Treat* is -0.311 and the interaction term of *Treat* and *HighRisk* is 0.112 with a standard error of 0.053. After adding additional controls, as shown in the column (2), the coefficient of interaction term is -0.105 that remains statistically significant at 5%. This means that the higher (lower) level of health risk, the weaker (stronger) the crowding-out effect of URBMI on the private medical insurance market. It is consistent with the Hypothesis IV.

**Table 6.** Mechanism of health risk

|  | (1) | (2) |
|---|---|---|
| *Treat* | -0.311*** | -0.317*** |
|  | (0.069) | (0.069) |
| *HighRisk×Treat* | 0.112** | 0.104** |
|  | (0.053) | (0.050) |
| Controls | N | Y |
| City FE | Y | Y |
| Year FE | Y | Y |
| $R^2$ | 0.811 | 0.834 |
| N | 3812 | 3500 |

*Notes:* The table reports the regression results from estimating Equation (5) when the explained variable is the log per capita premium income of private medical insurance. Standard errors in parentheses are double-clustered at city and year level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

*5.4 Alternative mechanisms*

We have provided some evidences of the potential mechanisms of the crowding-out effect of URBMI on the private medical insurance market. At the same time, there might be some other stories that support the null hypothesis that the URBMI has no significant crowding-out (or crowding-in)effect on the private medical insurance market. To make our results more convincing, we consider two of these stories that may contradict with our previous findings.

(1) Alternative mechanism: Consumption budget

First, the URBMI may increase the consumption budget, and people with more budget can afford more insurance service (Clark et al., 2008). Thus, it makes the cost of private health insurance seem more acceptable to

people with high budget or income (DeLeire et al., 2017). In other words, high income exerts a significantly positive effect on the intention to purchase private health insurance. The intention to purchase private health insurance shows a distinctly positive impact on actual purchase and the intention to purchase health insurance yield a significantly positive mediating effect on the crowding-out effect of URBMI (Mamun et al., 2021). It implies that the crowding-out effect of URBMI will be smaller in regions with higher average income levels.

To test this alternative mechanism, we present estimates from the following model:

$$y_{i,t} = \alpha + \theta Treat_{i,t} + \varphi HighInc_i \times Treat_{i,t} + \beta x_{i,t} + \mu_i + \lambda_t + \epsilon_{i,t} \quad (6)$$

where $HighInc_i$ refers to a dummy variable which turns one when the value of income per capita is larger than the median of the income level and zero for the opposite.

Table 8 shows the results from estimating Equation (6). The coefficient of interaction item inColumn 1 is -0.033 with a standard error of 0.057 which is statistically insignificant.After adding additional controls, the coefficient remains statistically insignificant. As for private medical and illness insurance, no matter adding additional control variables or not, the coefficients of interaction item also are statistically insignificant.These results show that the impacts of URBMI on the private health insurance market and its segments are similar in groups with high- or low-income level. Hence, the alternative mechanism of increasing budget is unlikely to be true.

**Table 8.** Alternativemechanism: Consumption budget

|  | (3) | (4) |
|---|---|---|
| *Treat* | -0.244*** | -0.289*** |
|  | (0.053) | (0.050) |
| *HighInc×Treat* | -0.021 | 0.022 |
|  | (0.059) | (0.060) |
| *Control variables* | N | Y |
| *City FE* | Y | Y |
| *Year FE* | Y | Y |
| $R^2$ | 0.809 | 0.838 |
| *N* | 3812 | 3500 |

*Notes:*The table reports the regression results from estimating Equation (6) when the explained variable is the log per capita premium income of private medical insurance.Standard errors in parentheses are double-clustered at city and year level. $^*$ $p < 0.10$, $^{**}$ $p < 0.05$, $^{***}$ $p < 0.01$.

(2) Alternative mechanism: Financial Literacy

Second, public insurance programs may increase people's financial literacy (Frijns et al., 2013), and a better financial literacy is associated with

a higher demand for financial products including commercial insurance (Bauhoff et al., 2020). In this way, the URBMI could crowdin effect on the private medical insurance market. The mechanism of financial literacy indicates that the crowding-out effect of URBMI will be smaller in regions with higher financial literacy level. To test this opposite story to our results, we run the following regression shown in Equation (7):

$$y_{i,t} = \alpha + \theta Treat_{i,t} + \varphi HighLiteracy_i \times Treat_{i,t} + \beta x_{i,t} + \mu_i + \lambda_t + \epsilon_{i,t} \tag{7}$$

where we measure the financial literacy using the ratio of population with high education in a given city. $HighLiteracy_i$ refers to a dummy variable which equals to one when the ratio of population with high education is larger than the median value and zero otherwise.

**Table 9.** Alternativemechanism: Financial Literacy

|  | (3) | (4) |
|---|---|---|
| *Treat* | -0.137[**] | -0.170[***] |
|  | (0.054) | (0.051) |
| *HighLiteracy×Treat* | -0.215[***] | -0.216[***] |
|  | (0.067) | (0.063) |
| *Control variables* | N | Y |
| *City FE* | Y | Y |
| *Year FE* | Y | Y |
| $R^2$ | 0.823 | 0.844 |
| *N* | 3812 | 3500 |

*Notes:*The table reports the regression results from estimatingEquation (7) when the explained variable is the log per capita premium income of private medical insurance.Standard errors in parentheses are double-clustered at city and year level. [*] $p < 0.10$, [**] $p < 0.05$, [***] $p < 0.01$.

Results from estimating Equation (7) are presented in Table 9. No matter adding control variables or not, the coefficient on the interaction term is negative, and it is statistically and economically significant. This means that the crowding-out effect of URBMI is significantly larger in regions with higher financial literacy level, which is contradicted with the implication from the financial literacy mechanism.

Although this result indicates that the financial literacy mechanism is unlikely to hold, it does imply something about the heterogeneous effects of URBMI among cities with different levels of financial literacy. Private health insurance in China has a substantially lower cost-performance ratio than URBMI. Given that the theory assumes that financial literacy improves returns on savings (Lusardi and Mitchell, 2014), only people with a sufficient level of financial knowledge can differentiate this information, though. In other words, assuming there is no adverse selection, if people have a higher (lower) level of financial literacy, they can make wiser (more

unreasonable) choices on public and private health insurance and choose URBMI to replace private health insurance in the context of China.

## 6. Conclusion

To estimate the net welfare of public health insurance, it is necessary to analyze its causal effect on private health insurance. We estimate this effect based on a quasi-experiment of URBMI in China. We find that URBMI has a negative but statistically insignificant effect on private health insurance. Furthermore, we distinguish between private health insurance into private medical and illness insurance, and the results show that URBMI has a significant crowding-out effect on the former one but not on the latter one. This could help to explain the mixed findings in the literature about the relationship between public and private health insurance. Regarding on potential mechanisms, we demonstrate that the crowding-out effect of URBMI on the private medical insurance is through lowering the medical price, increasing the accessibility to health services and improving individuals' health status.

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# 生存预期与中老年人商业医疗保险需求：感性与理性之别

徐睿、王晓军、孟辉、周明[1]

**摘要：**

通过在主、客观维度基础上考虑时间因素，本文论述了感性和理性视角下生存预期的差别，并检验了其对我国中老年人商业医疗保险需求的影响：感性生存预期易受外界冲击影响，引发正向连带效应；理性视角下则具有稳定性。相关影响通过居民社交积极性和生活满意度进行传导，家庭养老观念和遗赠动机会对其产生抑制效应。此外，现代化的生活环境、特定的人生阶段和充裕的物质条件是生存预期对中老年人商业医疗保险需求影响显著的重要条件。

**关键词：** 感性，理性，生存预期，保险需求

## 一、引言与文献回顾

改革开放四十多年来，我国经济迅速发展，其中，健康领域的改革发展取得显著成就，人民的健康水平和身体素质在持续提高。同时，人口老龄化、生态环境及生活方式变化等，也给维护和促进人民健康带来一系列新的挑战。从家庭角度来看，由于健康本身不能交易也不可储蓄和跨期配置，因此健康风险是现代家庭金融中最具有特殊性的风险因素。从年龄分布角度来看，健康风险的威胁对象主要在中老年人群体，对这种潜在风险威胁的反馈，一般被具象为对自身生命存续的预期。这种反馈在已有文献中有多种名称，例如"生存预期"、"健康预期"、"预期寿命"、"自评健康余命"等等，但整体含义相近，为简便起见，本文统一将其称为"生存预期"。

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随着家庭财富积累程度不断变大，中老年家庭成员的生存预期逐渐成为了家庭财富管理的重要考量因素，因为它既直接对家庭的财富分配和资金花费决策过程产生影响（何兴强和史卫，2014）[1]，也对潜在的医疗费用支出产生影响，进而间接影响家庭财富的积累进程（丁梦和冯宗宪，2020）[2]。已有研究表明，居民的健康状态在出现一系列波动后（例如患有角膜炎等疾病），其对于医疗保险的关注度和相应保费支出显著增加（周华林和王向楠，2018）[3]，可见生存预期差的居民劳动收入稳定性较差，风险厌恶水平较高（刘潇等，2014）[4]，持有风险金融资产的概率要低（雷晓燕和周月刚，2010）[5]，这显然不利于家庭财富的正常运作和稳定增值。综上可知，健康风险主要来源于家庭的人力资本（尤其是中老年人），并通过"生存预期"表现出来，其对于家庭资金的释放有显著的约束性，是影响家庭金融发展的潜在因素，不利于社会经济运行的长久平稳。

因为医疗保险的覆盖可以使得家庭减少预防性储蓄量，优化资产配置和消费结构，显著增加家庭的资本活力并扩大内需（周钦等，2015；王美娇和朱铭来，2015）[6][7]，所以，积极参加包括商业疾病保险、医疗保险、医疗意外保险和长期看护保险在内的商业健康保险是居民防范健康风险的重要手段。当然，社会医疗保险和商业医疗保险的发展并非矛盾，已有研究表明，社会医疗保险制度的发展通常对商业健康保险起到推动和促进作用（王向楠，2011）[8]：政府可以考虑通过提高保险筹资水平来提高医保报销比例，并且大力发展商业医疗保险（吴庆跃和周钦，2015；王稳和孙晓珂，2020）[9][10]，尽管这一过程可能会存在城乡地域差异（彭浩然等，2017；易行健等，2019）[11][12]；另外，整合城乡居民基本医保制度虽然会增加财政压力，但也会显著提高居民商业健康保险需求（胡祁和朱铭来，2021）[13]。可见，在重视社会医疗保险制度发展的基础上，鼓励和促进商业医疗保险的发展不失为缓解健康风险的重要途径之一。

在已有文献中，对于居民健康风险的定义与量化大多是在主观视角下进行展开（例如黄国桂，2020、2021；邹小芃和杨芊芊，2019；王记文，2021）[14]~[17]，虽然这为直观的健康风险比较和预测提供了便利，但是从主观视角展开（其中，刘潇等（2014）[12]曾将主观视角和客观视角结合起来），难免会受到人群对健康信息和所处环境了解程度的局限（黄匡时，2018）[18]：在当下时间节点，受访主体的主观感受（例如情绪、思维方式）往往会受到各类外界信息的冲击，此时的自我评判偏向于感性视角；但是如果放眼整个生命周期，受访主体的主观感受则会不自觉地与外界信息冲击相剥离，从而产生更为理性的评判[1]。所以同一个人在不同时间状态下对于

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1 以被动失业这一负面信息冲击为例，受访主体在当下时间节点评价自身的健康状况时，会将这一负面信息冲击导致的消极情绪带入其中，得出的自评是感性的；但是在十年后，让同一受访主体再一次对当时的自身健康状况进行回顾，他往往会将当时的

同一项事务，做出的评判和选择是不同的。另外，当人们在日常工作生活中，往往是将主观和客观两个角度糅合起来考虑问题的，比如反应居民生理和心理状况的客观维度指标也会对居民的自评健康产生影响（杨珏和李建新，2017）[19]，因此仅仅从主观角度出发略显单薄和片面，也不太符合实际情况，所以在前人基础上，本文加入对时间状态的考量，将传统的主观、客观因素重新排列组合，尝试从感性和理性的全新视角进一步诠释生存预期概念的内涵，对已有研究成果进行拓展。在此基础上，本文就不同视角下的生存预期对于中老年人商业医疗保险需求的特异化影响进行了对比研究，并尝试性地提出了两个潜在的影响路径。

此外，虽然目前放开了二胎甚至三胎生育，但是曾经实施时间较长的计划生育政策深刻地改变了我国家庭的子女养育和财富传承理念，传统的家庭养老模式受到一定的冲击，自我养老的观念愈发为大众所接受。不同养老观念间的博弈和遗赠动机的存在同样会对我国中老年人商业医疗保险需求造成潜在影响，这为现有研究所忽视，所以本文将之作为调节因子加入模型框架，进一步分析感性和理性视角下的生存预期对我国中老年人商业医疗保险的需求影响，对前人的研究进行丰富。

最后，不同社交圈的认知特点会影响个体在融入社区的过程中获得的信息内容；而具有不同财富认知和生命历程的个体对家庭运行和自身健康的重视程度也有所不同，两者都会影响其对商业医疗保险的需求。因此，本文还从社交圈认知、财富水平和个体生命历程三个方面对上述基准模型进行了异质性分析，进一步挖掘了生存预期的两个视角对我国中老年人商业医疗保险需求影响的差异。

## 二、研究设计

本文采用 2018 年中国健康与养老追踪调查（CHARLS）的全国调查数据进行实证检验与分析。该数据库由北京大学国家发展研究院主持构建，北京大学中国社会科学调查中心负责执行，抽样收集了全国除宁夏回族自治区、西藏自治区和海南省之外的 28 个省（直辖市、自治区）45 岁及以上中老年家庭的数据，内容涵盖了家户成员基本信息、身体与精神健康状况、养老医疗保险覆盖、工资收入与资产分布等各个方面。

（一）变量设定

1.被解释变量

2020 年 6 月 1 日正式实施的《中华人民共和国基本医疗卫生与健康促进法》第八十三条指出：“国家鼓励发展商业健康保险，满足人民群众多样化健康保障需求。”商业医疗保险属于商业健康保险的重要组成部分，能够

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身体状态放在整个生命周期中进行冷静比较和考量，得出的自评结果会和十年前有所不同，偏向于理性。

对我国居民的健康管理各环节形成补充，因此本文的核心被解释变量为：我国中老年人购买商业医疗保险的意愿，用 com_med 表示；并以问卷中"你是否购买了商业医疗保险"这一问题作为数据整理依据：如果受访者购买了商业医疗保险，则取值为1，否则取值为0。

2.解释变量

与已有研究一致，本文的核心解释变量为生存预期，用 surv_exp 表示；不同的是，本文进一步将自评主体的时间状态考虑在内，从而将主观视角进一步细分为：回顾性主观自评、即时性主观自评和展望性主观自评，三者分别代表：站在当下时点对过去的回顾总结、自身对当下情境的瞬时反馈和站在当下时点对未来的展望预测。如上文所述，在充分考虑时间状态的条件下，主观视角中的感性因素和理性因素得以挖掘：一般而言，人们总是对回顾和展望性评判更具理性考量，而对当下情境的瞬时反馈则会不自觉地倾注更多感性因素；整体来看，前二者属于"延时性"主观自评，是对后者的修正和批判。另外，由于理性的评判同时是以自身的客观状态为基础，所以本文在刘潇等（2014）[12]提出的主客观视角结合的方法论上，将核心解释变量的结构进一步完善为感性视角和理性视角，其中，感性视角 perc_exp 是对即时性主观自评内容的提炼，包括了生理和心理两个方面；理性视角 rati_exp 则包含了对延时性主观自评内容的提炼和自评主体的客观状态得分（详见图1）。
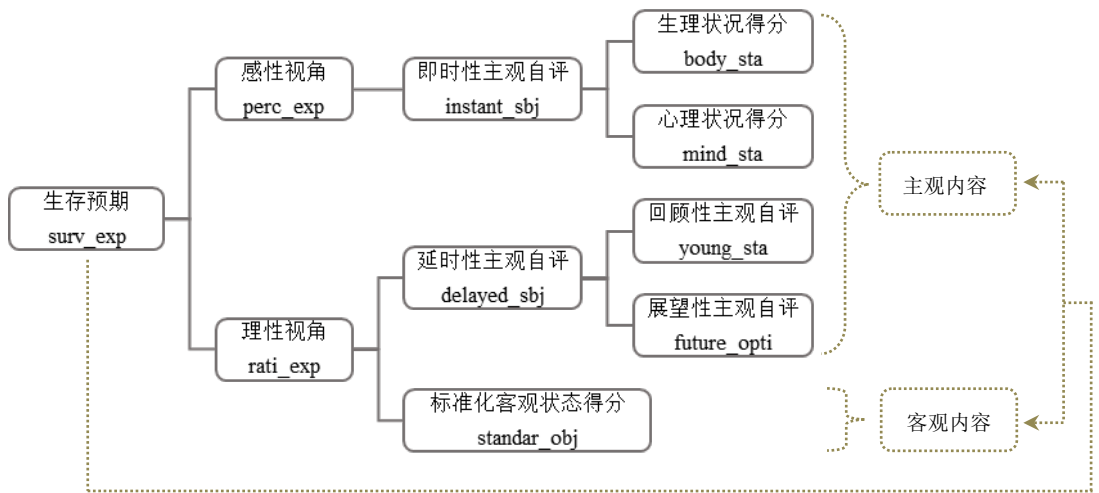


图 1 解释变量结构层次图

从图1可以看出，已有研究成果更多是将视角聚焦在即时性主观自评中的生理方面，而在考虑时间状态后，本文所构建的感性与理性视角则能够较为全面地概括中老年人生存预期的内涵；同时，从感性和理性角度进

行思考和评价也更加贴合人们在日常工作生活中的习惯。

对于感性视角，我们参考已有文献，整理 DA002 问题"您认为您的健康状况怎样？"作为生理状况自评的代理变量，记作 body_sta[1]；整理 DC004 问题"您觉得自己现在的记忆力怎么样？"作为心理状况自评的代理变量，记作 mind_sta，取值同前。为方便分析，上述两个变量取均值得到即时性主观自评变量 instant_sbj，即感性视角下生存预期。

对于理性视角下的延时性主观自评，我们整理 DA048 问题"您 15 岁之前（包括 15 岁）的身体状况怎么样？"作为回顾性主观自评的代理变量，记作 young_sta；整理问题 DA081"对于居民未来生存年数乐观的衡量"作为展望性主观自评的代理变量，记作 future_opti[2]。为方便分析，两者取均值得到延时性主观自评变量 delayed_sbj。

对于理性视角下的标准化客观状态得分，本文吸纳刘潇等（2014）[12]、邹小芃和杨芊芊（2019）[16]提出的思路，对 CHARLS 问卷中受访者的个人身体状况问题进行了系统提取和得分规定[3]，并使用熵权法对其依次去量纲和计算权重，得到最终得分变量 standar_obj。

最后，为方便分析，将延时性主观自评变量 delayed_sbj 和标准化客观状态得分 standar_obj 取均值，得出理性视角下生存预期；而居民总的生存预期则也是感性生存预期和理性生存预期的均值。

3.控制变量

对于控制变量，本文处理如下：借鉴潘炜迪和袁辉（2021）[21]，本文将婚姻状况设置为有偶且同居（包括暂时分居）取值 1，无偶或分居取值 0，以体现降低生活成本对个人保险需求的影响作用；考虑到我国交通信息网络的完善，本文认为居住地因素逐渐成为取代居民户口的商业保险需求影响因素，并纳入控制变量；对于个人资产部分，本文将其分为现金资产、银行存款、是否理财和是否拥有住房公积金进行控制；参考雷晓燕和周月刚（2010）[13]、周慧珺等（2020）[22]，将家庭月平均消费水平而非个人收入作为控制变量[4]；摔倒可能会造成的髋关节骨折[1]是我国中老年人遭受意

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1 对于问题中的选项"很好"、"好"、"一般"、"不好"、"很不好"，本文依次取值 1~5；对于选项"不知道"，本文认为此类受访者对长寿乐观持有中庸态度，因此同样取值 3。
2 该问题依据年龄将受访者分为了 9 组，并分别询问其未来至少存活至特定岁数的可能性，其选项设置分别为：几乎不可能（取值 1）、不太可能（取值 2）、有可能（取值 3）、很可能（取值 4）、简直一定（取值 5）。
3 主要包括残疾种类数量统计、慢性病诊断数量统计、慢性病病情得分、疼痛得分、日常生活能力（ADL）得分、抑郁得分和认知能力得分共 46 个指标。其中，以饶育蕾和朱锐（2014）[20]构建的认知能力指标为基础，我们认为认知能力得分包括了受访者的简易精神状态检查、回忆能力得分（包括即时、延迟回忆得分）、逻辑推理能力得分和计算能力得分 4 大部分，与心理状况自评相对应，属于纯客观得分。
4 消费往往包含了个人心理偏好，同时考虑到中老年人的消费动机更多从家庭角度出发，而家庭的消费偏好往往会影响个人对于健康的重视程度，本文用家庭月平均消费

外伤害死亡的主要原因，因此将其作为生活中遭受各类冲击情况的代理变量归入控制变量，以使得回归结果更加稳健；另外，本文还对典型不良嗜好进行控制：其中，是否抽烟变量细分为1.仍然抽烟、2.目前已戒烟和3.从不抽烟三个选项；考虑到中老年人往往会认为适当的饮酒有利于身心健康，因此本文将是否饮酒变量改进为1.喝烈性酒、2.不喝烈性酒、3.从不喝酒三个选项。除此之外，本文照例将年龄、性别、受教育程度等常见人口学特征纳入控制。

对以上变量剔除缺失值和异常值后，获得有效截面样本10078个，具体描述性统计结果详见表1[2]。

表1主要变量的描述性统计

| 变量 | 均值 | 标准差 | 最小值 | 最大值 |
|---|---|---|---|---|
| 是否拥有商业医疗保险 | 0.05 | 0.22 | 0.00 | 1.00 |
| 总体生存预期 | 2.26 | 0.51 | 0.80 | 4.00 |
| 感性视角 | 2.51 | 0.73 | 1.00 | 5.00 |
| 即时性主观自评（生理） | 3.11 | 1.01 | 1.00 | 5.00 |
| 即时性主观自评（心理） | 1.92 | 0.79 | 1.00 | 5.00 |
| 理性视角 | 2.00 | 0.47 | 0.61 | 3.00 |
| 延时性主观自评 | 3.24 | 0.87 | 1.00 | 5.00 |
| 标准化客观状态得分 | 0.77 | 0.17 | 0.14 | 1.00 |
| 性别 | 0.53 | 0.50 | 0.00 | 1.00 |
| 年龄 | 59.79 | 9.12 | 20.00 | 108.00 |
| 居住地 | 2.45 | 0.85 | 1.00 | 4.00 |
| 受教育水平 | 4.02 | 1.81 | 1.00 | 11.00 |
| 婚姻状况 | 0.90 | 0.30 | 0.00 | 1.00 |
| 信仰 | 1.91 | 0.28 | 1.00 | 2.00 |
| 现金资产 | 6.90 | 1.66 | 0.00 | 13.13 |
| 银行存款 | 8.10 | 2.69 | 0.00 | 14.81 |
| 投资状况 | 0.04 | 0.19 | 0.00 | 1.00 |
| 住房公积金 | 0.04 | 0.20 | 0.00 | 1.00 |
| 是否抽烟 | 2.26 | 0.88 | 1.00 | 3.00 |

水平取代个人消费作为控制变量。另外，为使得数据更加平稳，本文对现金资产、银行存款和家庭月平均消费这三个控制变量分别进行了对数化处理。

1 虽然中老年人摔倒遭遇骨折的概率不足10%，大概率不会严重影响当事人的生命质量，但是心理阴影难医。

2 本文还针对受访者的不同性别、不同居住地以及是否摔倒过三个视角进行了分类描述统计，具体结果详见本文附录。

| 是否喝烈性酒 | 2.39 | 0.82 | 1.00 | 3.00 |
|---|---|---|---|---|
| 是否摔倒过 | 0.18 | 0.38 | 0.00 | 1.00 |
| 家庭月平均消费 | 7.23 | 1.04 | 0.00 | 11.92 |

如表 1 所示，被解释变量的均值为 0.05，说明商业医疗保险在我国中老年人群体中的普及率较低。对于解释变量而言，感性视角下的生存预期均值为 2.51，标准差为 0.73，而理性视角下则分别为 2.00 和 0.47，说明相比于理性视角，受访者在感性视角下的生存预期可能更加"乐观"，另外，由于感性视角属于即时性的自我评判，故而其波动性也较大；上述比较也体现了依据不同时间状态对主观自评进行划分的必要性。进一步来看，受访者的生理状况自评均值在 3.11，而心理状况自评均值却在 1.92，且标准差远小于前者得分，可见我国中老年人的心理健康水平自我感觉和身体健康水平并非同步，同时，以往研究仅仅通过受访者的生理状况自评来分析问题也是不全面的，有必要将二者结合起来（即文中所说的感性视角）重新审视。与感性视角的纯主观内容不同，理性视角包含了主观和客观两个方面。延时主观自评包含了受访者的回顾和展望，其均值在 3.24，且波动较小；而受访者当下状态则通过熵权法计算得分，最大值为 1，均值接近 0.8，可见受访者在接受采访时身体的客观健康程度较好；此外，它也可以看作是对延时主观自评的补充。从受访者个人特征来看，样本平均年龄在 60 岁左右，婚姻状况较为稳定，多有所陪伴，平均受教育水平在小学学历以上；同时，受访者在接受采访时普遍开始戒烟，同时对于喝酒的态度更多是趋向于小酌怡情，说明对自身的健康问题关心程度较高。

（二）模型构建

本文使用二元因变量"你是否购买了商业医疗保险"来诠释中老年人的商业医疗保险购买意愿。其中，问卷受访者已经拥有商业医疗保险的样本量占总样本的比重为 4.86%，属于"稀有事件"范畴，根据 King 和 Zeng（2001）[23]和 Allison（2012）[1]，本文使用 Firth-Logit 模型[2]进行基础回归，设连接函数 $F(\mathbf{X}, \boldsymbol{\beta})$ 是逻辑分布的累积分布函数，则有

$$\Pr(\mathrm{com\_med}=1|\mathbf{X})=F(\mathbf{X}, \boldsymbol{\beta})=\frac{\exp(\mathbf{X}'\boldsymbol{\beta})}{1+\exp(\mathbf{X}'\boldsymbol{\beta}+\varepsilon)}$$

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1 Allison 的观点详见其博客
https://statisticalhorizons.com/logistic-regression-for-rare-events/
2 与 Logit 模型的区别在于，Firth-Logit 模型采用惩罚极大似然估计法（Penalized Maximum Likelihood Estimation）对系数向量 β 进行估计，以减小"稀有事件偏差"，从而得到有限、一致的参数估计结果（Heinze 和 Schemper，2002）[24]。

其中，ε 是残差项，而 $\mathbf{X'\beta}$ 包含了核心解释变量、控制变量和常数项。如果令 $Y = \ln\left(\dfrac{p}{1-p}\right)$，其中，$p = Pr\left(com\_med = 1\middle|\mathbf{X}\right)$，则基础回归模型如式（1）~（3）：

$$Y = \alpha_1 + \alpha_2 \cdot surv\_exp + \sum \varphi \cdot CVs + \varepsilon_1 \quad （1）$$

$$Y = \beta_1 + \beta_2 \cdot perc\_exp + \sum \delta \cdot CVs + \varepsilon_2 \quad （2）$$

$$Y = \gamma_1 + \gamma_2 \cdot rati\_exp + \sum \phi \cdot CVs + \varepsilon_3 \quad （3）$$

在式（1）基础上，本文还分别构建了调节效应模型如式（4）、中介效应模型如式（5）~（6）：

$$\begin{aligned} Y = \eta_1 + \eta_2 \cdot surv\_exp + \eta_3 \cdot var_{modra} + \\ \eta_4 \cdot surv\_exp \times var_{modra} + \sum \kappa \cdot CVs + \varepsilon_4 \end{aligned} \quad （4）$$

$$var_{media} = \lambda_1 + \lambda_2 \cdot surv\_exp + \sum \vartheta \cdot CVs + \varepsilon_5 \quad （5）$$

$$Y = \mu_1 + \mu_2 \cdot surv\_exp + \mu_3 \cdot var_{media} + \sum \omega \cdot CVs + \varepsilon_6 \quad （6）$$

其中，$var_{modra}$ 表示调节变量；$var_{media}$ 表示中介变量。对于式（4）~（6）中的核心解释变量 surv_exp，本文同样会从感性视角（perc_exp）和理性视角（rati_exp）分别进行分析对比。

## 三、实证结果

（一）基准回归结果分析

表 1 报告了受访者的生存预期及其两个视角分别对商业医疗保险需求程度的影响，核心解释变量的结果显著性较强。其中，第（1）、（3）、（5）列作为对比组，展示了没有加入控制变量时的回归结果；在加入了控制变量后，结果如（2）、（4）、（6）列所示。整体来看，在其他条件不变时，当受访者的总体生存预期改善 1 个单位，个人对商业医疗保险需求的优势比增加 63.07%，该结果在 1%水平下显著。

从感性视角来看，生存预期与受访者的商业医疗保险需求呈现正向关系。具体而言，保持其他条件不变，如果外界情境的变动对居民个体产生一定冲击，导致其对于当前环境下自身生存状态的看法趋于悲观，即使自

身的生理指标与之前无异，其对于商业医疗保险的需求也会变得降低。可见，在考虑了心理因素后，我国中老年人产生的负面情绪波动非常容易引发连带效应，不利于中老年人的健康管理。当然，如果受访者的感性生存预期得到修复和改善，例如在遭遇地震或洪涝侵袭后，随着灾后重建工作在当地有续展开，未来可期的心态会使得受灾群众对于购买商业医疗保险的需求出现增长。从传统观念角度来讲，类似于"癌症病人较为忌讳谈论死亡和财富分配，而常人却更愿意了解配置健康保险的得失"，如果受访者自我判断改善，健康的心理暗示加强，他才不会抗拒外界对于商业医疗保险的推广以及健康管理概念的宣传。

表2中老年人生存预期与商业医疗保险购买意愿

|  | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| 生存预期 | 0.915*** (0.09) | 0.489*** (0.10) | | | | |
| 感性视角 | | | 0.517*** (0.06) | 0.287*** (0.07) | | |
| 理性视角 | | | | | 0.956*** (0.11) | 0.426*** (0.12) |
| 性别 | | -0.018 (0.14) | | -0.014 (0.14) | | 0.002 (0.14) |
| 年龄 | | 0.039 (0.06) | | 0.051 (0.06) | | 0.036 (0.06) |
| 年龄平方 | | -0.001* (0.00) | | -0.001** (0.00) | | -0.001* (0.00) |
| 居住地 | | -0.074 (0.06) | | -0.081 (0.06) | | -0.077 (0.06) |
| 受教育水平 | | 0.125*** (0.03) | | 0.131*** (0.03) | | 0.133*** (0.03) |
| 婚姻状况 | | -0.387** (0.19) | | -0.391** (0.19) | | -0.388** (0.19) |
| 信仰 | | 0.111 (0.18) | | 0.106 (0.18) | | 0.093 (0.18) |
| 现金资产 | | 0.087*** (0.03) | | 0.090*** (0.03) | | 0.092*** (0.03) |
| 银行存款 | | 0.033* (0.02) | | 0.034* (0.02) | | 0.035* (0.02) |

| | | | | | | |
|---|---|---|---|---|---|---|
| 投资状况 | | 1.029*** (0.15) | | 1.027*** (0.15) | | 1.046*** (0.15) |
| 住房公积金 | | -0.276 (0.18) | | -0.262 (0.18) | | -0.277 (0.18) |
| 是否抽烟 | | 0.169** (0.07) | | 0.169** (0.07) | | 0.177** (0.07) |
| 是否喝酒 | | -0.165** (0.07) | | -0.169** (0.07) | | -0.178*** (0.07) |
| 是否摔倒 | | 0.098 (0.13) | | 0.083 (0.13) | | 0.063 (0.13) |
| 家庭消费 | | 0.315*** (0.06) | | 0.319*** (0.06) | | 0.314*** (0.06) |
| 常数项 | -5.138*** (0.22) | -6.680*** (1.92) | -4.337*** (0.17) | -6.637*** (1.94) | -4.972*** (0.24) | -6.346*** (1.92) |

注：括号里为稳健标准误。*代表 $p < 0.10$，**代表 $p < 0.05$，***代表 $p < 0.01$；样本量N=10078；本文将系数通过公式 $exp(\hat{\beta}_i)$ 转化为优势比进行分析。

从理性视角看，生存预期与受访者的商业医疗保险需求同样呈现正向关系。与感性视角不同的是，其他条件不变的情况下，理性视角下生存预期的变动带来商业医疗保险需求的优势比变动无疑是巨大的；但从另一个方面来说，相比于感性视角下生存预期很容易因受到外界冲击的影响而变动，理性视角下生存预期的变动并不是件容易或经常发生的事，具有稳定性。这也解释了为什么现实中会有很多人受身边因素影响，"一时冲动"为自己配置了商业医疗保险，但过了一段时间后会后悔甚至放弃续保的现象。

值得注意的是，按照本文的指标构建逻辑，保险市场中常见的逆向选择问题存在于理性生存预期视角下。因为当居民理性生存预期出现恶化时，往往意味着自身年轻时健康状况不佳，各项生理指标出现了恶化，或者未来生存年乐观开始削弱，而这些特征或心理变动都是只有当事人自己才知晓的信息。此时，信息的不对称以及理性人特征反而会提高当事人的商业健康保险需求，因此和本文回归结果并不一致，如何利用微观数据进一步给出具体的实证依据则是下一步的研究方向。

经分析可知，中国居民对于商业医疗保险的需求变动是同时受到自身感性因素和理性因素的共同影响，当感性或者理性视角下生存预期趋于乐观，居民对于上述两种商业保险的需求优势比会上升，反之则会下降。从系数对比来看，如果变动量相同，感性视角下生存预期对商业医疗保险需求优势比的影响小于理性视角，即如果一个整体健康水平较高的居民的感

性视角下生存预期出现恶化的同时，其理性视角下生存预期出现改善，虽然前者会抑制居民对于商业保险的购买欲望，但是受两者综合影响后，居民对于上述两种商业保险的需求依然是上升的。

（二）稳健性检验

为保证结果的稳健性，参考陈强（2014）[25]，使用 CLog-Log 模型对基础回归重新进行分析，该模型主要使用非对称的极值分布作为连接函数 $F(x, \beta)$，即：

$$Pr(y = 1|x) = F(x, \beta) = 1 - exp\left(-e^{x'\beta}\right)$$

这使得在 CLog-Log 模型中，事件发生概率 $p$ 趋向于 1 的速度要更快，与本文所提到的稀有事件的发生相呼应。

在指标构建上，考虑到前文主要采用均值法将四个维度集合成总的生存预期，可能造成信息的冗余和重复，此处重新采用主成分分析法（PCA）进行信息浓缩和降维，重新构造居民生存预期指标并放入模型进行分析，以检验指标构建的合理性。

商业保险在中国的发展时间较短，在受教育水平不同的人群中，现代风险分摊意识和传统"家庭养老"观念的结合程度不尽相同；另外，感性视角下的生存预期是一种瞬时的心理乐观状态变动，而不同文化素养水平的居民在面临同样的外界冲击时其心理的波动程度是不同的，即受访者很可能是因为自身受教育水平较高而直接影响了其对于商业险的需求。综合上述二者，本文认为，对于问卷受访者而言，大专或本科及以上为高学历水平，在排除此类数据的干扰后，重新观测感性生存预期对于居民商业保险需求的影响，以验证基础回归结果的稳健型。

对于理性生存预期而言，该指标的构造来源于主观和客观两个维度，参考周烁和张文韬（2021）[26]，我们基于受访者的认知得分记录，剔除填写问卷真实性不高的样本。本文中，认知能力得分的取值范围在 0~43 分之间，样本数据均值为 23.50 分。我们设定，认知得分低于 10%分位数 10.5分的样本会出现回答与实际情况不符的情况，当高于此分数时，问卷数据真实性较高。

表 3 的第（1）、（3）、（5）列展示了上述 4 种方式下的稳健性检验结果，为重点突出，此处省略了模型控制变量回归结果的展示。不难看出，在更换了基本计量模型、使用主成分分析法重新构造总生存预期指标，以及剔除高学历水平和低认知得分样本后，参数估计结果依然具有极强的显著性，与本文基准回归结果保持了一致。当然，如果对感性视角生存预期剔除认

知得分过低样本、对于理性视角预期得分剔除高学历样本，其结果依然具有稳健性，此处不再展示。

表 3 生存预期与商业医疗保险购买意愿的稳健性检验结果

| | （1） | （2） | （3） | （4） | （5） | （6） |
|---|---|---|---|---|---|---|
| | CLog-Log | EProbit | CLog-Log | EProbit | CLog-Log | EProbit |
| 生存预期 | 0.290[***]<br>(0.06) | 2.049[***]<br>(0.31) | | | | |
| 感性视角 | | | 0.299[***]<br>(0.07) | 1.437[***]<br>(0.01) | | |
| 理性视角 | | | | | 0.385[***]<br>(0.12) | 2.023[**]<br>(0.80) |
| 控制变量 | Y | Y | Y | Y | Y | Y |
| 样本量 | 10078 | 10078 | 9800 | 10078 | 9067 | 10078 |

注：括号里为稳健标准误；*代表 p＜0.10，**代表 p＜0.05，***代表 p＜0.01；此处分析省略了控制变量的展示。下表同。

最后，尽管表 2 显示我国中老年人的生存预期与商业医疗保险需求存在显著的正向相关关系，但是两类潜在的内生性问题可能导致回归结果的偏差：遗漏变量问题和反向因果问题。对于第一类情况，本文除了基于已有研究加入尽可能多的控制变量，还依据现实环境变化和实际研究问题对常规的控制变量进行了改进以求结果的准确；另外，本文使用熵权法将问卷中反映受访者个人身体状况的 46 种问题集成在居民生存预期的理性视角下，尽可能地解决遗漏变量问题。

对于第二类情况，由于其导致的内生性问题相对复杂，我们尝试使用工具变量法来解决。参考邹小芃和杨芊芊（2019）[16]，我们选取受访者父母寿命的均值作为工具变量，以检测核心解释变量可能存在的内生性问题。理由在于，对于受访者而言，其年龄在 45 周岁以上，基本能够见证其父母辈中年和老年的生命历程，其健康状况和（预计）寿命长度往往对受访者在评价自身生存预期时起到参考和修正作用，但是，父母年龄的均值和受访者的保险需求没有直接的联系，因此可以作为工具变量。由于感性视角生存预期为有序取值，而其他两个为连续取值，故而选择 Eprobit 方法，其结果如表 3 第（2）、（4）、（6）列所展示，边际影响显著为正，与基准回归结果一致。经过一系列检验可以看出，本文基准回归的核心结论是稳健的，接下来本文将从家庭成员和受访者个人生活状态角度对上述结果进行进一步的分析。

## 四、进一步分析

### （一）养老观念与遗赠动机

广义来讲，包括购买商业医疗保险在内的健康管理行为都属于"自我养老"范畴，而受传统思想文化影响，我国中老年人主要秉持"家庭养老"的观念 [1]，正常来讲，子女数量越多，中老年人的家庭养老可靠性预期就会越强。可以说，子女的数量和行为往往会对受访者的商业医疗保险购买决策以及健康管理意愿产生重要的协同影响。此外，受访者对于家庭成员尤其是子女的遗赠动机也是值得考虑的重要因素。Kopczuk 和 Lupton（2007）[30]认为，父母的遗赠倾向来源于利己主义动机，给予子女财富能增加其自身的效用。根据张艳和杨德才（2021）[31]的总结，为了能够提高子女的福利水平和幸福感，父母甚至不惜省吃俭用以留下更多遗产，该行为主要受我国儒家思想的影响，因此遗赠动机的成因往往是"给予的快乐"，即父母对子女福利的无私关注，而非为换取子女的赡养（Dalgaard 和 Jensen，2009）[32]。

综上而言，无论是从家庭养老的角度，还是从遗赠动机的角度，受访者的子女都应当是分析个人商业医疗保险需求乃至家庭资产配置的重要因素。参考 Laitner 和 Sonnega（2010）[33]，受访者自身的遗赠动机可以通过其子女数量来展现。因此，本文用受访者子女数量作为家庭养老可靠性预期和遗赠动机的代理变量，将其与生存预期组成交叉项进行调节效应分析，具体结果如表 4，为避免多重共线性影响，并使得系数效应一致，交叉项测量的变量均已做中心化处理。

表 4 子女数量的调节效应结果

|  | (1) | (2) | (3) |
|---|---|---|---|
| 生存预期 | 0.336[***] (0.12) | | |
| 子女数量*生存预期 | -0.190[**] (0.08) | | |
| 感性视角 | | 0.213[***] (0.08) | |

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1 一般而言，我国的养老模式分为：社会养老、家庭养老和自我养老（陈友华，2012；陈东和张郁杨，2015）[27][28]，其中，社会养老对于传统的家庭养老模式存在有限的替代效应（张川川和陈斌开，2014）[29]。此处的家庭养老包含了中老年人在遭遇健康危机（如生病、摔倒、意外、日常起居困难等）时，子女对其的照顾护养和经济支持。

| | | | |
|---|---|---|---|
| 子女数量*感性视角 | | -0.078<br>(0.06) | |
| 理性视角 | | | 0.248$^*$<br>(0.13) |
| 子女数量*理性视角 | | | -0.273$^{***}$<br>(0.09) |
| 子女数量 | -0.097$^{**}$<br>(0.05) | -0.107$^{**}$<br>(0.05) | -0.102$^{**}$<br>(0.05) |
| 控制变量 | Y | Y | Y |
| 常数项 | -6.410$^{***}$<br>(1.99) | -6.478$^{***}$<br>(2.03) | -6.165$^{***}$<br>(1.98) |

表 4 模型（1）分析了家庭子女因素对于受访者总的生存预期与商业医疗保险需求关系的调节效应。结果显示：受访者子女数量与受访者总的生存预期的交互项系数在 5%水平下显著为负，这说明随着受访者子女数量的增加，其家庭养老的可靠性预期不断上升，同时受访者的遗赠动机也越强烈，这将会削弱总的生存预期变动时受访者对于购买商业医疗保险的动机和需求程度。表 4 模型（2）分析了家庭子女因素对于受访者感性视角下生存预期与商业医疗保险需求关系的调节效应，其结果并不显著。而表 4 模型（3）分析了家庭子女因素对于受访者理性视角下生存预期与商业医疗保险需求关系的调节效应，其结果显示：受访者子女数量与受访者理性视角下的生存预期的交互项系数在 1%水平下显著为负，削弱效应强烈。综合分析来看，在生存预期变动时，家庭子女数量对于受访者的商业医疗保险需求变化起到削弱作用，且这种削弱作用主要来自于从理性视角，这意味着家庭养老观念在我国中老年人心中根深蒂固的影响和地位，也表明着遗赠动机的存在，使得我国中老年人对于子女福利的关注始终优先于对自身健康预期的关注。

（二）社交程度与生活感受

从家庭成员方面来看，它对于受访者生存预期与商业医疗保险需求关系的影响主要体现在调节效应上，即属于居民需要事前充分考虑的重要因素。从个人角度来说，在充分考虑了各项指标后，个人行为与心理的变化往往构成了受访者商业医疗保险需求变动的影响途径。

1.社交积极性

人拥有永恒的社会属性，居民在社交过程中交换信息、提高认知、完善自我并实现价值。当然，随着 5G 技术的进步，居民日常进行社交的方式有了更加多元化的发展，例如浏览新闻短视频、参加在线讲座等新方式为居民带来了自由、灵活的社交新体验。可以说，居民的社交积极性对保持

其自身认知和生活方式与时俱进具有重大意义。

表 5 生存预期、社交积极性和商业医疗保险需求

| | （1） | | （2） | | （3） | |
| | （1a） | （1b） | （2a） | （2b） | （3a） | （3b） |
| | OLS | CLog-Log | OLS | CLog-Log | OLS | CLog-Log |
| 生存预期 | 0.228*** (0.02) | 0.413*** (0.10) | | | | |
| 社交积极性 | | 0.183*** (0.03) | | | | |
| 感性视角 | | | 0.119*** (0.02) | 0.241*** (0.07) | | |
| 社交积极性 | | | | 0.186*** (0.03) | | |
| 理性视角 | | | | | 0.235*** (0.03) | 0.357*** (0.11) |
| 社交积极性 | | | | | | 0.187*** (0.03) |
| 控制变量 | Y | Y | Y | Y | Y | Y |
| 中介效应占比 | 9.18% | | 8.41% | | 10.96% | |

在此基础上，本文尝试通过受访者参与社交的途径数量来衡量居民日常社交的积极性，并将其放入基准回归框架中进行中介效应分析。根据DA056问题"您过去一个月是否进行了下列社交活动（可多选）"所提供的11个选项进行统计，可以得出受访者在过去一个月参与社交途径的数量。表5汇报了具体结果。其中，（1a）、（2a）和（3a）列的因变量是受访者社交积极性，使用OLS进行回归；（1b）、（2b）和（3b）列的因变量是受访者购买商业医疗保险意愿，使用CLog-Log模型进行回归。回归结果显示，无论是感性视角还是理性视角，在其他条件不变情况下，受访者生存预期的改善可以显著地提高其本人的社交积极性，进而带来信息获取渠道的拓展以及自身认知水平的提高，受访者主动进行健康管理的意识会在和他人交流与对比中得到加强与改进，因此他对于商业医疗保险的需求会加强。可见，居民的社交积极性在中老年人生存预期对于购买商业医疗保险的意愿影响中具有显著的中介效应。

表 6 生存预期、生活满意度和商业医疗保险需求

| | （1） | | （2） | | （3） | |
|---|---|---|---|---|---|---|
| | （1a） | （1b） | （2a） | （2b） | （3a） | （3b） |
| | OLogit | CLog-Log | OLogit | CLog-Log | OLogit | CLog-Log |
| 生存预期 | 1.344*** | 0.399*** | | | | |
| | (0.04) | (0.11) | | | | |
| 生活满意度 | | 0.138** | | | | |
| | | (0.07) | | | | |
| 感性视角 | | | 0.831*** | 0.226*** | | |
| | | | (0.03) | (0.07) | | |
| 生活满意度 | | | | 0.156** | | |
| | | | | (0.07) | | |
| 理性视角 | | | | | 1.019*** | 0.345*** |
| | | | | | (0.05) | (0.12) |
| 生活满意度 | | | | | | 0.169*** |
| | | | | | | (0.07) |
| 控制变量 | Y | Y | Y | Y | Y | Y |
| 中介效应占比 | 31.73% | | 36.45% | | 33.30% | |

2.生活满意度

从个人的心理角度变化来说，其向外最直观的表达就是居民对于生活的满意度。根据DC028问题"总体来看，您对自己的生活是否感到满意？"整理生活满意度变量，对一点也不满意赋值为1；不太满意赋值为2；比较满意赋值为3；非常满意赋值为4；极其满意赋值为5，并使用Ordered-Logit模型和CLog-Log模型进行中介效应的回归分析。

表6汇报了具体结果，其中，（1a）、（2a）和（3a）列的因变量是受访者生活满意度；（1b）、（2b）和（3b）列的因变量是受访者购买商业医疗保险意愿。回归结果显示，无论是感性视角还是理性视角，在其他条件不变情况下，受访者生存预期的恶化都会显著地降低其自身的生活满意度看法，进而会对商业医疗保险的需求产生抑制，这与基准回归部分所述相契合。当然，如果受访者的生存预期得到改善，未来可期的心态会改善对自己生活满意度的评价。受访者自我判断良好，健康的心理暗示加强，为使自己在可预见的未来依旧能够获得较高的幸福感，他就不会抗拒外界对于商业医疗保险的推广和宣传，同时对于商业医疗保险的需求也会增强。可见，

居民的生活满意度在中老年人生存预期对于购买商业医疗保险的意愿影响中同样具有显著的中介效应。

## 五、异质性分析

### （一）居住地段与环境认知

相比于户籍差异，居民居住地的不同往往代表了不同的社交圈，而如上文所分析，不同社交圈各异的认知特点会影响当地住户在融入社区的实践过程中获得的信息内容，这更会对居民个人的商业医疗保险需求产生不同的影响。在 CHARLS 问卷中，居民的居住地类别主要包括城市、城乡结合部（城郊）、乡镇和特殊区域，由于特殊区域样本量极小，我们只对前三类进行分样本回归，结果如表 7 所示。

表 7 居民生存预期对商业医疗保险需求的影响：居住地分类

| 模型：CLog-Log | 全样本 | 城市 | 城郊 | 乡镇 |
|---|---|---|---|---|
| 生存预期 | 0.489*** | 0.713*** | 0.689** | 0.254* |
|  | (0.10) | (0.16) | (0.30) | (0.15) |
| 感性视角 | 0.287*** | 0.392*** | 0.286 | 0.187* |
|  | (0.07) | (0.10) | (0.19) | (0.10) |
| 理性视角 | 0.426*** | 0.706*** | 0.899*** | 0.119 |
|  | (0.12) | (0.19) | (0.31) | (0.16) |
| 控制变量 | Y | Y | Y | Y |
| 样本量 | 10078 | 2318 | 951 | 6758 |

注：表格仅展示核心解释变量系数。下表同。

对于居住在城市的受访者而言，无论是感性视角还是理性视角，他们的生存预期都会显著地对于自身购买商业医疗保险意愿产生正向的影响，且系数本身均大于总体样本，显示出城市居民较高的财富水平和更加注重自身健康保障的认知特点。对于居住在城郊的受访者而言，虽然整体上，他们的生存预期会显著地对于自身购买商业医疗保险意愿产生正向的影响，但是其贡献力主要来自于其理性角度，这主要是因为市区内医疗资源云集，但是物价高、人口密集，看病的时间成本较高，而商业医疗保险往往具有免排队通道、VIP 预留病房和线上复诊等附加管家式服务，这对于生活节奏快但离城市中心较远的城郊居民具有较大价值。对于生活在乡镇的受访者而言，其整体生存预期的显著性影响则主要来自感性视角；从理性视角而言，闲余资金并不充裕加上医疗资源的相对匮乏，以及"小病

不医"的陈旧思维都导致了其生存预期的变动无法显著地影响个人商业医疗保险需求。

从整体上看，从城市到城郊再到乡镇的显著性是依次下降的，这表明生活环境的不同对于居民的健康认知确实会产生不同的影响：越是充裕的生活条件和环境，越是能够给居民带来积极的生存预期得分，进而增加居民对于商业保险的购买需求。

（二）消费水平与财富认知

随着社会经济的发展，我国居民的消费内容和选择得到了极大的丰富，其具体的消费行为和决策除了能够反映个人的财富和收入水平，还包含了居民个人的心理偏好和效用最大化。同时，考虑到中老年人的消费动机更多从家庭角度出发（即整体的效用最大化），而家庭的消费结构和偏好通常也会影响个人对于日常健康管理和消费决策的具体内容。所以，我们对受访者及其家庭的月平均消费水平按照 50%分位数（1500 元）和 75%分位数（3000 元）进行了分组。其中，消费在 1500 元以下可以定义为低收入人群，消费在 1500 元~3000 元区间可以定义为中收入人群，而 3000 元以上则属于高收入人群。回归结果如表 9 所示。

表 9 居民生存预期对商业医疗保险需求的影响：消费水平分类

| 模型：CLog-Log | （1） | （2） | （3） |
|---|---|---|---|
| 消费水平区间 | 1500 及以下 | (1500,3000] | 3000 以上 |
| 生存预期 | 0.349* | 0.386** | 0.629*** |
| | (0.19) | (0.17) | (0.17) |
| 感性视角 | 0.266** | 0.243** | 0.297** |
| | (0.12) | (0.11) | (0.12) |
| 理性视角 | 0.140 | 0.299 | 0.737*** |
| | (0.22) | (0.19) | (0.19) |
| 控制变量 | Y | Y | Y |
| 样本量 | 5104 | 3350 | 1624 |

如表 9 所示，对于家庭月平均消费水平在 1500 元及以下的受访者而言，当感性生存预期改善时，他们当然希望未来能够获得更好的商业医疗保障服务，因此对于商业医疗保险的需求会显著上升；但是对于这类受访者而言，他们本身属于低收入群体，衣食、水电等基本消费占据了家庭支出的大头，因此从理性视角来说，其自身生存预期变得乐观时，他们的注意力可能并不在自身的健康管理和医疗保障上，而是将对子女未来发展意义重

大的教育培训需求放在首位，故而回归结果并不显著；而对于中高收入群体来说，由于其家庭整体的消费内容已经满足日常需求，且消费弹性低于低收入群体，因此当生存预期发生改善，他们更愿意把眼光投向包括健康管理在内的提升生活安全感和幸福感的事情上，因此其商业医疗保险需求会出现显著上升，且消费能力越强，对于商业医疗保险的需求上升越多。

（三）人生阶段与责任认知

处于不同人生阶段的人群通常需要承担不同的个人，以及家庭责任，因此会具有不同的风险偏好和认知水平。另外，其人生进程和财富积累状态也有所不同，面对不同的外界冲击和自身身体或心理上的健康变化，处在人生不同阶段的居民的感性和理性生存预期也会产生不同的变化，因此有必要对 CHARLS 问卷中的受访者就不同年龄段进行分组回归分析。综合世界卫生组织的年龄划分标准以及我国《老年人权益保障法》相关内容，本文将受访者划分为中年人（45 岁~60 岁）和老年人（61 岁及以上）这 2 类人群，并在此基础上进一步细致区分，其分段回归结果如表 8 所示。

表 8 居民生存预期对商业医疗保险需求的影响：年龄段分类

| 模型：CLog-Log | 中年人 | | | 老年人 | |
|---|---|---|---|---|---|
| 具体年龄区间 | [45,50] | [51,55] | [56,60] | [61,65] | 65 以上 |
| 生存预期 | 0.613*** (0.18) | 0.035 (0.18) | 0.451 (0.30) | 1.142*** (0.30) | 0.490 (0.30) |
| 感性视角 | 0.340*** (0.11) | 0.053 (0.12) | 0.276 (0.21) | 0.634*** (0.20) | 0.192 (0.21) |
| 理性视角 | 0.573*** (0.21) | -0.060 (0.21) | 0.339 (0.28) | 1.180*** (0.34) | 0.720** (0.35) |
| 控制变量 | Y | Y | Y | Y | Y |
| 样本量 | 1690 | 2258 | 1590 | 1722 | 2802 |

根据分组回归结果来看，50 岁以下的中年人其生存预期与个人商业医疗保险需求具有显著的正向关系。但是随着年纪的增长，这种显著关系消失了，即 51 岁至 60 岁这段接近退休的时间中，受访者购买的商业医疗保险的意愿与自身生存预期变动并不相关。这从侧面反映出，51~60 岁年龄段受访者并没有将精力重心放在自身的健康管理上，而是更加关注个人事业上升阶段最后的冲刺，因为这能够为受访者背后的家庭（尤其是子女的成长）带来切实的幸福感提升。这再一次契合了上文"遗赠动机对于我国中老年人主动进行健康管理存在一定程度的抑制作用"的结论。只有帮助

这个年龄段的人群树立更合理、更科学的健康管理与财富保值升值关系的理念,才能合理地引导他们通过商业保险手段进一步减小健康和养老风险。

对于老年人来说,由于 60 岁之后普遍开始了退休生活,此时他们对于自身的健康管理愈发重视,因此总体上来说,其生存预期与购买商业医疗保险意愿显著正相关。不过由于市面上的商业医疗保险、健康保险的普遍首次承保年龄上限在 65 周岁,因此显著性主要集中在 61~65 岁区间内。由此可见,不同年龄段的群体,其各自的感性和理性生存预期对于商业保险需求影响的显著性是有差异的,在显著前提下其影响程度也各不相同。

## 六、研究结论

通过在已有主观和客观维度的基础上考虑时间因素,本文阐释了感性和理性视角下生存预期的差别,并使用 Firth-Logit 模型和 CLog-Log 模型依次检验和对比了二者对我国中老年人商业医疗保险需求的影响。

总体来看,居民的生存预期对于其购买商业医疗保险的意愿存在显著的正向的影响。其中,感性视角的生存预期易受外界冲击影响,引发正向连带效应;理性视角下则具有稳定性。但是,在其他条件不变的情况下,理性视角下生存预期变动带来居民商业医疗保险需求的变动却远超过感性视角。此外,在生存预期变动时,传统的家庭养老观念和遗赠动机对于受访者的商业医疗保险需求变化起到抑制作用,且这种削弱作用主要来自于从理性视角。而机制分析表明,社交积极性和生活满意度是重要的效应传导途径。最后,本文从居住地、年龄段和消费水平三个方面对上述结果进行了异质性分析。可知,现代化的生活环境、特定的人生阶段和充裕的物质条件是生存预期对中老年人商业医疗保险需求影响显著的重要条件。

依据本文研究成果,商业医疗保险的产品设计和推广销售应当更加考虑到不同人群不同的感性和理性生存预期特点,并设定更具有针对性和更加人性化的服务保障内容。例如,进一步扩展保单持有者每年健康体检的项目内容,以帮助客户提高对自身身体状况的了解,改善其理性视角下的生存预期,增加用户与产品粘性。同时,保险行业应当继续拓宽城乡结合部和乡镇的保险知识信息流通渠道,针对细分市场目标人群的特点,进行多样化的健康知识普及和风险分摊概念科普,合理地引导居民通过保险手段减小健康风险。当然,按照本文的指标构建逻辑,保险市场中常见的逆向选择问题存在于理性生存预期视角下,如何利用微观数据进一步给出具体的实证依据则是下一步的研究方向。

健康是促进人的全面发展的必然要求,是经济社会发展的基础条件。提高商业医疗保险参保意愿是稳定中老年人生存预期的重要手段之一,同时在维护家庭金融运行的稳定性方面具有重要意义。对于年轻人而言,这有助于缓解养老负担,提高生育意愿,增进生活幸福感;对于国家而言,

大力发展各类商业健康保险可以降低个人自付比例，缓解公共筹资比例扩张压力，减轻财政负担，具有非常重要的政策和现实意义。

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# 后记

　　2022 年 10 月 28 日至 29 日，由北京大学经济学院和北京大学中国保险与社会保障研究中心（CCISSR）主办的"北大赛瑟（CCISSR）论坛·2022（第十九届）"在北京大学以"线下+线上"方式举行。本届论坛的主题是"共同富裕与现代化国家建设"。

　　在 10 月 29 日上午大会的"主旨演讲"阶段，中国银保监会副主席梁涛以"推动中国式保险业现代化 为实现共同富裕保驾护航"为题，全国社会保障基金理事会副理事长陈文辉以"关于共同富裕与高质量发展的几点思考"为题，中国社会保险学会会长、人社部原副部长胡晓义以"共同富裕议题下的养老金水平"为题，分别发表了精彩的大会主旨演讲。在大会"主题演讲"阶段，嘉宾们围绕"共同富裕与现代化国家建设"发表了演讲。人力资源社会保障部农村社会保险司司长刘从龙以"在实施共同富裕战略中推进城乡居民基本养老保险高质量发展"为题，中国发展研究基金会秘书长方晋以"共同富裕：公平与效率的统一"为题，国家卫生健康委员会国家卫生发展研究中心副主任甘戈以"新时代公立医院高质量发展"为题，中国保险学会会长董波以"进一步推动保险业高质量发展 更好地服务中国式共同富裕现代化建设"为题，北京大学中国保险与社会保障研究中心主任郑伟以"中国式现代化对社会保障的新要求"为题，分别发表了精彩的大会主题演讲。演讲引起了线下线上参会代表的强烈反响和广泛好评。

　　在 10 月 29 日下午，"2022 中国保险业发展圆桌论坛"暨《中国保险业发展报告 2022》新书发布会举行。来自北京大学、清华大学、中国社科院、中国保险资产管理业协会、中国精算师协会、研究中心理事单位、研究报告支持单位等机构的嘉宾发表了演讲。郑伟代表报告撰写组介绍了刚刚出版的《中国保险业发展报告 2022》的基本背景和主要内容。接着，中国保险资产管理业协会执行副会长曹德云、清华大学中国保险与风险管理研究中心主任陈秉正、中国社科院金融研究所保险研究室主任郭金龙、中国人民健康保险公司副总裁李晓峰、中央财经大学保险学院院长李晓林、泰康保险集团副总裁黄新平、恒安标准养老保险公司董事长万群、上海商涌科技公司首席执行官邢静、国宝人寿保险公司监事长张锐、南开大学卫生经济与医疗保障研究中心主任朱铭来、对外经贸大学保险学院院长助理王国军、中国精算师协会会长王和、中国出口信用保险公司首席经济学家王稳、北京工商大学保险研究中心主任王绪瑾、中国人民大学财政金融学院保险系主任魏丽，分别发表了精彩演讲。随后，北京大学中国保险与社会保障研究中心名誉主任孙祁祥围绕论坛主题发表了精彩的闭幕演讲。

　　在大会之前的 10 月 28 日下午，北大赛瑟（CCISSR）论坛"专题学术

论坛"举行。本届论坛共设六个学术分论坛。山东财经大学于文广教授主持第一分论坛"保险与经济社会",北京大学陈凯副教授主持第二分论坛"保险业务经营",南开大学朱铭来教授主持第三分论坛"长期护理保险",中央财经大学杨再贵教授主持第四分论坛"老龄化与养老保险",北京大学姚奕副教授主持第五分论坛"健康保险",中南财经政法大学胡宏兵教授主持第六分论坛"保险与共同富裕"。二十余篇入选论文的作者在专题学术论坛上宣读了自己的论文,二十余位参会代表对论文进行了评议,大家就相关问题进行了热烈的交流和讨论。。

为了让更多的同仁分享 2022 年第十九届"北大赛瑟(CCISSR)论坛"的成果,我们将大会演讲和部分专题学术研讨论文结集成册,希望大家能从中获得一些有益的信息与启迪。

"北大赛瑟(CCISSR)论坛"已经成功举行了十九届,得到了许多方面人士的高度评价。回想起来,当初我们创办这个论坛时就是想在国内保险、社会保障和风险管理领域搭建一个规范、稳定、各方受益的学术交流平台。万事开头难,经过这几年的摸索和实践,"北大赛瑟(CCISSR)论坛"初步找到了一条既与国际接轨又符合中国背景的学术交流模式,并且得到了各方面的高度关注、肯定、鼓励和支持,对此我们倍感欣慰。

感谢本届论坛的大会演讲嘉宾的杰出贡献!感谢六个学术分会场的各演讲人对本届论坛的精彩奉献!感谢北京大学中国保险与社会保障研究中心的各理事单位对中心各项活动的热心支持和积极参与!我们相信,在社会各界的关心和支持下,"北大赛瑟(CCISSR)论坛"能够不辱使命,越办越好!

北大赛瑟(CCISSR)论坛组委会
2022 年 12 月 20 日于北京