近五年科研成果清单及代表性成果原件

申请人：王天宇 北京大学 应届经济学博士

科研成果清单：

<u>发表论文</u>

1. Xuezheng Qin, Tianyu Wang, Castiel Zhuang. “Intergenerational Transfer of Human Capital and Its Impact on Income Mobility: Evidence from China.” *China Economic Review*, in Press (DOI: 10.1016/j.chieco.2014.10.005).
2. 王天宇，彭晓博. 社会保障对生育意愿的影响——来自新型农村合作医疗的证据.《经济研究》2015(2).
3. 李绍荣，王天宇. 最优经济政策的性质——以货币政策为例.《经济理论与经济管理》2015(4).
4. 邱牧远，王天宇，黄娅娜. 社会网络的策略性生成：一个文献综述. 《经济问题探索》2015(5).

<u>工作论文</u>

5. Xi Chen and Tianyu Wang. Does Money Relieve Depression? Regression Discontinuity Evidence from the Pension Eligibility Age

   （Fifth Biennial Conference of American Society of Health Economics, Los Angeles, U.S.A, 2014,

   NBER Summer Institute, Boston, U.S.A 2014）
6. 秦雪征，王天宇. 人力资本的代际传导对社会流动性的影响：理论与实证的探究.《经济研究》工作论文.
7. 王天宇，邱牧远. 延迟退休年龄，养老金缺口与青年就业.
8. 王天宇，邱牧远. 放开计划生育政策能扭转人口老龄化趋势吗？——基于多期世代交叠模型的分析.
9. 邱牧远，王天宇. 计划生育，经济增长与收入不平等.
10. 黄娅娜，王天宇. 中国存在退休消费之谜吗？
11. 彭晓博，王天宇. 医疗保险与健康不平等——来自新农村合作医疗的证据.
12. 王天宇，冯驰. 创新，模仿与监管——对电子产品市场“山寨”现象的分析.

国家社会科学基金资助期刊

总第569期 第50卷

经济研究

ECONOMIC RESEARCH JOURNAL

高培勇：论完善税收制度的新阶段

杨汝岱：中国制造业企业全要素生产率研究

王天宇
彭晓博：社会保障对生育意愿的影响：来自新型农村合作医疗的证据

付明卫等：国产化率保护对自主创新的影响
——来自中国风电制造业的证据

第二届国家期刊奖

2013 · 中国百强报刊

2015 2

经济研究 （月刊） 2015年第2期（总第569期第50卷） 2月20日出版

目　录

创刊时间：
1955年4月
封面题字：
郭沫若
主管单位：
中国社会科学院
主办单位：
中国社会科学院经济研究所
主编：
裴长洪
社长：
王　诚
编辑者：
经济研究编辑部
值班编辑：
唐寿宁
出版者：
经济研究杂志社
地址：
北京阜外月坛北小街2号
邮政编码：
100836
电话：
(010)68034153
传真：
(010)68034985
网址：
www.erj.cn
www.cesgw.cn
电子信箱：
erj@cass.org.cn
定价：
45.00元

印刷者：
北京盛通印刷股份有限公司
国内总发行：
北京报刊发行局
国内邮发代号：
2-251
订购处：
全国各地邮局
国外总发行：
中国国际图书贸易总公司
国外发行代号：
M16
广告经营许可证：
京西工商广字0165号
封面及广告设计：
王红梅
常年法律顾问：
陆康（重光律师事务所）
本刊不以任何形式收取版面费
举报电话：010-63094651

社会保障对生育意愿的影响：来自新型农村合作医疗的证据*

王天宇　彭晓博

内容提要：社会保障制度对个体生育决策和总和生育率的深远影响早已被大量研究所证实，但由于人口政策对生育行为的限制，鲜有文献对中国情境进行探讨。本文考察了新型农村合作医疗制度的建立对居民生育意愿的影响。基于两期家庭决策模型的分析表明，带有补贴的新农合会对生育数量产生两种方向相反的效应：收入效应和挤出效应，前者导致生育意愿的提高，后者导致生育意愿的下降。利用中国健康与营养调查（China Health and Nutrition Survey，CHNS）2000—2009 年的数据，本文发现挤出效应占主导地位，参加新农合使居民想再要孩子的意愿降低了 3%—10%。据此，本文认为社会保障体系建设的持续推进将为放松人口政策提供空间，实现从强制少生到自愿适度生育的转变。

关键词：社会保障　生育意愿　计划生育　新型农村合作医疗

一、引　言

自上世纪 80 年代以来，中国政府实施了严格的计划生育政策，期望以此控制人口数量，提高人口素质。这一政策取得了显著效果，据估计，在过去 30 年中，中国共少生人口超过 4 亿，总和生育率由 20 世纪 70 年代初的 5.8 下降到 1.8 左右（国家人口发展战略研究课题组，2007）。但近年来，该政策引起了越来越多的争议：一是控制人口数量的目标和维护公民的生育权之间可能存在矛盾；二是计划生育政策对人口结构的改变所引发的工资上涨可能导致中国失去廉价劳动力优势；三是计划生育政策可能使青年人背负越来越沉重的养老负担。能否找到合适的替代政策工具，用经济的、自愿的方式控制人口数量，是摆在决策者面前的一道难题。本文试图探讨社会保障，尤其是农村医疗保障，是否能成为解决这一难题的备选答案。

在没有生育政策限制的社会里，人们生育的原因大致可以分为五种：（1）意外事件，在有效避孕措施发明之前，生育往往是不可控的；（2）繁衍责任，繁衍是一切物种的本能，人类也不例外，中国传统文化中重视家族的延续与壮大，强调“不孝有三，无后为大”，即可视为繁衍本能的体现；（3）情感需要，父母将孩子视为自己的情感寄托，从亲子关系中得到快乐；（4）社会压力，当周围的同辈群体都生育时，没有孩子的夫妇会受到质疑，承受压力；（5）经济工具，通过家庭内部的代际转移，孩子可以发挥储蓄和保险的功能。生育的经济因素是本文关注的重点。

在缺乏有效财富储蓄和保值途径的古代社会，生育子女并在子女成年后获得供养（通常是每个月固定数额的返还），从而达到平滑消费、维持生命的目的，这是生育的第一种经济功能——储蓄。当父母衰老后，如果身患重疾，子女通常要倾尽所有，挽救父母生命，这是生育的第二种经济功能——保险。这两种经济功能的实现，都必须依靠社会的隐性契约，隐性契约的执行则主要依靠道

*　王天宇，北京大学经济学院，邮政编码：100871，电子信箱：wangtianyu@ pku. edu. cn；彭晓博，中国社会科学院研究生院、中国社会科学院经济研究所，邮政编码：100836，电子信箱：xiaobopeng1212@ 163. com。王天宇感谢北京大学才斋奖学金的资助，彭晓博感谢国家社会科学基金重大项目（13&ZD042）和国家社科基金项目（14BGL145）的资助。感谢匿名审稿人提出的宝贵意见，文责自负。

德和习俗。中国的传统生育观念可以归纳为两点:养儿防老和重男轻女。养儿防老是典型的经济动机。重男轻女也可归结为经济上的考虑:女儿出嫁后经济、人身都已不再独立,很难在父母养老上做出和儿子相同的贡献。这两种生育观念都体现了上述两种经济功能。

生育意愿中的经济因素得到了中外大量生育调查的确认。例如,2002 年农村居民生育意愿调查显示,生育目的是“养儿防老”的占 28.0%(姚从容等,2010);Nugent(1985)对多个跨国观念调查做了总结,发现在世界各地养老均是生育的主要动机之一。

建立完善的社会保障体系,对生育意愿中经济动机的影响是多方向的。一方面,社会保障对“养儿防老”具有替代作用,家庭由“自我保险”模式转入“社会安全网”模式(秦雪征,2010),从而对最优生育数量产生负向影响;另一方面,社会保障放松了低收入家庭的预算约束,收入效应使得最优生育数量上升。目前的研究,无论是理论还是实证,大都支持社会保障对生育的替代作用。

1949 年以来,中国的社会保障体系建设主要集中于城市,尤其是城市中的国有部门,包括公费医疗、退休金等制度的建立。农村的社会保障主要依靠旧的农村合作医疗(以下简称“旧农合”)。旧农合始于 20 世纪 50 年代,60、70 年代得到了迅速发展,高峰期覆盖率曾达到 90%。70 年代末经济改革后,随着以生产队为基本单位的集体农业生产方式被家庭联产承包责任制所取代,旧农合也逐渐解体。到 1985 年,旧农合的覆盖率仅为 5% 左右。后经恢复,2003 年的覆盖率接近 20%。①新型农村合作医疗(以下简称新农合)从 2003 年开始试点,此后得到迅速推广。到 2008 年末,新农合基本上实现了全覆盖,成为中国农村居民的基本医疗社会保障制度。表 1 显示了从 2004 年到 2012 年的新农合开展情况。我们将新农合的特点总结为以下几点:(1)以家庭为单位自愿参加;(2)由中央财政和地方财政共同补贴,地方补贴力度差异很大;(3)以保大病、保住院为主;(4)统筹到县。以上几个特点中,以家庭为单位有利于克服逆向选择问题,也在一定程度上避免了参保的内生性问题;财政补贴强化了社会保障的收入效应,从而使得正向影响成为可能;保大病和保住院正是对养儿防老功能的替代;统筹到县相比旧农合的统筹到村、乡增强了社会保障网络对疾病风险,尤其是对传染病疫情的抵抗能力。

表 1　　2004—2012 年新型农村合作医疗开展情况

指标	2004	2005	2006	2007	2008	2009	2010	2011	2012
开展县(市、区)个数	333	678	1451	2451	2729	2716	2678	2637	2,566
参合人口数(亿)	0.80	1.79	4.10	7.26	8.15	8.33	8.36	8.32	8.05
参合率(%)	75.2	75.7	80.7	86.2	91.5	94	96	97.5	97.5

注:参合数据来自《中国卫生统计年鉴 2011》,《中国卫生统计年鉴 2012》,《中国卫生统计年鉴 2013》和《2013 年我国卫生和计划生育事业发展统计公报》。

中国的“一胎政策”从 1980 年开始推行,其强制性导致最优生育数量不再可测,但生育意愿仍然是可测的。生育意愿一般定义为在没有政策限制的情况,对生育数量、子女性别的期望。生育意愿无法代替总和生育率这一指标,但在一定程度上反映了人们的最优生育决策。近年来,中国开展的生育意愿调查主要有“2001 年全国计划生育与生殖健康调查”、“2002 年全国城乡居民生育意愿调查”、“2006 年全国人口和计划生育抽样调查”等(姚从容等,2010)。由北卡罗来纳州立大学人口中心主持的“中国健康与营养调查”(China Health and Nutrition Survey, CHNS)从 1989 年开始追踪调查中国妇女的生育观念和生育史,为我们的研究提供了宝贵的数据素材。

本文使用 CHNS 数据,评估了从 2000 年到 2009 年这一时间段内,新型农村合作医疗制度的建立对农村妇女(农村家庭)生育意愿的影响。二元选择模型和泊松模型的回归结果都表明,参合降

① 根据各年度《中国卫生统计年鉴》整理。

低了农村家庭的生育意愿。用 2000 年和 2004 年的数据进行双重差分(Difference-In-Difference)和双重差分-倾向得分匹配(DID-Matching)估计的结果表明,我们的结论是稳健的。

本文接下来的安排是:第二节简要回顾与本研究相关的文献;第三节建构了一个简单的理论模型说明新农合对生育意愿的影响机制;第四节介绍所用数据和计量方法,并对相关变量进行描述性统计分析;第五节报告主要回归结果并进行稳健性检验;第六节总结全文,并提出相应的政策建议。

二、文献综述

经济学中对生育和人口的关注可以追溯至古典时代。早在 1798 年,Malthus 的人口理论就从宏观和历史的视角对生育行为、人口增长和社会发展之间的关系做出了大胆预测。但在新古典范式内为生育决策分析建立起正式微观基础则要归功于 Becker(1960),他首次在效用最大化的框架下讨论了生育数量、生育质量、人力资本投资等与生育行为密切相关的家庭决策。Becker 的标准模型(Becker,1960;Becker & Lewis, 1973;Becker & Tomes,1976)中利他机制是单向的。父母将孩子视为消费品,生育的主要动机是孩子的数量和质量带来的直接满足感。后续的双向利他(two-sided altruism)模型(Kimball,1987;Hori & Kanaya,1989)则强调孩子对父母的收入返还,生育的动机由消费变为消费和投资。无论是单向利他模型还是双向利他模型,上述研究都将生育视为家庭的统一决策。以 Manser & Brown(1980)为开端的另一些研究将家庭内部的一系列决策视为内部谈判的结果。Eswaran(2002)的模型将生育数量视为夫妻的纳什谈判解,强调双方的谈判力量(Bargaining Power)在生育决策中的作用。两种生育模型都得到了一些实证研究的支持(Hotz et al.,1997;Holger,2006)。

针对社会保障对生育数量影响的理论研究主要是在家庭效用最大化的新古典框架下展开的。Barro & Becker(1989)在单向利他假设下初步探讨了社会保障税率的提高对生育数量的负向影响,后续研究大多得到了类似结论(Nishiruma & Zhang,1995;Rosati, 1996; Wigger,1999)。另一些研究在双向利他的假设下考察社会保障水平的提高对生育的影响(Ehrlich & Lui,1998; Baldwin & Jones,2002),发现了更为显著的替代作用。必须要指出的是,上述研究的结论都是在现收现付制(Pay-As-You-Go)或完全累积制(Fully Funded)的社会保障制度下得出的,并未考虑到类似于新农合这样针对特定人群、享受国家补贴的社保政策所带来的收入效应。

社会保障与生育关系的经验研究可分为两类。一是以国家或地区为观测值的跨国研究,数据形式为横截面或面板。Holm(1975)以包含 67 个国家的跨国数据研究了社会保障项目对总和生育率的影响,得出了显著的负影响结论,作者认为社会保障对生育的影响强度并不亚于教育、收入等传统因素。Boldrin et al.(2005)对 104 个国家 1997 年的社会保障税率①与总和生育率之间关系的考察也得出了类似结论。Holmqvist(2011)专门针对撒哈拉以南非洲国家 1960—2007 年的数据作了研究,发现带补贴养老金制度的引入使得每个妇女少生 0.5—1.5 个孩子。二是以某一国家或地区长时段数据为基础的时间序列研究。例如,Cigno et al.(2002)对德国时间序列数据的研究表明,社会保障覆盖率的提高对生育产生了负向影响;针对 1931—1984 年间的意大利时间序列数据的研究则表明人均养老金数额每增加 10%,总和生育率下降 0.02,这一效果是相当微弱的(Cigno & Rosati,1992)。

与本文相关的另一类文献涉及到新农合对农村居民工作、消费、迁徙等行为的影响。甘犁等(2010)应用 CHNS 数据研究证明政府对新农合的补贴将撬动 2.36 倍的农村居民消费增长。白重恩等(2012)用 2003—2006 年的农村定点观察数据所做的分析也肯定了新农合对拉动农村居民消费的正向作用,并认为这种影响在收入较低或健康状况较差的家庭中更强。Bai & Wu(2014)则在

① 社会保障税率的定义为社会保障占 GDP 的百分比。

理论模型中将新农合对消费影响分解为收入效应和预防性储蓄的改变两种效应，并提供了经验证据。秦雪征和郑直(2011)用 CHNS 数据证明了新农合在异地参与和就诊方面的歧视政策导致了劳动力流动的"枷锁效应"和"拉回效应"。齐良书(2011)用 2003—2006 年覆盖全国 30 个省区的微观面板数据，发现新农合在降低贫困发生率、增加中低收入农户收入和降低村庄内不平等方面发挥了重要作用。从 2003 年试点以来，新农合在促进农村居民健康、保障农民病有所医等方面的实施效果一直有所争议(朱恒鹏,2009)，但作为中国农村居民覆盖面最广的社会保障体系，其对参保人群的重大经济行为和家庭决策的深刻影响，已经得到了上述研究的肯定。

三、理论模型

我们考虑一个两期模型。该模型在 Bai & Wu(2014)的框架下引入生育的养老功能。第一期父母取得确定性收入 y_1，选择消费 c_1，储蓄 s，生育子女的数量 n，做出新农合参保决策 $I=\{0,1\}$，并且支付子女的抚养费用 $\phi(n)$。假设 $\phi'(n)>0$，$\phi''(n)\leqslant 0$，即抚养子女的边际成本递减或不变，有一定的规模效应。购买保险的费用为 e。储蓄利率为 r。

第二期父母取得确定性收入 y_2 和储蓄本利 $s(1+r)$，选择消费 c_2，并支付医疗保险和子女转移支付之余的医药费缺口。我们假定父母在第二期的医疗费用为随机变量 M。医保的报销额度为 $r(M)$。除了医保报销之外，子女也会承担一部分的医疗费用，我们假定子女承担的医疗费用为 $f(n,M)$。假定 $f_M(n,M)\geqslant 0$，$f_n(n,M)\geqslant 0$，即子女承担的医药费随着总医疗费用的增加和子女数量的增加而增加。

父母的终生效用为 $u(c_1,c_2,n)$，消费和子女都被视为正常品，边际效用为正。子女数量直接进入效用函数，对应着"多子多福"的生育动机。

父母的效用最大化问题表述如下：

$$\underset{c_1,c_2,n,I}{Max}\ E_1 u(c_1,c_2,n) \tag{1}$$

$$s.t.\quad c_1+s+\phi(n)+eI=y_1 \tag{2}$$

$$c_2+M-Ir(M)-f(n,m)=y_2+(1+r)s \tag{3}$$

当医疗保险不存在($I=0$)时的最优生育数量为 $n(0)$。当有保险覆盖，且父母选择参保($I=1$)时的最优生育数量 $n(1)$。

假设两者的差 $\Delta(n)=n(1)-n(0)$，则可将 $\Delta(n)$ 分解成下列形式：

$$\Delta(n)=\underbrace{h(e-E(r(M)))}_{\text{收入效应}}+\underbrace{n(1)-n(0)-h(e-E(r(M)))}_{\text{医疗保险对生育的挤出效应}}。\tag{4}$$

其中，$h(\cdot)$为收入效应，是保险的期望支付 $E(r(M))$和保费 e 的差额的增函数。若保费 e 为公平保费，则 $e=E(r(M))$，收入效应 $h(0)=0$。

下面我们在一个不失一般性的例子中探讨 $\Delta(n)$的符号和比较静态。

我们选用对数形式，可分离的效用函数，即

$$u(c_1,c_2,n)=\ln(c_1+\rho n)+\beta E(\ln c_2) \tag{5}$$

其中，ρ 是孩子数量相对于消费的权重，β 为主观折现因子。

抚养子女的边际成本为常数 k，即 $\phi(n)=nk$。

父母在第二期以概率 p 患上大病，需支付医疗费用 m；以概率 $1-p$ 保持健康，需支付医疗费用 0。若父母患病且购买了新农合，则父母可报销 $1-\alpha$ 比例的医药费。此处的假设对应新农合保大病不保门诊的政策。假设 $f(n,m)=\tau nm$，即子女承担的医疗费用随着子女数量线性增长。我们同时假设 $\tau n<\alpha$，即子女承担的部分并不足以支付保险未能覆盖的医疗费用。首先，农村地区收入偏低，子女的收入水平不足以应付父母的大病开支(癌症等致命疾病和糖尿病等慢性病往往伴随着

巨额医疗费用);其次,新农合对大病的报销比例不高,仍然有相当大比例[①]的医疗费用需要由患者自行支付。我们将储蓄利率 r 设为 0,一方面,这是简化模型的需要,另一方面,农村缺乏储蓄投资渠道,储蓄养老的收益率低于养儿防老,这一假定是合理的。另两个假定是 $k>\rho$ 和 $k<\rho+\tau pm$,前者可以理解为养育子女在第一期所能取得的回报无法抵偿边际抚养成本,后者意味着子女作为消费品和投资品的总回报大于抚养成本。

由一阶条件可解出最优生育数量:

$$n(0)=\frac{(k-\rho-\tau mp)\beta p\tau(y_1+y_2)+(k-\rho)[p(k-\rho)-\tau mp-(1-p)\beta p\tau m]}{(k-\rho)\tau[(1-\beta)(k-\rho)-\tau m(1+\beta-2\beta p)]} \tag{6}$$

$$n(1)=\frac{(k-\rho-\tau mp)\beta p\tau(y_1+y_2-e)+(k-\rho)[\alpha p(k-\rho)-\tau mp-\alpha(1-p)\beta p\tau m]}{(k-\rho)\tau[(1-\beta)(k-\rho)-\tau m(1+\beta-2\beta p)]} \tag{7}$$

生育数量的改变为:

$$\Delta(n)=\underbrace{\frac{[(1-\alpha)pm-e]\beta(k-\rho-\tau pm)}{(k-\rho)[(1-\beta)(k-\rho)-\tau m(1+\beta-2\beta p)]}}_{\text{新农合对生育的收入效应}}+\underbrace{\frac{-(1-\alpha)(k-\rho-\tau m)(k-\rho+\beta\tau pm)}{(k-\rho)\tau[(1-\beta)(k-\rho)-\tau m(1+\beta-2\beta p)]}}_{\text{新农合对生育的挤出效应}} \tag{8}$$

因为目前新农合由中央财政和地方财政大力补贴,因此保费 e 小于公平保费 $(1-\alpha)pm$,收入效应为正。新农合对生育的挤出效应则为负,一部分子女养老功能由社会保障代替。收入效应和挤出效应的符号相反,使得总体效果 $\Delta(n)$ 的符号并不明确。检验 $\Delta(n)$ 的符号是本文实证部分的首要任务。

对 $\Delta(n)$ 做简单的比较静态分析可知:$\partial\Delta(n)/\partial e<0$,即新农合降低生育意愿的效果会随着保费的提高而提高,其作用原理是保费的提高减弱了收入效应;$\partial\Delta(n)/\partial\alpha>0$,即新农合降低生育意愿的效果会随着自付比例的提高而降低,自付比例的提高同时减弱了收入效应和挤出效应,但收入效应的改变更大。

当然,在新农合覆盖时,家庭必须比较间接效用函数 $V(n(0))$ 和 $V(n(1))$ 才能做出是否参合的决策。但是由于保费 e 远小于公平保费 $(1-\alpha)pm$,我们有理由相信在相当宽松的参数区间内,参合是最优决策。从新农合覆盖地区极高的参合率和极低的退合率来看,这种假定也是合理的。

四、数据和实证方法

(一)数据

本文的数据来自中国健康与营养调查(CHNS),该调查由北卡罗来纳州立大学人口中心开展实施,已经于 1989、1991、1993、1997、2000、2004、2006、2009 和 2011 年展开了 9 次调查,以此为基础形成了 4400 个家庭及其所在社区的面板数据,追踪了过去 20 年中国城镇和农村居民的营养、健康以及经济状况的变化。该调查已经被众多针对新型农村合作医疗的研究所采用,例如 Lei & Lin (2009)、甘犁等(2010)和马双等(2010)。本文选取 2000—2009 年共计四次的调查数据进行研究。

CHNS 个人问卷询问了是否参加农村合作医疗,2009 年以前未区分旧农合与新农合;社区分卷

① 新农合的报销比例因就诊医院级别和就诊形式(门诊、住院、大病)而有所差异。以住院补偿为例,镇卫生院、二级医院、三级医院的报销比例分别为 60%、40%、30%。

中询问了本社区是否开展农村合作医疗以及开始时间。仿照其他研究者的处理方法（Lei & Lin，2009；秦雪征、郑直，2011），将所在县（市）有社区开展农村合作医疗并且开始时间晚于2003年视为该县（市）已经开展了新农合，将个人参加合作医疗同时所在县（市）已经开展新农合的观测值视为加入新农合。这种处理方法只适用于2004年和2006年，在2004年以前的调查中，所有参加合作医疗的均为旧农合；在2009年的调查中，所有参加合作医疗的均为新农合。由于新农合的参合者在有些地区并不必须拥有农村户口①，我们的样本包括CHNS所有农村居民点的调查对象。另外，我们注意到退出新农合的现象极少发生，在整个样本中，仅有54个家庭有过退出新农合的行为。我们从样本中删除了这一小部分退合家庭。同时，也删去了仅仅出现在2000年的家庭。

CHNS从1991年到2009年都对生育观进行了调查。调查对象包括在婚、丧偶和离婚的所有52岁以下妇女。调查面向三类人群：正在怀孕妇女；无子女且未怀孕妇女；有女子且未怀孕妇女。每类人群各有两个问题：第一个问题是如果可以自由选择生育数量，是否还想要孩子；第二个问题是如果还想要孩子，想要孩子的数量是多少。我们将生育意愿归纳为两个变量：一是二元变量——是否还想要孩子；二是计数变量——还想要几个孩子。在计数变量中，不想要孩子等同于还想要孩子的数量为0。表2显示了生育意愿在不同年龄群体中的分布。

表2　不同年龄组别已婚妇女生育意愿

年龄组	还想要的孩子数量				
	0	1	2	3	大于3
25岁以下	190	125	48	2	2
25—30	473	150	34	4	0
30—35	838	122	22	2	2
35—40	1139	88	20	3	1
40—45	1263	37	10	3	0
45岁以上	1487	15	7	3	1

除此之外，我们选取其他对生育意愿有影响的因素作为控制变量，主要包括以下几方面：一是妇女自身变量，如婚姻状况、年龄、教育程度、自评健康②、职业、收入③和是否有其他健康保险；二是与丈夫有关的变量，包括兄弟姐妹数量、教育程度、自评健康、职业、收入等；三是家庭层面变量，主要包括已有子女数量和夫妻双方共有几位老人在世。

表3根据受访者的保险状况报告了样本的描述统计特征，包括全体样本、参合者样本和非参合者样本的均值及标准差，同时我们计算了参合者和非参合者各变量的均值差异。由表3可知，有839个（47.9%）样本参加了新农合。

从表3来看，8%的参合者还想再要一个孩子，低于非参合者（10%）；同时，参合者还想要孩子的数量平均为0.11，而非参合者为0.125。这两个结果表明，参合者的生育意愿低于非参合者，这也在一定程度上预示了我们的实证分析结果。

参合家庭的妇女，在年龄、收入、教育和私人部门工作比例上都显著高于非参合家庭妇女，但在自评健康、兄弟姐妹数量和拥有其他健康保险的比例上则低于非参合家庭妇女。

参合家庭的丈夫，在收入和私人部门工作比例上也显著高于非参合家庭的丈夫，但在兄弟姐妹数量、教育上则与非参合家庭无显著差异。从生育孩子数量来看，参合群体只有不到一个（0.714），显著低于非参合群体。参合群体和非参合群体的家庭养老负担基本相当，每个家庭夫妻双方共有2.3个在世的老人。

① 例如，安徽省阜阳市规定农、林、牧、渔场工人均可参加新农合。

② CHNS2009年问卷中的自评健康一项略有调整，改为"你认为现在的生活怎么样"，答案分为很好、好、中等、差和很差五个等级。此问题出现在问卷"疾病史"部分。我们询问了CHNS官方，得到反馈确认此问题与往年自评健康相似。为保证数据取值一致，我们将"差"和"很差"合并为"差"。

③ 收入已经按照CHNS发布的折算指数统一折算为2009年水平。我们根据全年总收入和全年劳动时间计算了小时工资。由于劳动时间的统计误差较大，我们去除了妇女及其丈夫小时工资最高的1%观测值。

表 3 主要变量描述性统计

变量	(1) 全样本	(2) 参合者	(3) 非参合者	(4) (2)—(3)
关键变量:				
是否想要孩子	0.093 (0.290)	0.080 (0.271)	0.100 (0.300)	-0.020 (0.013)
还想要几个孩子	0.119 (0.475)	0.110 (0.479)	0.125 (0.472)	-0.016 (0.022)
是否参加新农合(1=是 0=否)	0.425 (0.494)	— —	— —	
育龄妇女变量:				
年龄	39.661 (7.315)	40.337 (6.849)	39.161 (7.614)	1.176*** (0.333)
结婚次数大于 1	0.001 (0.023)	0.000 (0.000)	0.001 (0.030)	-0.001 (0.001)
兄弟姐妹数量	3.708 (1.837)	3.584 (1.850)	3.801 (1.824)	-0.218** (0.084)
小时工资对数	1.976 (1.428)	2.344 (1.409)	1.706 (1.382)	0.638*** (0.064)
正规教育年限	6.373 (3.266)	6.625 (3.202)	6.185 (3.296)	0.440** (0.148)
自评健康:好	0.132 (0.339)	0.109 (0.311)	0.150 (0.357)	-0.041** (0.015)
自评健康:一般	0.470 (0.499)	0.418 (0.493)	0.508 (0.500)	-0.090*** (0.023)
从事农业工作	0.571 (0.495)	0.674 (0.469)	0.495 (0.500)	0.179*** (0.022)
在国有部门工作	0.136 (0.343)	0.020 (0.141)	0.221 (0.415)	-0.200*** (0.015)
在私营部门工作	0.123 (0.328)	0.144 (0.352)	0.107 (0.309)	0.038* (0.015)
有其他健康保险	0.070 (0.256)	0.014 (0.119)	0.112 (0.316)	-0.098*** (0.011)
丈夫变量:				
兄弟数量	1.701 (1.383)	1.751 (1.441)	1.665 (1.340)	0.085 (0.063)
姐妹数量	1.732 (1.321)	1.749 (1.335)	1.716 (1.311)	0.034 (0.060)
小时工资对数	2.195 (1.575)	2.580 (1.548)	1.914 (1.536)	0.666*** (0.070)

续表 3

变量	(1) 全样本	(2) 参合者	(3) 非参合者	(4) (2)—(3)
正规教育年限	8.007 (2.654)	8.100 (2.727)	7.939 (2.600)	0.161 (0.121)
自评健康：好	0.173 (0.378)	0.126 (0.333)	0.207 (0.405)	-0.080*** (0.017)
自评健康：一般	0.501 (0.500)	0.471 (0.499)	0.523 (0.500)	-0.052* (0.023)
从事农业工作	0.529 (0.499)	0.648 (0.478)	0.442 (0.497)	0.206*** (0.022)
在国有部门工作	0.167 (0.373)	0.039 (0.195)	0.260 (0.439)	-0.221*** (0.016)
在私营部门工作	0.214 (0.410)	0.245 (0.430)	0.192 (0.394)	0.052** (0.019)
家庭变量：				
一共生育的孩子数量(包括怀孕)	1.284 (1.094)	0.714 (1.033)	1.707 (0.935)	-0.993*** (0.045)
养老负担(夫妻双方在世的父母数量)	2.382 (1.336)	2.370 (1.369)	2.393 (1.313)	-0.023 (0.061)
	1974	839	1135	

注：2000、2004、2006 和 2009 年的观测值分别为 250、669、524 和 531。其中，2000、2004、2006 和 2009 四个年度的参合比例分别为 0%、9.2%、54.8% 和 96.2%；想再生孩子的妇女比例分别为 8.1%、10%、8.2% 和 9.4%。

（二）实证方法

1. 基本回归模型

首先，我们将生育意愿视为二元变量，用面板线性概率模型、Logit 模型和 Probit 模型分别探测参合对于生育意愿的影响。我们用 ferwill 代表生育意愿，NCMS 代表是否参合，x 代表控制变量，则三种模型可以表示为：

$$\Pr(ferwill_{it} = 1 \mid NCMS_{it}, x_{it}) = \begin{cases} \tau NCMS_{it} + x_{it}'\beta, & \text{线性概率模型} \\ G(\tau NCMS_{it} + x_{it}'\beta), & Logit\ \text{模型} \\ \Phi(\tau NCMS_{it} + x_{it}'\beta), & Probit\ \text{模型} \end{cases} \qquad (9)$$

其中，$\Pr(ferwill_{it}=1 \mid NCMS_{it}, x_{it})$ 为控制参合状态和其他协变量之后生育意愿为 1 的条件概率，$G(\cdot)$ 为 Logistic 分布的累积分布函数，$\Phi(\cdot)$ 为正态分布的累积分布函数。

其次，我们将生育意愿视为计数变量（记为 numferwill），即“在无生育政策限制下，想再生几个孩子”。计数变量适用于泊松模型，泊松模型也为其他生育决策研究者所采用（Wang & Famoye, 1997）。但标准的泊松模型存在两个问题。第一，生育意愿数据零值过多，超过 80% 的家庭表示不想再生孩子。我们将“再生几个孩子”的决策视为两部决策，首先决定是否再生孩子，其次决定再生几个孩子。第一步可用 Logit 模型估计。第二，标准泊松模型假定 $Var(numferwill) = E(numferwill)$，而在实际应用中经常出现 Over-dispersion 现象，即 $Var(numferwill) = \sigma E(numferwill)$，$\sigma$ 为大于 1 的参数。我们用负二项回归（negative binomial regression）将 σ 变为待估参数，解决 over-dispersion 问题。这两个解决方案的合并,即是 ZINB(zero-inflated negative binomial)模型。模型方程可由下式表示:

$$\Pr(numferwill_{it} = 0) = \omega_{it} + (1 - \omega_{it})\exp(-\lambda_{it}) \tag{10}$$

$$\Pr(numferwill_{it} = n) = (1 - \omega_{it})\frac{\exp(-\lambda_{it})\lambda_{it}^{n}}{n!}, n \geqslant 1 \tag{11}$$

$$\lambda_{it} = \exp(x_{it}'\beta) \tag{12}$$

$$\omega_{it} = \frac{\exp(z_{it}'\gamma)}{1 + \exp(z_{it}'\gamma)} \tag{13}$$

其中,ω_{it}为家庭 i 在 t 期第一步选择不再生孩子的概率,由 z_{it}决定。

2. 内生性问题的处理

上述模型很可能存在内生性问题。内生性的来源一方面是保险市场常见的逆向选择,即参合者具有不可观测的异质性,这些个人特质影响生育率,从而导致计量方程的扰动项和自变量相关(Cohen & Siegelman,2010);二是反向因果关系,生育很可能对社保覆盖率有影响(Ehrlich & Kim,2007)。

(1)工具变量

为了处理参合的内生性问题,我们用家庭所在的行政村是否开展了新农合作为工具变量,类似的做法为大多数新农合的研究文献所采用(Lei & Lin,2009;甘犁等,2010)。值得指出的是,社区是否开展新农合本身也是二元变量,用一个二元变量做另一个二元变量的工具变量,可能因为多个二元变量的变异性(Variation)不够而导致识别困难。为此,我们采用双变量 Probit 模型,联合估计生育意愿决定方程和参合方程,以考察工具变量结果的稳健性。

(2)DID 和 DID-Matching

我们也用双重差分法(Difference-in-Difference)对这一问题进行了估计。在处理组和控制组的选择上,本文将 2003 年之前视为政策实施前,2003 之年后视为政策实施后。据此划分,2000 年属于政策实施前,2004、2006 和 2009 年属于政策实施后。新农合未覆盖的家庭和新农合覆盖但未参合家庭均视为控制组,将新农合覆盖且参合的家庭视为控制组。这一划分方法也被马双和甘犁(2009)、甘犁等(2010)所采用。为了让 DID 结果更加可信,我们对政策实施后的观测值仅保留了 2004 年份。具体回归方程设定如下:

$$ferwill_{it} = \beta_0 + \delta After_t + \eta Treat_i + \phi After \cdot Treat + x_{it}'\beta + \varepsilon_{it} \tag{14}$$

其中,*After* 表示时间 t 在政策实施后,*Treat* 表示个体 i 处于控制组,交叉项系数 ϕ 是我们关注的双重差分效果。x 为控制变量。

虽然 DID 可消除处理组和控制组之间不随时间变化的不可观测异质性,但如果处理组和控制组的初始条件存在显著差异,则 DID 的估计仍有偏,此时可与匹配方法一同使用。本文准备采用倾向得分匹配和双重差分结合的方法(Difference-in-Difference with propensity score matching)。倾向得分匹配方法的基本原理是寻找一个基于倾向得分度量的、和处理组相似反事实组(counterfactual group)。运用 Matching 估计平均处置效应,就是为每个处理对象寻找相匹配的对比对象,在处理组内求其结果之差的均值。对 DID-matching 的详细介绍可以参考 Abadie(2005)。

五、实证结果和稳健性检验

(一)实证结果

从表 4 可以看出,在将生育意愿视为二元变量时,新农合降低了生育意愿。这一作用是显著的,但在数值上并不大。线性概率的随机效应和固定效应估计结果表明,参加新农合后想再生孩子

的愿望降低了3.7%;Probit和Logit的估计结果更小一些,在其余各变量均值处,参加新农合的边际效果为2%—3%。

表4　　面板二元选择变量回归结果

自变量	(1) 线性概率(RE)	(2) 线性概率(FE)	(3) Probit	(4) Logit
参加新农合	-0.037** (0.015)	-0.037* (0.022)	-0.374*** (0.130)	-0.739*** (0.249)
育龄妇女变量:				
年龄	-0.014*** (0.001)	-0.013* (0.007)	-0.087*** (0.009)	-0.167*** (0.017)
结婚次数大于1	0.019 (0.256)	0.031 (0.052)	-1.751 (231.551)	-6.285 (435.791)
兄弟姐妹数量	-0.001 (0.004)	-0.001 (0.008)	0.015 (0.028)	0.029 (0.054)
小时工资对数	-0.013** (0.006)	-0.019** (0.008)	-0.089* (0.046)	-0.176** (0.088)
正规教育年限	0.002 (0.002)	0.007 (0.006)	0.024 (0.019)	0.043 (0.038)
自评健康:好	-0.035 (0.022)	0.016 (0.028)	-0.331* (0.177)	-0.654* (0.344)
自评健康:一般	0.003 (0.014)	0.022 (0.021)	-0.087 (0.114)	-0.150 (0.216)
从事农业工作	0.008 (0.021)	0.011 (0.026)	0.153 (0.173)	0.251 (0.329)
在国有部门工作	-0.071** (0.033)	0.025 (0.068)	-0.551* (0.292)	-1.093* (0.564)
在私营部门工作	-0.032 (0.026)	-0.033 (0.042)	-0.071 (0.214)	-0.128 (0.407)
有其他健康保险	-0.018 (0.025)	-0.003 (0.031)	-0.192 (0.220)	-0.318 (0.430)
丈夫变量:				
兄弟数量	-0.011** (0.005)	-0.003 (0.009)	-0.086** (0.041)	-0.138* (0.081)
姐妹数量	-0.008 (0.005)	-0.005 (0.011)	-0.052 (0.040)	-0.083 (0.078)
小时工资对数	-0.005 (0.005)	0.006 (0.009)	-0.044 (0.042)	-0.104 (0.081)
正规教育年限	0.001 (0.002)	-0.000 (0.005)	0.005 (0.023)	0.006 (0.043)

续表 4

自变量	(1) 线性概率(RE)	(2) 线性概率(FE)	(3) Probit	(4) Logit
自评健康:好	0.028 (0.021)	0.034 (0.027)	0.183 (0.160)	0.367 (0.305)
自评健康:一般	-0.006 (0.015)	-0.001 (0.021)	-0.021 (0.121)	-0.048 (0.231)
从事农业工作	0.024 (0.026)	0.017 (0.034)	0.144 (0.226)	0.317 (0.435)
在国有部门工作	0.038 (0.035)	-0.057 (0.059)	0.346 (0.309)	0.678 (0.599)
在私营部门工作	0.037 (0.028)	0.067 (0.043)	0.208 (0.236)	0.424 (0.453)
家庭变量:				
一共生育的孩子数量	-0.033*** (0.007)	-0.029** (0.011)	-0.397*** (0.071)	-0.827*** (0.144)
养老负担	-0.010* (0.006)	-0.024** (0.010)	-0.048 (0.046)	-0.101 (0.092)
常数项	0.785*** (0.063)	0.684** (0.303)	2.880*** (0.473)	5.775*** (0.912)
参合边际效应	-0.037	-0.037	-0.031	-0.026
观测值	1974	1974	1974	1974

注:(1)***、** 和 * 分别代表在 1%、5% 和 10% 的水平显著;(2)括号中为标准误;(3)第三列和第四列报告的是估计系数。

生育意愿随着年龄的增长和已有子女数量的增加而显著下降,这和表 2 中关于不同年龄组别已婚妇女生育意愿的统计描述相符,也与其他实证研究结果一致(例如,陈字、邓昌荣,2007);妇女的小时工资对数显著降低了生育意愿。丈夫的兄弟数量对生育意愿的负向作用也十分显著,这体现了男权社会中妇女生育意愿受到丈夫家庭传宗接代任务的影响。

表 5 面板二元变量 IV 估计结果

自变量	(1) 线性概率(FE) 第二阶段	(2) 线性概率(RE) 第二阶段	(3) 线性概率 第一阶段
参加新农合	-0.097** (0.038)	-0.050*** (0.019)	
所在社区开展了新农合			0.800*** (0.013)
参合边际效应	-0.097	-0.050	
观测值	1974	1974	1974

注:(1)***、** 和 * 分别代表在 1%、5% 和 10% 的水平显著;(2)括号中为标准误;(3)其他控制变量的选取与表 4 相同,为了节约篇幅,未予报告。

为了处理参合的内生性,我们用社区是否开展新农合作为参合的工具变量,估计线性概率固定效应和随机效应模型,估计结果由表 5 给出结果。两种估计方法得到的系数绝对值均有增大,随机效应模型中新农合对生育意愿的作用效果为 -0.050,固定效应则为 -0.097。其他变量的结果未发生太大变化。第一阶段回归中,社区开展新农合对于家庭参合的回归系数为 0.80,且十分显著,说明新农合的社区覆盖与家庭参合之间的高度相关。工具变量对应的 F 值为 3769.71,大于文献中推荐的判

别弱工具变量的标准(Stock et al. ,2002),不存在弱工具变量问题。

在将生育意愿视为计数变量时,参合对于生育意愿的降低效果也是显著的。在 ZINB 模型中,似然比检验未能拒绝 alpha = 0 的原假设,即模型不存在 over-dispersion 问题。于是我们不再估计 over-dispersion 参数 σ,模型退化为 ZIP(zero-inflated Poisson)。ZIP 模型的 Vuong 检验 P 值为 0.019,表明 ZIP 相比于普通泊松模型更加合适。我们计算了 ZIP 模型的边际效应,参加新农合使得想再要孩子数量的预测值降低了 0.018 个。具体结果见表 6。

表 6　泊松模型回归结果

自变量	(1) ZINB	(2) ZIP
参加新农合	-0.413* (0.225)	-0.414* (0.221)
参合边际效应	-0.019	-0.019
观测值	1974	1974

注:(1)***、** 和 * 分别代表在 1%、5% 和 10% 的水平显著;(2)括号中为标准误;(3)其他控制变量的选取与表 4 相同,为了节约篇幅,未予报告;(4)第一阶段中我们选用了妇女年龄、收入、丈夫兄弟数量、姐妹数量、教育年限等关键变量,为了节约篇幅,未予报告。

无论是将生育意愿视为二元变量还是计数变量,参加新农合对于生育意愿的降低效果都是显著的,但在数值上并不大。这一方面和新农合的保险范围、保险强度较弱有关;另一方面也可能是我们理论模型中所强调的收入效应在起作用。但由于数据的限制,目前我们还没有发现对收入效应进行直接检测的可行方法。

(二)稳健性检验

1. 分样本回归结果

新农合对于生育意愿的影响效果,在不同群体中是异质的。探讨该作用效果在不同群体中的差异,有助于更精准地进行政策干预。我们按照已婚妇女的年龄和收入进行了分样本回归。结果表明:

新农合对生育意愿的降低作用,主要发生在年轻的育龄妇女组群中。表 7 报告了将育龄妇女按照是否大于 35 岁分为两组的线性概率模型回归结果(仍然使用社区参合作为工具变量),发现年轻组的边际效果为 -0.113,年长组为 -0.021。一般而言,生育行为发生的高峰期是 20—30 岁。目前处于这一年龄区间育龄妇女的生育行为和生育意愿,决定了中国未来的人口结构。如果试图通过社会保障体系的建立达到控制人口数量的目的,政策的目标人群也应该定位于此。

新农合对生育意愿的降低作用,主要发生在低收入的育龄妇女群体中。我们将育龄妇女的年收入(统一折算为 2009 年价格水平)按照小于 3000 元、3000—6000 元、6000—9000 元和大于 9000 元分为四个子样本,分别进行线性概率模型估计。表 8 给出的估计结果可以清晰地看到新农合的系数绝对值随着收入的提高而降低,且后三组的系数均不显著。

表 7　按年龄分样本回归结果

自变量	(1) 年龄小于等于 35 岁	(2) 年龄大于 35 岁
参加新农合	-0.113** (0.055)	-0.021 (0.016)
观测值	534	1440

注:(1)***、** 和 * 分别代表在 1%、5% 和 10% 的水平显著;(2)括号中为标准误;(3)控制变量和表 4 相同,为节省篇幅,未予报告;(4)因为年龄会随调查时间而变化,用固定效应将导致观测值太少,此处我们采用随机效应(RE)估计量,并采用社区是否参合作为工具变量。

表 8　按收入分样本回归结果

自变量	(1) 0—3000	(2) 3000—6000	(3) 6000—9000	(4) 9000 +
参加新农合	-0.073** (0.034)	-0.033 (0.035)	-0.054 (0.048)	-0.025 (0.038)
观测值	669	491	289	525

注:(1)本表格中(1)—(4)列子样本分别代表育龄妇女年收入按 2009 年物价水平折算后小于 3000 元人民币、3000—6000 元、6000—9000 元和高于 9000 元;(2)***、** 和 * 分别代表在 1%、5% 和 10% 的水平显著;(3)括号中为标准误;(4)其余控制变量和表 4 相同,为节省篇幅,未予报告;(5)考虑到分成四个样本后能够形成面板的观测值较少,同一家庭可能因为在两次调查中收入变化而落入不同的子样本,此处我们采用随机效应(RE)估计量,并采用社区是否参合作为工具变量。

2. 对内生性的再讨论

本小节中，我们使用 Bivariate Probit、DID 和 DID-matching 三种方法对对内生性问题进行再讨论，结果表明，新农合对生育意愿的影响是稳健的。

Bivariate Probit 的模型的估计结果由表 9 给出。该模型中新农合对生育意愿的边际效果较小，仅为 -0.013，但在 1% 的水平上是显著的。

表 9　Bivariate Probit 模型回归结果

自变量	(1)	(2)
	生育意愿	参加新农合
参加新农合	-0.428*** (0.157)	
所在社区开展了新农合		2.855*** (0.109)
参合边际效应	-0.013	
观测值	1974	1974

注：(1) ***、** 和 * 分别代表在 1%、5% 和 10% 的水平显著；(2) 括号中为标准误；(3) 方程(1)和方程(2)中的解释变量分别同表 5 中 IV 模型两个阶段的解释变量相同，为节省篇幅，未予报告。

在双重差分(DID)分析中，我们使用 2004 年新参合者作为处理组(57 个观测值)，未参合者为控制组(862 个观测值)，2000 年为政策实施之前，2004 年为政策实施之后。结果表明，处理组(参合者)的平均处理效应(ATT)为 -0.118，且十分显著。说明新农合使得参合者的生育意愿显著降低了 11.8%。

我们同时使用 DID-matching 方法进行了分析，具体采用核匹配对比法(带宽 h = 0.2)，对处理组和控制组的定义与 DID 相同。匹配显著降低了实验组和控制组的(可观测)差异，匹配后，大部分控制变量的 t 值都小于 1.65。该方法得到的平均处理效应(ATT)为 -0.074，表明新农合使得参合者的生育意愿下降了 7.4%，这和 DID 的结果基本一致，也和前文几类实证方法的结果形成呼应，在一定程度上证实了结果的稳健性。

3. 对养儿防老机制的检验

表 10　加入男孩交叉项的回归结果

	(1)	(2)
参加新农合	-0.121** (0.055)	-0.298** (0.136)
已有一个男孩	-0.147 (0.138)	-0.161 (0.141)
参加新农合 × 已有一个男孩		0.217* (0.114)
观测值	1369	1369

注：(1) 本表格的样本为已经有孩子的家庭；(2) ***、** 和 * 分别代表在 1%、5% 和 10% 的水平显著；(3) 括号中为标准误；(4) 其余控制变量和表 4 相同，为节省篇幅，未予报告；(5) 模型采用固定效应。

社会保障对生育的替代作用，取决于生育中经济动机的存在。我们认为，按照中国的文化习俗，养老的任务主要由儿子承担。这意味着，与有男孩的家庭相比，没有男孩家庭的生育动机更为强烈，社保对于生育的影响效果也应该更大。表 10 报告了以已经有孩子的家庭为样本的回归结果：新农合系数为负，已有一个男孩系数为负，两者的交叉项为正，新农合的总体边际效果仍然为负，符合上述作用机制。并且，已经有孩子的样本新农合对生育的影响大于全样本结果，这也符合我们的预期：生育动机除了养老之外，还有情感、传宗接代等等，妇女生育第一胎的意愿不会受到社会保障太大的影响，社保影响最大的是已经有了孩子的家庭的再生育意愿。

六、结论和展望

过去的十年里，中国逐步建立起覆盖城乡，包括养老保险和医疗保险在内的全方位社会保障体系。社会保障制度的建立对居民生活产生了巨大影响，养儿防老等传统社会的生育观念正在改变。

本文研究了新型农村合作医疗对生育意愿的影响。基于 Becker 家庭决策模型的理论分析表明，带有补贴的新农合会产生两种效应：一方面新农合对家庭预算约束的放松提升了家庭的生育意愿，另一方面新农合对养儿防老功能的替代降低了家庭的生育意愿。实证结果显示，后一种作用占主导地位，参合使得家庭生育意愿平均降低了 3%—10%。该效果主要是由年轻、低收入的育龄妇

女生育意愿的改变所引起的。并且,用 DID 和 DID-matching 等多种方法进行的估计证明了这一结果的稳健性。

由于数据的限制,我们无法考虑新农合补贴额度和报销比例的地区差异,这使得收入效应的确认较为困难。如果收入效应存在,那么我们有理由相信:将农村医保体系由国家财政支付大量补贴的新农合模式转为完全依靠个人缴费的无补贴模式,生育意愿的降低会更为显著。

我们的研究结果具有明显的政策含义:计划生育并非控制人口的唯一途径,建立健全社会保障体系,可以间接调控人口数量,实现从强制少生到社保覆盖下的自愿适度生育。

参考文献

国家人口发展战略研究课题组,2007:《国家人口发展战略研究报告》,《人口研究》第 1 期。

白重恩、李宏彬、吴斌珍,2012:《医疗保险与消费:来自新型农村合作医疗的证据》,《经济研究》第 2 期。

马双、臧文斌、甘犁,2010:《新型农村合作医疗保险对农村居民食物消费的影响分析》,《经济学(季刊)》第 10 卷第 1 期。

马双、张劼,2011:《新型农村合作医疗保险与居民营养结构的改善》,《经济研究》第 5 期。

齐良书,2011:《新型农村合作医疗的减贫、增收和再分配效果研究》,《数量经济技术经济研究》第 8 期。

秦雪征,2010:《社会安全网、自我保险与商业保险:一个理论模型》,《世界经济》第 10 期。

秦雪征、郑直,2011:《新农合对农村劳动力迁移的影响:基于全国性面板数据的分析》,《中国农村经济》第 10 期。

姚从容、吴帆、李建民,2010:《我国城乡居民生育意愿调查研究综述:2000—2008》,《人口学刊》第 2 期。

陈字、邓昌荣,2007:《中国妇女生育意愿影响因素分析》,《中国人口科学》第 6 期。

朱恒鹏,2009:《新医改研究综述》,《经济学动态》第 10 期。

Abadie, A., 2005, "Semiparametric Difference-in-Differences Estimators", *Review of Economic Studies* 72, 1—19.

Attane, I., 2002, "China's Family Planning Policy: An overview of Its Past and Future", *Studies in Family Planning* 33, 103—113.

Bai, C., and Wu, B., 2014, "Health Insurance and Consumption: Evidence from China's New Cooperative Medical Scheme", *Journal of Comparative Economics* 42, 450—469.

Barro, R., and Becker, G., 1989, "Fertility Choice in a Model of Economic Growth", *Econometrica* 57, 481—501.

Becker, G., 1960, "An Economic Analysis of Fertility", in Ansley C., ed., Demographic and Economic Change in Developed Countries, Princeton, NJ: Princeton University Press, 209—240.

Becker, G., and Lewis, G., 1973, "On the Interaction between the Quantity and Quality of Children", *Journal of Political Economy* 81, 279—288.

Becker, G., and Tomes, N., 1976, "Child Endowments and the Quantity and Quality of Children", *Journal of Political Economy* 84, 143—162.

Boldrin, M., Nardi, M., and Jones, L., 2005, "Fertility and Social Security", Federal Reserve Bank of Minneapolis Research Department Staff Report 359.

Cigno, A., and Rosati, F., 1992, "The Effects of Financial Markets and Social Security on Saving and Fertility Behaviour in Ttaly", *Journal of Population Economics* 5, 319—341.

Cigno, A., Casolaro, L., and Rosati, F., 2002, "The Impact of Social Security on Saving and Fertility", *Finanz Archiv/Public Finance Analysis* 59, 189—211.

Cohen, A., and Siegelman, P., 2010, "Testing for Adverse Selection in Insurance Markets", *Journal of Risk and Insurance* 77, 39—84.

Eswaran, M., 2002, "The Empowerment of Women, Fertility, and Child Mortality: Towards a Theoretical Analysis", *Journal of Population Economics* 15, 433—454.

Ehrlich, I., and Lui, F., 1998, "Social Security, the Family and Economic Growth", *Economic Inquiry* 36, 390—409.

Holger, S., 2006, "Bargaining over Fertility in Rural Ethiopia", Proceedings of the German Development Economics Conference.

Holm, C., 1975, "Social Security and Fertility: An Intergenerational Perspective", *Demography* 12, 629—644.

Hotz, V., Klerman, J., and Willis, R., 1997, "The Economics of Fertility in Developed Countries", *Handbook of Population and Family Economics* Vol 1 Part A, 275—347.

Holmqvist, G., 2011, "Fertility Impact of High-coverage Public Pensions in Sub-Saharan Africa", *Global Social Policy* 11, 152—174.

Hori, H., and Kanaya, S., 1989, "Utility Functionals with Non-paternalistic Intergenerational Altruism", *Journal of Economic Theory* 49, 241—265.

Kimball, M., 1987, "Making Sense of Two-sided Altruism", *Journal of Monetary Economics* 20, 301—366.

Manser, M., and Brown, M., 1980, "Marriage and Household Decision-Making: A Bargaining Analysis", *International Economic Review* 21, 31—44.

Malthus, T., 1798, "An Essay on the Principle of Population, as It Affects the Future Improvements of Society. with Remarks on the Speculations of Mr. Godwin, M. Condorcet, and other Writers", London: J. Johnson.

Nishimura, K., and Zhang, J., 1995, "Sustainable Plans of Social Security with Endogenous Fertility", *Journal of Population Economics* 47, 182—194.

Nugent, J., 1985, "The Old-age Security Motive for Fertility", *Population and Development Review* 11, 75—97.

Rosati, F., 1996, "Social Security in a Non-altruistic Model with Uncertainty and Endogenous Fertility", *Journal of Public Economics* 60, 283—294.

Rosenbaum, P., and Rubin, D., 1983, "The Central Role of the Propensity Score in Observational Studies for Causal Effects", *Biometrika* 70, 41—55.

Wang, F., 2012, "Family Planning Policy in China: Measurement and Impact on Fertility", MPRA Paper No. 42226.

Wang, W., and Fayome, F., 1997, "Modeling Household Fertility Decisions with Generalized Poisson Regression", *Journal of Population Economics* 10, 273—283.

Wigger, B., 1999, "Pay-as-you-go Financed Public Pensions in a Model of Endogenous Growth and Fertility", *Journal of Population Economics* 12, 625—640.

The Impact of Social Security on Fertility Desire: Evidence from the New Rural Cooperative Medical Scheme

Wang Tianyu[a] and Peng Xiaobo[b]

(a: School of Economics, Peking University;

b: Graduate School, Chinese Academy of Social Sciences)

Abstract: It has been demonstrated by a great deal of research that social security has profound influence on individual fertility decision and total fertility rate. However, relevant study about China is still rare due to the limitation (distortion) of one-child policy on fertility behavior. In this paper, we investigate the effect of the New Rural Cooperative Medical Scheme (NCMS) on fertility desire of China's rural residents. Based on two-period family decision model, we find that the NCMS with subsidy from government has both income effect and crowd-out effect. The former increases the fertility desire while the latter depresses it. Armed with the 2000—2009 longitudinal data from China Health and Nutrition Survey (CHNS), we find that the crowd-out effect outweighed the income effect. That is, enrolling in the NCMS makes the fertility desire of women at child bearing age decrease by 3%—10%. Supported by both theoretical and empirical evidence, we think the establishment and popularization of health insurance can provide some opportunity for the relaxation of the one-child policy.

Key Words: Social Security; Fertility Desire; One-Child Policy; New Rural Cooperative Medical Scheme

JEL Classification: I13, I18, H40

(责任编辑:詹小洪)(校对:梅　子)

China Economic Review xxx (2014) xxx–xxx

Contents lists available at ScienceDirect

China Economic Review

Intergenerational transfer of human capital and its impact on income mobility: Evidence from China

Xuezheng QIN *, Tianyu WANG [1], Castiel Chen ZHUANG [2]

School of Economics, Peking University, Beijing 100871, China

ARTICLE INFO

Article history:
Received 25 March 2014
Received in revised form 20 October 2014
Accepted 20 October 2014
Available online xxxx

JEL classification:
J25
O14
J62

Keywords:
Human capital
Income mobility
Intergenerational transmission
China

ABSTRACT

This paper analyzes theoretically and empirically the impact of intergenerational transmission of human capital on the income mobility in China. We use a three-period overlapping-generations (OLG) model to show that the human capital transfer plays a remarkable role in determining the parent-to-offspring investment in human capital and the intergenerational elasticity of income. We then estimate a simultaneous equations model (SEM) using the 1989–2009 China Health and Nutrition Survey (CHNS) data to verify our theoretical predictions. The results show that (i) human capital, measured by health and education, is directly transmitted from one generation to the next, reflecting the parent-induced inequality of development opportunities among offspring in China; (ii) the estimated intergenerational income elasticity increases from 0.429 to 0.481 when the direct transfer of human capital is accounted for, suggesting that omitting this mechanism would overestimate China's income mobility. Our findings provide policy implications on strengthening human capital investments among the disadvantaged groups, reinforcing reforms that promote equality of opportunity, and improving the efficiency of labor markets in China.

1. Introduction

Social mobility, or the changeability of one's socio-economic status, is an important measure of the equality of a society and an essential determinant of a country's sustainability in economic growth. Since the 1980s, China's unprecedentedly rapid economic growth has been accompanied by emerging social problems such as rising income inequality and lack of social mobility. These problems challenge the continuous development of China's economy and threaten to drag the country into the middle income trap. In recent years, several studies that look at the intergenerational correlation of income have consistently found (i) internationally low income mobility (an important dimension of social mobility) in China, and (ii) the sustaining income inequality is to a high extent explained by the inequality of opportunity (Deng, Bjorn, & Li, 2012; Gong, Leigh, & Meng, 2012; Zhang & Eriksson, 2010). Table 1 summarizes the estimated intergenerational elasticity of income in major countries, among which China ranks higher than most of the developed and some of the developing countries, indicating a low level of intergenerational income mobility. This phenomenon is not only against the moral principal of social equity by giving poor people too few opportunities to improve their economic status, but also leads to decreased incentives among individuals to invest in human capital and thus negatively affects the long run economic efficiency (Moaz & Moav, 1999).

Based on the theory of human capital (Becker, 1993), one of the root causes of low intergenerational mobility is that human capital (including health and education) remains invariant through intergenerational transmission. In particular, some studies point out that

* Corresponding author. Tel.: +86 10 6275 7237; fax: +86 10 6275 4237.
E-mail addresses: qin.econpku@gmail.com (X. Qin), wangtianyu@pku.edu.cn (T. Wang), zogcee@gmail.com (C.C. Zhuang).
[1] Tel.: +86 138 1178 9779.
[2] Tel.: +86 188 0140 6111.

http://dx.doi.org/10.1016/j.chieco.2014.10.005
1043-951X/

Please cite this article as: Qin, X., et al., Intergenerational transfer of human capital and its impact on income mobility: Evidence from China, *China Economic Review* (2014), http://dx.doi.org/10.1016/j.chieco.2014.10.005

Table 1
International comparison of intergenerational income elasticity.

Country	Elasticity	Data	Source
U.S.	0.45–0.48	PSID	Lee and Solon (2009)
Canada	0.23	Income Tax Information	Corak and Heisz (1999)
U.K.	0.37	Combined Dataset	Nicoletti and Ermish (2007)
Italy	0.33	SHIW	Piraino (2007)
France	0.41	FQP	Lefranc and Trannoy (2005)
Germany	0.24	G-SOEP	Vogel (2008)
Australia	0.25	Combined Dataset	Leigh (2007)
Finland	0.20	Administrative Register Information	Pekkarinen et al. (2009)
Norway	0.25	Administrative Register Information	Nilsen et al. (2008)
Denmark	0.14	Administrative Register Information	Hussain et al. (2008)
Peru	0.50	LSMS	Grawe (2004)
Malaysia	0.40	MFLS	Grawe (2004)
Pakistan	0.18	LSMS	Grawe (2004)
Brazil	0.52	PNAD	Dunn (2007)
China	0.45	CHNS	Zhang and Eriksson (2010)

Notes: (i) Some information in the table is quoted from Deng et al. (2012) and Björklund and Jäntti (2009); (ii) all coefficients represent father–son income elasticity; (iii) PSID stands for Panel Study of Income Dynamics; (iv) SHIW (Survey on Household Income and Wealth) is sponsored by Bank of Italy; (v) FQP (Formation, Qualification, Profession) is conducted by INSEE (Institut National de la Statistique et des Études Économiques); (vi) G-SOEP stands for German Socio-Economic Panel; (vii) LSMS (Living Standard Measurement Survey) is conducted by the Word Bank; (viii) MFLS stands for Malaysian Family Life Survey; (ix) PNAD stands for Pesquisa Nacional por Amostra de Domicilios; (x) CHNS stands for China Health and Nutrition Survey.

the low human capital accumulation of the disadvantaged population is consistently inherited from the previous generations, thus limiting the ability of the poor to improve their income (Yao & Zhao, 2007). To a large extent, China's recent experience has mirrored the above findings. In the health care sector, the traditional rural Cooperative Medical Scheme (CMS) and urban Labor Insurance Scheme (LIS) gradually collapsed during the marketization and privatization reforms in the 1980s and 1990s, leading to persistent deterioration in the access to basic health care (Blumenthal & Hsiao, 2005); consequently, inequality in health capital expands through intergenerational transmission, and people in poor-health groups see little chance to substantially improve their health status. From the perspective of education, the inequality in educational opportunities is also aggravated during China's education reforms in the 1990s that are characterized by industrialization and enrollment expansion in the higher education sector (Liu, 2006); specifically, well-educated parents who are generally in high socio-economic status manage to obtain better access to high-quality educational resources for their children (Li, 2006). Given the above stylized facts, an important theoretical and practical question is: whether and to what extent does the intergenerational persistence of human capital contribute to the decreasing income mobility in China? A valid answer to the above question is not only meaningful to economic theorists, but can also provide policy guidance to China's current reforms in education, health care and income redistribution.

This paper is among the first to study the impact of human capital transmission across generations on the intergenerational income mobility in China from both theoretical and empirical perspectives. In the theoretical analysis, we incorporate the mechanisms of direct and indirect transmission of human capital into a classical three-period overlapping-generations (OLG) framework based on Becker and Tomes (1979), the results of which provide a new way to solve for parents' optimal investment in their children's human capital and to calculate the implied elasticity of income across generations. In the empirical analysis, we use the 1989–2009 China Health and Nutrition Survey (CHNS) data to investigate the intergenerational transmission of income, health and education based on a simultaneous equations model, the results of which show remarkable transmission of human capital (health and education) across generations and that the parent–offspring income elasticity tends to be underestimated (i.e., income mobility tends to be overestimated) if the direct transfer of human capital is not accounted for. Our findings provide a new approach to estimate the intergenerational income elasticity, which facilitates deeper understanding on the root causes of the lack of social mobility in China.

The paper proceeds as follows: Section 2 briefly reviews the relevant literature; Section 3 lays out the theoretical model and its extension; Section 4 describes our empirical strategy and data; Section 5 presents the estimation results and robustness tests; Section 6 concludes the paper.

2. Literature review

The intergenerational income elasticity, estimated by the regression coefficient of parents' logarithmic permanent income on their children's, is commonly used as a measure of social mobility in the economic literature. Limited by data availability and measurement accuracy, the traditional estimates of intergenerational income elasticity tend to be biased downwards; an example is Solon (1999), which estimates that the elasticity in the U.S. is only 0.2. Facilitated by large databases such as the Panel Study of Income Dynamics (PSID), the National Longitudinal Survey of Youth (NLSY) and the National Child Development Study (NCDS), recent studies are able to obtain more accurate estimates for different countries. For example, based on NLSY, NCDS and Nordic register data, Jäntti et al. (2006) conduct a cross-country study on the intergenerational elasticity of income, and find that the estimated elasticity is 0.517 in the U.S., 0.306 in the U.K., and less than 0.2 in three Nordic countries (Norway, Finland and Denmark). Vogel (2008) studies the intergenerational mobility in Germany by calculating permanent incomes based on the German Socio-Economic Panel (G-SOEP) data, and concludes that the elasticity is about 0.24.

In developing countries, empirical literature on this topic is relatively scant due to the often and abrupt social structural changes and the lack of reliable datasets. For example, China as the largest developing country has experienced dramatic social and economic transformation. Before the 1980s, measuring the intergenerational income mobility in China seemed barely meaningful. This is because the demand and supply of labor were strictly controlled by the government under the centrally planned economy, and thus the labor market does not exist and the marginal value of labor cannot be reflected by wages. However, after the economic reforms in the 1980s, the labor market in China has gradually developed and taken its role in determining the market wages, the income elasticity therefore becomes an important indicator to measure the social mobility and the equity of income distribution across generations. Several studies have discussed this issue: Zhang and Eriksson (2010) estimate the correlation of parental household income and children's individual income using the 1989–2006 CHNS data, and conclude that the elasticity is about 0.45; based on the 2004 Chinese Urban Household Education and Employment Survey (UHEES), Gong et al. (2012) explicitly take the impacts of life cycle and measurement errors into account in their estimation, and report that in urban China the father–son elasticity of income is about 0.6 and the father–daughter elasticity is up to 0.97, which is extremely high compared with the other related studies (e.g., Guo & Min, 2007 use the same data but get an estimate of 0.32 for father–son elasticity); based on the data from China Household Income Projects (CHIP) and Chinese General Social Survey (CGSS), Chen and Yuan (2012) find that the income elasticity in China fell dramatically at first and remained stable afterwards during 1988–2005, and that the estimated elasticity in urban areas (about 0.4) is higher than that in rural areas (about 0.3); Sun, Huang, and Hong (2012) also use the 2006 CGSS data and find that the labor migration plays an important role in reducing the intergenerational persistence of income in China.

Compared with the estimation of intergenerational elasticity, more attention is deserved on the mechanisms giving rise to the resemblance in income between parents and children. The commonly identified channels include, e.g., inheritance of wealth or family businesses, allocation of political rents, institutional perseverance and the environmental and behavioral influences by social cohorts (Borjas, 1992). Among all the contributing factors, the intergenerational transmission of human capital undoubtedly plays an important role: on the one hand, since one's income is largely determined by his human capital accumulation, parents' human capital can influence their financial capability to invest in children's human capital and thus influence their children's future income (we call this the indirect transmission mechanism); on the other hand, parents' human capital can be directly transferred to their children through genetic or non-genetic channels, which also influences the future income of their children (we call this the direct transmission mechanism).

Intergenerational investment in human capital within families (the indirect mechanism) is first studied by Becker and Tomes (1979). They discuss parents' trade-offs among consumption, financial investment and human capital investment in their children under the altruistic utility assumption. Their model contributes to the better understanding of the determination of income mobility across generations and sparks many follow-up studies. For example, Loury (1981) introduces the credit constraints to Becker's model and concludes that under the constraints, parents in low-income families can hardly borrow enough money to invest in children's human capital, contributing to the decreases in the intergenerational mobility. Moaz and Moav (1999) and Nakamura and Murayama (2011) also discuss this mechanism but under the assumption that human capital stock is discrete (uneducated vs. educated). Based on the above theoretical findings, some empirical studies also try to assess the possibility of improving the intergenerational income mobility through public policy tools, including public education policies (Restuccia & Urrutia, 2004) and income taxation policies (Zhu & Vural, 2012).

The second mechanism, in comparison, mainly refers to the direct transfer of human capital (such as education and health) across generations through the genetic channels (e.g. natural inheritance of cognitive abilities and health conditions) or behavioral channels (e.g. children learn from their parents by observing their activities). This direct transmission mechanism has drawn increasing attention in the recent literature. For example, using the NLSY data, Akbulut and Kugler (2007) show that height, weight (or BMI) and other health indicators of parents are highly correlated with those of their children. Eriksson, Bratsberg, and Raaum (2005) also demonstrate the strong similarity of health conditions between parents and offspring using the data from Denmark. However, the above studies did not address the endogeneity issue in their estimation, thus are mainly concerned with the intergenerational correlation (rather than causality) of human capital. In an effort to identify the causality, some studies use special samples to explicitly eliminate the influences of unobserved factors on the health capital of both generations. For example, Currie and Moretti (2007) control the influences of family background and genetic endowment using a sample of sibling mothers obtained from the Californian individual birth record data (assuming that sisters share similar family background and genes), and find that the low-birth-weight mothers also tend to have low-birth-weight children, suggesting a causal impact of health status across generations. Similarly, Royear (2009) identifies the causality using a confidential dataset of twins based on the 1960–2002 Californian birth records. Other studies also discover the intergenerational link of the health-related habits such as smoking, drinking and physical inactivity. For example, Loureiro, Sanz-de-Galdeano, and Vuri (2010) use data from the Birth Household Panel Survey (BHPS) to find that children's smoking behaviors are significantly influenced by their parents'.

Compared with the intergenerational transmission of health, the estimation on the direct transfer of educational attainment is more complicated, as the indirect transmission mechanism (parents' investment in children's education) can be a compounding factor. Nevertheless, several recent studies have demonstrated the genetic transmission of cognitive abilities: Black, Devereux, and Salvanes (2009) and Björklund, Eriksson, and Jäntti (2010) use data from the Swedish military enlistment tests to estimate the correlation in Intelligence Quotient (IQ) between fathers and sons, and find a high correlation coefficient around 0.32; based on the G-SOEP data, Anger and Heineck (2010) show that cognitive skills that are based on past learning are more strongly transferred from parents to children than those related to innate abilities. On the other hand, some recent studies also aim to identify the non-genetic transmission of education using special data samples or exogenous instruments. For example, Plug (2004) uses a sample of adoptees from the Wisconsin Longitudinal Survey (WLS) and finds that the adoptive father–child correlation of schooling is strong but the adoptive mother–child correlation is not significant; Chevalier (2004) and Black, Devereux, and Salvanes (2005) identify the non-genetic effects by using the education reforms in Britain and Norway as natural experiments, and they both find that only the mother–son transmission of schooling is significant while the other three pairs (i.e., mother–daughter, father–son, and father–daughter) are not.

3. Theoretical model

3.1. The benchmark model

We start by formulating an OLG model in which all individuals go through three life stages: dependent children, working young adults and retired old adults. We assume an individual accumulates human capital in childhood; he works and earns income in young adulthood while at the same time make decisions on consumption, savings for old age and investment in their children; in old adulthood, the individual retires and consumes on his savings. For simplicity, the benchmark model only accounts for parents' investment in their children's human capital (the indirect transmission) without considering the direct human capital transfer across generations.

In the above context, the expected utility of a representative working young adult can be expressed as follows:

$$EU_t^p = u(c_t^p) + \beta u(c_{t+1}^p) + \beta\alpha E_t u(y_{t+1}^c) \tag{1}$$

where superscript p denotes parent, and superscript c denotes child. $u(\cdot)$ is the instantaneous utility function, with $u'(\cdot) > 0$ and $u''(\cdot) < 0$. Working young adults earn an income of y_t^p in period t and make decisions on current consumption c_t^p, old-age savings s_t^p, and offspring human capital investment e_t^p. Following the altruistic assumption in the literature (Becker & Tomes, 1979; Solon, 2004), the adults also derive utility from their children's expected income in period $t+1$, y_{t+1}^c, when the children themselves become young adults, with α being the altruistic parameter and β being the discount factor. In period $t+1$, the working young adults retire and spend their deposits $(1+R_t)s_t^p$ on consumption c_{t+1}^p.

We assume human capital is the engine of economic growth, and one's income y_t is determined by his human capital accumulation h_t:

$$y_t^p = (h_t^p)^\gamma, \quad 0 < \gamma < 1 \tag{2}$$

where $0 < \gamma < 1$ ensures a diminishing marginal return of human capital on income. Furthermore, one's human capital stock depends on his innate ability A_t^c and the investment e_t^p from their parents, thus

$$h_{t+1}^c = A_t^c (e_t^p)^{1-\sigma}, \quad 0 < \sigma < 1 \tag{3}$$

where $0 < \sigma < 1$ ensures that the human capital investment is also subject to diminishing marginal returns. For simplicity, we assume that one's innate ability is randomly determined as follows:

$$\ln A_t^c = \ln \bar{A} + \varepsilon_t \tag{4}$$

where $\ln \bar{A}$ is the mean (constant) value of $\ln A_t^c$, and $\varepsilon_t \sim N(0, v^2)$ is the random shock following a normal distribution. Since the parents may not know their children's innate abilities, we assume they form expectations on $\ln A_t^c$ based on their own innate abilities A_t^p and make investment decisions accordingly.

Following the literature convention (e.g., Solon, 2004), we simplify the instantaneous utility function using the logarithmic form, thus the optimization problem becomes:

$$\underset{c_t^p, s_t^p, e_t^p, c_{t+1}^p}{Max} \; EU_t^p = \ln c_t^p + \beta \ln c_{t+1}^p + \beta\alpha E_t \ln y_{t+1}^c \tag{5}$$

$$s.t. \quad c_t^p + s_t^p + e_t^p \le y_t^p \tag{6}$$

$$c_{t+1}^p \le (1+R_t) s_t^p \tag{7}$$

$$y_{t+1}^c = (h_{t+1}^c)^\gamma \tag{8}$$

Please cite this article as: Qin, X., et al., Intergenerational transfer of human capital and its impact on income mobility: Evidence from China, *China Economic Review* (2014), http://dx.doi.org/10.1016/j.chieco.2014.10.005

where Eqs. (6) and (7) are parents' budget constrains in period t and $t+1$ respectively (both are binding constraints), and Eq. (8) is the income determination equation. The associated first-order conditions (F. O. C.) are:

$$\frac{c_t^p}{c_{t+1}^p}\beta(1+R_t)=1 \tag{9}$$

$$\frac{1}{c_t^p}=\alpha\beta\gamma(1-\sigma)\frac{1}{e_t^p} \tag{10}$$

where Eqs. (9) and (10) respectively indicate that in the optimal equilibrium, parents' marginal utility of consumption in period t equals that in period $t+1$, and that parents' marginal utility of consumption equals the marginal utility return of human capital investment in their children. The resulting closed-form solution of e_t^p, derived by Eqs. (6), (7), (9) and (10), is:

$$e_t^p=y_t^p\frac{\alpha\beta\gamma(1-\sigma)}{1+\beta+\alpha\beta\gamma(1-\sigma)}. \tag{11}$$

Through standard comparative static analyses, we have the following proposition:

Proposition 1. In a stable growth-equilibrium regime, parents' optimal human capital investment e is positively influenced by parents' income y, the altruistic parameter α, the discount factor β, and the parameter γ in the income deterministic function, but is negatively influenced by the parameter σ in the human capital accumulation function.

Proof. Take the derivatives of e with respect to y, α, β, γ, and σ, respectively, we have:

$$\begin{aligned}
&\frac{\partial e_t^p}{\partial y_t^p}=\frac{\alpha\beta\gamma(1-\sigma)}{1+\beta+\alpha\beta\gamma(1-\sigma)}>0,\ \frac{\partial e_t^p}{\partial\alpha}=y_t^p\frac{(1+\beta)\beta\gamma(1-\sigma)}{[1+\beta+\alpha\beta\gamma(1-\sigma)]^2}>0\\
&\frac{\partial e_t^p}{\partial\beta}=y_t^p\frac{\alpha\gamma(1-\sigma)}{[1+\beta+\alpha\beta\gamma(1-\sigma)]^2}>0,\ \frac{\partial e_t^p}{\partial\gamma}=y_t^p\frac{(1+\beta)\alpha\beta(1-\sigma)}{[1+\beta+\alpha\beta\gamma(1-\sigma)]^2}>0\\
&\frac{\partial e_t^p}{\partial\sigma}=-y_t^p\frac{(1+\beta)\alpha\beta\gamma}{[1+\beta+\alpha\beta\gamma(1-\sigma)]^2}<0.
\end{aligned}$$

The proposition is intuitive and reasonable. Parents in high-income families are more financially capable, thus are more likely to invest in their children's human capital; if the altruistic parameter α increases, the marginal utility return on human capital investments will also increase and thus motivates parents to invest more in their children; the increase of the discount factor β indicates that children's future income becomes more important to parents and thus their incentive to invest grows (although parents' future consumption also becomes more important and will partially decrease the incentive); the increase of the parameter γ will lead to a higher marginal return of human capital on income, thus also contributes to parents' higher incentive to invest; however, the increase of the parameter σ will decrease the marginal return of human capital investment, therefore depressing the incentives to invest.

Since our main focus is on the intergenerational income mobility rather than parents' optimal investment decisions, we substitute Eqs. (3), (4) and (8) into Eq. (11) to form the intergenerational income transmission function as follows:

$$\ln y_{t+1}^c=\gamma\ln\overline{A}+\gamma(1-\sigma)\ln y_t^p+\gamma(1-\sigma)\ln\frac{\alpha\beta\gamma(1-\sigma)}{1+\beta+\alpha\beta\gamma(1-\sigma)}+\gamma\varepsilon_t. \tag{12}$$

Following the conventional functional form in the intergenerational income regressions (e.g. Jäntti et al., 2006), Eq. (12) can be simplified to:

$$\ln y_{t+1}^c=\eta_0+\eta_1\ln y_t^p+u_t \tag{13}$$

where the intercept η_0, the income elasticity η_1 and the residual u_t are given as follows:

$$\eta_0=\gamma\ln\overline{A}+\gamma(1-\sigma)\ln\frac{\alpha\beta\gamma(1-\sigma)}{1+\beta+\alpha\beta\gamma(1-\sigma)} \tag{14}$$

$$\eta_1=\gamma(1-\sigma) \tag{15}$$

$$u_t=\gamma\varepsilon_t. \tag{16}$$

In Eq. (13), the intergenerational income elasticity η_1 is the key parameter of interest, and its functional form in Eq. (15) suggests that the elasticity is effectively determined by the marginal return of human capital on income and the rate of return on parent-to-offspring human capital investment. When these two returns increase, parents' incentive to invest in their children's human capital grows, which in turn leads to a higher intergenerational income elasticity and lower income mobility.

3.2. The extended model

The above benchmark model only accounts for the indirect transmission of human capital through the parent-to-offspring investment. In the extended model, we further introduce the direct transmission mechanism and discuss its impact on income mobility. The model extension is motivated by several recent studies that characterize the direct transfer of human capital across generations independent of the investment channel, i.e. human capital can be directly transmitted through genetic inheritance and non-genetic influences. For example, Hertz et al. (2007) estimate the intergenerational persistence of educational attainment in 42 countries and report that the intergenerational correlations between parents' and children's schooling remain surprisingly high and stable around 0.4 to 0.6 for the past fifty years. Chevalier, Denny, and McMahon (2009) report similar results through a multi-country study of intergenerational educational mobility and find a positive relationship between intergenerational income mobility and return to education. Based on these findings, the human capital accumulation function can be modified to:

$$h_{t+1}^c = A_t^c \left(h_t^p\right)^{\sigma} \left(e_t^p\right)^{1-\sigma}, \quad 0 < \sigma < 1 \tag{17}$$

where $(h_t^p)^{\sigma}$ represents the direct transfer of human capital from parents to their offspring, which operates in conjunction with the indirect transmission channel e_t^p in determining h_{t+1}^c through the Cobb–Douglas function, with the parameters σ and $(1-\sigma)$ reflecting their relative contribution.

Replacing Eq. (3) with Eq. (17) in the optimization, the above modification does not change the optimal solution for e_t^p. However, the intergenerational income transmission function will be changed into the following form:

$$\ln y_{t+1}^c = \gamma \ln \bar{A} + \gamma\sigma \ln h_t^p + \gamma(1-\sigma) \ln y_t^p + \gamma(1-\sigma) \ln \frac{\alpha\beta\gamma(1-\sigma)}{1+\beta+\alpha\beta\gamma(1-\sigma)} + \gamma\varepsilon_t \tag{18}$$

which can be further simplified to:

$$\ln y_{t+1}^c = \theta_0 + \theta_1 \ln y_t^p + v_t \tag{19}$$

where the intercept θ_0, the income elasticity θ_1 and the residual v_t are given as follows:

$$\theta_0 = \gamma \ln \bar{A} + \gamma(1-\sigma) \ln \frac{\alpha\beta\gamma(1-\sigma)}{1+\beta+\alpha\beta\gamma(1-\sigma)} \tag{20}$$

$$\theta_1 = \sigma + \gamma(1-\sigma) \tag{21}$$

$$v_t = \gamma\varepsilon_t. \tag{22}$$

Eq. (21) suggests that the intergenerational income elasticity is determined by not only the two marginal returns but also the relative contribution of the two mechanisms (i.e., the direct transfer and indirect investment of human capital). Furthermore, a comparison of Eq. (15) in the benchmark model and Eq. (21) in the extended model leads to the following proposition:

Proposition 2. Ignoring the direct parent-to-offspring transfer of human capital will result in an underestimation of the intergenerational income elasticity and an overestimation of the intergenerational income mobility.

As the main prediction of our theoretical model, Proposition 2 will be verified empirically using China's nationally representative data in the following sections.

4. Estimation methods and data

4.1. Regression models

In the empirical analysis, we aim to verify the above theoretical propositions by estimating and comparing the intergenerational income elasticities based on a basic Ordinary Least Squares (OLS) model and a simultaneous equations (SE) model, which in turn correspond to the benchmark model and the extended model in Section 3 respectively.

Following the previous literature on intergenerational income elasticity (Jäntti et al., 2006; Solon, 2002)[3], the basic OLS model is specified as follows:

$$\ln y^c = \eta_0 + \eta_1 \ln y^p + \eta_2 edu^c + \eta_3 health^c + \eta_4 X + u_t \tag{23}$$

where y^c is the children's yearly income, and y^p represents parents' permanent income as measured by their average income in all sample years; edu^c indicates children's educational attainment (measured by years of schooling) and $health^c$ represents their health status (measured by height, BMI, presence of chronic diseases or self-rated health), which jointly reflect children's human capital stock. Eq. (23) can be considered as an empirical approximation of Eq. (13), with the parameter η_1 (i.e. the intergenerational income elasticity) being specified by Eq. (15) and the parameters η_2 and η_3 (i.e. the influence of human capital on income) being specified by Eq. (2). Vector X includes a set of control variables such as parents' age and occupation and children's age, gender and regional characteristics (urban/rural, coastal/inland). Consistent with the benchmark theoretical model, the above OLS model does not account for the direct transfer of human capital across generations.

In the SE model, we explicitly introduce the direct human capital transfer and study its impact on the intergenerational income elasticity. Specifically, the parent-to-offspring transmission of income, education and health can be specified as follows:

$$\ln y^c = \theta_0 + \theta_1 \ln y^p + \theta_2 edu^c + \theta_3 health^c + \theta_4 X + v_t \tag{24}$$

$$edu^c = \alpha_0 + \alpha_1 edu^p + \alpha_2 \ln y^p + \alpha_3 health^c + \alpha_4 X + \varepsilon_t \tag{25}$$

$$health^c = \gamma_0 + \gamma_1 health^p + \gamma_2 \ln y^p + \gamma_3 edu^c + \gamma_4 X + \mu_t \tag{26}$$

where Eqs. (24)–(26) are theoretically motivated by Eqs. (19) and (17), and they jointly form a system of simultaneous estimation equations characterizing the determination of children's income, education and health respectively. In Eqs. (25) and (26), we also consider the interaction of socio-economic factors to account for the income–education–health gradient as verified by the previous literature (Smith, 2004). Specifically, in Eq. (25), children's years of schooling (edu^c) are influenced by not only their parents' years of schooling edu^p (the direct transfer of education), but also their parents' income and their own health status; this is because parents' income may influence their financial capability of investing in children's education (the indirect transmission mechanism; see Taubman (1989) and Plug and Vijverberg (2005) for empirical support), and children's education may also be closely related to their own health status (education–health gradient; see Behrman (1996) and Glewwe, Jacoby, and King (2001) for empirical references). Similarly, in Eq. (26), children's health status ($health^c$) is influenced by their parents' health ($health^p$), their parents' income and their own educational attainment (empirical support can be found in Eriksson, Pan, & Qin, in Press; Goode, Mavromaras, & Zhu, 2014). Additionally, we also control the demographic and regional factors (as denoted by vector X) in all the three transmission equations.

In the above system, the interaction between children's education and health might cause endogeneity in Eqs. (25) and (26) due to reversed causality, i.e., $E[health^c \cdot \varepsilon_t] \neq 0$ in Eq. (25) and $E[edu^c \cdot \mu_t] \neq 0$ in Eq. (26).[4] To avoid the endogeneity bias, we use $health^p$ in Eq. (26) as the instrumental variable (IV) for $health^c$ in Eq. (25), and use edu^p in Eq. (25) as the IV for edu^c in Eq. (26). Consequently, we identify the SE model through Indirect Least Squares (ILS) by substituting the fitted values of edu^c and $health^c$ into Eq. (24) based on the IV estimation of Eqs. (25) and (26). The validity of this identification strategy depends on whether the IVs satisfy the "power condition" and "exclusion restriction". The former means that the IVs ($health^p$ and edu^p) should be highly correlated with their corresponding endogenous variables ($health^c$ and edu^c), and the latter means that the IVs should be uncorrelated with the corresponding residuals (ε_t and μ_t). As mentioned in Section 2, the "power condition" is widely supported by the empirical literature that shows strong correlation between parents' health (education) and their children's health (education). Moreover, this correlation will also be formally tested on our data using the F statistics for the significance of the IVs in the first stage regressions associated with Eqs. (25) and (26) (see Section 5.2). On the other hand, the "exclusion restriction" indicates that parents' health status should not directly influence their children's years of schooling, and parents' education should not have a direct impact on their children's health, i.e. there should not be a cross-transmission of human capital between the two dimensions (education and health). While there is continued debate regarding the validity of the above argument (see Currie, 2011 for a review on the "nature–nurture interaction" hypothesis), much recent literature has provided strong support for the "exclusion restriction". For example, based on the Demographic and Health Surveys in 22 developing countries, Desai and Alva (1998) find that parental education is not significantly correlated to infant mortality rates and children's height-for-age z scores; McCrary and Royer (2011) use the confidential natality data from Texas and California to find that mothers' years of schooling have no significant impact on infants' health status (birth weight); based on the NCDS data, Lindeboom, Liena-Nozal, and Klaauw (2009) also find no evidence of cross-transmission of human capital. For our

[3] In the previous studies, the OLS regression may exclude edu^c, $health^c$ or X; for comparison, we estimate the elasticity using different combinations of the explanatory variables in the basic OLS model (see Table 3).

[4] The endogeneity in Eqs. (25) and (26) may also be caused by the unobserved factors (such as genetic inheritance of cognitive ability and health condition) that influence both the human capital of children and of their parents. However, due to data limitation, we are not able to control such endogeneity using special samples or instrumental variables, which is an admitted limitation of our study.

Please cite this article as: Qin, X., et al., Intergenerational transfer of human capital and its impact on income mobility: Evidence from China, *China Economic Review* (2014), http://dx.doi.org/10.1016/j.chieco.2014.10.005

data, the "exclusion restriction" condition is empirically verified through a direct test of the cross-transmission between education and health by re-specifying Eqs. (25) and (26) in the SE estimation (see Section 5.3).

4.2. Data and sample description

The data used in this paper are from the China Health and Nutrition Survey (CHNS), which is jointly sponsored by the Carolina Population Center at the University of North Carolina and the Chinese Center for Disease Control and Prevention. CHNS has so far included nine waves (1989, 1991, 1993, 1997, 2000, 2004, 2006, 2009 and 2011). In each round of surveys, it uses a multistage random cluster process to draw a sample of about 4400 households with a total of 15,000 to 19,000 individuals in nine provinces (Guangxi, Guizhou, Heilongjiang, Henan, Hubei, Hunan, Jiangsu, Liaoning and Shandong) that vary substantially in geographic and economic characteristics. In each province, four counties are randomly selected based on a weighted sampling scheme; within each county, "communities" are randomly selected as secondary sampling units (SSU) which represent urban neighborhoods or rural villages. The household survey collects detailed information on the respondent's socio-demographic characteristics, health status, nutrition intake, medical care utilization, etc. The community survey provides information on the local hygiene, public service infrastructure, health care resources, insurance coverage, etc.

We use the adult sample from the 1989–2009 CHNS household survey data and match each father–child pair using the family relationship information (considering the low female labor participation rate in rural China, we do not consider the mother–child pairs). The sample restriction criteria are then applied as follows. First, we exclude children under 25 years old as they may still be

Table 2
Sample summary statistics of select variables.

Variable	Definition	Overall	Rural	Urban	Son	Daughter
Children's variables						
cincome	Yearly income	8943.3	8397.4*	9706.1	9163.3	8087.8
	(Yuan)	(12224.3)	(12608.1)	(11634.9)	(12864.3)	(9297.2)
cedu	Education	9.949	9.347*	10.789	9.900	10.136
	(years)	(3.342)	(2.613)	(4.005)	(3.192)	(3.870)
cheight	Height	166.7	166.6	166.9	168.9*	158.1
	(cm)	(7.399)	(6.946)	(7.993)	(6.120)	(5.424)
csrh	Self-rated health	2.912	2.890	2.948	2.908	2.928
	(1–4)	(0.708)	(0.710)	(0.704)	(0.718)	(0.662)
coverweight	Overweight	0.279	0.264	0.299	0.306*	0.172
	(1 = yes)	(0.448)	(0.441)	(0.458)	(0.461)	(0.378)
cchronic	Chronic diseases	0.0213	0.0113*	0.0351	0.0230	0.0143
	(1 = have)	(0.144)	(0.106)	(0.184)	(0.150)	(0.119)
cage	Age	31.90	31.07*	33.06	32.10*	31.14
	(years)	(6.183)	(5.052)	(7.332)	(6.227)	(5.963)
cgender	Gender	0.205	0.141*	0.294	–	–
	(1 = female)	(0.404)	(0.348)	(0.456)	–	–
Fathers' variables						
fincome	Avg yearly income	9062.4	8214.6*	10246.8	8945.2	9518.1
	(Yuan)	(8287.0)	(7837.9)	(8746.9)	(7921.2)	(9580.3)
fedu	Education	8.354	7.587*	9.425	8.205*	8.935
	(years)	(3.486)	(2.730)	(4.095)	(3.334)	(3.975)
fheight	Height	165.1	164.7*	165.7	165.1	165.1
	(cm)	(6.039)	(6.095)	(5.923)	(6.147)	(5.609)
fsrh	Self-rated health	2.665	2.612*	2.753	2.649	2.738
	(1–4)	(0.752)	(0.735)	(0.772)	(0.754)	(0.738)
foverweight	Overweight	0.380	0.326*	0.455	0.388	0.348
	(1 = yes)	(0.486)	(0.469)	(0.498)	(0.488)	(0.477)
fchronic	Chronic diseases	0.212	0.156*	0.290	0.208	0.226
	(1 = have)	(0.409)	(0.363)	(0.454)	(0.406)	(0.419)
fage	Age	60.41	58.33*	63.30	60.39	60.46
	(years)	(7.586)	(6.616)	(7.907)	(7.662)	(7.300)
fstate	Job sector	0.424	0.403*	0.453	0.399*	0.520
	(1 = state-owned)	(0.494)	(0.491)	(0.498)	(0.490)	(0.501)
Environmental variables						
Urban	Urban/rural	0.417	–	–	0.371*	0.599
	(1 = urban)	(0.493)	–	–	(0.483)	(0.491)
Coast	Costal/inland	0.410	0.403	0.420	0.405	0.430
	(1 = costal)	(0.492)	(0.491)	(0.494)	(0.491)	(0.496)
Sample size		1364	795	569	1085	279

Notes: (i) Data source: 1989–2009 China Health and Nutrition Survey (CHNS). (ii) * denotes 10% significance level in t-test for urban–rural or son–daughter differences; (iii) BMI = weight (kg) / squared height (sq. m); one is overweight when BMI ≥25; (iv) self-rated health 1–4 are "poor, fair, good and excellent", respectively; (v) all income are measured in 2009 *Yuan*.

in school (especially for those who receive higher education) and thus may not have a stable income. Second, we exclude individuals with no income or those whose income values are missing, as they cannot be used for income elasticity estimation. Third, we exclude observations with missing information on education, health, father's education and health, as well as other control variables. Our final study sample thus contains 1364 father–child pairs, among which urban residents account for 41.7% and the father–son relationship accounts for 79.5%. Table 2 reports the sample descriptive statistics of the main variables, which are analyzed as follows.

(1) *Income*. CHNS collects household and individual income data on a yearly basis. The annual income is calculated from seven main sources: home gardening, collective and household farming, raising livestock or poultry, collective and household fishing, small handicraft and small commercial household business, pension gratuity and wages paid by formal sectors. An interpolation method is applied for observations with missing information in one or some (but not all) of the income sources. We accept this official interpolation method, but drop the individuals with negative or missing total income values. We then convert all income measures into 2009 *Yuan* and apply the natural logarithm transformation. To smooth the impact of unobserved shocks to one's permanent income, we follow the literature convention (Mazumder, 2005; Solon, 1992) and take the average values of fathers' annual income in all sample waves as a proxy measure of fathers' permanent income. Although we cannot apply the same treatment to children as most of them do not have multiple-wave data on income, their yearly income may still serve as a valid proxy for permanent income, because (i) the life cycle impact on income is partially controlled by children's age in the regressions; (ii) recent studies on intergenerational income elasticity show that the single-year income can be a reliable proxy for one's permanent income for individuals aged between 30 and 40 (Bohlmark & Lindquist, 2006; Haider & Solon, 2006), which is indeed the case for a significant portion of our sample children; (iii) since children's income is the dependent variable in the regression, the measurement errors in this variable will not cause the error-in-variable problem and the associated estimation bias. As shown in Table 2, fathers' average permanent income is 9062.4 *Yuan*, while children's average income is 8943.3 *Yuan* for the full sample. Meanwhile, the incomes of both generations are higher in the urban areas compared to the rural areas, but no significant difference is found between the son sample and the daughter sample.

(2) *Education*. CHNS collects individuals' educational information by directly asking their years of formal schooling and the highest level of education attainment. Based on this information, we construct the continuous variables *ceduyear* and *feduyear* to measure children's and fathers' educational human capital stock. Table 2 indicates that the children's average years of schooling is 9.9 for the full sample; the urban residents are significantly better educated than their rural counterparts (10.8 vs. 9.3), and the female children sample has higher average education levels than the son sample (10.1 vs. 9.9). Possibly because of the Great Cultural Revolution and other historical factors, the sample fathers' average years of schooling is only 8.4, which is substantially lower than their offspring; moreover, the urban–rural gap in education is larger for fathers (9.4 vs. 7.6) than for their offspring.

(3) *Health*. CHNS has rich and comprehensive information on individual health and nutrition. Following the literature convention, we use the following variables to measure an individual's health capital stock: (i) Height. Largely determined by childhood nutrition intake and medical conditions, height is commonly used in health economics as a measure of a person's long-term health, especially in developing countries (Strauss & Thomas, 1998). In our sample, the mean of children's height is 166.7 cm, which is about 2-cm taller than that of their fathers', reflecting an improvement of nutrition and health status in the past decades. (ii) Self-rated health (SRH). CHNS asks respondents to rate their health status since the 1997 wave,[5] with "Excellent, Good, Fair and Poor" as possible answers (valued sequentially as 4, 3, 2 and 1).[6] In our sample, children's average SRH is 2.9, which is greater than that of their fathers' (2.7); meanwhile, the proportion of children who choose "Excellent" and "Poor" is smaller than fathers, reflecting not only the improvement but also a more concentrated distribution of health in the younger generation. (iii) Overweight. This dummy variable is generated based on the Body Mass Index (BMI) calculated from respondents' self-reported height and weight (BMI = weight (kg) / height (m)^2). BMI is widely used as a measure of health status among adults; according to the WHO recommended thresholds, a person is defined as overweight when his BMI is greater than or equal to 25. As shown by Table 2, 27.9% of children and 38.0% of parents in our sample are overweight, and the prevalence of overweight in the son sample is higher than that in the daughter sample. (iv) Presence of chronic diseases. This dummy variable is generated based on Xie (2011), i.e., if one reports in the disease history that he suffers from hypertension, diabetes, stroke or transient ischemic attack, asthma (either told or diagnosed by a doctor) or other chronic conditions (e.g., heart disease or chest pain) during the past 4 weeks, then the respondent is considered to have chronic diseases. Table 2 shows that the prevalence of such diseases is 2.13% among children and 21.2% among fathers, with higher prevalence found in the urban sector. Among the above health indicators, we use height as the main measure, and use the others for robustness check in Section 5.3.

(4) *Other control variables*: From Table 2, the sample average age for children is 32 (rural children are younger), which is close to the optimal age for the permanent income approximation (Bohlmark & Lindquist, 2006; Haider & Solon, 2006). Fathers who work in the state-owned sector (including government and state-owned enterprises) account for 42.4% of the sample, and the ratio is higher in urban areas than rural areas (45.3% vs. 40.3%). Additionally, our regressions also control for fathers' age (averages 60.4 for the whole sample), urban status (averages 41.7%) and whether living in the coastal region (averages 41.0%).

[5] The question for SRH has been slightly changed since the 2009 wave, and becomes "How do you rate your life at present" (U420), with "Excellent, Good, Fair, Bad, Very bad" as possible answers. CHNS has officially confirmed the comparability of this question with the previous ones on self-rate health. For consistency, we combine "Bad" and "Very bad" into "Poor" as the lowest SRH value.

[6] Some studies prefer to use the ordered-probit or other sequential choice models on such variables. However, Ferrer-i-Carbonell and Frijters (2004) point out that traditional OLS regression on such discrete psychometric indicators may give more precise standard error estimation compared to the non-linear models. For robustness test purpose, we use both OLS and ordered-probit for the evaluation of self-reported health, and get consistent results.

Table 3
Estimation results of the OLS model.

Dependent variable: log(cincome)				
Variable	(1)	(2)	(3)	(4)
log(fincome)	0.536***	0.457***	0.487***	0.429***
	(0.034)	(0.035)	(0.035)	(0.035)
cedu	–	–	0.034***	0.029***
	–	–	(0.007)	(0.007)
cheight	–	–	0.013***	0.010**
	–	–	(0.004)	(0.004)
cage	–	0.069**	–	0.071**
	–	(0.033)	–	(0.033)
$cage^2$	–	−0.001*	–	−0.001*
	–	(0.000)	–	(0.000)
cgender	–	−0.058	–	0.058
	–	(0.062)	–	(0.077)
fage	–	−0.016	–	−0.015
	–	(0.040)	–	(0.038)
$fage^2$	–	0.000	–	0.000
	–	(0.000)	–	(0.000)
fstate	–	−0.463***	–	−0.438***
	–	(0.057)	–	(0.058)
Urban	–	0.061	–	0.006
	–	(0.055)	–	(0.055)
Coast	–	0.238***	–	0.202***
	–	(0.051)	–	(0.052)
Constant	4.116***	4.112***	1.998***	2.319*
	(0.304)	(1.144)	(0.624)	(1.374)
Sample size	1364	1364	1364	1364
R-square	0.185	0.251	0.205	0.261

Notes: (i) Data source: 1989–2009 China Health and Nutrition Survey (CHNS). (ii) ***, ** and * denote 1%, 5% and 10% significance levels, respectively; (iii) household-level clustered standard errors are in parentheses; (iv) please refer to Table 2 for definitions of variables.

5. Empirical results

5.1. Results of the OLS model

Table 3 presents the OLS regression results based on Eq. (23). Model (1) has a simple univariate specification that focuses on the intergenerational income elasticity without considering the impacts of other covariates. The estimated income elasticity is 0.536, which means a 10% increase in fathers' income will on average lead to a 5.36% increase in their children's income. The estimate is relatively high compared to other countries,[7] reflecting the low intergenerational income mobility and the persistently high income inequality in China. To check the robustness of the linear specification of the simple OLS model, we also perform a non-parametric estimation on the relationship between children's and fathers' log-income based on the (Epanechnikov) kernel regression (see Fig. 1). The results indicate that the parametric and non-parametric estimations basically coincide with each other in areas where the sample data are densely populated, but the two estimations somewhat diverge in the boundary areas where the data points are sparse.

To further control the impact of personal and environmental factors on children's income, Model (2) introduces the variables in vector *X* into the regression, and sees the estimated elasticity being reduced to 0.457, which is close to the estimation in the comparable studies (e.g., Zhang & Eriksson, 2010). Models (3) and (4) focus on the impact of children's human capital accumulation, and add children's education and health status to the estimation of Models (1) and (2) respectively. The results suggest that the estimated intergenerational income elasticity decreases from 0.536 to 0.487 in the simple OLS model after the human capital impact is accounted for. For the multivariate OLS model that controls the vector *X*, the estimated elasticity further decreases to 0.429, which means that a 10% increase in fathers' income will on average lead to a 4.29% increase in their children's income. These findings pinpoint the non-negligible role of children's human capital stock in the intergenerational income transmission process. In the following, we will use the estimated elasticity of the multivariate OLS model (0.429) as the benchmark to compare with the SE model estimates.

Table 3 also reveals the influence of other factors on children's income. For example, the coefficient estimate of education is 0.029 (significant at 1% level), which means the return to one-year increase in formal schooling is a 2.9% increase in the yearly income. This estimate is consistent with the prior studies (e.g., Sun, 2004 finds that the returns to schooling in China during the 1980s and 1990s are 3–5%), but somewhat lower than the more recent estimates (e.g., Deng & Ding, 2013 find that the returns are higher than 7% in both rural and urban China after 2000), possibly because a large portion of our sample comes from the pre-2000 period. As another human capital indicator, health (measured by height) also has a positive effect on income (the coefficient is 0.010 and significant at 5% level),

[7] According to Table 1, the estimated elasticity is between 0.45–0.48 in the U.S., 0.3–0.4 in the Western European countries, 0.2 in the Nordic countries, and 0.4–0.5 in the Latin American countries.

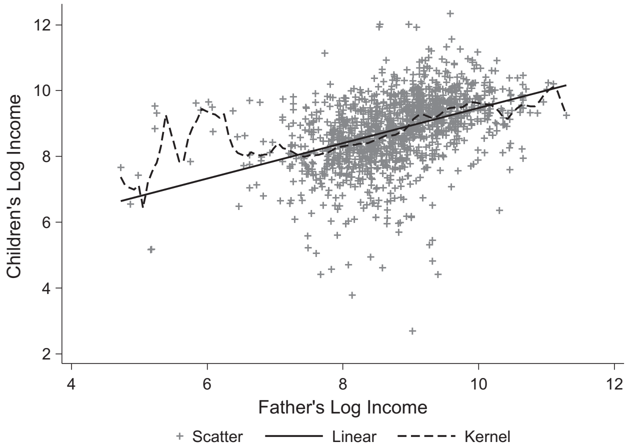

Fig. 1. Parametric and non-parametric estimations of intergenerational income elasticity. Notes: (i) The figure presents the predicted relationship between children's and fathers' logarithmic income based on parametric and non-parametric regressions using the same sample drawn from the 1989–2009 China Health and Nutrition Survey (CHNS); (ii) the solid (straight) line represents a linear regression; (iii) the dashed line represents a kernel regression using the Epanechnikov kernel function when the optimal bandwidth is derived by minimizing the mean integrated squared error (MISE).

which is in line with the findings in the relevant literature (e.g., Becker, 2007; Currie & Madrian, 1999). Moreover, the impact of children's age on their own income has an inverted U shape, reflecting the diminishing marginal return of age (which can also be considered as the potential working experience). With regard to regional differences, children's income in coastal areas is on average 20.2% higher than that in inland areas (significant at 1% level), but the difference between the urban and rural areas is not significant when other factors are considered.

5.2. Results of the SE model

The SE model illustrated by Eqs. (24)–(26) characterizes the more complicated mechanisms of intergenerational transmission. To identify the model, we first use the IV techniques to estimate Eqs. (25) and (26), and then substitute the fitted values of education and health into Eq. (24) to apply the ILS estimation. The standard errors and the associated statistical inference of the ILS estimates are based on the bootstrap procedure. Table 4 reports the main results of the SE model.

We first focus on the equations of human capital transmission. Models (2) and (4) report the second-stage estimation of the IV regressions for the transmissions of education and health, respectively. Model (2) suggests a significant direct transfer of education across generations: after controlling the endogeneity of parental education, a 1-year increase in fathers' formal schooling will averagely lead to a 0.310-year increase in children's years of schooling (when other factors are controlled, similarly hereinafter). As predicted by the human capital theory, well-educated parents tend to pay more attention to their children's cultivation, thus the high educational attainment can be transferred from parents to offspring through genetic and non-genetic channels. This finding is also confirmed by other empirical studies, e.g., Hertz et al. (2007) estimate the intergenerational correlations of schooling in 42 countries and find a strong persistence of educational attainment: the coefficient is around 0.6 in African countries, 0.46 in the U.S., about 0.4 in the Western European countries, and lowest in the Nordic countries. Model (4) shows that the father–child transmission of health is also evident: a 1-cm increase in fathers' height will on average lead to a 0.485-cm increase in children's height. This intergenerational transmission in health indicators is also empirically supported by other studies (e.g., Currie & Moretti, 2007; Eriksson et al., 2005; Eriksson et al., in Press).

In addition, Models (2) and (4) also demonstrate the impact of fathers' income on children's education and health: a 10% increase in fathers' income will lead to a 0.039-year increase in children's formal schooling (significant at 1% level) and 0.03-cm increase in children's height (not statistically significant). This to some extent reflects the indirect human capital transmission mechanism, i.e., parents' education can influence their financial capability (measured by income) of investing in their children and thus indirectly impact their children's human capital. Apart from the above factors, children's schooling is also significantly correlated with their living locations and fathers' employment sectors, while their height is significantly correlated with their age, gender, living locations, fathers' age and working sectors.

Furthermore, Models (3) and (5) report the corresponding first-stage estimations of the IV regressions. As expected, children's heights and schooling are significantly correlated with their corresponding IVs—fathers' heights and schooling—with F values of 426.88 and 139.98, which far exceed the recommended threshold ($F \geq 10$) in the literature (Stock, Wright, & Yogo, 2002), suggesting that the IVs are not likely to be weak.

Next, we turn our attention to the estimation of intergenerational income transmission. Model (1) shows the ILS estimation results of Eq. (24), which indicates that the estimated income elasticity will increase from 0.429 (in the benchmark OLS model) to 0.481 (in the SE model) after accounting for the direct transfer of human capital. This in turn suggests that ignoring the direct human capital transmission tends to result in an overestimation of parent-to-offspring income mobility, a finding that is consistent with the

Please cite this article as: Qin, X., et al., Intergenerational transfer of human capital and its impact on income mobility: Evidence from China, *China Economic Review* (2014), http://dx.doi.org/10.1016/j.chieco.2014.10.005

Table 4
Estimation results of the simultaneous equations model.

	(1)	(2)	(3)	(4)	(5)
	Income	Education	1st Stg Edu	Health	1st Stg Hlth
Variable	log(cincome)	cedu	cheight	cheight	cedu
log(fincome)	0.481***	0.394***	0.320*	0.305	0.393***
	(0.043)	(0.106)	(0.172)	(0.194)	(0.106)
fedu	–	0.310***	0.012	–	0.310***
	–	(0.026)	(0.043)	–	(0.026)
fheight	–	–	0.485***	0.485***	−0.001
	–	–	(0.023)	(0.023)	(0.014)
cedu	−0.043	–	–	0.038	–
	(0.026)	–	–	(0.137)	–
cheight	0.017*	−0.002	–	–	–
	(0.010)	(0.030)	–	–	–
cage	0.077**	0.138	−0.525**	−0.531**	0.139
	(0.034)	(0.119)	(0.194)	(0.192)	(0.119)
cagesq	−0.001**	−0.002	0.006**	0.006**	−0.002
	(0.000)	(0.002)	(0.003)	(0.003)	(0.002)
cgender	0.133	−0.140	−11.153***	−11.148***	−0.120
	(0.124)	(0.395)	(0.340)	(0.338)	(0.209)
fage	−0.023	−0.141	0.564**	0.570**	−0.142
	(0.042)	(0.150)	(0.246)	(0.244)	(0.151)
fagesq	0.000	0.001	−0.004*	−0.004*	0.001
	(0.000)	(0.001)	(0.002)	(0.002)	(0.001)
fstate	−0.465***	−0.591***	−0.981***	−0.959***	−0.589***
	(0.061)	(0.185)	(0.299)	(0.304)	(0.183)
Urban	0.084	0.901***	1.140***	1.106***	0.899***
	(0.064)	(0.195)	(0.309)	(0.347)	(0.190)
Coast	0.204***	0.119	1.464***	1.460***	0.117
	(0.056)	(0.189)	(0.287)	(0.286)	(0.176)
Constant	1.534	6.411	75.911***	75.674***	6.272
	(2.004)	(6.415)	(8.063)	(7.977)	(4.949)
Sample size	1364	1364	1364	1364	1364
R-square	0.254	0.186	0.559	0.560	0.186
IV 1st Stage F	–	–	426.88***	–	139.98***

Notes: (i) Data source: 1989–2009 China Health and Nutrition Survey (CHNS). (ii) ***, ** and * denote 1%, 5% and 10% significance levels, respectively; (iii) bootstrap standard errors (1000 replications) are reported for model (1), household-level clustered standard errors are for other models; (iv) IV 1st Stage F is the F statistics for the significance of the IVs in the first stage regressions; (v) please refer to Table 2 for definitions of variables.

theoretical Proposition 2. To formally test the statistical significance of the difference between the two models, we conduct a Hausman test by treating the main coefficients of the benchmark OLS model as efficient estimators and those of the SE model as consistent estimators, and find that the P value is about 0.008, suggesting the null hypothesis that these two sets of coefficients are identical can be rejected at 1% level. With regard to the other control variables, the SE model (Model (1) in Table 4) gives similar estimation compared to the OLS model (Model (4) in Table 3). Overall, the above results indicate that the intergenerational income elasticity is underestimated by the traditional model (OLS) that leaves out the direct transmission of human capital, thus China's income mobility can be even poorer than what the recent studies have shown (e.g. Chen & Yuan, 2012; Zhang & Eriksson, 2010).

5.3. Robustness tests

In addition to the above main results, we will also check the robustness of the empirical specifications by running regressions on various sub-samples, using alternative health indicators, and testing the cross-transmission of human capital.

Table 5
Sub-sample estimates of intergenerational income elasticity.

	(1)	(2)	(3)	(4)	(5)
	Overall	Son	Daughter	Rural	Urban
OLS model	0.429***	0.415***	0.464***	0.409***	0.421***
	(0.035)	(0.041)	(0.067)	(0.047)	(0.050)
SE model	0.481***	0.468***	0.520***	0.501***	0.463***
	(0.043)	(0.051)	(0.075)	(0.069)	(0.059)
Sample size	1364	1085	279	795	569

Notes: (i) Data source: 1989–2009 China Health and Nutrition Survey (CHNS). (ii) ***, ** and * denote 1%, 5% and 10% significance levels, respectively; (iii) for the OLS model, household-level clustered standard errors are in parentheses; (iv) for the SE model, bootstrap standard errors are in parentheses; (v) all regressions control for the individual- and region-level characteristics.

Please cite this article as: Qin, X., et al., Intergenerational transfer of human capital and its impact on income mobility: Evidence from China, *China Economic Review* (2014), http://dx.doi.org/10.1016/j.chieco.2014.10.005

Table 6
Estimated intergenerational income elasticity using alternative health indicators.

	(1) Height	(2) SRH	(3) Overweight	(4) Chronic Diseases
OLS model	0.429***	0.369***	0.428***	0.433***
	(0.035)	(0.036)	(0.035)	(0.035)
SE model	0.481***	0.426***	0.510***	0.488***
	(0.043)	(0.050)	(0.043)	(0.044)
Sample size	1364	1364	1364	1364

Notes: (i) Data source: 1989–2009 China Health and Nutrition Survey (CHNS). (ii) ***, ** and * denote 1%, 5% and 10% significance levels, respectively; (iii) for the OLS model, household-level clustered standard errors are in parentheses; (iv) for the SE model, bootstrap standard errors are in parentheses; (v) all regressions control for the individual- and region-level characteristics.

First, we estimate both the OLS and the SE models in different sub-samples based on gender and regional stratifications. Table 5 summarizes the estimated elasticities for the samples of father–son, father–daughter, father–child in rural areas and father–child in urban areas. The results suggest that the father–daughter elasticity is higher than the father–son elasticity in both the OLS and SE models, which is consistent with the prior findings (e.g., Gong et al., 2012; Raaum et al., 2007). This result in turn indicates that women are at disadvantage in the intergenerational transmission in China as they lack the socio-economic resources to break the intergenerational persistence of income. More interestingly, we also find that the urban–rural differences of the father–child elasticity are reversed in the OLS and SE models: the elasticity in urban areas is slightly larger than that in rural areas (0.421 vs. 0.409) in the OLS model, while the difference is "0.463 vs. 0.501" in the SE model, suggesting that the income mobility in urban areas is actually higher than that in rural areas after considering the direct transmission of human capital. One possible reason is that the allocation of labor resources is more market-oriented and thus one's socio-economic status is more changeable in urban China. This result also corresponds to the fact that the Gini coefficient in rural China has been persistently higher than the urban one (Sutherland & Yao, 2011), reflecting that income is more unequally distributed in China's rural sector. Lastly, we find that the estimated elasticity of the SE model is robustly larger than that of the OLS model in each sub-sample, which validates the main conclusion drawn from our empirical analysis.

Next, we test the sensitivity of health measures by using alternative indicators of individual health status in the SE regressions. Table 6 reports the estimated father–child income elasticity using height, SRH, overweight status and presence of chronic diseases. In summary, we find that the elasticity estimates are quantitatively consistent across different health indicators, and that the SE model gives robustly larger estimates than the OLS model in all regressions.

Lastly, we discuss the cross-transmission of human capital. In the labor economics literature, the continued debate on the "nature–nurture interaction" focuses on whether there is a reciprocal effect between the inborn factors (e.g. genes) and the acquired traits (e.g. family environment) in the process of intergenerational transmission (Currie, 2011). In the current context, the existence of nature–nurture interaction would suggest a cross-transmission of human capital, i.e. fathers' health (nature) would impact children's education (nurture), and vice versa (see Fig. 2 for the illustrated mechanisms). In the above SE model, we simply omit the cross-transmission of human capital in both the education and health regressions. Here we estimate the cross-transmission for robustness check purpose by adding fathers' health into Eq. (25) and fathers' education into Eq. (26).[8] Table 7 reports the results of the new model, which provides no supporting evidence on the "nature–nurture interaction": the correlation is not statistically significant between fathers' height and children's schooling and between fathers' schooling and children's height, respectively. This finding on the non-existence of cross-transmission is also confirmed by other empirical studies (e.g., Lindeboom et al., 2009; McCrary & Royer, 2011), and in turn provides validating support for the "exclusion restriction" condition for our original IV strategy. Furthermore, the estimated income elasticity remains statistically significant and quantitatively similar with the main model, which further verifies the robustness of our previous findings.

6. Conclusions

Social mobility is an important measure of the degree to which a society gives equal opportunities to its members, and it largely determines the sustainability of a country's economic development. However, looking back upon history, the lack of social mobility has been a malady in China for millennia. Since the Wei dynasty (220–265), social mobility had been deficient for a long time due to the nine-rank system, where only the rich and powerful could be selected as candidates for the privileged administrative officials. After the Sui dynasty (581–619), the imperial examination system started to give the poor a chance to join the upper classes, but it was not sufficient to fundamentally remedy the lack of social mobility for various reasons.[9] In modern times before the 1980s, a free labor market was still absent in China under the centrally planned regime where the demand and supply of labor were strictly controlled by the government. Since the 1980s, the labor market has gradually developed and taken its role in determining the market wages,

[8] With the addition of variables, Eqs. (25) and (26) cannot be identified using the original IV strategy and thus will be estimated using OLS with their fitted values employed in the ILS estimation of Eq. (24).

[9] For example, some people in the lower class were banned from taking the imperial examination.

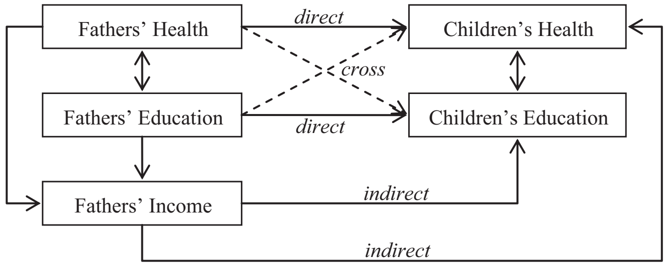

Fig. 2. Mechanisms of intergenerational transmission of human capital. Notes: The solid lines represent the direct and indirect transmission paths of human capital, and the dashed lines represent the (hypothesized) paths of cross-transmission of human capital.

enabling the measurement of social mobility by estimating the intergenerational elasticity of income. However, under the influences of the one-child policy, the household registration (*Hukou*) system, the reforms of social security and the enrollment expansion of higher education, the mechanisms of income transmission in China become more and more complicated. Nevertheless, it has been commonly recognized that the continuous development of China's economy is increasingly challenged by the lack of social mobility despite its impressive growth record during the past decades. In recent years, several empirical studies have shown that the intergenerational elasticity of income in China is internationally high, implying that the society does not give equal economic opportunities to all its members. In this paper, we aim to address the root causes of the problem by studying the impact of parent-to-offspring transmission of human capital on the intergenerational income mobility.

In the theoretical analysis, we incorporate the direct and indirect transmission of human capital into a classical three-period overlapping-generations (OLG) framework, which enables us to analyze the determination mechanism of parents' optimal investment in their children's human capital and the implied income mobility across generations. In this framework, parents' investment behavior is not only motivated altruistically by the return on human capital investments in terms of children's future income (the indirect human capital transmission), but also influenced by the efficiency of the direct transfer of human capital across generations. The most important conclusion derived from the theoretical model is that the intergenerational elasticity of income will be underestimated if we omit the direct transfer of human capital. In the empirical analysis, we verify the theoretical prediction using the nationally representative CHNS data and a simultaneous equations model accounting for the intergenerational transmission of income, health and education, comprehensively. The results show that the direct transfer of education and health is significant with coefficients of 0.310 and 0.485 respectively, reflecting that the opportunities in education and health are not equally/randomly distributed, i.e., children's development opportunities depend greatly on their parents' human capital. Meanwhile, after controlling the direct transfer of human capital across generations, the estimated income elasticity increases to 0.481, which is higher than that given by the benchmark OLS model (0.429). This finding is consistent with the theoretical prediction, suggesting that the traditional estimations in the literature, without considering the direct transmission mechanism of human capital, may have painted an over-optimistic picture on China's income mobility. Our sub-sample analyses further indicate that the intergenerational income elasticity is higher for the father–daughter pairs than the father–son pairs and for the urban areas than the rural areas, suggesting that the vulnerable groups (such as women and rural residents) are put into further disadvantage in development by the intergenerational persistence of socio-economic status.

Table 7
Robustness test results assuming cross-transmission of human capital.

	(1)	(2)	(3)
	Income	Education	Health
Variables	log(cincome)	cedu	cheight
log(fincome)	0.471***	0.380***	0.279
	(0.041)	(0.106)	(0.173)
fedu	–	0.310***	−0.021
	–	(0.026)	(0.045)
fheight	–	−0.020	0.485***
	–	(0.017)	(0.023)
cedu	−0.032	–	0.104**
	(0.024)	–	(0.044)
cheight	0.019**	0.039**	–
	(0.009)	(0.017)	–
Sample size	1364	1364	1364

Notes: (i) Data source: 1989–2009 China Health and Nutrition Survey (CHNS). (ii) ***, ** and * denote 1%, 5% and 10% significance levels, respectively; (iii) model (1) is estimated by Indirect Least Squares, and models (2) & (3) are estimated by Ordinary Least Squares; (iv) bootstrap standard errors are in parentheses; (v) all regressions control for the individual- and region-level characteristics; (vi) please refer to Table 2 for definitions of variables.

Our findings provide several policy implications for remedying the lack of social mobility and promoting the equity in opportunities in China: (i) given the important role of human capital transmission in determining the offspring socio-economic status, emphasis should be put on promoting the accumulation of human capital in the impoverished areas and disadvantaged groups through "equal opportunity" reforms (e.g., encouraging the enrollment of high schools and colleges in rural areas, facilitating access to basic health care among migrant workers, etc.); the interaction between income mobility and human capital transmission suggests that such reforms are conducive to alleviating the negative impact of income inequality and may help China avoid the middle-income trap. (ii) Since the intergenerational transmission of education and health is remarkable, the above "equal opportunity" reforms promise to not only facilitate the human capital accumulation for the current generation but also improve it by the "multiplying effect" for future generations, which is beneficial to the long-run economic growth of China; likewise, the current education and health system reforms by the Chinese government (such as quality improvement in the rural primary and secondary schools and the establishment of a universal health insurance system) are also likely to play a crucial role in maintaining the sustainable and harmonious social development in China in the long run. (iii) Our theoretical analysis also suggests that the intergenerational elasticity of income depends to a large extent on the marginal return of human capital investment on income; thus, in order to improve social mobility, it is important to further enhance the efficiency of labor market institutions towards better allocation of human resources. To achieve this goal, the following policy reforms are recommended: gradually alleviating the restrictions on labor migration under the household registration (*Hukou*) system, which gives the poor a chance for better labor market returns through migration; phasing out the attachment of social welfares (e.g. health insurance and housing funds) with one's working location so as to prevent the "job lock" and facilitate the regional mobility of labor force; finally, reforming the wage determination mechanisms to allow wage rates to reflect the marginal productivity of human capital, which in turn gives more incentives for human capital investment in both the private and public sectors.

Acknowledgment

This study is supported by the National Natural Science Foundation of China (71103009), Ministry of Education of China (12JZD036), and Beijing Higher Education Young Elite Teacher Project (YETP0039). We are grateful to two anonymous referees for their helpful comments and we thank UNC Carolina Population Center for providing the data. The authors are responsible for all remaining errors.

References

Akbulut, M., & Kugler, A.D. (2007). *Inter-generational transmission of health status in the U.S. among natives and immigrants.* Mimeo. University of Houston.
Anger, S., & Heineck, G. (2010). Do smart parents raise smart children? The intergenerational transmission of cognitive abilities. *Journal of Population Economics*, *23*(3), 1255–1282.
Becker, G.S. (1993). *Human capital: A theoretical and empirical analysis with special reference to education.* University of Chicago Press.
Becker, G.S. (2007). Health as human capital: Synthesis and extensions. *Oxford Economic Papers*, *59*(3), 379–410.
Becker, G.S., & Tomes, N. (1979). An equilibrium theory of the distribution of income and intergenerational mobility. *Journal of Political Economy*, *87*(6), 1153–1189.
Behrman, J. (1996). The impact of health and nutrition on education. *The World Bank Research Observer*, *11*(1), 23–37.
Björklund, A., & Jäntti, M. (2009). Intergenerational income mobility and the role of family background. In Wiemer Salverda, Brian Nolan, & Tim Smeeding (Eds.), *Handbook of Economic Inequality*. Oxford, UK: Oxford University Press.
Björklund, A., Eriksson, T., & Jäntti, M. (2010). IQ and family background: Are associations strong or weak? *The B.E. Journal of Economic Analysis & Policy*, *10*(1) (Article 2).
Black, E., Devereux, J., & Salvanes, K. (2005). Why the apple doesn't fall far: Understanding intergenerational transmission of human capital. *American Economic Review*, *95*(1), 437–449.
Black, S., Devereux, P., & Salvanes, K. (2009). Like father, like son? A note on the intergenerational transmission of IQ scores. *Economic Letters*, *105*(1), 138–140.
Blumenthal, D., & Hsiao, W. (2005). Privatization and its discontents — The evolving Chinese health care system. *New England Journal of Medicine*, *353*, 1165–1170.
Bohlmark, A., & Lindquist, M. (2006). Life-cycle variations in the association between current and lifetime income: Replication and extension for Sweden. *Journal of Labor Economics*, *24*(4), 879–896.
Borjas, G. (1992). Ethnic capital and intergenerational mobility. *Quarterly Journal of Economics*, *107*(1), 123–150.
Chen, L., & Yuan, Z. (2012). Trends and mechanisms of intergenerational income mobility in China. *The Journal of World Economy*, *72*(6), 115–131 (in Chinese).
Chevalier, A. (2004). Parental education and child's education: A natural experiment. *IZA discussion paper no. 1153*.
Chevalier, A., Denny, K., & McMahon, D. (2009). A multi-country study of intergenerational educational mobility. In Peter Dolton, Rita Asplundh, & Erling Barth (Eds.), *Education and inequality across Europe*. London: Edward Elgar.
Corak, M., & Heisz, A. (1999). The intergenerational earning and income mobility of Canadian men: evidence from longitudinal income tax data. *The Journal of Human Resources*, *34*(3), 504–533.
Currie, J. (2011). Inequality at birth: Some causes and consequences. *American Economic Review*, *101*(3), 1–22.
Currie, J., & Madrian, B. (1999). Health, health insurance and the labor market. *Handbook of Labor Economics*, *3*, 3309–3416 (Part C).
Currie, J., & Moretti, E. (2007). Biology as destiny? Short- and long-run determinants of intergenerational transmission of birth weight. *Journal of Labor Economics*, *25*(2), 231–264.
Deng, Q., Bjorn, G., & Li, S. (2012). Intergenerational income persistency in urban China. *IZA discussion paper no. 6907*.
Deng, F., & Ding, X. (2013). The long-term trends in rates of returns to education in China. *Statistical Research*, *30*(7), 39–47 (in Chinese).
Desai, S., & Alva, S. (1998). Maternal education and child health: Is there a strong causal relationship? *Demography*, *35*(1), 71–81.
Dunn, C. E. (2007). The intergenerational transmission of lifetime earnings: Evidence from Brazil. *The B.E. Journal of Economic Analysis & Policy*, *7*(2) (Article 5).
Eriksson, T., Bratsberg, B., & Raaum, O. (2005). Earnings persistence across generations: Transmission through health? *Memorandum 35/2005*. Oslo University, Dept. of Economics.
Eriksson, T., Pan, J., & Qin, X. (2014s). The intergenerational inequality of health in China. *China Economic Review*. http://dx.doi.org/10.1016/j.chieco.2014.06.005 (in Press).
Ferrer-i-Carbonell, A., & Frijters, P. (2004). How important is methodology for the estimates of the determinants of Happiness? *The Economic Journal*, *114*(497), 641–659.
Glewwe, P., Jacoby, H., & King, E. (2001). Early childhood nutrition and academic achievement: a longitudinal analysis. *Journal of Public Economics*, *81*(3), 345–368.
Gong, H., Leigh, A., & Meng, X. (2012). Intergenerational income mobility in urban China. *Review of Income and Wealth*, *58*(3), 481–503.
Goode, A., Mavromaras, K., & Zhu, R. (2014). Family income and child health in China. *China Economic Review*, *29*, 152–165.

Please cite this article as: Qin, X., et al., Intergenerational transfer of human capital and its impact on income mobility: Evidence from China, *China Economic Review* (2014), http://dx.doi.org/10.1016/j.chieco.2014.10.005

Grawe, N. D. (2004). Intergenerational mobility for whom? The experience of high- and low-earnings sons in international perspective. In M. Corak (Ed.), *Generational Income Mobility in North America and Europe*. Cambridge, UK: University Press (Chapter 4, 58-89).
Guo, C., & Min, W. (2007). Research on the relationship between education and intergenerational income mobility of Chinese urban household. *Educational Research*, *29*(6), 3–14 (in Chinese).
Haider, S., & Solon, G. (2006). Life-cycle variation in the association between current and lifetime earnings. *American Economic Review*, *96*(4), 1308–1320.
Hertz, T., Jayasundera, T., Piraino, P., Selcuk, S., Smith, N., & Verashchagina, A. (2007). The inheritance of educational inequality: International comparisons and fifty-year trends. *The B.E. Journal of Economic Analysis & Policy*, *7*(2) (Article 10).
Hussain, M. A., Munk, M. D., & Bonke, J. (2008). How sensitive is Intergenerational earnings mobility to different measures. *Danish National Centre for Social Research, Working Paper* 07:2008.
Jäntti, M., Bratsberg, B., Røed, K., Raaum, O., Naylor, R., Österbacka, E., et al. (2006). *American exceptionalism in a new light: A comparison of intergenerational earnings mobility in the Nordic countries, the United Kingdom and the United States.* IZA discussion papers No. 1938.
Lee, C. I., & Solon, G. (2009). Trends in intergenerational income mobility. *The Review of Economics and Statistics*, *91*(4), 766–772.
Lefranc, A., & Trannoy, A. (2005). Intergenerational earnings mobility in France: Is France more mobile than the US? *Annales d'Economie et de Statistique*, 57–77.
Leigh, A. (2007). Intergenerational mobility in Australia. *The B.E. Journal of Economic Analysis & Policy*, *7*(2) (Article 7).
Li, Y. (2006). Institutional change and educational inequality: Mechanisms in educational stratification in urban China (1966–2003). *Social Sciences in China*, *27*(4), 97–109 (in Chinese).
Lindeboom, M., Liena-Nozal, A., & Klaauw, B. (2009). Parental education and child health: Evidence from a schooling reform. *Journal of Health Economics*, *28*(1), 109–131.
Liu, J. (2006). Expansion of higher education in China and inequality in entrance opportunities: 1978–2003. *Society*, *26*(3), 158–209 (in Chinese).
Loureiro, M., Sanz-de-Galdeano, A., & Vuri, D. (2010). Smoking habits: Like fathers, like son, like mother, like daughter? *Oxford Bulletin of Economics and Statistics*, *72*(6), 717–743.
Loury, G. (1981). Intergenerational transfers and the distribution of earnings. *Econometrica*, *49*(4), 843–867.
Mazumder, B. (2005). Fortunate sons: New estimates of intergenerational mobility in the United States using social security earnings data. *The Review of Economics and Statistics*, *87*(2), 235–255.
McCrary, J., & Royer, H. (2011). The effect of female education on fertility and infant health: Evidence from school entry policies using exact date of birth. *American Economic Review*, *101*(1), 158–195.
Moaz, Y., & Moav, O. (1999). Intergenerational mobility and the process of development. *Economic Journal*, *109*(458), 677–697.
Nakamura, T., & Murayama, Y. (2011). Education cost, intergenerational mobility, and income inequality. *Economic Letters*, *112*(3), 266–269.
Nicoletti, C., & Ermisch, J. F. (2007). Intergenerational earnings mobility: changes across cohorts in Britain. *The B.E. Journal of Economic Analysis & Policy*, *7*(2) (Article 9).
Nilsen, Ø. A., Vaage, K., Aakvik, A., & Jacobsen, K. (2008). *Estimates of intergenerational elasticities based on lifetime earnings.* IZA discussion papers No. 3709.
Pekkarinen, T., Uusitalo, R., & Kerr, S. (2009). School tracking and intergenerational income mobility: Evidence from the Finnish comprehensive school reform. *Journal of Public Economics*, *93*(7), 965–973.
Piraino, P. (2007). Comparable estimates of intergenerational income mobility in Italy. *The B.E. Journal of Economic Analysis & Policy*, *7*(2) (Article 4).
Plug, E. (2004). Estimating the effect of Mother's schooling on children's schooling using a sample of adoptees. *American Economic Review*, *94*(1), 359–368.
Plug, E., & Vijverberg, W. (2005). Does family income matter for schooling outcomes? Using adoptees as a natural experiment. *Economic Journal*, *115*(506), 879–906.
Raaum, O., Bratsberg, B., Røed, K., Osterbacka, E., Eriksson, T., Jantti, M., et al. (2007). Marital sorting, household labor supply and intergenerational earnings mobility across countries. *The B.E. Journal of Economic Analysis and Policy*, *7*(2) (Article 7).
Restuccia, D., & Urrutia, C. (2004). Intergenerational persistence of earnings: The role of early and college education. *American Economic Review*, *94*(5), 1354–1378.
Royear, H. (2009). Separated at birth: US twin estimates of the effects of birth weight. *American Economic Journal: Applied Economics*, *1*(1), 49–85.
Smith, J. (2004). Unraveling the SES: Health connection. *Population and Development Review*, *30*, 108–132.
Solon, G. (1992). Intergenerational income mobility in the United States. *American Economic Review*, *82*(3), 393–408.
Solon, G. (1999). Intergenerational mobility in the labor market. *Handbook of Labor Economics*, *3*, 1761–1800 (Part A).
Solon, G. (2002). Cross-country differences in intergenerational earnings mobility. *Journal of Economic Perspectives*, *16*(3), 59–66.
Solon, G. (2004). A model of intergenerational mobility variation over time and place. In Miles Corak (Ed.), *Generational income mobility in North America and Europe*. Cambridge, UK: Cambridge University Press.
Stock, J.H., Wright, J.H., & Yogo, M. (2002). A survey of weak instruments and weak identification in generalized method of moments. *Journal of Business and Economic Statistics*, *20*(4), 518–529.
Strauss, J., & Thomas, D. (1998). Health, nutrition, and economic development. *Journal of Economic Literature*, *36*(2), 766–817.
Sun, Z. (2004). *Schooling cost, return and household decision in rural China.* Beijing: Beijing Normal University Press (in Chinese).
Sun, S., Huang, W., & Hong, J. (2012). Why free labor migration is so important? — Based on the perspective of intergenerational income mobility. *Economic Research Journal*, *58*(5), 147–159 (in Chinese).
Sutherland, D., & Yao, S. (2011). Income inequality in China over 30 years of reforms. *Cambridge Journal of Regions Economy and Society*, *4*(1), 91–105.
Taubman, P. (1989). Role of parental income in educational attainment. *American Economic Review*, *79*(2), 57–61.
Vogel, T. (2008). *"Reassessing intergenerational mobility in Germany and the united states: The impact of differences in lifecycle earnings patterns", SFB discussion paper 2006-055.* Berlin: Humbolt University.
Xie, E. (2011). Economic impact of Chinese residents' chronic diseases. *World Economic Papers*, *55*(3), 74–86 (in Chinese).
Yao, X., & Zhao, L. (2007). Intergenerational income mobility and the channels of intergenerational transmission in China: 1989–2000. *LEPP working paper* (in Chinese).
Zhang, Y., & Eriksson, T. (2010). Inequality of opportunity and income inequality in nine Chinese provinces, 1989–2006. *China Economic Review*, *21*(4), 607–616.
Zhu, G., & Vural, G. (2012). Intergenerational effect of parental time and its policy implications. *Working paper*.

延迟退休、就业和福利

王天宇 邱牧远
（北京大学经济学院，北京，100871）

摘要：随着人口老龄化进程的加剧，中国的养老金缺口问题日益突出。延迟退休年龄作为呼之欲出的政策工具，对就业市场的负面影响一直为社会各界担忧。本文在一个世代交叠模型中，引入劳动力市场的搜寻和匹配机制，分析延迟退休对就业和福利的影响。我们利用中国微观数据对模型进行了校准。数值模拟表明，推迟退休年龄有助于缓解养老金财政压力，但效果会随着退休年龄提高幅度的增大而逐渐减弱。延迟退休对就业和福利的影响呈现出较大的年龄异质性。退休年龄组因为养老金领取数额的增加而得到福利改进；临退休年龄组因为失业率大幅上升而福利受损；青年群体的失业率没有太大改变，但会因为劳动力市场就业形势的严峻而付出额外的搜寻努力，从而导致福利的降低。

关键词：延迟退休 就业 养老金缺口 劳动参与率 搜寻匹配

一、引言

近年来，随着人口预期寿命的增加和生育率的放缓，中国社会正逐步迈向老龄化，养老金缺口危机日益显现。虽然从统计数据来看，目前城镇职工养老保险年度收支在全国层面仍能保持一定的结余（见附录），但当前的基金收入中包含一部分国家财政补贴和参保扩面带来的增收[1]，而且十几个省份已经开始出现赤字（郑秉文，2013）。考虑到中国的人口结构和经济增速放缓的可能，未来的养老金财政压力不容小视。据世行一项估算，从 2001 年到 2075 年，中国养老金缺口可能达 9 万亿人民币，目前对于中国养老金缺口的估算，最乐观的也认为缺口将达到 3 万亿人民币。[2]最新测算（刘学良，2014）表明，从 2010-2050 年城镇职工养老保险的财政缺口所形成的隐性债务，折现到 2010 年高达 52.3 万亿。如何解决中国的养老金危机？在众多方案中，延迟退休年龄被认为可操作性强、效果迅速，得到众多专家学者的青睐[3]，并得到官方的认可[4]。然而，对该政策可能带来的负面效果，尤其是对就业造成的影响，也有很多学者表示了担忧。[5]本文试图在一般均衡框架下，引入劳动者和厂商的双向搜寻机制，探讨延迟退休政策对就业和福利的影响。

Samuelson（1958）和 Arron（1965）指出，在一个现收现付的社保体系中，养老金负担取决于人口增长、工资增长、工作年限和预期寿命。假设人口增长和工资增长恒定，则当预期寿命逐渐增加时，工作年限也应逐渐增加，从而使得养老金蓄水池达到“流入”和“流出”的动态平衡。预期寿命的提高是全球普遍趋势，因此，很多国家都已经调整或计划调整退休年龄，以应对养老金支出压力。[6]例如，日本已经在 2001 年至 2013 年间将男性退休年龄从 60 岁逐步调高至 65 岁，澳大利亚在 2014 年将女性退休年龄从 63 岁调高到 65 岁，并将在

---

[1] 以 2011 年为例，根据郑秉文（2012）的测算，扣除财政补贴 2272 亿元和扩大覆盖面导致的“补缴”1511 亿元，则结余仅为 347 亿元。

[2] 见凤凰网专题报道 http://finance.ifeng.com/news/special/tuixiu2012/

[3] 见新浪财经报道 http://finance.sina.com.cn/hy/20141211/124421051999.shtml?wbf=more

[4] 2013 年 18 届三中全会《中共中央关于全面深化改革若干重大问题的决定》指出，要研究制定渐进式延迟退休年龄政策。

[5] 见网易 http://money.163.com/10/0920/00/6H00SPSC00252G50.html。

[6] OECD 国家退休年龄和政策较为详尽的数据可以参见 http://www.oecd.org/employment/emp/ageingandemploymentpolicies.htm

2017-2023 年间将男女退休年龄都逐步调至 67 岁。[7]Galasso（2008）通过对政策偏好的模拟表明，随着人口老龄化进程的加剧，欧洲所有国家都将通过推迟退休年龄的政策决议。

我国现行法定退休年龄为男职工 60 岁，女职工 50 岁，女干部 55 岁。该设定来自计划经济年代的一系列法规[8]。由于人均预期寿命的增加、教育年限变长和男女劳动能力差异的缩小，该退休年龄已经很难与时代发展相适应，一方面整体退休过早，另一方面男女差异过大。同时，中国的养老金体系，尤其是城镇养老金体系，是在国企急速改革的背景下建立起来的，未能充分考虑到长寿风险和财政支付能力（赵耀辉，徐建国，2001）。

从国际比较来看，中国目前的人均预期寿命和美国、德国等西方发达经济体相比大约落后 4-5 岁，但退休年龄，尤其是女性退休年龄，却落后 7-10 岁（见附录）。即便是和印尼、韩国等亚洲国家相比，中国也呈现出男女退休年龄差距过大的问题（Giles，2011）。值得指出的是，中国存在着普遍的提前退休现象（封进、胡岩，2008；李实等，2008）。据测算，在 2000-2006 年间，男性在 60 周岁之前退休的达到 54%，女性在 50 周岁和 55 周岁之前退休的分别达到 30%和 73%（封进、胡岩，2008）。在提前退休的人群中，有相当一部分在退休之初就开始领取养老金，进一步加重了养老金负担（郭席四，2005）。

从上述实际出发，延迟退休年龄似乎是充分利用人力资本、促进男女就业权利平等、填补养老金缺口的必然选择。然而，退休年龄作为劳动力市场上的重要制度，具有牵一发而动全身的效果，延迟退休的影响也是多维度的。

首先，延迟退休可能会对各年龄层面的就业造成冲击。对于青年群体，我国每年新增劳动适龄人口不低于 1000 万，高校毕业生超过 700 万，如果新增劳动力因为延迟退休而无法得到就业机会，将会引发严重的社会问题。延迟退休意味着老龄劳动力供给量的增加，这是否会对青年就业造成挤出效应，已有的少量经验证据大多倾向于老年劳动力供给并未减少青年的就业，甚至促进了青年的就业（Gruber and Milligan，2008；Munnell and Wu，2013；张川川、赵耀辉，2014）。理论上，这一问题至少取决于三个因素：一是各年龄层劳动力的替代程度，二是企业对劳动力供给变化的反应，三是劳动力市场的运行机制。对于老年群体，目前中国已经存在大量的内退、预退和退休前几年无所事事等现象，说明临退休人员的劳动供给意愿不足或就业机构对老年劳动力的需求不足，延迟退休年龄是否会进一步加剧这种现象？已经有文献指出，养老金制度的改革对老年劳动者的劳动供给影响最大（French and Jones，2012）。如果延迟退休换来的是老年劳动参与率的下降，则政策效果会大打折扣。

其次，延迟退休固然会改变总人口中缴费/收益人群的比例，但由于劳动力供给增加，可能会导致均衡工资的下降，如果养老保险采用按照工资固定比例缴纳，则会对缴费总额产生影响。而当前发放的养老金社会统筹部分的基数则是根据劳动者退休前的当地平均工资制定的，这意味着延迟退休政策在实行初期不一定能对养老金财政恶化起到缓解作用。即便在长期中，工资的变化已经收敛到一个新的均衡，每期养老金的支付总量也可能因为养老金支付方案对老龄劳动力的激励效应而出现非线性的变化。

第三，延迟退休政策即便能够缓解养老金财政缺口，对一项公共政策成功与否的评价也必然要涉及到社会福利。延迟退休可能引起各年龄层人群福利的复杂变化。例如，在现行退休年龄下原本应该退休，但在延退下必须继续工作的群体可能会因为休息权被剥夺而福利受损，但退休后群体却可能因为养老金发放的增多（取决于养老金机制设定）而福利改善。也

[7] 对于 OECD 国家的退休年龄调整计划，可参见《卫报》的报道：
http://www.theguardian.com/news/datablog/2010/jun/24/retirement-ages-oecd-countries

[8] 主要是 1951 国务院颁布的《中华人民共和国劳动保险条例》；1955 国务院颁布的《关于国家机关工作人员退休处理暂行办法》；1978 全国五届人大第二次会议通过的《国务院关于安置老弱病残干部的暂行办法》和《国务院关于退休、退职的暂行办法》。

就是说，延迟退休可能引起各个年龄层人群之间的利益冲突，政策制定者不能不考虑这一点。法国近年来提高退休年龄引起的抗议风潮以及政策的最终失败值得引起我们的重视。

本文试图在一个带有劳动力市场搜寻机制的多期世代交叠模型中探讨上述复杂作用机制，对延迟退休的政策效果做出“事前评估”。我们用中国数据对模型进行了校准，用校准后的模型模拟延迟退休 1-5 岁时养老金缺口、就业和福利的变化。同时，我们也将延迟退休的政策效果和其替代选项——提高缴费比例和降低发放基数的政策效果做了对比。

本文余下部分的安排是：第二节对已有文献进行梳理，介绍我们的模型和已有研究的关系；第三节给出模型设定并定义稳态均衡；第四节利用中国数据校准模型；第五节对延迟退休的效果进行模拟；第六节对延迟退休和其他政策的效果做对比，深入讨论延迟退休对劳动力市场和福利的作用机制，第七节总结全文并对指出未来的研究方向。

二、文献综述

由于世界主要发达国家都在经历不同程度的人口老龄化，普遍面临着养老金兑付压力问题，养老金体系改革在过去二十年间一直是公共经济学领域研究热点（一个较为详尽的综述可以参考邹铁钉和叶航，2013）。我们重点关注那些考察以延迟退休作为缓解养老金财政危机手段的研究。值得指出的是，现有文献多数针对 OECD 国家，这些国家中有一些只有领取养老金的最早年限，但并没有法定退休年龄。我们对于提高养老金最早领取年限的研究给予同样的关注。

针对尚未开展的延迟退休或提高养老金领取年龄政策，进行效果评估，属于“事前评估”（ex ante evaluation）（Wolpin，2013），其基本研究框架是生命周期动态选择模型，退休制度的改革会影响个体生命周期中的消费、工作和储蓄决策，引起复杂的效应。利用生命周期动态选择模型研究退休、消费和劳动力供给的代表文献有 Heyma（2004）、French（2005）和 Van de Klaauw & Wolpin（2008）等等，但这些文献没有明确地针对提高退休年龄和养老金领取年限做出评估。Haan & Prowse（2014）针对德国的社保体系构建了生命周期模型分析延迟养老金领取年龄的政策效果。他们的估计结果表明，延迟养老金领取年龄 3.76 年，便可以在未来 40 年内抵消人均预期寿命增加带来的养老金财政危机，但这一改变会一起劳动力供给的剧烈变化。只有当老年人可以找到工作时，改革才会体现出整体福利增进的效果。

Haan & Prowse（2014）的研究，没有刻画企业行为。然而，如果我们无法了解企业的招聘行为，就无法获知延迟退休对劳动力需求的影响。另外，Haan & Prowse（2014）的框架属于局部均衡，但延迟退休可能在一般均衡层面上引起产品价格、工资、利率等宏观指标的变化，这些宏观因素又进一步作用于个体选择和企业行为。但普通的一般均衡框架下的世代交叠模型中对劳动力市场始终是出清的，无法回答当前的改革语境下最为关心的失业问题。把劳动力市场上的搜寻机制嵌入 OLG 框架的尝试始于 Galor & Lach（1990），他们在两期模型中研究了劳动力市场的信息显示和摩擦性失业之间的关系；Heel（2003）的多期 OLG 模型允许个体在生命周期的每一年做出搜寻和就业决策，探讨了短期和长期失业保险对有摩擦的劳动力市场的不同影响；Lugauer（2012）在与 Heel（2003）类似的设定下展示了人口结构对经济周期的作用。

在动态一般均衡、个体生命周期和搜寻-匹配这一框架下探讨延迟退休政策效果的努力始于 Keuschnigg & Keuschnigg（2004），他们构建了与本文类似的模型研究奥地利的养老金改革，发现提高退休年龄引起了劳动力市场搜寻强度的增加，从而减少了失业。Fisher & Keuschnigg（2010）注意到了延迟退休对劳动力市场影响的异质性，他们的模型表明延迟退休导致中年人福利损失最大。

另一类文献针对已经发生的延迟退休改革做出事后评估（ex post evaluation）。例如，奥地利政府在 2000-2010 年间把男女退休年龄分别从 55 岁和 60 岁提升为 58.25 岁和 62 岁。

Staubli & Zweimüller（2012）发现，改革后 55-58.25 岁的女性和 60-62 岁的男性就业概率仅仅上升了了 6.8%和 10.1%，改革仅仅起到了微弱的效果。French & Jones（2012）通过对养老金改革实证研究的总结，也发现养老金改革政策对老年人的劳动供给影响远大于对青年人的影响。

国内相关研究大多采用精算模型和人口统计方法对延迟退休对养老金财政支出和劳动供给影响做预测。例如，余立人（2012）认为，延迟退休年龄固然改变了每一年的养老金积累，但由于退休后剩余寿命变短，个人账户每个月的发放数额也会增加，再加上利率和平均工资增长率的不确定影响，延迟退休对于养老金支付能力的影响方向是不确定的。曾益等（2013）指出从现在开始延迟退休只能在短期内减小政府补贴额度，长期来看仍然无法摆脱养老金财政困境。精算模型对于养老金财政的短期预测效果很好，但是因为没有经济主体的行为模型作为支撑，无法研究劳动供给行为和福利变化，更无法探讨一般均衡下退休政策的复杂传导机制，长期预测效果的可信性也因此降低。

原新和万能（2006）通过对退休年龄和老年劳动参与率的国际比较，指出中国目前应设法提高老龄劳动参与率，而非推迟退休年龄。姚远等（2012）对逐渐延迟退休引起的抚养比和劳动力供给变化进行了预测，指出中国出生人口的几次高峰均在农村，延迟退休能够增加的 60-65 岁城镇劳动力数量可能极为有限。汪泽英（2013）的测算表明，提高法定退休年龄能有效缓解 2024 年后的劳动力短缺现象，但短期内会引起失业率的上升。

因为希望能够刻画延迟退休引起的个体行为和企业行为变化，同时考虑失业问题和养老金财政，我们的模型框架采用带有劳动力市场摩擦的一般均衡模型，与 Keuschnigg & Keuschnigg（2004）类似。但是，Keuschnigg & Keuschnigg（2004）的建模背景是现收现付制养老金，我们的模型则采用中国的统账结合养老金体系。

三、模型设定

本文在 Heer(2003)的基础上，在 75 期的世代交叠模型中讨论退休年龄的延迟对缓解养老金财政压力的作用以及对就业的影响。经济中的主体一共有三类，分别是个体、厂商和政府。接下来，我们将分别给出模型中人口动态、家户行为、企业行为和养老金体系的设定，并刻画劳动力市场上的搜寻-匹配过程，给出工资的决定机制。

（一）人口动态

假设经济的人口增长率 η 是外生的。[9]个体在 25 岁时进入劳动力市场，并参与经济。工作 TW 年后退休。退休后，其能够存活的最大年限为 TR，最长寿命为 $J = TW + TR$。退休年龄 TW 由政府选定，对个体来说是外生变量。为了讨论长寿风险，我们假设个体从年龄 j 进入到 $j+1$ 的概率为 sp^j。

由于我们并不讨论进入劳动力市场之前的个体行为，所以将 21 岁设为第一期。

令 ψ^j 代表年龄为 j 的个体在总人口中的比例，则稳态时 ψ^j 满足：

$$\psi^{j+1} = \frac{\psi^j \cdot sp^j}{1+\eta} \tag{1}$$

$$\sum_{j=1}^{J} \psi^j = 1 \tag{2}$$

在稳态中，任何一期的横截面上人口分布都是相同的。

（二）个体行为

[9] 如果各年龄层上的人口死亡率不随时间而变化，即 $\left\{sp_t^j\right\}_{j=1}^{J}$ 不随 t 而变化，则人口增长率等同于人口出生率。

家户在前TW期能够在劳动力市场中工作，并在$TW+1$期退休。家户在工作期内的状态有两种：失业或就业。当处于就业状态时，家户无弹性的供给一单位的劳动并得到相应的税后工资收入。在退休之后，家户领取相应的养老保险金。

令$\varepsilon=0$代表个体在当期失业，$\varepsilon=1$ 代表个体在当期就业，$\varepsilon=2$代表个体当期退休。在工作期限内，给定当期状态，工人在下一期是否就业是一个随机变量。具体而言，当工人本期状态为就业时，会以一个外生的概率κ离职；当工人本期属于失业状态时，以概率$1-e^{-\pi_t s^z}$ 找到下一期的工作，π_t和z是外生参数。找到工作的概率随着寻找工作强度s增大而增大。我们将在劳动力市场设定将进一步解释此公式的含义。工人在第一期寻找到工作的概率为p_0[10]， p_0对于工人也是外生的[11]。

我们用下标代表时间期数，上标代表年龄。个体的效用取决于三个因素：消费c，工作带来的损失D^w和找工作带来的损失$D^s\cdot s_t^{j+1}$。因此在t期年龄为j的个体在当期的效用为[12]：

$$u_t^j=\frac{(c_t^j)^{1-\sigma}-1}{1-\sigma}-D^w\cdot \mathrm{I}\{\varepsilon_t^j=1\}-D^s\cdot s_t^{j+1}\cdot \mathrm{I}\{\varepsilon_t^j=0\} \tag{3}$$

其中，消费的效用符合 CRRA 形式，劳动力供给的变化只发生在广延边际（extensive margin），如果当期失业，则找工作行为发生在期末，一个单位的搜寻努力带来成本D^s。$\mathrm{I}\{\cdot\}$为工作状态的示性函数。遵循 Lucas and Stocky（1989），我们可以将家户的最大化问题写成递归的形式，即：

$$\underset{\{c,s\}}{Max}\, W_t^j(a,\varepsilon)=u_t^j+\beta\cdot sp^{j+1}E_t[\mathrm{W}_{t+1}^{j+1}(a',\varepsilon')]$$

$$s.t.\, a_t^{j+1}+c_t^j=(1+r_t)a_t^j+y(\varepsilon_t^j) \tag{4}$$

$$a_t\geq 0 \tag{5}$$

其中，a为资产，r为利率。$\boldsymbol{c}$和$\boldsymbol{s}$分别代表个体的终生消费、搜寻向量，即$\boldsymbol{c}=[c_t^j,c_{t+1}^{j+1},\cdots]$，$\boldsymbol{s}=[s_t^j,s_{t+1}^{j+1},\cdots]$。个体在当期的收入流$y(\varepsilon_t)$是当期状态$\varepsilon$的函数，其定义如下：

$$y(\varepsilon_t^j)=\begin{cases} beq_t, & ,\varepsilon=0 \\ (1-\tau_t^1)w_t^j+beq_t & ,\varepsilon=1 \\ pen_t^j+beq_t & ,\varepsilon=2 \end{cases} \tag{6}$$

其中，τ_1为养老金个人缴费比例，pen_t^j为养老金收入。其中， beq_t为偶然遗产（accidental bequest）， 为养老金个人缴费比例。我们假设在每一期政府会把所有死亡人口的遗产平分给所有在世个体。这是一个技术性假定，对模型结果影响不大，类似的设定在文献中较为常见。求解上述消费者最大化问题，可以得到失业工人搜寻努力程度的一阶条件：

$$\mathrm{e}^{-\pi\cdot s^z}s^{z-1}=\frac{D^s}{\beta\cdot \mathrm{z}\cdot\pi\cdot\Delta W_{t+1}} \tag{7}$$

其中，$\Delta W_{t+1}=[W_{t+1}(a_{t+1},\varepsilon=1)-W_{t+1}(a_{t+1},\varepsilon=0)]$，代表下期工作和失业之间的值函数之差，可以被视为搜寻带来的收益。对（7）式两边求导可得$\partial s/\partial\Delta W_{t+1}>0$。即搜寻的努力程度是下期工作与失业值函数之差的增函数。这一点我们在分析延迟退休和其他政策时会反复用到。同时，折现因子越大、工作的负效用越小、找到工作的概率越大，则搜寻的努力程度越大。

（三）厂商行为

在劳动力市场上存在着同质性的厂商，其生产函数为柯布道格拉斯形式，即：

$$y=Ak_t^{\alpha}n_t^{1-\alpha} \tag{8}$$

[10] 该求职行为发生于进入劳动力市场之前的一期期末。

[11] 在退休后有$s=0$。

[12] 为了避免消费出现负值，我们在数值模拟时将（3）式中的c改为$\bar{c}+c$，$\bar{c}$为常数。

其中A为全要素生产率，k_t为期初的资本存量，n_t为期初雇佣劳动的数量。在期末，厂商雇佣的劳动力会以一个外生的概率κ离职。厂商下期创造h_t个新职位，每个新职位所需要支付的货币成本为p_c。每个职位可以以概率q_t 招收到新的工人，对于厂商来讲，q_t是外生的。因此厂商在每一期选择雇佣的劳动力数量满足：

$$n_{t+1}=q_t h_t+(1-\kappa)n_t \tag{9}$$

厂商的资本存量满足：

$$k_{t+1}=i_t+(1-\delta)k_t \tag{10}$$

其中，i_t为投资，δ为资本折旧率。

给定厂商期初的资本存量k_t和雇佣劳动的数量n_t，厂商的值函数为：

$$V(k_t,n_t)=\max_{i_t,h_t}\ Ak_t^{\alpha}n_t^{1-\alpha}-w_t(1+\tau_2)n_t-p_c h_t-i_t+\frac{1}{1+r_{t+1}}V(k_{t+1},n_{t+1}) \tag{11}$$

其中，τ_2为企业缴纳的养老保险比率，这一比率同样是由国家强制规定的。

厂商的雇佣行为由下列一阶条件决定：

$$\frac{r_t+\kappa}{q_t}+(1+\tau_2)w_t=A(1-\alpha)k_t^{\alpha}n_t^{-\alpha} \tag{12}$$

（四）劳动力市场和均衡工资率的决定

我们用一个简单的搜寻-匹配机制来刻画带有摩擦的劳动力市场。

假定由于搜寻和匹配机制的存在，在t期年龄为j资产为a，并于失业状态的工人付出的搜寻成本为$s_t^j(a,\varepsilon=0)$，而资产为k、雇佣劳动力为n_t的代表性厂商创造的新职位为h_t。

我们假设失业工人寻找工作的概率π是每一期厂商创造的新职位h_t和适龄劳动力中失业人口u_t的函数，其表达式为：

$$\pi_t=\mu h_t^{\varsigma}u_t^{1-\varsigma} \tag{13}$$

工资由厂商和当期新被动力通过纳什谈判的方式确定。具体来说，如果能够匹配成功，则产生由厂商和工人共享的“剩余”。工人的“剩余”来自其匹配成功所获得的工资。

在稳态时，我们假设一个“代表性工人”完成搜寻后，接受工作的价值为W^E，不接受工作的价值为W^U，则W^E和W^U可以表示为：

$$W^E=\left(1-\tau_1\right)wu'\left(\bar{c}^e\right)-D^w+\beta\overline{sp}\left[\left(1-\kappa\right)W^E+\kappa W^U\right] \tag{14}$$

$$W^U=-D^s\bar{s}+\beta\overline{sp}\left[\bar{p}W^E+\left(1-\bar{p}\right)W^U\right] \tag{15}$$

需要指出的是，（15）式中的搜寻成本$D^s\bar{s}$发生在下一期初，而工资和补贴等收益则发生在期末。

假设企业设立职位后，与“代表性工人”完成匹配的边际收益为V^f，未完成匹配造成职位在当期空缺的边际收益为V^v，则：

$$V^f=A(1-\alpha)k^{\alpha}n^{-\alpha}-w\left(1+\tau_2\right)+\frac{1}{1+r}\left[\left(1-\kappa\right)V^f+\kappa V^v\right] \tag{16}$$

$$V^v=0 \tag{17}$$

一个典型的纳什谈判问题可以表述如下：

$$\underset{w}{Max}\left(W^E-W^U\right)^{\lambda}\left(V^f-V^v\right)^{1-\lambda} \tag{18}$$

其中，W^E-W^U是工人接受工作协议的净收益，V^F-V^V是企业与该工人达成工作协议的净收益，λ是工人的谈判力量（Bargaining Power）。

经过求解简单的极大值问题，可以得到谈判解：

$$w=\psi_1 A(1-\alpha)k^{\alpha}n^{-\alpha}+\psi_2 D^w+\psi_3 D^s\bar{s} \tag{19}$$

（19）式中各参数的值分别为：

$$\psi_1=\frac{\lambda}{1+\tau_2} \tag{20}$$

$$\psi_2=\frac{(1-\lambda)}{(1-\tau_1)u'(\overline{c}^e)} \tag{21}$$

$$\psi_3=-\frac{\left[(1-\beta)-\beta\overline{sp}(1-\overline{p}-\kappa)\right](1-\lambda)\overline{s}}{\left[1-(1-\overline{p})\beta\overline{sp}\right](1-\tau_1)u'(\overline{c}^e)} \tag{22}$$

从（19）式可以看出，工人的边际生产率越高，对企业的价值越大，因此企业越害怕失去工人，因此谈判出来的工资越高；工作带来的负效用越大，接受该工作对工人境遇的改善越小，需要更高的货币补偿，因此谈判得到的工资越高；找工作带来的负效用越大，工人在谈判中会处于更不利的地位（不接受工作就要投入到下一期的搜寻），从而降低了谈判工资。

（五）养老金体系

按照中国目前的城镇职工养老保险制度，我们假设经济中的养老保险为个人账户+社会统筹模式。由于模型中每一期的长度为一年，我们将实际制度做简化后换算成年。

个人账户即劳动者每期缴纳工资的τ_1部分[13]，按照经济中各期利率自动滚存，在退休后每年领取个人账户在记账利率r_p下的累积总额除以计发期数。除了个人缴纳部分之外，企业需每期为雇员缴纳工资的τ_2比例。退休者每期领取的社会统筹部分为退休前工资的τ_3比例乘以工作年限[14]。个人账户部分为个人账户累积总额除以计发期数T_p。一个退休者每年领取的养老金为

$$pen_i=\tau_3 w\sum_{t=1}^{TW}1_{\varepsilon_t=1}+\frac{1}{T_p}\sum_{t=1}^{TW}1_{\varepsilon=1}\tau_1 w\left(1+r_p\right)^{TW-t+1} \tag{23}$$

给定社会养老保险的替代率θ，我们有下列养老金预算平衡公式：

$$\theta\sum_{j=TW+1}^{T}\psi^j w=\tau\sum_{j=1}^{TW}\psi^j n^j w+D \tag{24}$$

其中，D为养老金年度缴费和支出之差，即养老金财政。

（六）稳态均衡

给定政府的政策$\Omega=\{\theta,TR\}$，一个稳态均衡包括厂商的值函数$V(k_t,n_t)$，家户的值函数$\{W^j(\mathrm{a},\varepsilon)\}_{j=1}^T$以及消费、失业人口的搜寻努力程度和资产的政策方程$\{c^j(a,\varepsilon)\}_{j=1}^T\}_{j=1}^T$，$\{a'(a,\varepsilon)\}_{j=1}^T$，$\{s^j(a,\varepsilon)\}_{j=1}^{TW}$，资产和就业状态在各年龄之间的分布$\{\phi^j(a,\varepsilon)\}_{j=1}^T$以及价格$w$和$r$，使得：（1）个体的政策方程最大化家户终身效用；（2）企业的生产、雇佣行为最大化当期利润；（3）商品市场、资本市场出清；（4）工资由纳什谈判解（14）决定；（5）各年龄段的就业、失业人口比例符合人口动力学方程。对稳态均衡的详细定义参见附录。

四、模型校准

本文的模型没有显示解，因此在文中我们需要进行数值模拟。在模拟之前，我们必须对模型中的参数进行校准。我们将所有参数分为四组，第一组是个体效用函数、企业行为和劳动力市场的参数；第二组是退休年龄；第三组是养老金制度中的参数；第四组是人口动态参数。

在所有参数中，除了工作的效用损失D^w，搜寻职位的效用损失D^s和企业设置新职位的成本p_c之外，其他参数均根据文献或统计数据直接设定。在附录（五）中，我们给出了参数设定的详细依据。

---

[13] 城镇职工养老保险有个人账户缴费基数的最高上限，本文忽略这一点。

[14] 城镇职工养老保险规定当地上年度在岗职工月平均工资和本人指数化月平均缴费工资的平均值，本文将其调整为个人工资的 1%。

对于这三个需要校准的参数，我们选定城镇职工养老保险设立时的失业率、劳动收入占比和养老金缺口作为校准目标。中国官方发布的城镇登记失业率偏低，且缺少变化，我们综合 Xue & Zhong（2003）、Giles et al.（2005）和 Knight & Xue（2004）从多个微观调查中计算的失业率，将 1997 年的失业率设置为 10%。对于劳动收入占比，我们综合白重恩和钱震杰对于中国劳动收入占比变化的系列研究（白重恩、钱震杰，2009，2010），将其设定为 40%。对于养老金缺口，我们假设 1997 年养老保险刚刚建立时缺口为零。需要说明的是，在校准时我们将人口增速也取为 1997 年时的值（1%）。但在下一节的政策模拟分析中，我们将人口增速调整为当前值（0.5%），养老金缺口由此产生。

校准结果为模型失业率为 12%，劳动收入占比为 39.8%，养老金缺口为 1%[15]。

我们把所有参数总结成下表：

表 1 主要参数取值

参数名称	参数含义	数值
个体参数		
σ	风险规避系数	2
β	折现因子	0.9
D^w	工作的效用损失	1.29
D^s	搜寻的效用损失	0.12
企业参数		
A	生产函数	1
α	资本弹性	0.35
δ	资本折旧率	0.1
p_c	设置新职位的成本	0.698
κ	劳动力离职率	0.0835
劳动力市场参数		
μ	匹配函数	1
λ	议价能力	0.5
退休年龄参数		
TW	工作的最后一期	56
养老金参数		
τ_1	个人缴费占工资比例	8%
τ_2	企业缴费占工资比例	20%
τ_3	社会统筹养老金发放基数	1%
T_p	计发月数	见附录五
r_p	个人账户记账利率[16]	0%
人口参数		
η	人口增长率	1%，0.5%
sp^j	年龄 j 的生存概率	见附录五

五、延迟退休的政策效果分析

在本节中，我们利用上一节校准的模型，将人口增速设定为 0.5%（2013 年人口增速为 0.492%）。模拟退休年龄从 57 岁开始依次增加 1 岁，直至 62 岁的政策效应。需要交代的是，本节中所作的分析，如不加特殊说明，均指稳态均衡之间的比较。

[15] 我们定义的养老金缺口为赤字额度与需支付额度的比。

[16] 目前，个人账户的记账利率普遍偏低，大多在 2%至 3%之间。考虑到通货膨胀，实际利率几乎为零或为负值。为了数值算法上的简便，此处我们将其设置为 0。

在改变退休年龄TW时，对于养老金个人账户，我们按照个人账户计发期数表，找到相应的新的退休年龄对应的计发期数。其余的养老金制度规则不变。

（一）延迟退休的总体效应

首先，我们讨论延迟退休对劳动力市场、福利和宏观经济的总体影响，表 3 列出了延迟退休 1-5 年的数值模拟结果。

表 2 延迟退休的总体效果

	退休年龄	57（基准）	58	59	60
人口和宏观经济	工作与退休人口比	1.464	1.572	1.690	1.818
	总产出	1.000	1.022	1.043	1.065
	养老金缺口	4.63%	1.66%	-1.05%	-4.04%
	养老金替代率	50.73%	53.05%	55.47%	58.18%
劳动力市场	搜寻努力	1.000	0.983	0.969	0.955
	失业率	8.19%	8.49%	8.80%	9.13%
	工资	1.000	1.001	1.002	1.005
社会福利	消费	1.000	1.022	1.044	1.065
	当期效用	1.000	1.020	1.038	1.056
	值函数	1.000	1.003	1.032	1.062
	资产 Gini 系数	0.381	0.380	0.376	0.370

注：（1）表中的工作与退休人口比指的是工作年龄段的人口和退休年龄段人口的比，我们考虑了 65 岁后各年龄层的死亡概率，计算的是期望人口比；产出是指经济体一期的总产出；养老金缺口定义为当期缺口额度/当期应付总额；失业率是指经济中的总体失业率，定义为就业人口/工作年龄段的人口；工作搜寻是指全体工作年龄段人口搜寻努力程度的总和；工资是指未扣减养老保险费用之前的工资；消费是指经济体一期的总消费；效用和值函数分别指所有年龄层个体的平均效用（值函数）按照人口比重的加权求和；资产 Gini 系数是指全社会各年龄层人口（包括退休人口）的资产 Gini 系数。

（2）对于工作搜寻、消费、效用、值函数和产出等不容易解释数值经济含义的变量，我们将基期正规化为 1，汇报延迟退休 1-5 岁时各变量相对于基期的变化。

（3）养老金缺口为正意味着有财政赤字，缺口为负意味着有财政盈余。

（4）替代率指的是在整个工作年限内没有失业的人每年的养老金相对于工资w的比例。

当退休年龄被提高时，首先改变的是工作年龄段与退休年龄段的人口比例。当延迟退休 5 岁时，这一比例会从目前的 1.464：1 上升到 1.818：1。虽然工作年龄段的人口未必实现充分就业，但总的来讲，工作人口会增多，由此带来产出的增长和养老金缺口的下降。根据我们的计算，当延迟退休 2 岁时，养老金缺口会消失。

但是，延迟退休对养老金财政危机的作用是逐渐衰减的。提高退休年龄固然会使潜在的缴费群体增加，但仍然存在着两个相反方向的作用：一是由于企业用工调整成本的存在，延迟退休会引起失业率的上升，失业人口在我们的模型中并不缴费，但仍然领取社会统筹部分的养老金；二是我国的城镇职工养老保险个人账户计发月数设定对于晚退休有一定的奖励[17]，这意味着随着退休年龄的提升，每一期每个个体个人账户的领取额度加速提升。这种奖励机制在很多国家的养老金方案中都存在，但在延迟退休时，奖励政策对养老金财政的影响会被放大，制定延迟退休政策时需要对个人账户的发放方案做细致的考虑与调整。这一点上，我们与余立人（2012）的分析是一致的。

延迟退休后，养老金替代率上升了。一方面，工作年限增加意味着社会统筹部分的乘数增大；另一方面，工作年限的增加对应着个人账户计发年数的降低和每个月发放数额的增加。替代率是个体每一期做出工作搜寻、在整个生命周期中配置消费的重要影响因素。在后文的分析中也会反复用到。

[17] 例如，对照计发月数表，当退休年龄从 57 岁提高到 58 岁时，计发月数从 158 下降为 152，少了 6 个月，但当退休年龄从 61 提高到 62 时，计发月数从 132 下降到 125，少了 7 个月。

当退休年龄被提高时，适龄劳动人数增加，每一期在寻找工作的人会增多。对企业来说，由于调整成本的存在，不可能相对于劳动力的增加等比例设置职位。个体的劳动力供给和企业的劳动力需求在搜寻-匹配的市场机制中，综合作用结果是失业率的上升，退休年龄每提高一岁，失业率上升约 0.3 个百分点。

随着退休年龄的提高，平均搜寻努力程度呈下降趋势。在后文的分析中，我们可以看到，这是由青年人因为劳动力市场供给增加而增大搜寻力度和中老年因为工作期延长、搜寻边际收益变低和替代率增加而降低搜寻力度共同作用的结果。

在延迟退休时，工资略有上升。当延迟退休 2 岁时，工资相对于未延迟退休时上升了 0.2%。从传统的劳动力供给-需求角度来看，这有悖于常理。但是，在搜寻-匹配机制下，工资是由劳资双方通过纳什谈判决定的。在纳什谈判解（19）-（22）中，工资受到工人的边际产量、工作的负效用和找工作的负效用影响。D^w 和 D^s 的系数 ψ_2 和 ψ_3 都与工作人口的边际平均消费有关。延迟退休通过影响边际产量、消费和搜寻努力程度间接影响工作，作用机制十分复杂。

在对福利的度量中，我们选取了消费、当期效用、值函数和资产 Gini 系数几个指标。在我们的模型中，消费相对于总产出的比例基本上是稳定的。因为延迟退休会带来总产出的增加，所以消费也随之增加。

给定岗位数量和匹配函数，个体搜寻努力程度的增加或降低不会改变总体就业率。但是，降低搜寻工作的努力则会带来效用的增进。延迟退休导致的生命周期中工作时间的变长则会导致效用损失。退休前和退休后收入的增加导致消费增加则使效用增进。从表 4 可以看出，各个年龄层当期效用的加总呈现出随着退休年龄增加而增加的趋势，但增加的幅度在递减。这一趋势是容易理解的：搜寻努力的减弱十分微小，对福利可以忽略不计；延长工作时间对福利的负面影响是以 D^w 线性形式进入效用函数的；养老金增加则通过提升消费增进福利的，消费的边际效用递减。

如果从整个生命周期的角度来度量福利，则每个人在特定年龄 j 上的福利应该用值函数 $W_j\left(a,\varepsilon\right)$ 来度量。相应地，社会福利应当由每个年龄层上的平均福利以人口比例加权来度量。从数值模拟的结果来看，值函数表达的社会福利同样随着延迟退休年龄的增加而增加。

社会-经济地位的不平等对于人的福利有重要影响（Alesina et al. 2004; Theloudis，2011）在我们的模型中，工资并没有异质性，但因为历史工作状态的不同，会出现资产的异质性。为了衡量延迟退休对不平等的影响，我们构建了资产的基尼系数。随着退休年龄的推后，不平等程度减弱了。到养老金缺口补齐时，资产的基尼系数由基期的 0.381 降至 0.376。为了理解其中的作用机制，我们可以设想工作人口是没有失业的，不平等可以分解为工作组和退休组的组内不平等以及组间不平等。工作人口一直大于退休人口，工作时的收入也高于退休收入。退休年龄的延迟，提高了工作组和退休组之间人口的比重，同时提升了退休后的收入（由于养老金累积和计发月数增加），这两点都降低了组间不平等，而组内不平等没有改变，总体作用方向是降低了不平等。

（二）延迟退休的异质性效应

在本节中，我们考察延迟退休在不同年龄上的异质性效果。由于各项效果对于政策力度的增加基本上是单调的，我们只就能够填平缺口的延迟退休 2 岁进行分析。

首先，我们考察搜寻努力程度。图 2 展示了延迟退休 2 岁时各年龄层搜寻努力和失业率的变化。从左侧图形来看，面对延迟退休政策，青年群体的搜寻努力与基期相比有小幅上升，在 35 岁左右两者持平，35 岁后搜寻努力小于基期。无论是基期还是延迟退休 4 年时，接近退休的老年人都较为消极，搜寻努力收敛于零。对于刚刚进入劳动力市场的个体，因为几乎没有财产，如果失业便只能依靠政府发放的偶然遗产生活。因此，面对延迟退休带来的劳动力供给量增加，青年的反应是极力增加搜寻努力程度。对于接近退休的失业老年人，因为找工作难度的加大，需要付出比以前更多的搜寻努力才能再次上岗，考虑到即将领取养老金（尤其是当养老金替代率升高时），重新找工作的激励不足。

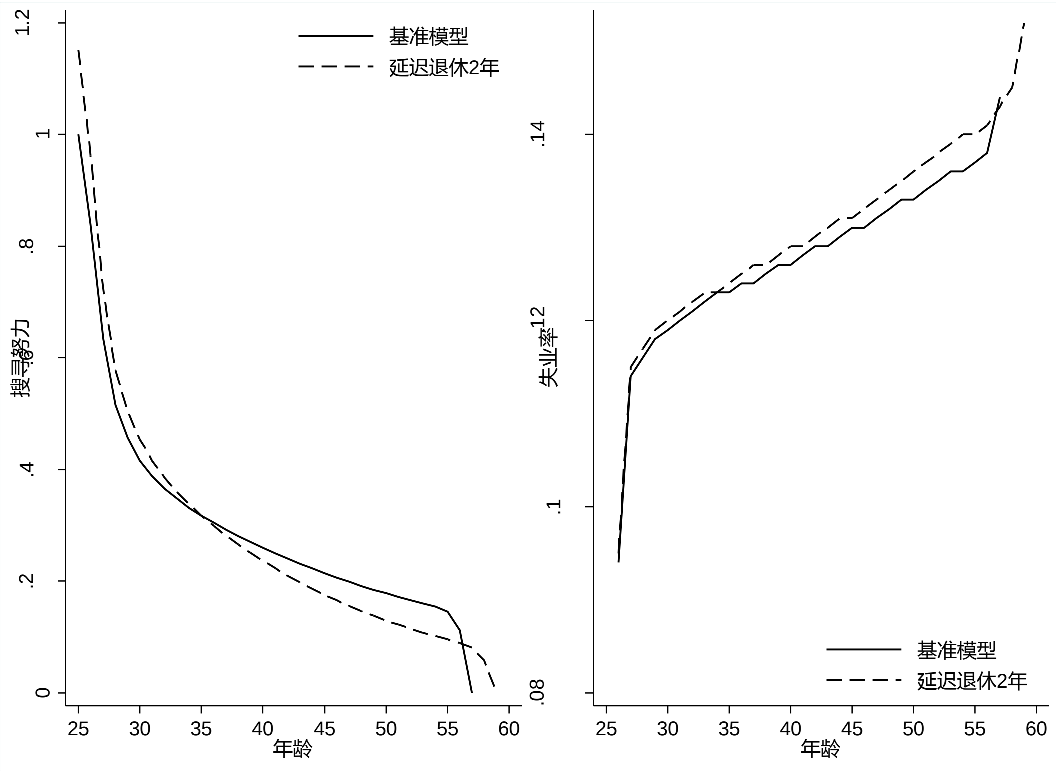

图 1 延迟退休对搜寻努力和失业的影响

由于搜寻努力程度加大，青年群体的失业率只有略微增加，这一结果与 Keuschnigg & Keuschnigg（2004）是一致的。但随着年龄的升高，失业率的上升愈发严重。当延迟退休至 59 岁时，59 岁年龄组的失业率到达 15.2%。这并不符合我们对中国劳动力市场的直觉，但是，我们的模型对劳动力供给的设定是在广延边际上（extensive margin），即劳动者只能选择是否供给劳动力。在现实中，劳动力的供给变化可能发生在集约边际（intensive margin）上，即我们通常看到的临近退休人员处于半离职状态。在 OECD 国家，养老金改革的一大问题就是如何避免老年人劳动参与率的下降。在现实中，多数老年人由于在工作技能上处于弱势，实际劳动参与率的下降可能更为严重。French（2005）根据 PSID 数据计算的 60 岁左右健康老年人的劳动参与率是 70%左右，非健康老年人劳动参与率为 25%左右。

我们已经说过，搜寻努力程度的提升不会增加总体就业率，但相对搜寻努力程度却决定了各年龄层之间失业的分布。下图展示了延迟退休两岁时不同年龄搜寻努力程度和失业率之间的关系。可以看到，随着年龄的增加，搜寻努力逐渐从下降，失业率则逐渐上升。

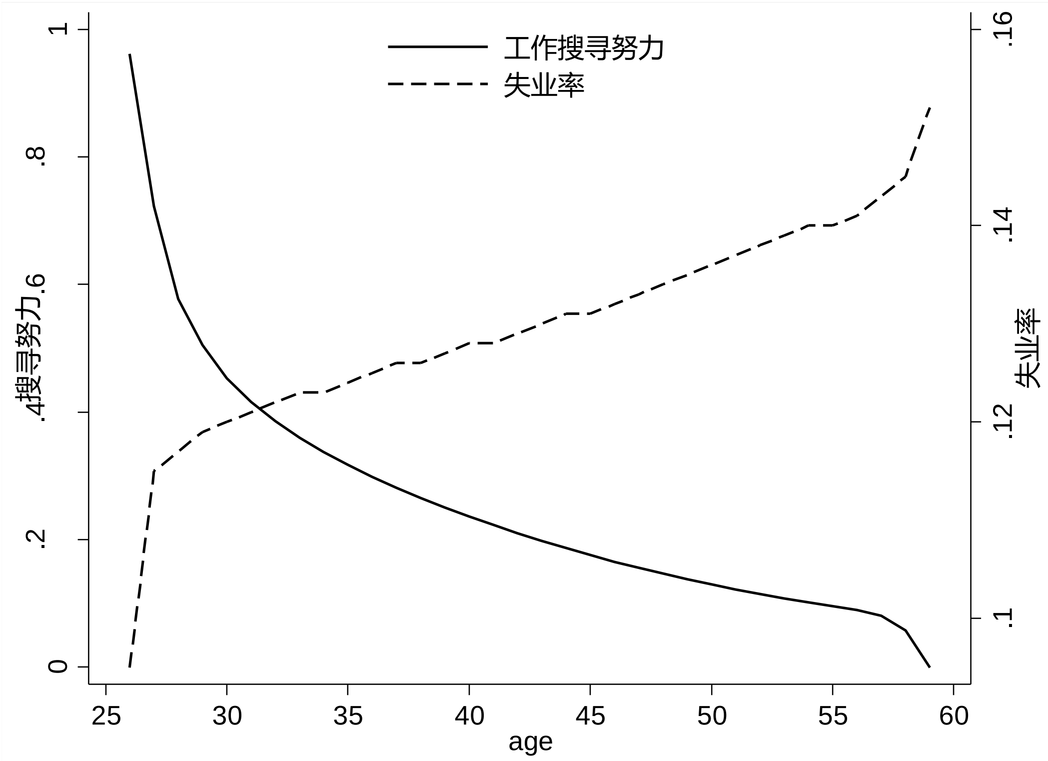

图 2 搜寻努力程度与失业率的关系

延迟退休对于不同年龄层福利的影响较为复杂。

首先，我们来看消费的改变。从图 3 来看，在整个生命周期中，延迟退休都导致了消费的增加。这一变化，可以从两个方面来理解。一是体现了延迟退休"把蛋糕做大"的效果。从表 2 来看，延迟退休 2 年使经济中的总消费提高了约 4%。劳动人口的增加，带来总产出的增加，从而带来消费的增加。这一效果在青年期不明显，因为这一时期资产很少，面临较强的借贷约束。在中年到老年阶段，收入增加带来的累积效果开始显现。二是体现了延迟退休的"分蛋糕"效果。如表 2 所示，延迟退休提高了替代率，使工作期的储蓄率下降，从而带来消费的增加。消费的增加在青年期并不多，因为这一时期资产很少。在中年到老年阶段，收入增加带来的累积效果开始显现。

在我们的模型中，个体的最大存活年龄是 100 岁。但是在 80 岁之后，存活的概率已经很低，个体不可能预留太多的资源用于临终消费。所以，延迟退休和基准模型的消费差异在预期寿命下达到最大，此后逐步缩小。

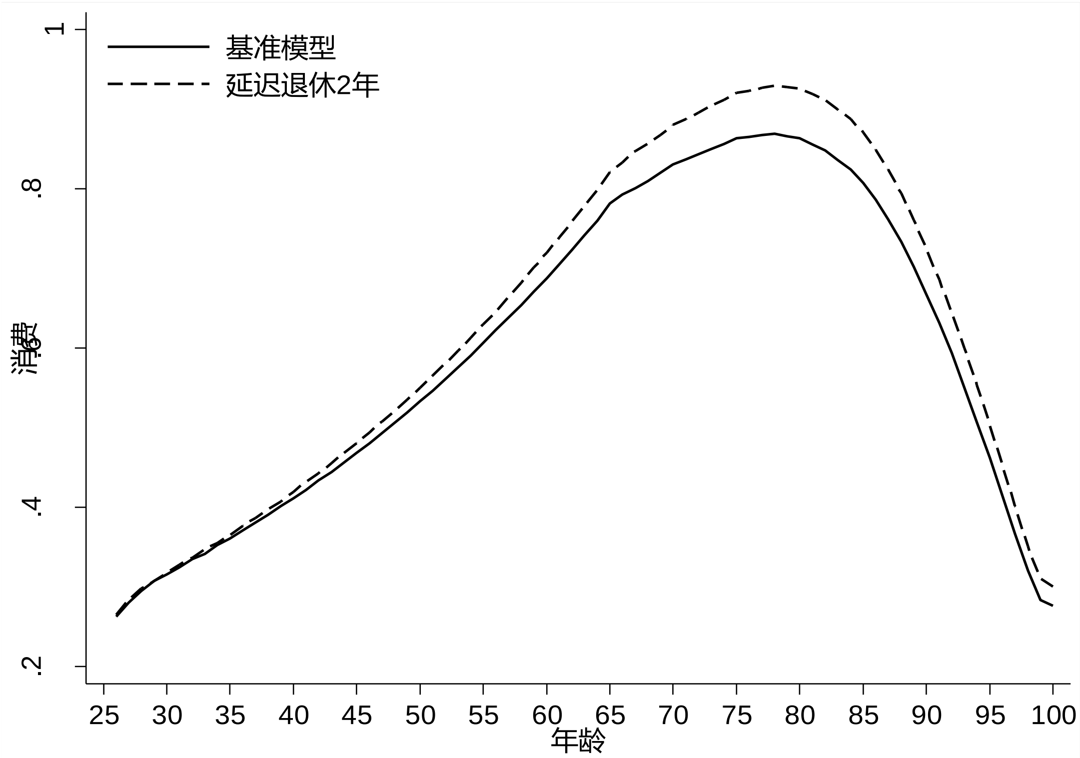

图 3 延迟退休对消费的影响

工作搜寻的负效用、工作的负效用和消费共同构成了效用函数。我们从图 4 可以看到，就当期效用而言，无论是在基准模型还是延迟退休模型中，效用函数的走势都在退休年龄上出现了拐点，即从退休开始上升。这是因为，在决定效用函数的三要素中，最优消费在临近的几期内是平滑的，工作搜寻力度在退休前也已经逐步变为 0，但工作带来的负效用会在退休后立刻变为 0，因此引起了效用的跳跃。但是，比较生命周期上各期之间的效用并没有太大意义，我们关注的是同一年龄上的代表性个体，在基本模型和延迟退休模型中的效用差异。图中可以看出，从最初进入劳动力市场一直到基本模型中的退休年龄（57），延迟退休的效用始终不低于基本模型。在青年期，两者的效用基本持平，这是因为一方面延迟退休后消费增加，另一方面青年人因为就业市场的竞争变得激烈而付出了额外的搜寻努力。

在基准模型中应该退休，但在延迟退休模型中需要再工作两年的 57-59 群体的效用是受损的，这是工作带来的负效用的直接后果。在对延迟退休的讨论中，这一群体的休息权是否能得到保障也是政策争论的热点。

对于年龄超过延迟退休后的退休年龄（59 岁）的群体，延迟退休后效用始终大于基准模型。这可以理解为消费上升的后果。

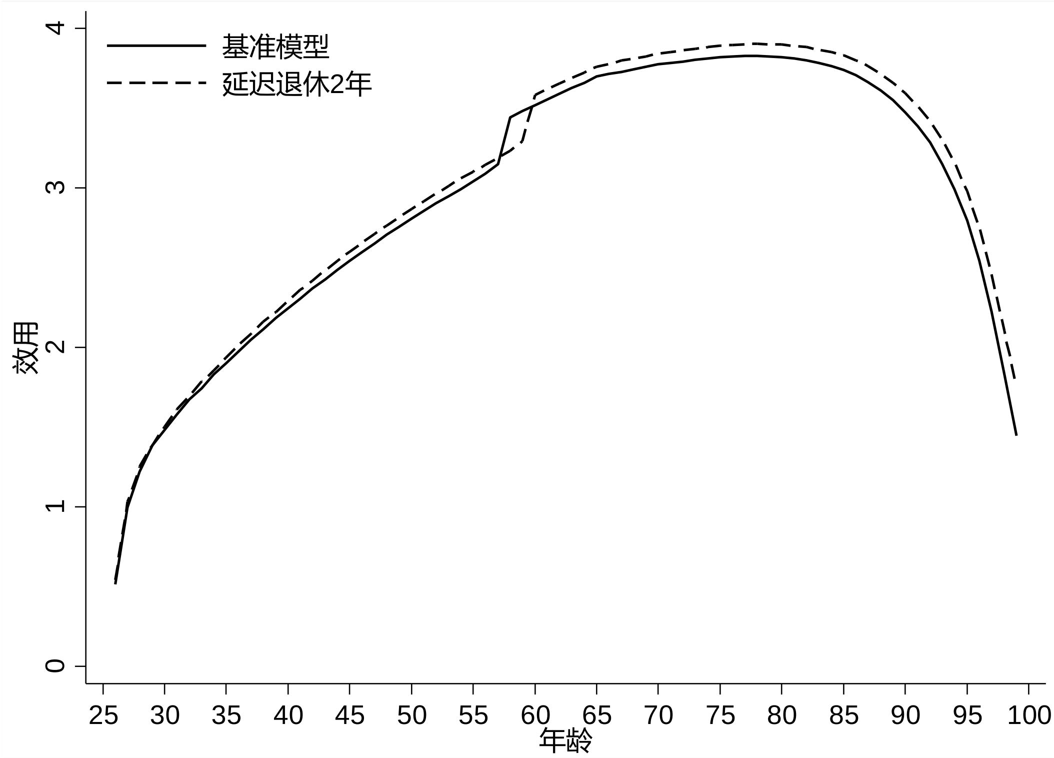

图 4 延迟退休对即期效用的影响

在个体生命周期的任意时点上，对其福利的最准确度量都应该是剩余生命周期中效用函数的折现值。这一价值可以构成不同年龄层的人支持或反对养老金改革政策的基础。下图展示了延迟退休对不同年龄层个体值函数的影响。从该图中可以清晰地看到延迟退休对各年龄层的福利影响都是正的，但却展现出了明显的年龄异质性。

在 59 岁之后，延迟退休下的值函数明显高于基准模型。延迟退休政策对于这一部分人来讲，可以说是完全的正效应。他们会享受养老金替代率升高带来的消费上升。大约从 52 岁开始，直到退休年龄，值函数的提升并不明显，这说明该政策对邻近退休的中老年人福利增进最小，这与 Fisher & Keuschnigg（2010）的分析一致。对照图 2 的失业率分布，这种效应不难理解：在提高退休年龄时，临近退休者一旦失业，会发现重新找工作的难度变大，找工作的热情衰减，搜寻力度降低，无法与年轻人竞争，只能退出劳动力市场，等待达到退休年龄后领取退休金。因为最后几期并未参与工作，退休后领取的社会统筹部分一定会减少，刚刚进入劳动力市场的青年群体和中年群体值函数变化较小，随着退休年龄的升高略有上升。

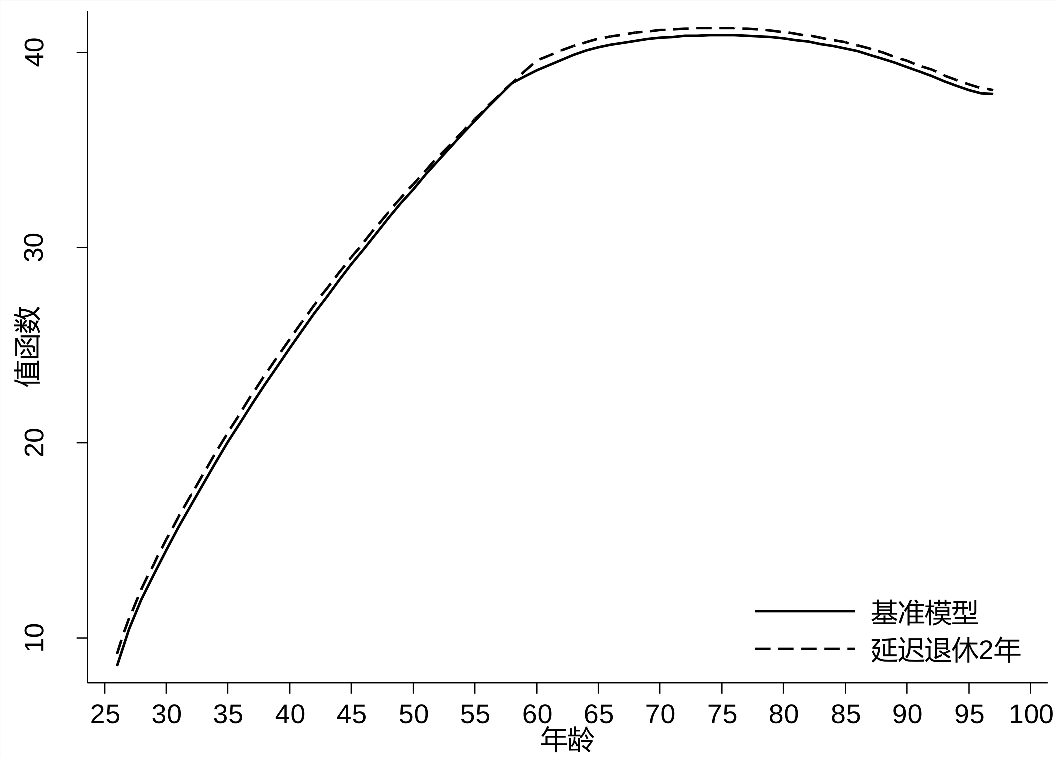

图 5 延迟退休对值函数的影响

上述对消费、当期效用和值函数的分析表明，延迟退休政策总体上来讲会引起福利的改进，对退休后老年人的福利改进最大，对原政策下应该退休但新政策下要继续工作的人福利改进最小。

（三）延迟退休和其他政策的比较

我们认为，如果不改变统账结合的养老保险模式，则解决财政危机的方法无非是开源节流。 短期内，“开源”可以由扩大养老金征缴人群（即扩面）来实现，但在长期中，开源的唯一途径是提高养老金税率。“节流”即降低养老金开支，除杜绝提前领取外，节流的方法只能是降低养老金的支付系数。而延迟退休则改变了缴费人群和收益人群的比例，属于既开源又节流。本小节中，我们将填平养老金缺口的三种改革方案——延迟退休 2 年、提高企业养老金税率至 21%和降低社会统筹养老金发放乘数至 0.95%进行比较分析。下表展示了三种方案和基准模型之间整体效果的比较。

表 3 三种政策的总体效果对比

	方案	基准模型	延迟退休两年	企业税率提升至 21%	发放乘数降至 0.95%
人口和宏观经济	工作与退休人口比	1.464	1.572	1.464	1.464
	总产出	1.000	1.022	0.998	1.003
	养老金缺口	4.63%	1.66%	0.00%	-0.82%
	养老金替代率	50.73%	53.05%	50.73%	49.13%
劳动力市场	搜寻努力	1.000	0.983	1.001	0.995
	失业率	12.79%	12.89%	12.80%	12.77%
	工资	1.000	1.001	0.998	1.002
社会福利	消费	1.000	1.022	0.993	0.992
	当期效用	1.000	1.020	0.984	0.995
	值函数	1.000	1.003	0.957	0.960

资产 Gini 系数	0.381	0.380	0.380	0.383

注：（1）表中的工作与退休人口比指的是工作年龄段的人口和退休年龄段人口的比，我们考虑了 65 岁后各年龄层的死亡概率，计算的是期望人口比；产出是指经济体一期的总产出；养老金缺口定义为当期缺口额度/当期应付总额；失业率是指经济中的总体失业率，定义为就业人口/工作年龄段的人口；工作搜寻是指全体工作年龄段人口搜寻努力程度的总和；工资是指未扣减养老保险费用之前的工资；消费是指经济体一期的总消费；效用和值函数分别指所有年龄层个体的平均效用（值函数）按照人口比重的加权求和；资产 Gini 系数是指全社会各年龄层人口（包括退休人口）的资产 Gini 系数。

（2）对于工作搜寻、消费、效用、值函数和产出等不容易理解数值经济含义的变量，我们将基期正规化为 1，汇报延迟退休 1-3 岁时各变量相对于基期的变化。

（3）养老金缺口为正意味着有财政赤字，缺口为负意味着有财政盈余。

（4）替代率指的是在整个工作年限内没有失业的人每年的养老金相对于工资 w 的比例。

从上表可以看出，三种政策中，只有延迟退休会改变工作与退休的人口比。也正因为如此，延迟退休对产出的刺激作用也最大，降低发放乘数也能微弱地刺激产出，提高企业税率会降低产出。三种政策中，只有延迟退休能够在填补养老金缺口的同时提高养老金替代率，提高企业税率不会影响替代率，降低发放乘数则会降低替代率。

三种政策当中，延迟退休对劳动力市场的影响最大，具体而言，会提高整体失业率 0.1 个百分点，同时减弱找工作的平均努力程度。提升企业税率也会微弱地拉升失业率，降低发放乘数则会微弱地降低失业率。

三种政策当中，提高养老金税率对企业的影响最大。企业面对增加的社保税赋，会削减工作岗位，导致失业上升，工资下降。青年期和中年期劳动者因此而提高搜寻努力程度，进一步降低了福利。

从三种福利指标来看，延迟退休起到了提升整体福利的作用，尤其是消费。提高企业税率对平均消费的负面作用小于降低发放乘数，但对当期效用的负面影响则大于发放乘数，原因是提高税率提升了总体的工作搜寻努力程度，对效用产生了负影响。

从不同年龄层来看，延迟退休的影响作用发生于生命周期的各个时点上，增加企业税率主要是对工作期起作用，降低发放基数则主要是对退休期起作用。延迟退休对各个年龄层的福利影响都基本为正，增加企业税率主要是降低了青年期的消费和效用，降低发放基数则降低了老年期的消费和效用（见图 6）。

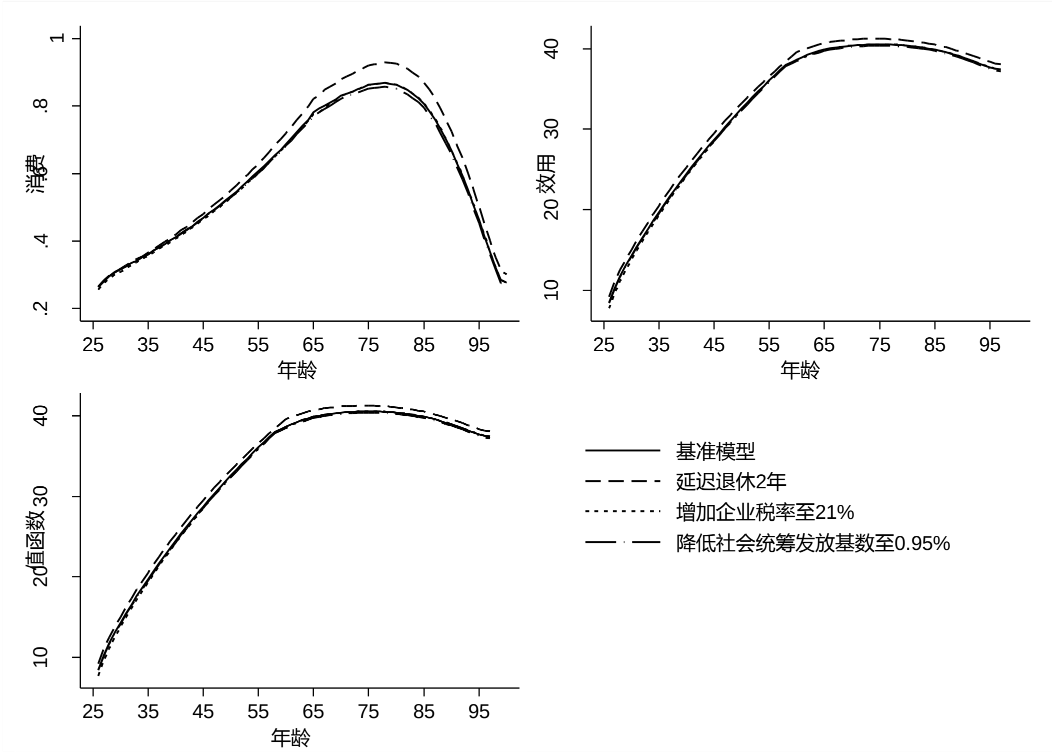

图 6 三种政策对福利影响的对比分析

如果我们可以认为，进入劳动力市场第一期的终生效用折现值代表了政策对整个生命周期的总体影响，则我们将 25 岁基准模型的值函数设为 0，延迟退休 2 年、增加企业税率和降低发放的值函数分别为 0.904、-0.748 和 0.022。我们的模型中，第一期进入劳动力市场人口的就业率完全是由新进入劳动力市场的人数与退休人数之比决定的，与其它各期不同。如果我们考察 26 岁时的值函数，则四者分别为 8.443、9.162、7.726 与 8.396。可以认为，降低 0.5%的发放标准对福利的影响不大，提高企业税率 1%对福利有较大的负面影响，而延迟退休则会对福利有较大的增进作用。

六、结论及政策建议

本文在一个包含劳动力市场搜寻-匹配机制的 75 期生命周期模型中，展示了法定退休年龄的提升对劳动力市场和福利的复杂影响。

当退休年龄提高时，一方面养老金的缴费年限增加，领取年限减少，对放缓财政压力的作用为正；另一方面个人账户计发期数规则对老年劳动力的奖励作用被放大，对放缓财政压力的作用为负。两者的综合作用导致延迟退休的财政效果随着退休年龄的提高幅度而逐步减弱。同时，退休年龄的提升使整个经济体中的劳动力要素投入增加，提高了产出，对社会的总体福利有改进作用。当退休年龄提高时，整个人口中适龄劳动力年龄段比例会增大，导致劳动力市场上供求状况和工人与企业谈判力量的变化，均衡工资下降，失业率上升，但后者的影响在不同年龄层中存在异质性。青年群体由于找工作努力程度的增加，失业率变化不大；临退休的中老年群体失业后会因为剩余工作年限较少和找工作难度的加大而表现得较为消极，搜寻努力程度增加不大，从而导致失业率上升。国际文献的实证研究（ French & Jones，2012）也发现养老金改革对不用年龄层的影响是异质的， 本文的结果和文献基本一致。

我们的模型中，个体每一期的福利取决于消费、找工作付出的成本和工作的成本。延迟退休对福利的影响即是对这三方面的综合作用。从即期效用来看，退休后个体因为养老金的增加而得到福利改进，青年群体因为付出更大的搜寻努力程度而福利下降，临退休中老年群体因为失业上升消费下降福利恶化。但是，生命周期每个时点上的个体，应该考虑到未来期

福利的折现。就剩余生命周期的整体福利而言，退休年龄升高时，青年群体的状况会变好，临退休中老年群体的状况仍然会恶化。

同时，我们也比较了延迟退休和另两种能够补齐养老金缺口的政策——提高企业社保税率和降低个人养老金发放基数的福利效应。我们对三种改革方案的福利分析阐释了养老金改革政策选择的一个浅显却经常被遗忘的道理：老龄化带来的问题，本质上是劳动参与率的下降，任何期望仅仅调整企业和个人税费关系或养老金运营管理的改革都是某种程度的“不劳而获”，长期来看，最终会导致福利的下降。

该研究的政策含义是十分丰富的。虽然延迟退休会减轻财政压力，但这种财政效果会随着退休年龄提升幅度的增加而逐渐减弱，并很可能存在着拐点。在现行制度下，提高 1-2 岁是较为稳妥的选择。

政策的制定还必须考虑到对不同年龄段的异质性效果。目前，由于历史原因，我国老年人口的劳动技能与青年群体存在着较大差距，55-65 岁年龄段的工作者很可能已经无法适应岗位需要。在我们的同质模型中，延迟退休造成的劳动力市场竞争加剧使老年人的劳动参与意愿大幅下降，考虑到现实中老年群体在技能上的落后，延迟退休的福利效果很可能大打折扣。如果能考虑让老年群体在某个区间内自主选择弹性退休，并制定相应的非全额养老金领取规则，劳动力市场的资源配置效率将会得到提升。

延迟退休会导致大部分年龄层工作搜寻力度的提高，尤其是资产较少的青年群体。而搜寻力度的普遍增加并不能带来整体就业率的上升，是一种福利净损失。如果我们能考虑在延迟退休出台之前，先建立健全失业保险制度，使劳动者不必对暂时性失业过于担心，社会的整体福利会得到改进。

参考文献

[1] 郑秉文 主编. 《中国养老金发展报告 2012》经济管理出版社. 2012.

[2] 郑秉文 主编. 《中国养老金发展报告 2013:社保经办服务体系改革》经济管理出版社. 2013.

[3] 封进, 胡岩. 中国城镇劳动力提前退休行为的研究[J]. 中国人口科学, 2008, (4):88-94.

[4] 李实, 史泰丽, 古斯塔夫森. 中国城镇失业、提前退休和性别收入差距的变化. 《中国居民收入分配研究(III)》第 13 章. 北京师范大学出版社. 2008.

[5] 郭席四. 我国转型期提前退休行为的危害及其控制[J]. 统计与决策, 2005:48-49.

[6] 张川川，赵耀辉. (2014). 老年人就业和年轻人就业的关系: 来自中国的经验证据. 世界经济 (5), 74-90.

[7] 邹铁钉, 叶航. 养老保险体制改革[J]. 统计研究, 2014, (2).

[8] 原新, 万能. 缓解老龄化压力,推迟退休有效吗?[J]. 人口研究, 2006, (4).

[9] 姚远, 原新, 史佳颖等. 退休年龄调整:为何如此纠结?[J]. 人口研究, 2012, (6).

[10] 汪泽英. 提高法定退休年龄政策研究.[M]. 中国经济出版社. 2013.

[11] 封进，韩旭，陈文澜，章诗颖，陶潇潇. 变革中的养老金制度：记账式账户制. 《比较》2014 年第 3 期。

[12] 白重恩，钱震杰. 国民收入的要素分配：统计数据背后的故事[J]. 经济研究，2009, (3):27-41.

[13] 白重恩, 钱震杰. 劳动收入份额决定因素:来自中国省际面板数据的证据[J]. 世界经济, 2010, (12).

[14] 吴江，田小宝. 人力资源蓝皮书：中国人力资源发展报告（2011-2012）. 社会科学文献出版社. 2012.

[15] Aaron, H. (1966). The social insurance paradox. *Canadian journal of economics and political science*, 371-374.

[16] Samuelson, P. A. (1958). An exact consumption-loan model of interest with or without the social contrivance of money. *The journal of political economy*, 467-482.

[17] Giles,J., Wang, D. & Cai, W. (2011). The Labor Supply and Retirement Behavior of China’s Older Workers and Elderly in Comparative Perspective. *IZA discussion paper No. 6088.*

[18] Gruber, J., & Milligan, K. (2010). Do Elderly Workers Substitute for Younger Workers in the United States?. In *Social Security Programs and Retirement around the World: The Relationship to Youth Employment* (pp. 345-360). University of Chicago Press.
[19] Munnell, A., & Wu, A. (2013). *Do Older Workers Squeeze Out Younger Workers?* (No. 13-011). Stanford Institute for Economic Policy Research.
[20] Wolpin, K. I. (2013). *The limits of inference without theory*. MIT Press.
[21] French, E., & Jones, J. (2012). Public pensions and labor supply over the life cycle. *International Tax and Public Finance*, *19*(2), 268-287.
[22] Heyma, A. (2004). A structural dynamic analysis of retirement behaviour in the Netherlands. *Journal of Applied Econometrics*, *19*(6), 739-759.
[23] French, E. (2005). The effects of health, wealth, and wages on labour supply and retirement behaviour. *The Review of Economic Studies*, *72*(2), 395-427.
[24] Van der Klaauw, W., & Wolpin, K. I. (2008). Social security and the retirement and savings behavior of low-income households. *Journal of Econometrics*,*145*(1), 21-42.
[25] Keuschnigg, C., & Keuschnigg, M. (2004). Aging, labor markets, and pension reform in Austria. *FinanzArchiv: Public Finance Analysis*, *60*(3), 359-392.
[26] Fisher, W. H., & Keuschnigg, C. (2010). Pension reform and labor market incentives. *Journal of Population Economics*, *23*(2), 769-803.
[27] Haan, P., & Prowse, V. (2014). Longevity, life-cycle behavior and pension reform. *Journal of Econometrics*, *178*, 582-601.
[28] Galor, O., & Lach, S. (1990). Search unemployment in an overlapping-generations setting. *International Economic Review*, 409-419.
[29] Lugauer, S. (2012). Demographic change and the great moderation in an overlapping generations model with matching frictions. *Macroeconomic Dynamics*, *16*(05), 706-731.
[30] José J. Canals and Steven Stern, "Empirical Search Models", in Search Theory and Unemployment, eds., Steven Woodbury and Carl Davidson, Kluwer Academic Publications, 2002.
[31] Stokey, Nancy L., and Robert E. Lucas, Jr., with Edward C. Prescott, Recursive Methods in Economic Dynamics, Harvard University Press, 1989.
[32] French, E. (2005). The effects of health, wealth and wages on labor supply and retirement behavior. Review of Economic Studies, 72, 395–427.
[33] Eckstein, Zvi, and Gerard J. Van den Berg. "Empirical labor search: A survey." *Journal of Econometrics* 136.2 (2007): 531-564.
[34] Xue, J., & Zhong, W. (2003). Unemployment, poverty and income disparity in urban China. *Asian Economic Journal*, 17, 383– 405.
[35] Knight, J., & Xue, J. (2006). How high is urban unemployment in China? *Journal of Chinese Economic and Business Studies*,4(2), 91-107.
[36] Giles, J., Park, A., & Zhang, J. (2005). What is China's true unemployment rate?. *China Economic Review*, *16*(2), 149-170.
[37] Borowczyk-Martins, D., Jolivet, G., & Postel-Vinay, F. (2011). *New estimates of the matching function*. Technical report, mimeo.
[38] Heer, B. (2003). Employment and welfare effects of a two-tier unemployment compensation system. *International Tax and Public Finance*, *10*(2), 147-168.
[39] Petrongolo, B., & Pissarides, C. A. (2001). Looking into the black box: A survey of the matching function. *Journal of Economic literature*, 390-431.
[40] Alesina, A., Di Tella, R., & MacCulloch, R. (2004). Inequality and happiness: are Europeans and Americans different?. Journal of Public Economics, 88(9), 2009-2042.
[41] Theloudis, A. (2011). From income and consumption inequality to economic welfare inequality: the role of labor supply.
[42] Staubli, S., & Zweimüller, J. (2012). Does raising the retirement age increase employment of older workers? *University of Zurich Department of Economics Working Paper*, (20).

附录

一、近年来城镇职工养老保险的参保和收支情况

表 4 近年来城镇职工养老保险参保和收支情况

指标	2013	2012	2011	2010	2009	2008	2007
参保人数(万人)	32218.4	30426.8	28391.3	25707.3	23549.9	21891.1	20136.9
在职参保人数(万人)	24177.3	22981.1	21565	19402.3	17743	16587.5	15183.2
退休参保人数(万人)	8041	7445.7	6826.2	6305	5806.9	5303.6	4953.7
基金收入(亿元)	22680.4	20001	16894.7	13419.5	11490.8	9740.2	7834.2
基金支出(亿元)	18470.4	15561.8	12764.9	10554.9	8894.4	7389.6	5964.9
当期结余(亿元)	4210	4439.2	4129.8	2864.6	2596.4	2350.6	1869.3
累计结余(亿元)	28269.2	23941.3	19496.6	15365.3	12526.1	9931	7391.4

数据来源：国家统计局

二、各国退休年龄比较

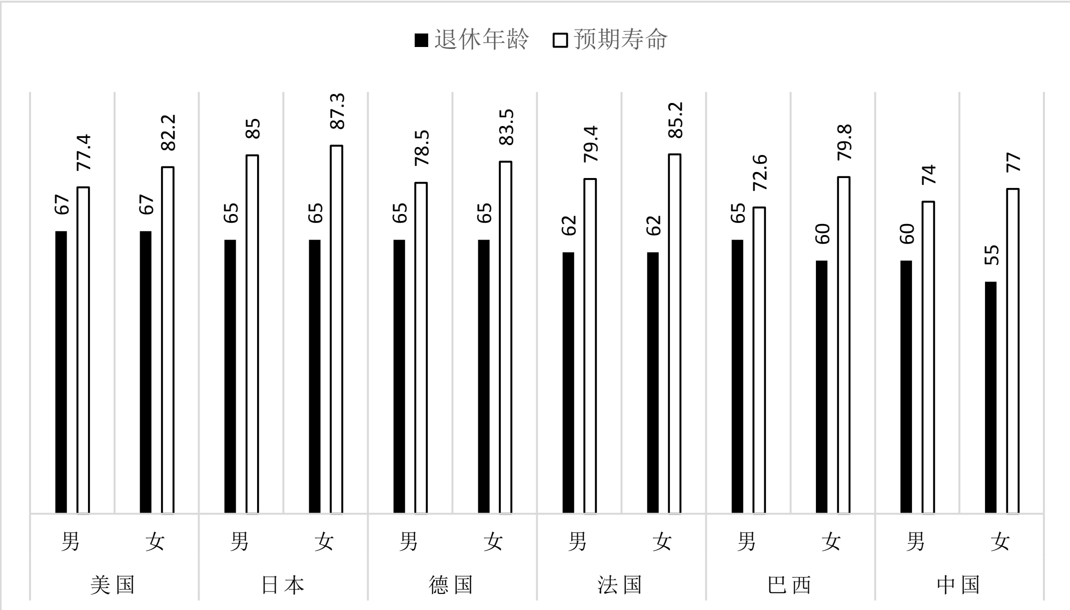

图 7 OECD 各国法定退休年龄和预期寿命

注释：[1] 数据来自世界卫生组织、OECD 统计数据和各国政府网站，图中所示退休年龄和预期寿命均为 2012 年数据;

[2] 一些国家的退休年龄是劳动者可以灵活选择的（例如美国），图中所列退休年龄是指能够领取全额养老金的年龄。

三、劳动力市场上的就业状态定义

我们将总人口定义为 1。令 $\phi_t^j(a,\varepsilon)$ 代表 t 期年龄为 j 的群体内不同资产和就业状态的分布，则相应各个年龄段就业人数 n_t^j，就业总人数 n^j 和失业人口总数 u^j 定义如下：

$$n_t^j=\psi_t^j\sum_a\phi_t^j(a,\varepsilon=1) \tag{25}$$

$$n_t=\sum_{j=1}^{TW}[\psi_t^j\sum_a\phi_t^j(a,\varepsilon=1)] \tag{26}$$

$$u_t \equiv \sum_{j=1}^{TW} \psi_t^j - n_t \tag{27}$$

其中，失业率定义为失业人口同小于退休年龄的总人口之间的比例。但是由于我们无法刻画工人在未进入经济之前的效用函数，所以假定第一期工人的工作来源是接替临退休工人的工作，而非劳动力市场上的搜寻。这样一来，第一期工人的失业率很大程度上取决于经济体中处于第一期和工作末期的人口比例。

退休人口占总人口的比例 n_R 为：

$$n_t^R = \sum_{j=TW+1}^{T} \psi_t^j \tag{28}$$

四、稳态均衡的定义

1、给定价格和政府政策，个体的政策方程 $\{c^j(a,\varepsilon)\}_{j=1}^T, \{\mathrm{a}'(\mathrm{a},\varepsilon)\}_{j=1}^T, \{\mathrm{s}(\mathrm{a},\varepsilon)\}_{j=1}^T$ 最大化家户的终身效用。

2、商品市场出清，即：

$$Ak^{\alpha} n^{1-\alpha} = i + h + \sum_j \psi^j \sum_{a,\varepsilon} c^j(a,\varepsilon)\phi^j(a,\varepsilon) \tag{29}$$

3、家户资产的回报率等于厂商的总剩余，即：

$$r\sum_{a,\varepsilon,j} a\cdot\psi^j\phi^j(a,\varepsilon) = Ak^{\alpha}n^{1-\alpha} - wn - h - i \tag{30}$$

4、新进入劳动力市场个体（j=1）的就业概率同本期就业人口中退休的比例为：

$$p^0 = \frac{\psi^{TW}(1-\kappa)\sum_a \phi^{TW}(a,\varepsilon=1)}{\psi^1} \tag{31}$$

5、工资率 w 满足厂商和工人纳什定价的一阶条件（14）。

6、匹配的就业岗位总量等于失业人口总量，即：

$$qh = (1-\kappa)n \tag{32}$$

7、投资等于资本折旧，即：

$$\delta k = i \tag{33}$$

8、各年龄就业人数、就业和失业总人数由式（10）-（12）决定。

五、模型参数设定

（一）个体、企业和劳动力市场参数

（1）个体效用函数中的参数设定

效用函数中共有四个参数，分别是风险规避系数 σ、折现因子 β、工作的效用损失 D^w 和搜寻职位的效用损失 D^s。对于前两个参数，我们根据已有文献直接设定。

（2）企业参数设定

在本模型中，TFP 增长率设置为 0。这一设定使生产函数中的参数 A 的取值不影响最后的结果，我们将其设置为 1。资本折旧率设置为 0.1。资本-产出弹性设置为 0.4。

工人离职率在不同行业、不同性质的企业中相差很大。目前，对离职率中国尚无官方统计或严谨的学术估算。我们综合多份离职调查报告将离职率 κ 设置为 0.08。

（3）劳动力市场参数设定

劳动力市场的参数即搜寻-匹配模型中的参数。对搜寻-匹配模型的估计，需要长时段的个体和企业数据（Canals & Stern, 2002; Eckstein & Van den Berg, 2007），目前中国尚无法进行。我们通过对已有文献的总结，直接给出这些参数，最后会尝试检验模型结果在不同参数下的稳定性。

工人的谈判能力 λ 设置为 0.5。在议价过程中，工人处于越强势的地位，λ 越大。在 $\lambda = 0.5$ 时，我们实际上假定了工人和企业拥有对称的谈判能力。这一设定来自 Heer（2003）。

匹配函数（matching function）中 μ 的作用相当于生产函数中的 A，当 μ 不随时间变化时，其数值并不影响最后的结果，我们将其设置为 1。ζ 的作用相当于生产函数中的 α，此参数使用不同数据的估计结果相差很大，基本在 0.35-0.75 之间（Petrongolo & Pissarides，2001；Borowczyk-Martins，2011）我们将其设置为 0.5。

（二）退休年龄的设定

其中，退休年龄的设定尤为关键。目前，中国的法定退休年龄为男 60 岁，女工人 50 岁，女干部 55 岁。由于本文的研究与城镇职工养老保险紧密相连，城镇职工养老保险的参保对象一般可归为“女干部”，养老保险的法定领取年龄也是 55 岁，因此可以忽略女工人。此外，由于我们的模型中没有区分性别和职业属性，必须设定统一的退休年龄。综合多种针对中国实际退休年龄的报告，我们将文中的法定退休年龄设置为 57 岁。[18]

（三）养老金参数的设定

按照现行的城镇职工养老保险制度，我们将个人缴纳的工资比例 τ_1 设定为 8%，企业缴纳的工资比例 τ_2 设定为 20%，退休后每月领取的社会统筹养老金基数 τ_3 设定为 1%。

目前，城职保个人账户的计发期数按月计算，给定的退休年龄范围为 40 岁至 70 岁。退休年龄和计发月数并非线性关系，而是对晚退休有一定的奖励。我们将该表转换为以年为单位，算出 T_p，具体结果如下：

表 5 个人账户计发期数

退休年龄	计发月数	T_p	退休年龄	计发月数	T_p
40	233	19.42	56	164	13.67
41	230	19.17	57	158	13.17
42	226	18.83	58	152	12.67
43	223	18.58	59	145	12.08
44	220	18.33	60	139	11.58
45	216	18.00	61	132	11.00
46	212	17.67	62	125	10.42
47	208	17.33	63	117	9.75
48	204	17.00	64	109	9.08
49	199	16.58	65	101	8.42
50	195	16.25	66	93	7.75
51	190	15.83	67	84	7.00
52	185	15.42	68	75	6.25
53	180	15.00	69	65	5.42
54	175	14.58	70	56	4.67
55	170	14.17			

（四）人口参数的设定

人口参数包括人口增长率和各年龄段人口的生存概率。

根据国家统计局数据[19]，1997 年和 2013 年人口自然增长率分别为 1.006%和 0.49%。我们取整设定为 1%和 0.5%。

由于在 65 岁之前死亡是一个极小概率事件，而在可能的退休期内（考虑退休延迟）死亡将对模型中的企业行为带来复杂影响，我们将 65 岁之前的死亡概率设为 0。

[18] 见《人力资源蓝皮书：中国人力资源发展报告（2011-2012）》、《中国养老金发展报告 2012》。

[19] 见国家数据网站 http://data.stats.gov.cn/workspace/index;jsessionid=2C10AB60C244B62BC7469BCEDFC88904?m=hgnd。人口指标由 1982，1990，2000 和 2010 人口普查以及各年度抽样调查推算得到。

我们寿命上限设置为 100 岁。这意味着，我们必须把一张正常的生命表压缩在 100 年之内，同时从 65 岁开始有死亡发生。我们的做法是，保持生命表中从 65 岁到 99 岁各年龄的相对死亡概率不变，将 100 岁到 101 岁的生存概率设置为 0，同时将 65 岁之前的总体死亡率均摊至 65-100 岁，使得新生命表计算的预期寿命等于 2010 年第六次人口普查的预期寿命 74.83 岁。

另外，目前的生命表大多采用分性别计算，而我们的模型中个体没有性别之分。我们按照 2000 年第五次人口普查的男女出生比例 117：100 为男女生存概率设置权重，计算不分性别的生存概率。具体生存概率如下：

表 6 各年龄条件生存概率

年龄	生存概率	年龄	生存概率	年龄	生存概率
65	0.93072	77	0.906096	89	0.823705
66	0.929653	78	0.902347	90	0.811636
67	0.928459	79	0.898193	91	0.798451
68	0.927128	80	0.893595	92	0.784075
69	0.925645	81	0.888508	93	0.768442
70	0.923993	82	0.882885	94	0.751485
71	0.922152	83	0.876673	95	0.733149
72	0.920103	84	0.869818	96	0.713388
73	0.917821	85	0.862261	97	0.692171
74	0.91532	86	0.853938	98	0.669483
75	0.912532	87	0.844784	99	0.645338
76	0.90948	88	0.83473	100	0

六、数值模拟步骤

（1）给定基准社保政策Ω；

（2）设定工资、利率和失业工人成功就业的概率的初始值；

（3）计算家户的最大化问题，得到政策方程：由于家户没有遗产动机，设定 J+1 期所有状态的值函数为 0，利用前向迭代得到各期家户在不同状态下的值函数，进而得到政策方程；

（4）给定初始值 p_0，并根据个体初始没有资产的条件得到状态在各期的分布状况，根据式（19）迭代计算出 p_0，更新 p_0 直到收敛；

（5）根据式（10）计算均衡的资本量k；

（6）根据式（33）计算均衡的投资数量；

（7）根据式（9）计算厂商空职位在下期匹配的概率q；

（8）根据式（12）计算稳态时厂商每期新创造的职位数量h；

（9）根据式（28）计算新的就业概率$\left\{sp_t^j\right\}_{j=1}^J$；

（10）根据式（12）计算新的均衡工资；

（11）根据企业利润对资本的一阶条件计算新的均衡利率；

（12）更新工资、利率和匹配概率直到收敛。