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**283/26**

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Kuan-Ming Chen, Maxwell Kellogg, Louis Kuan-Ju Tseng

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**JEL Codes:** J14, D10

**Keywords:** Long-Term Care, Family Welfare, Fertility, Health Shocks, Caregiving

**Recommended Citation:** Kuan-Ming Chen, Maxwell Kellogg, Louis Kuan-Ju Tseng (2026): Family Trajectories and the Burden of Care in the Aftermath of Old-Age Health Shocks. RFBerlin Discussion Paper No. 283/26

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# Family Trajectories and the Burden of Care in the Aftermath of Old-Age Health Shocks<sup>\*</sup>

Kuan-Ming Chen<sup>†</sup>      Maxwell Kellogg<sup>‡</sup>      Louis Kuan-Ju Tseng<sup>§</sup>

This Version: October 18, 2026

## Abstract

Elderly care needs are increasing in many rich and aging countries, yet the costs that they impose on families remain poorly understood. The economics literature has focused on how caregiving causes families to substitute from work toward informal care provision. But many families cannot afford to work less when care needs arise, so the costs of care may surface elsewhere. Consistent with this, we show that sudden declines in the health of elderly Taiwanese adults disrupt the life cycles of their adult sons: sons die at higher rates, exhibit worse health, and marry and have children at lower rates. Effects are largest in families least able to absorb care needs, which points to the burden of care as a key mechanism. Using reforms that expanded eligibility to hire migrant caregivers, we find evidence that live-in caregiver hiring can attenuate mortality and employment effects. Taking our quasi-experimental evidence to a flexible life cycle framework suggests that the welfare gains *to adult sons* from caregiver hiring can be large, driven mainly by improved survival.

**Keywords:** Long-Term Care, Family Welfare, Fertility, Health Shocks, Caregiving.

**JEL Codes:** J14, D10

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<sup>\*</sup>We gratefully acknowledge Taiwan's Ministry of Health and Welfare and Ministry of Labor for data access. We thank Su-Yuan Chan, Shiau-Fang Chao, Manasi Deshpande, Ming-Jen Lin, Edwin Leuven, Francis Wong, and Jonas Hjort for valuable comments and guidance, Liang-Cheng Chen, Po-Kang Hsiao, and Wei-Lun Lo for outstanding research assistance. Chen acknowledges the support from the National Science and Technology Council grant NSTC 114-2628-H-002-019 and the Yushan Fellow Program by the Ministry of Education (MOE) MOE-112-YSFSL-0003-001-P1. Chen thanks the support from the E. Sun Academic Award. All views expressed and any remaining errors are our own; they do not reflect the views of the Ministry of Health and Welfare or the Ministry of Labor.

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# 1 Introduction

Across the developed world, populations face rapid aging and rising rates of severe health problems, putting strain on a shrinking base of potential caregivers. The measurable costs of aging-related health problems are already large in many countries: long-term care (LTC) spending averaged 1.8 percent of GDP over OECD countries in 2021 (OECD, 2023). Yet formal services and safety nets rarely cover the physically demanding, day-to-day labor that LTC requires. Spouses and children instead tend to provide this care informally. Because this care flows through families rather than markets, existing evidence on its consequences has focused on the margin easiest to observe: how health shocks affect relatives’ earnings and labor supply. But when social support is limited and families have few working-age adults across whom to balance both income and care needs, substitution away from employment is a luxury that many simply cannot afford. Instead, families may be forced to borrow against their futures by neglecting investments in the health, marriage, and fertility of the next generation.

This paper characterizes the burden of aging-related health problems on younger relatives, the mechanisms by which they arise, and how public policy can mediate them. Our identification strategy is a staggered difference-in-differences design, in which otherwise similar “later-shocked” families serve as control units for families that suffer shocks earlier in time. We take the design to rich registry and health care utilization data on the population of Taiwan, where we concentrate on effects for adult sons of the elderly due to cultural and data quality considerations. Our analysis first focuses on strokes as they represent severe and acute health events for which timing is plausibly exogenous (Chandra and Staiger, 2007; Doyle, 2011; Fadlon and Nielsen, 2019, 2021). We document modest declines in the aggregate employment of adult sons after an elderly parent’s stroke (0.75 percentage points), comparable to or smaller than responses documented elsewhere (Frimmel et al., 2025; Rellstab et al., 2020; Ramirez Lizardi et al., 2025; Bergeot et al., 2024).

But parental strokes also derail adult sons from their pre-shock life cycle trajectories, with particularly stark consequences for their mortality. Adult sons’ all-cause mortality increases by 0.56 deaths per thousand per year, or 19 percent relative to the mortality experienced by their otherwise-similar peers whose parents later suffer a stroke. About 10 percent of the mortality effect comes from higher suicide rates, which increase by 35 percent relative to rates among those peers. Beyond increasing mortality, parental strokes cause adult sons to have 2.6 percent fewer children over the next five years, driven by those younger than 34 just before their parental stroke. These younger sons have 0.008 fewer children than their

peers over that five-year interval—a non-negligible effect in light of Taiwan’s total fertility rate of 0.98 in 2021 (which has since fallen below 0.9). While our design is not well-suited for estimating long-term effects, all estimated effects persist over the first 8 post-stroke years that we examine individually.

We next present evidence that the burden of informal care provision is a key driver of the adverse effects of a parental stroke on adult sons. Mortality effects are heavily concentrated among sons in families less able to absorb the time demands of care needs: those where sons have no identifiable siblings, and where sons have lower levels of schooling.

Turning from strokes to parental health shocks more broadly, we find similar effects on employment, mortality, and depression after hip fractures and even dementia diagnoses. To gauge the broader burden, we rescale our stroke estimates by the stroke’s effect on parental functional ability, according to our analysis of the Taiwan Longitudinal Study on Aging (TLSA) data. This yields an estimate of the marginal effect of functional limitations, which we can project onto the average level of functional ability among old-age parents more broadly. The result is a back-of-the-envelope estimate of the total burden that existing parental functional limitations place on adult sons, linearly extrapolating from the effects of a stroke. The exercise suggests that this total burden is substantial. For instance, the average functional limitations of one parent, drawn randomly from the distribution of older-age Taiwanese, can explain 0.28 deaths per thousand adult sons. By comparison, overall mortality rates experienced by men in their 40s, which range from 2 to 5 deaths per thousand per year.

We next examine whether those adverse effects of a stroke are attenuated by the hiring of live-in formal caregivers. Formal caregiving in Taiwan is performed almost exclusively by migrant workers, who first came to Taiwan in meaningful numbers after the introduction of centrally-regulated migrant care labor markets in the 1990s. Rates of migrant caregiver hiring have since increased to the point that migrant care workers represent roughly 1 percent of Taiwan’s total population as of 2020.

The rise in caregiver hiring has been concentrated among LTC-needers age 80 or older, for whom the Taiwanese government has gradually relaxed eligibility criteria. Among families that suffer a stroke in this older age group, caregiver hiring rates rose from 29 percent for families suffering a stroke in 2007 to nearly 37 percent for those suffering one in 2018. Over the same period, the adverse effects of parental strokes on adult sons’ mortality and employment effectively vanished in these families. To identify a causal effect of caregiver hiring, we exploit changes in the diff-in-diff effect estimates for a parental stroke just before

versus after major Taiwanese reforms began in 2012. This triple-difference design suggests that the broader time-series correlation, if anything, *understates* how much caregiver hiring mitigates a parental stroke’s adverse effects on adult sons. We interpret this as evidence that targeted migrant labor policy can improve the health and well-being of the native working-age population. That a parental stroke’s adverse effects vanish precisely as caregiver hiring expands is also strong evidence that those effects were driven by the burden of informal care to begin with. Such patterns are not consistent with alternative explanations like shared family health risks, grief at a parent’s condition, or reverse causality.

To the extent that we can examine consequences for marriage and fertility, caregiver hiring seems to attenuate the adverse effects of a stroke on these outcomes as well. However, because existing reforms have targeted only LTC-needers aged 80 and older, we cannot verify whether this is true among the younger adult sons at more critical ages for family formation (who tend to have younger parents).

Having established that reform-induced caregiver hiring buffers adult sons against a parental stroke’s adverse effects, we proceed by quantifying sons’ private welfare gains. A natural upper bound on *ex-post* willingness to pay (WTP) for migrant caregiver hiring is the cost savings of hiring a migrant caregiver over an otherwise-similar domestic one. But *ex-post* WTP may be substantially lower than *ex-ante* WTP given that LTC insurance markets are incomplete and families are poorly self-insured. To judge *ex-ante* WTP, we instead start from the causal effects of caregiver hiring on survival and “quality of life” (i.e., depression), and ask how much adult sons ought to benefit from them those health-related gains net of hiring costs. Assessing these gains is difficult though, because it requires knowing how individuals value resources (including their own lives) across very different states of the world.

We sidestep this issue by introducing an alternative parameter of interest: adult sons’ WTP for migrant caregiver hiring *and* insurance against the non-hiring income effects of the parental shock. We can decompose this into two parts: *ex-ante* WTP for insurance against non-hiring income effects *fixing caregiving arrangements*, and *ex-ante* WTP for caregiver hiring given the income insurance. We take the latter term (which we call the partial WTP, or “PWTP”, for hiring) as a lower bound for the full *ex-ante* WTP for this hypothetical reform, and show how to estimate it with a framework adapted from Murphy and Topel (2006) for valuing health and longevity. Under that framework and given the income insurance, the PWTP for hiring directly weighs the net monetary costs of hiring against survival and quality-of-life gains, both of which can be read off the value of a statistical life (VSL).

Taking the framework to the data and an existing VSL estimate suggests that the PWTP for hiring is very large and dominated by the benefits from survival gains. The survival-related benefits over the average three-year hiring contract yield gains worth roughly \$372,000 in 2019 USD, which is an order of magnitude larger than the cost of hiring a caregiver (\$32,000 not including food and accommodation). But even before getting to the survival gains, 33 percent of the costs of hiring are offset by adult sons' labor force entry. Improvements in quality of life, if proxied by depression diagnoses, offset another 18 percent of the caregiver hiring costs.

We argue that a large private PWTP for hiring points to substantial *social* welfare gains as well, as we find no clear negative externalities from the reform. Direct competition with the domestic labor force—the externality which is often of greatest interest to economists in the context of immigration policy—is limited because Taiwan has never had a meaningful domestic labor market for formal caregivers whose work intensity, particularly around-the-clock availability, was comparable to that of migrant caregivers. Consistent with substantial welfare gains, the initial 2012 reform was quickly followed by another expansionary reform in 2015, and more recently a third in 2024.

**Related literature.** Our paper contributes to an active empirical literature on the consequences of health shocks for families, which has focused mainly on labor supply. We extend that literature in two directions: to the non-employment outcomes of relatives—their health, family formation, and mortality—and to shocks suffered by elderly parents, for which the burden of LTC is plausibly a key mechanism. This paper also speaks to how formal caregiving policies can mitigate the costs LTC needs impose on families. We show—and price—the welfare improvements that labor supply responses alone would miss.

Many studies have leveraged registry data and difference-in-differences designs to examine how health shocks affect one's own employment (Halla and Zweimüller, 2013; Heinesen and Kolodziejczyk, 2013; Garcia-Gomez et al., 2013; Gupta et al., 2015; Dobkin et al., 2018; Parro and Pohl, 2021) and, to a lesser degree, spousal employment (Fadlon and Nielsen, 2021; Garcia-Gomez et al., 2013; Jeon and Pohl, 2017).<sup>1</sup> Less is known regarding shocks suffered by old-age adults. The existing evidence suggests parental health shocks are connected with weak, if any, declines in the labor supply of adult children (Frimmel et al., 2025; Rellstab et al., 2020; Ramirez Lizardi et al., 2025), which may mask offsetting responses across children

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<sup>1</sup>See Bauer and Sousa-Poza (2015) for a review of earlier survey-based evidence. Other notable papers have focused on broader financial distress (e.g., Dobkin et al., 2018), family health behaviors (Fadlon and Nielsen, 2019), and inheritance following a relative's fatal shock (Druehl and Martinello, 2022).

(Arrieta and Li, 2023). However, the onset of *informal caregiving* spells is associated with comparatively sharper declines (Maestas et al., 2024).

Far less attention has been paid to non-employment spillovers. What evidence exists mostly concerns shocks within couples, where a spouse’s health shock or death worsens the survivor’s mental health and mortality (Arteaga et al., 2025; Fadlon et al., 2026); for parental shocks, it is limited to modest mental health responses (Glaser and Pruckner, 2023; Ramirez Lizardi et al., 2025) and shifts in living arrangements (Bergeot et al., 2024). A strand of the medical literature has found informal caregivers to be of unusually poor health (e.g., Cameron et al., 2016; Haley et al., 2015; Schulz and Beach, 1999;<sup>2</sup> see Adelman et al., 2014 for a review), but with small samples it has been unable to distinguish the burden of care from selection into caregiving or intra-family correlations in health (Rigby et al., 2009).

The literature on public policy interventions in care provision has rapidly developed in recent years. Relative to existing work on expansions in access to international care labor markets (Frimmel et al., 2025; Grabowski et al., 2026), we study effects on the health, well-being, and longevity of working-age relatives and price the welfare gains from them. Other existing work finds that expanding access to formal care more generally reduces sickness absence and increases labor supply among informal caregivers (Løken et al., 2017; Frimmel et al., 2025; Shen, 2024), and that LTC insurance operates on the same margin (Coe et al., 2023; Fu et al., 2017). Mommaerts (2025), meanwhile, shows with a dynamic life cycle model how informal care crowds out the demand for, and reduces the welfare value of, such insurance.

The paper proceeds as follows. Section 2 describes Taiwan’s socio-demographic and policy environment and develops a stylized model of LTC needs and caregiving in multi-generational families, generating a set of testable predictions. Section 3 describes our data sources and key measures. Section 4 presents our workhorse staggered difference-in-differences design and reports our main effect estimates. Section 5 extends these estimates to non-stroke health shocks and extrapolates them to approximate the broader consequences of parental functional limitations. Section 6 examines how the effects of a parental stroke have evolved as caregiver-hiring criteria have liberalized. Section 7 presents a framework through which we can learn about willingness to pay from that evidence, and Section 8 uses it to estimate the welfare value of caregiver hiring for families that suffer strokes. Section 9 concludes.

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<sup>2</sup>One exception is Roth et al. (2013), which found caregivers tend to have *lower* mortality than non-caregivers with similar observables.

## 2 Background

### 2.1 Institutional Background

**LTC Needs in Taiwan.** With persistently low fertility rates and increasing life expectancies, Taiwan has one of the fastest-aging populations in the world. From 2000 to 2020, the share of the population over the age of 65 rose from 9 to 16 percent; by the end of 2025 it crossed 20 percent, formally making Taiwan a “super-aged” society (Ministry of the Interior, 2026). The government projects that this share will approach 47 percent by 2070, as fertility rates that have hovered below one child per woman since 2020 continue to shrink each successive generation (National Development Council, 2024). LTC needs are a large and growing social problem: the Ministry of Health and Welfare estimates that 923,000 Taiwanese (roughly 4 percent of the population) suffer functional limitations requiring long-term care as of 2026, a figure projected to reach 1.3 million by 2035 (Ministry of Health and Welfare, 2025b).

Taiwanese families predominantly meet these needs through informal caregiving performed by younger relatives. Beyond informal care, roughly one-third of LTC-needers hire an in-home caregiver from abroad. To hire such a caregiver, a family must receive government authorization, at which point a migrant caregiver is permitted to enter Taiwan on a short-term visa, live with the host family, and provide around-the-clock caregiving assistance. Institutional care (such as nursing homes) also exists but is less common than these at-home arrangements, representing roughly one-fifth of LTC cases (Chan and Chen, 2026).

The prevalence of in-home care, either by family members alone or with the aid of a live-in caregiver, aligns with the norms of filial piety which are well-established in Taiwan and other rapidly-aging East Asian societies. Traditionally, the greatest expectations of care are imposed upon the LTC-needer’s first-born son and—if married—that son’s spouse. In a 2013 government survey of LTC-needing families, 26 percent identified a son or a daughter-in-law (16 percent and 10 percent respectively) as the primary caregiver (Department of Health, 2013). By comparison, 29 percent identified the LTC-needer’s spouse as the primary caregiver, 8 percent identified a daughter, and the remaining 37 percent relied on care by other relatives (e.g., mothers, fathers, or siblings), formal caregivers, and nursing homes.

**LTC needs and the Taiwanese Social Safety Net.** Whereas Taiwan is well-known for modern health care services and a generous universal health insurance system, it maintains a limited set of social assistance programs for families that suffer LTC needs. These programs

provide a collection of small in-kind transfers, including up to 90 hours per month of day care, transportation services to hospitals, and subsidies on assistive devices.

The most important expansion of this system during our data period is the “Long-Term Care 2.0” plan, launched in 2017, which subsidizes home visits, day care, transportation, and respite services delivered through a network of community care stations. The program is tax-financed (from earmarked tobacco and estate tax revenues rather than social insurance premiums) and its budget has grown rapidly, from roughly 5 billion NTD in 2016 to 93 billion NTD by 2025 (Ministry of Health and Welfare, 2025a). Even so, public LTC programs are best understood as complements to, rather than substitutes for, family care. Subsidized services are capped at modest weekly hours and do not provide the around-the-clock supervision that severe functional limitations demand; moreover, families employing a live-in migrant caregiver were excluded from the program’s main home-care services until after the end of our data period.<sup>3</sup> Despite this growth, Taiwan continues to devote far fewer public resources to LTC (roughly 0.4 percent of GDP per year) than most OECD countries do (roughly 1.4 percent of GDP per year in public spending, out of 1.8 percent overall), and (unlike Japan, Korea, Germany, or the Netherlands) it has repeatedly declined to adopt public LTC insurance, retaining a tax-financed model (OECD, 2023; Ministry of Health and Welfare, 2025a).

Given the lack of public insurance against LTC needs, one might expect a care-needing Taiwanese family to turn to the general social safety net. But as we describe in Appendix B, the general safety net also provides little support. While income support programs exist, they are subject to aggressive *family*-level means tests, accounting for the resources of one’s parents and children and potentially other relatives. Participation in these programs is consequently low, amounting to only 2.6 percent of the Taiwanese population (Ministry of Health and Welfare, 2018).

Taiwanese families can also insure themselves against LTC needs privately. But as we lay out in Appendix C, the private LTC insurance market is long-recognized to be dysfunctional, and take-up is low. This is not a problem unique to Taiwan; private LTC insurance markets across many rich countries have similarly low take-up and are seen as incomplete and distorted. For instance, patterns in payouts and premiums suggest policies in the United States

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<sup>3</sup>Eligibility for LTC 2.0 services is determined by a questionnaire-based needs assessment administered by government care managers, which is separate from the Barthel-index certification process for hiring migrant caregivers described below. Families employing a migrant caregiver gained access to respite and short-term replacement care in 2023, and to the program’s home- and community-care services only in 2026 under the successor “Long-Term Care 3.0” plan (Ministry of Health and Welfare, 2025b).

are far from actuarially fair, especially after taking into account rates of policy forfeiture (Brown and Finkelstein, 2007). Rejection rates are high, and this has been attributed to asymmetric information at retirement age, when contracts are typically signed (Hendren, 2013). Meanwhile, benefits face high implicit tax rates due to interactions with Medicaid (Brown and Finkelstein, 2008).

With a “threadbare” social safety net and incomplete private insurance markets, shrinking Taiwanese families may face very high private costs when LTC needs arise. While there is no systematic study on how LTC needs affect indicators of distress like depression, domestic abuse, suicide, or parricide risks within families, these concerns have received wide attention from the government and national media. The auditing branch of the government in Taiwan, the Control Yuan, has filed a formal “Corrective Measure” that requires the executive branch of the government to act toward preventing these incidents (Control Yuan, 2020). There is no official statistic on these incidents, but the Taiwan Association of Family Caregivers has documented 62 cases between 2018 and 2024 in which a family caregiver killed the relative in their care (roughly one every six weeks) and in about one third of these cases the caregiver attempted suicide thereafter (Taiwan Association of Family Caregivers, 2024). Concerns are often raised in the Legislative Yuan (Taiwan’s parliament) as well; a keyword search of its session minutes returns more than 40 instances when LTC-related suicides were discussed from 2016 to 2022.

**Caregiver Regulations in Taiwan.** Migrant caregivers play a crucial role in the LTC system. As shown in Figure 1, the number of migrant caregivers grew from nearly 0 in the early 1990s to well over 200,000 by 2019. Since then, migrant caregivers have represented roughly 1 percent of Taiwan’s entire population (Ministry of Labor, 2025). Virtually all of Taiwan’s formal live-in caregivers are migrant workers. Hiring a migrant caregiver in the early 2020’s costs a family roughly 335,000 NTD (10,800 USD), excluding the cost of accommodations, or roughly one quarter of average gross household income (1.4 million NTD in the early 2020’s; DGBAS, 2026).<sup>4</sup> While this would be a major expense for most

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<sup>4</sup>The main annualized caregiver hiring costs include: 240,000 NTD in wages, 29,500 NTD in health insurance fees, 24,000 NTD in “employment stability fees”, and 1,000 NTD for a residence permit. Caregivers are also due overtime pay if they work during their weekly day off, as many do. In addition, families must pay for flights and hire through an intermediary, which typically charged around 9,000 NTD in additional fees in 2025 (see, for example, <https://www.kdmanpower.com.tw/news-detail/143/>). The largest component, wages, has been fixed at 20,000 NTD per month since 2022; from 2015 until then it stood at 17,000 NTD (Ministry of Labor, 2022).

families, the volume of work hours that migrant caregivers provide is substantial: migrant-hiring families in a 2019 survey reported receiving on average 13.4 daily working hours, with one-third saying the caregiver worked *every day* of the surveyed month (Ministry of Labor, 2020, Tables 2 and 3). Hiring domestic labor to provide services of comparable quantity and quality at the national minimum wage would cost 735,000 NTD annually—roughly twice the cost of a migrant care worker.<sup>5</sup> In this way, the allowance of migrant caregiver hiring acts like a large subsidy on high-volume formal care provision for qualifying families.

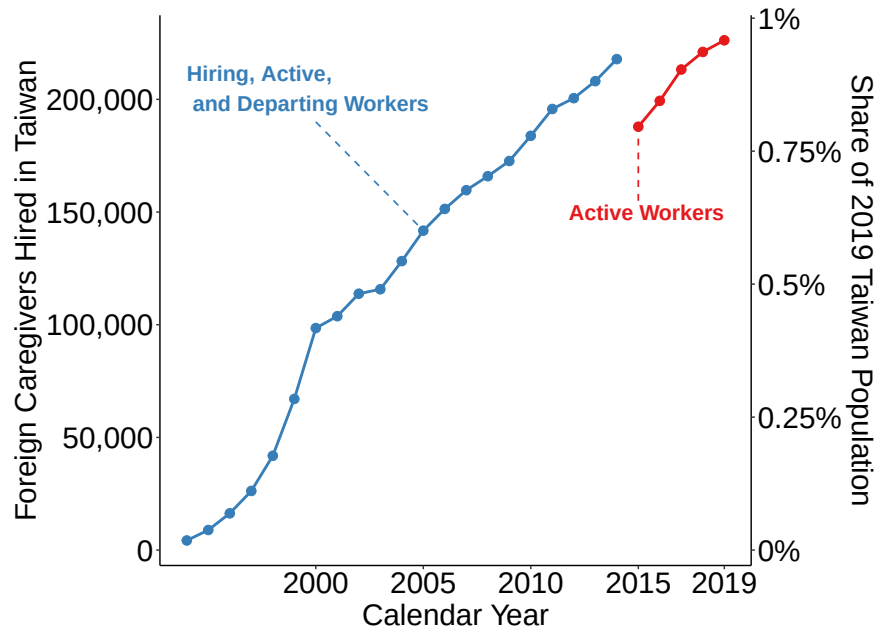


Figure 1: Number of Migrant Hired Caregivers

*Notes: The number of hired caregivers comes from the Ministry of Labor’s Yearbook of Labor Statistics: 1995–2019 from the 2024 edition (Ministry of Labor, 2025) and 1994 from the 2023 edition (Ministry of Labor, 2024). Through 2014, the reported counts include workers who had entered Taiwan but not yet completed hiring, and workers reported missing who had not yet departed; from 2015 onward, the counts include only workers actually under employment.*

The process of hiring a migrant caregiver is heavily regulated by the Taiwanese government. An individual that intends to hire a caregiver needs to obtain a certification of need issued by a doctor. The certification records the applicant’s Barthel index, which is a commonly-used measure that summarizes the number and severity of functional limitations

<sup>5</sup>We arrive at this calculation by taking the hourly minimum wage of 150 NTD, assuming 26 days of 13.3 hours of work, and accounting for an additional 18% in taxes, pension, and insurance fees. In reality, posted rates for domestic care workers are often above minimum wages. Intermediaries, news media, and government agency reports have quoted the annual cost of domestic care provision at 600,000-720,000 NTD throughout the 2010’s and early 2020’s, though the volume of care provision assumed in their calculations is unclear (Control Yuan, 2013; Li et al., 2009; Pension Fund Association, 2022; Silver CommonWealth, 2026).

suffered by an individual with a single numerical score (Mahoney and Barthel, 1965). An individual is approved if one’s Barthel index score is sufficiently low (meaning one has sufficiently many and severe limitations), where the score threshold below which one qualifies is higher (more lenient) at older ages. If an individual meets the eligibility criteria, she can formally start the hiring process.

To find a caregiver after receiving approval, families usually rely on intermediary firms that match them with foreign workers. A migrant caregiver receives a temporary work visa for the duration of her employment contract, and she resides in the household of the LTC-needer for this period. A contract is typically valid for three years, after which the LTC-needer must seek a renewal to continue receiving formal care. In most cases where the condition is severe, applying for renewal does not require a re-assessment of the LTC-needer’s health. The clear, formal ties between migrant caregivers and their host families limit the scope for undocumented immigration or the overstaying of visas through this system.

While this framework has been in place since 1992, two major reforms were implemented over our data period to expand access to formal caregiving. Both reforms relax the Barthel index thresholds that older-age LTC-needers must reach in order to receive approval for a migrant caregiver. The first reform, occurring in 2012, introduced the age gradient by relaxing the Barthel index threshold from a score of 35 to a score of 60 for LTC-needers who are 80 years of age or older (Council of Labor Affairs, 2012). The second reform, occurring in 2015, further relaxed the criteria so that any LTC-needer aged 85 or older is eligible to hire a migrant caregiver if her Barthel index is strictly less than 100 (Ministry of Labor, 2015).

The changes in the certification thresholds introduced by these reforms are substantial. To put them into clearer context, a Barthel score of 35 would *exclude* any LTC-needer with no bowel or bladder control issues, who can eat and use the toilet without assistance. By contrast, one can receive a score of 60 if one is unable to climb stairs, move 50 yards across a level surface, and transfer between a bed and chair.

More recently, a further legislative reform has continued this liberalization, lifting the Barthel Index requirement completely for LTC-needers age 80 and older and for those age 70 to 79 diagnosed with Stage II or more advanced cancer (Legislative Yuan, 2024). The exemption, passed in December 2024 and effective from August 2025, newly qualifies more than 600,000 elderly Taiwanese for caregiver hiring. While this reform is too recent to enter our data, it reflects continued political momentum toward liberalizing the caregiver hiring system. Opposition to this reform within the labor and health ministries was meaningful,

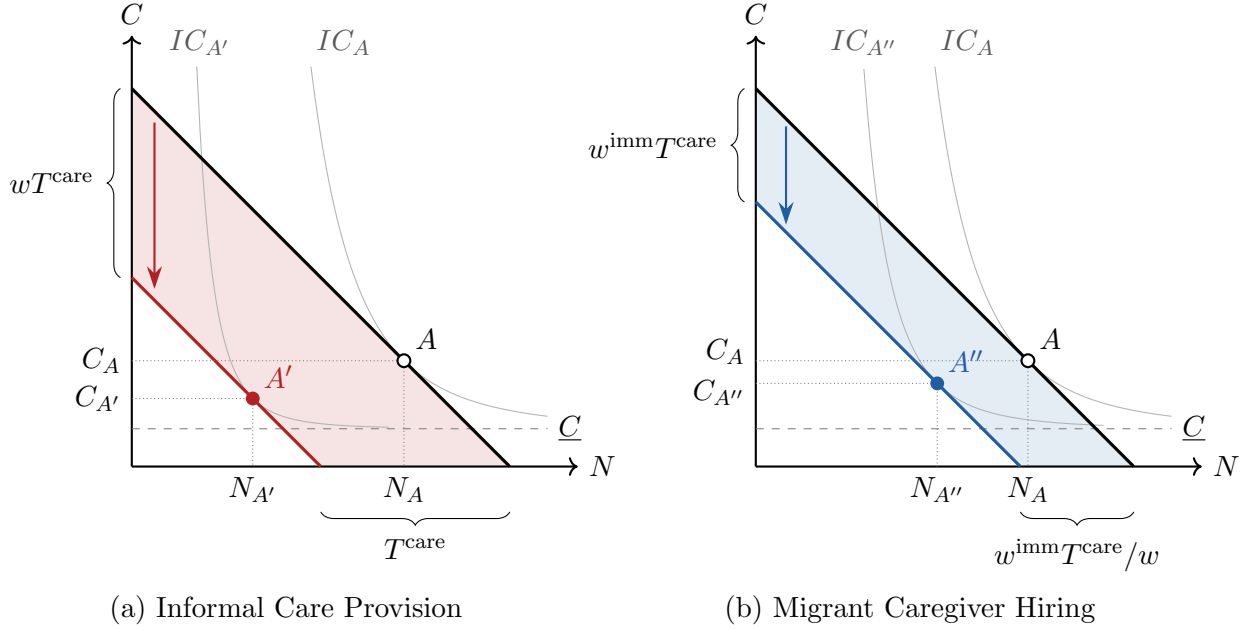


Figure 2: Budget Set in a Stylized Model with Family LTC Needs

*Notes:* Point  $A$  is the chosen bundle in the absence of long-term care needs. Point  $A'$  is the chosen bundle when care needs arise and informal care provision (at a time cost  $T^{\text{care}}$ ) is the only option for meeting them. Point  $A''$  is the chosen bundle when care needs arise and one can hire a migrant caregiver at total monetary cost  $w^{\text{imm}}T^{\text{care}}$ , where  $w^{\text{imm}}$  is less than the domestic prevailing wage  $w$ .

but based primarily on concerns that further expansion could outstrip the supply of migrant caregivers and have distributional consequences (Taiwan News, 2025).

## 2.2 Conceptual Framework

To formalize just one way in which care needs and caregiver hiring opportunities can seriously impact families, we present a stylized model of work and leisure which we view as a reasonable approximation of the true problem faced by many Taiwanese families. A work-capable adult (“worker”) with a dependent older relative (e.g., a parent) chooses home production ( $N$ ) and consumption ( $C$ ) to maximize utility. The worker trades off between these goods through a time endowment of  $T$  hours, which can be directly allocated to home production or to

funding consumption via work at wage  $w$ :<sup>6</sup>

$$\max_{C,N} U(C, N) \quad \text{s.t.} \quad T - T^{\text{care}} \mathbb{1}\{\text{Care Needs}\} = C/w + N$$

where preferences  $U$  are smooth and increasing in both  $C$  and  $N$ , with a subsistence point  $\underline{C}$  that the worker will maintain at any cost:

$$U_C(C, N) \rightarrow \infty \quad \text{as } C \rightarrow \underline{C}^+$$

We think of home production as capturing non-work time investments that facilitate good health or family development (meeting and marrying a partner, bearing and raising a child, etc.). Consumption is a composite good, and subsistence represents the point where the worker consumes only the food and shelter necessary for his immediate survival.

Panel a of Figure 2 demonstrates the comparative statics in this framework from introducing long-term care needs for the worker’s dependent relative. Those care needs require  $T^{\text{care}}$  hours of care which the worker can either perform himself or hire a *domestic* caregiver to perform at a wage similar to his own. In either case, the cost is the same: amounting to a leftward shift in the budget constraint by  $T^{\text{care}}$  or an equivalent downward shift of  $wT^{\text{care}}$ .

**Key insights.** This framework provides two key insights that will shape our empirical analyses. First, LTC needs will primarily cause workers in small families to reduce home production—not consumption—simply because they *cannot afford* to forgo their sole income source. The preferred bundle in the absence of care needs (bundle  $A$ ) is already close to the subsistence level of consumption. Thus, we expect to see the costs of LTC provision manifest not in reduced work or earnings (which fund consumption), but instead in outcomes generated by home production: marriage, fertility, and good health. Workers adjust on the home production margin not because doing so is costless, but because the alternative of reducing consumption would be even more costly.

Our second insight comes from considering how migrant caregiver hiring should affect these families. The option to hire a migrant caregiver, willing to work at an hourly wage

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<sup>6</sup>The framework we lay out here can be connected to the static problem of the child generation in the two-generation, non-cooperative model of LTC needs of Barczyk and Kredler (2018). Our framework is close to a special case of that model where the parent generation has limited income and savings, and no Medicaid coverage exists (so that informal caregiving or child-funded formal caregiving must occur). We view this as a reasonable approximation of the circumstances faced by many Taiwanese families. Unlike Barczyk and Kredler (2018), we distinguish between consumption and home production in order to make theoretical predictions regarding marriage, fertility, and mortality.

$w^{\text{imm}} < w$ , effectively reduces the cost of long-term care provision. In hiring the migrant caregiver, the worker’s budget for consumption and home production is shifted downward by only  $w^{\text{imm}}T^{\text{care}}$ , mitigating the adverse effects of care needs on the worker’s budget set and allowing the worker to enjoy a higher level of home production.

**Relevance.** While plausibly relevant for *some* families in many developed economies, we argue that this framework provides a good representation of the *average family’s problem* in Taiwan for three reasons. First, families in Taiwan tend to be self-reliant: we already established the “threadbare” nature of Taiwan’s social safety net in Section 2.1. Second, Taiwanese families tend to be small: like many developed countries, Taiwan has faced below-replacement fertility since the mid-1980s. Third, families tend to allocate a large fraction of their consumption toward basic necessities, suggesting that cutting back on work income would be difficult. Data from the Taiwan Survey of Family Income and Expenditure indicates that 31 percent of non-health consumption was dedicated toward food alone in 2015. By comparison, only 16 percent of non-health consumption goes to food in the United States, based on Consumer Expenditure Survey data.<sup>7</sup>

A second, complementary mechanism may extend this prediction even to higher-income families, though it falls outside the stylized model above: opportunities to mix caregiving with part-time employment may simply be unavailable. Part-time employment is exceedingly rare in Taiwan, representing less than 4 percent of all employment (Ministry of Labor, 2023). If small families cannot afford a worker forgoing employment *entirely*, then LTC needs will force even high-income workers to substitute away from home production instead.

### 3 Data

How LTC needs affect families is ultimately an empirical question. We take this question to data consisting of multiple linked Taiwanese administrative registries. Our main resource is the National Health Insurance Research Database (NHIRD), which contains administrative records from the universal health care insurance system. We link multiple datasets within the NHIRD to the Household Registry, which is an actively-maintained running register of all Taiwanese households. In this way we construct a dataset that contains rich

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<sup>7</sup>These statistics are based on food spending both at and outside home. However, the Survey of Family Income and Expenditure does not distinguish food spending outside the home from (comparatively smaller) spending on hotels. Food at home on its own accounts for 17 percent and 9 percent of total non-health spending in Taiwan and the United States respectively.

demographic information, family structure, employer-reported earnings records, and other economic variables, complete inpatient, outpatient, and drug records, migrant caregiver hiring, and statuses for disability and middle-low income programs. Our data covers virtually the whole population of more than 23 million individuals, and we construct a panel spanning from 2000 to 2020.

Creating family linkages is a key step in the process of constructing our dataset. We describe how we do this—combining health insurance records, household registries, and birth records—in Appendix D. Our matching process is not perfect, and struggles in particular with matching women who had already formed independent households by 2000. We therefore focus our main analysis on adult sons, for whom linkages are most reliable.

**Measurement of health shocks and LTC needs.** We detect negative health shocks to the elderly using medical record data drawn from the NHIRD. In our main analyses, we focus on the onset of a stroke. We isolate the first time in an individual’s medical history that an ICD code tied to the diagnosis or treatment of a stroke arises.<sup>8</sup>

We focus first on strokes because they are strongly linked to an abrupt change in care needs. In doing so, we follow a public health literature in which strokes have been a frequent focal event (Bickel et al., 2012; Dobkin et al., 2018; Fadlon and Nielsen, 2019, 2021). Using representative survey data collected every four years on elderly adults from the TLSA, Table 1 shows that the onset of a stroke is associated with sharper increases in many functional limitations than hip fractures—another regularly-studied health shock with ties to care needs.

### 3.1 Sample Construction

We construct our analytic sample by first identifying all individuals who suffer one of our identified health shocks—“LTC-needers” hereafter. For each LTC-needer, we identify the *first time* they experience a shock, and we take that first shock to be our event of interest.<sup>9</sup> We then merge this core dataset of health events with panel data on family members, based on family links that had been established at least one calendar year before the shock. Indexing links in this way is most relevant for spouses and children-in-law who can enter and exit the family over time through marriage and divorce.

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<sup>8</sup>The ICD-9 and ICD-10 codes we use are listed in Appendix E.

<sup>9</sup>When we study other conditions in addition to strokes, we handle sample construction separately for each condition to maintain consistency with the sample used in our main analysis. For instance, a person who suffers a stroke and later suffers dementia will appear in both the stroke and dementia samples, with different dates of onset.

Table 1: Changes in elderly functional limitations around strokes and other health events

|                              | Strokes |                  | Hip Fractures |                  |
|------------------------------|---------|------------------|---------------|------------------|
|                              | Pre     | Post-Pre         | Pre           | Post-Pre         |
| Bathing Difficulty           | 0.1988  | 0.69<br>(0.10)   | 0.2242        | 0.34<br>(0.09)   |
| Clothing Difficulty          | 0.1601  | 0.65<br>(0.10)   | 0.1858        | 0.34<br>(0.08)   |
| Eating Difficulty            | 0.0863  | 0.36<br>(0.07)   | 0.1074        | 0.15<br>(0.06)   |
| Ambulating Difficulty        | 0.1488  | 0.49<br>(0.08)   | 0.1681        | 0.31<br>(0.08)   |
| Walking Difficulty           | 0.1648  | 0.57<br>(0.09)   | 0.1824        | 0.35<br>(0.08)   |
| Incontinence                 | 0.1535  | 0.54<br>(0.09)   | 0.1727        | 0.40<br>(0.09)   |
| <b>Barthel Score (upper)</b> | 95.5288 | -16.50<br>(2.51) | 94.8197       | -8.56<br>(2.13)  |
| <b>Barthel Score (lower)</b> | 94.1641 | -19.95<br>(2.79) | 93.4208       | -10.89<br>(2.55) |

*Notes:* This table reports the differential change in functional limitations an elderly person reports in the last survey wave just before (“Pre”) versus first survey wave just after (“Post”) reporting the given health condition, compared to otherwise-similar peers in the TLSA. Individuals are exactly matched on birth year and sex. TLSA survey waves occur every four years. The final two rows use these limitation reports to impute a Barthel score. Limitation reports are binary in the TLSA, which is not sufficiently granular to exactly construct a Barthel score. Instead, we construct a lower bound by supposing any reported limitations are “severe,” and an upper bound by supposing they are “moderate”. See Chan and Chen (2026) for detailed construction.

We refine our sample of health shock events with three restrictions. First, our design requires that we observe pre-shock outcomes, so we exclude events that occur in the first year of our data. Second, we restrict our sample to cases where this health shock event occurs after the LTC-needer is age 50, out of concern that the few earlier-life shocks may be less severe, less random, or more foreseeable. Third, we exclude a health shock event if the LTC-needer’s spouse had already suffered a shock, to avoid potential endogeneity concerns. Our exercise takes the remaining health shock events, identifies the adult sons of the LTC-needers, and follows those adult sons over the entire data period, until either the end of the panel or the sons’ own deaths. We do not condition on the survival of the LTC-needer past their health shock.

### 3.2 Summary Statistics

Families that suffer a stroke are not a representative sample of the population of Taiwanese families. To contextualize the responses that we document and to describe how the burden of strokes is distributed across society, we first examine the observable characteristics of stroke-suffering families. We present summary statistics on these families in Table 2, comparing them to families which are never observed suffering a stroke.

In total, we observe almost 1.7 million unique families in which an elderly person suffers a stroke. Within those families, the average older generation consists of 1.49 adults (i.e., we find 0.49 spouses for each stroke-sufferer), who we link to an average of 1.73 younger relatives including daughters-in-law. We manage to link more sons (0.83 on average) than we do daughters (0.49 on average) consistent with the limitations of our matching procedure. Compared to the average family of an older-age Taiwanese, those who suffer a stroke tend to have similar average family sizes (3.22 immediate relatives versus 3.20). However, stroke-suffering families have fewer identifiable daughters (0.49 versus 0.66) but more identifiable daughters-in-law (0.41 versus 0.22). Table 2 also shows that the composition of stroke-suffering families is roughly similar to the composition suffering a hip fracture or dementia diagnosis, except that those suffering dementia tend to have slightly fewer adult children.

Table 2: Average Family Links to Old-Age Taiwanese Adults

|                    | Stroke Sample |                  | Hip Fracture | Dementia    |
|--------------------|---------------|------------------|--------------|-------------|
|                    | Never-Treated | Treated (Stroke) |              |             |
| Elderly Parents    | 1.48          | 1.49             | 1.46         | 1.45        |
| Sons               | 0.84          | 0.83             | 0.82         | 0.77        |
| Daughters          | 0.66          | 0.49             | 0.43         | 0.42        |
| Daughters-in-Law   | 0.22          | 0.41             | 0.45         | 0.44        |
| <b>Total</b>       | <b>3.20</b>   | <b>3.22</b>      | <b>3.16</b>  | <b>3.07</b> |
| Number of Families |               | 1,694,377        | 526,891      | 395,222     |

*Notes:* This table reports statistics on our main analytic sample of families as defined in Section 3.1. Averages among never-treated families are taken based on a 5% random sample.

Table 3: Covariate balance between sons in stroke-suffering families, their matched controls, and the broader population

|                              | Other Families | Analytic Sample Families |         |                   |
|------------------------------|----------------|--------------------------|---------|-------------------|
|                              | Never-Treated  | Control                  | Treated | Treated - Control |
| Parent's Age                 |                | 65.99                    | 67.01   | 1.02              |
| Married (pct)                | 49.80          | 50.08                    | 49.95   | -0.13             |
| Children per Hundred         | 91.18          | 92.39                    | 92.59   | 0.20              |
| Worked (at all)              | 0.74           | 0.74                     | 0.74    | 0.00              |
| Months Worked                | 8.32           | 8.31                     | 8.33    | 0.01              |
| Annual Earnings (1k)         | 301.60         | 297.35                   | 298.43  | 1.08              |
| Annual Hospital Bed Days     | 0.92           | 0.95                     | 1.03    | 0.08              |
| Annual Medical Spending (1k) | 11.92          | 12.15                    | 12.57   | 0.41              |
| Depression Diagnoses         | 13.18          | 13.66                    | 14.20   | 0.54              |
| Son's Age at Parent's Onset  |                | 43.85                    | 38.72   | -5.13             |

*Notes:* This table reports statistics on our main analytic sample of adult sons in families defined in Section 3.1, in event-year -1. Control individuals consist of individuals in later-treated families which have been matched to treated families on birth year, education level, and family structure. “Parent’s age” and “age at parent’s own onset” are omitted from the never-treated column because they are not defined in this sample by construction. Averages among never-treated families are taken from a random sample in the health insurance records. For variable definitions, see notes on Table 4.

## 4 Responses to a Parent’s Stroke

In this section, we estimate the average effect of health shocks in our data, integrating across the various caregiver hiring regimes that occurred over our sample period. We defer distinguishing the effects of health shocks under alternative caregiver hiring regimes to Section 6.

Our core parameters of interest are average treatment effects on the treated. Formally, let  $Y_{it}^c(1)$  denote the potential outcome an individual  $i$  would experience  $t$  years from calendar year  $c$ , having suffered a family health shock in year  $c$ . Let  $Y_{it}^c(0)$  denote the potential outcome that same individual would experience if the shock had not yet occurred by calendar year  $c+t$ .<sup>10</sup> Our first object of interest is the  $t$ -specific average treatment effect on the treated, which averages over effects for different shock cohorts  $c$ :

$$ATT(t) \equiv E_c [E [Y_{it}^c(1) - Y_{it}^c(0) \mid D_i^c = 1]], \quad (1)$$

where  $D_i^c$  is an indicator that equals 1 if person  $i$  belongs to cohort  $c$  (i.e., suffered a family health shock in year  $c$ ). Thus,  $ATT(t)$  is the average effect,  $t$  years after a health shock, among individuals whose families actually experienced one. We also examine the overall ATT that averages across all available post-shock event times  $t \geq 0$ :

$$ATT \equiv E_t [ATT(t)], \quad t \geq 0. \quad (2)$$

## 4.1 Empirical Strategy

In order to identify and estimate the effects of elderly health shocks, we rely on a staggered difference-in-differences design. This design compares trends around the calendar year of a health shock for individuals in affected families to trends among contemporaneous “control” individuals who are observably similar but belong to families that suffer health shocks at a later date. The validity of the design hinges on a common trends assumption: had one’s health shock occurred *later*, one’s outcomes would have continued evolving over time with the same trends observed among the actually-later-shocked families. This assumption is implied by—but is weaker than—the assumption that the timing of health shocks is random with respect to one’s outcomes.<sup>11</sup> That the timing of old-age strokes is plausibly random therefore makes strokes particularly amenable to this design.

Taking the design to our data, we estimate a series of two-by-two difference-in-differences regressions, each on different subsamples consisting of observations from a given event-year  $\ell$  and from a fixed pre-shock reference event-year  $\ell_0$ :

$$Y_{it}^c = \alpha_{ic} + \gamma_\ell \mathbb{1}\{t = \ell\} + \delta_\ell D_i^c \mathbb{1}\{t = \ell\} + \epsilon_{it}^c, \quad t \in \{\ell_0, \ell\}, \quad (3)$$

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<sup>10</sup>Note that the no-anticipation assumption is built into the untreated potential outcome notation; if  $c_0$  is the actual year of treatment assigned to person  $i$  (where  $c_0 > c + t$ ), note that  $Y_{it}^c(0)$  is not a function of  $c_0$ .

<sup>11</sup>If the timing of health shocks is fully random, we would see not just common pre-shock trends between shocked and later-shocked families, but common pre-shock *levels*.

estimated separately for each  $\ell \neq \ell_0$ , where  $Y_{it}^c$  denotes the outcome for person  $i$ ,  $t$  years away from calendar year  $c$ . The sub-sample contains individuals from “treated” families that suffer a health shock in calendar year  $c$  (for whom  $D_i^c = 1$ ), as well as otherwise similar individuals from later-treated families who serve as “controls” (for whom  $c$  denotes the year in which their treated peer’s family suffered a health shock). These later-treated controls are reweighted to exactly match the treated families on the distribution of observable characteristics.<sup>12</sup> Beyond the restrictions we already imposed on our analytic sample in Section 3.1, the characteristics on which we balance in all analyses are birth year (for person  $i$ ), education level (six categories: none, elementary, junior high, high school, college, above college), and family structure (LTC-needer has a spouse and zero vs. one vs. multiple children) in calendar year  $c - 1$ .<sup>13</sup> Our design restricts to relatives who survive up to  $t = -1$ , but allows both treated and control relatives to die in years  $t \geq 0$ , falling out of the data in subsequent years.

Under the common trends assumption above, the coefficients  $\delta_\ell$  recover  $ATT(\ell)$  from Equation 1 for the given outcome  $Y$ . Comparisons of  $ATT(\ell)$  across values of  $\ell$  need not reflect treatment effect dynamics alone, however. Later-treated cohorts contribute only to estimates at shorter horizons, so differences across  $\ell$  can also reflect heterogeneity in effects across cohorts. We construct overall ATT estimates (Equation 2) throughout the paper by averaging  $\hat{\delta}_\ell$  over post-shock event times  $\ell \geq 0$ .

## 4.2 Pre-Stroke Outcomes and Covariate Balance

If the timing of strokes is indeed conditionally random, then we should find no meaningful conditional differences across treatment cohorts in observable pre-stroke characteristics. Table 3 compares average pre-period characteristics of adult sons and their stroke-suffering parents to those of the later-treated controls. For reference, we also report average characteristics of adult sons whose parents are never observed suffering a stroke. By design, our treated and control samples of adult sons match exactly on their own age, birth year, and education which we omit from the table. They are nearly indistinguishable on average marital status—a non-mechanical result because we balance on marital status only for the LTC-needers themselves.

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<sup>12</sup>Our estimator amounts to the IPW estimator proposed by Callaway and Sant’Anna (2021).

<sup>13</sup>Note that family structure is time-varying; an individual can appear as a control unit for two different family structures, for different values of  $c$ , if that individual’s family structure changes between those two calendar years.

We do not balance on one’s own pre-period number of children, employment, earnings, or health. It is therefore reassuring to see that differences on these pre-stroke outcomes between treated and later-treated adult sons are economically negligible and much smaller than differences between treated and never-treated families. We also do not directly balance on the LTC-needing parent’s age, in order to keep the set of matching characteristics parsimonious. Even so, later-LTC-needing parents in the weighted control group are only one year younger on average than LTC-needing parents in the treated group—despite suffering their own strokes roughly five years later in calendar time (when their adult sons are on average age 43.85 for the later-treated versus age 38.72 for the treated). Given that we *do* balance exactly on age of the adult sons, this suggests we are exploiting primarily variation in the timing of a stroke across otherwise-similar families as intended, instead of comparing adult sons with very young parents in the control group to otherwise-similar adult sons with very old parents in the treated group.

### 4.3 Results

**Overall effects on mortality, health, and family formation.** We begin by reporting estimates for the overall ATT, pooling across all post-stroke years, in Column 1 of Table 4. Strokes are tied to substantial adverse effects on the health and well-being of adult sons, all of which are very precisely estimated. Mortality rates increase by an additional 0.56 deaths per thousand per year. Relative to mortality rates among later-treated peers of 3.05 deaths per thousand per year, this is a 19 percent increase. This includes more deaths classified explicitly as suicides, the risk of which increases by 0.053 per thousand per year—a 35 percent increase over the baseline of 0.151 suicides per thousand per year among later-treated peers. While suicides explain a modest 10 percent of the total mortality effects, stigma against classifying suicide as a cause of death means this estimate likely represents a lower bound on the mortality effect coming from intentional self-harm.

Other measures of health and family formation suggest modest adverse effects beyond mortality, relative to baseline means among later-treated peers. Depression diagnoses increase by 2.90 diagnoses per thousand (4.4 percent) and their medical costs increase by 1000 NTD per year (5.2 percent). Five years after the stroke, the probability that an adult son would be married falls by 1.28 per thousand and their fertility declines by 4.04 per thousand (a 2.6 percent decline in fertility over this period).

Because household registry links require family members to be alive in 2011, they are selected on survival up to that point. In Appendix Table A.1, we show our main estimates

are similar (and in the case of mortality, *larger* in magnitude) when we instead build families exclusively from health insurance records.

**Outcomes of older versus younger adult children.** How the costs of a parent’s stroke materialize for adult children should depend on the life cycle phase in which the shock occurred. Marriage and fertility tend to happen when adult children are relatively young, whereas mortality tends to be a greater risk at older ages. Moreover, a simple model of “suicidal” behavior would predict larger mortality responses to a negative shock among older people than younger people (e.g., Hamermesh and Soss, 1974). Motivated by this reasoning, we split our sample based on the life cycle phase of the adult son one year prior to the stroke event and estimate effects separately for each group. We group individuals into three age bins based on their age in the year before their parental stroke: i) 33 and younger, ii) between 34 and 48 years old, and iii) 49 and older. We report the results in Columns 4, 5, and 6 of Table 4.

Effects on all-cause mortality tend to increase in magnitude as adult sons progress through the life cycle. Annual all-cause mortality increases by 0.13 deaths per thousand among the youngest age group (an 11 percent increase in mortality compared to later-treated peers), but by 1.74 per thousand among the oldest one (a 26 percent increase in mortality compared to later-treated peers). We see roughly a 5 percent increase in medical spending across the age distribution: a 755 NTD per year increase among those 33 and younger versus 1823 NTD per year among those age 49 and older.

Decreases in marriage and fertility are concentrated in the earliest phase of the life cycle, suggesting that a parental stroke disrupts the household formation process. Among adult sons who are younger than 34 prior to the stroke, marriage rates decline by 2.8 per thousand (a 2 percent decline in net marriage rates) and cumulative fertility declines by 8.4 children per thousand five years later (a 3 percent decline in fertility). Marriage and fertility do not change much among older-age adult sons, for whom families have largely settled by the time of the health shock.

**Dynamic effects of a stroke.** While we have so far focused on average effects of a stroke integrating over post-stroke years, it can be informative to examine how these responses vary over time. We next plot in Figures 3 and 4 estimates for the “dynamic” average effects of a parent’s stroke (the separate event-time response coefficients  $\delta_\ell$  for each  $\ell$  in Equation 3) for several key outcomes: those related to family formation (marriage and fertility) in Figure 3 and those related to health and economic outcomes (employment, mortality, medical

spending, and depression) in Figure 4. These figures show *increases in magnitude* over time in the year-by-year average effects for all outcomes. This suggests that parental health shocks do not merely cause a slight shift in the timing of life events for adult sons, but instead a persistent trend break in them. Constructing these “dynamic” estimates in pre-stroke years also generally reveals no meaningful evidence of pre-trends, consistent with the idea that strokes are unanticipated random events.<sup>14</sup>

**Comparing the scale of labor and non-labor effects of a stroke.** The effects of a stroke on mortality are large in magnitude, relative to their effects on income and labor supply. Accumulating the year-by-year employment effects among surviving sons gives 6.6 lost work-years per thousand over the first five years after a stroke. Accumulating the year-by-year mortality effects over that period implies 3.0 years of lost life per thousand sons, counting only years *after* the year of death. If sons who died would otherwise have maintained their pre-stroke employment rate (0.74 workers per adult), these deaths alone account for 2.22 lost worker-years per thousand adults. Under that assumption, accounting for mortality-related losses would increase the *near-term employment effects alone* from a parental stroke by roughly one-third.

## 4.4 Evaluating Channels for Adult Children’s Responses

Long-term care needs are one channel by which parental health shocks may be linked to adverse outcomes for adult children. We next gauge their role by exploring effect heterogeneity across families that should be better-positioned versus worse-positioned to absorb sudden care needs. Since we cannot measure care needs or care provision directly, we rely on two proxies: i) the adult son’s educational attainment and ii) whether he has identifiable siblings. We summarize heterogeneity along these proxies in Table 5.

Overall, our results are consistent with the reasoning laid out in the conceptual framework of Section 2.2: effects on mortality, medical spending, and employment attenuate for sons who have siblings or higher levels of education. The gradient on mortality effects is

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<sup>14</sup>We do detect some evidence of anticipation for marriage and fertility responses (Figure 3b). We investigate this further in Appendix Figure A.1, where we repeat the event study plot after splitting the younger age group into three finer subpopulations based on their pre-stroke family status: unmarried adult sons, married but childless adult sons, and married adult sons with pre-existing children. The pre-trend in marriage disappears. Pre-trends in fertility come entirely from married adult sons with pre-existing children, who tend to have slightly fewer children than their later-treated peers in the year just before the stroke. Meanwhile, fertility effects are driven most by unmarried sons.

Table 4: Economic and Health Effects on Adult Sons of Elderly Parents' Strokes

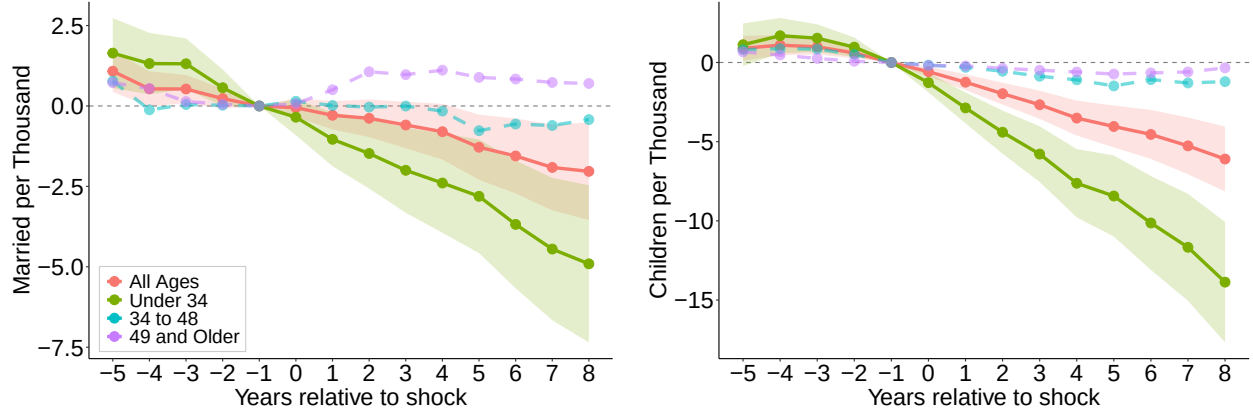
|                                                 | (1)             | (2)                             | (3)             | (4)                                      | (5)             | (6)             |
|-------------------------------------------------|-----------------|---------------------------------|-----------------|------------------------------------------|-----------------|-----------------|
|                                                 | Stroke          | Overall Effects<br>Hip Fracture | Dementia        | Effects of Stroke, by Age of Son<br>≤ 33 | 34 to 48        | ≥ 49            |
| <hr/> Health <hr/>                              |                 |                                 |                 |                                          |                 |                 |
| Deaths per 1k                                   | 0.56<br>(0.06)  | 0.81<br>(0.20)                  | 0.66<br>(0.23)  | 0.13<br>(0.03)                           | 0.53<br>(0.07)  | 1.74<br>(0.24)  |
| Suicides per 1k                                 | 0.05<br>(0.01)  | 0.09<br>(0.02)                  | 0.03<br>(0.05)  | 0.03<br>(0.01)                           | 0.05<br>(0.01)  | 0.12<br>(0.02)  |
| Depression per 1k                               | 2.90<br>(0.31)  | 1.98<br>(0.86)                  | 2.85<br>(1.12)  | 2.92<br>(0.36)                           | 2.99<br>(0.42)  | 2.56<br>(1.04)  |
| Medical Spending                                | 1000<br>(166)   | 1303<br>(373)                   | 729<br>(635)    | 755<br>(202)                             | 927<br>(213)    | 1823<br>(476)   |
| Hospital Days                                   | 0.13<br>(0.02)  | 0.24<br>(0.07)                  | 0.08<br>(0.09)  | 0.12<br>(0.02)                           | 0.14<br>(0.03)  | 0.11<br>(0.05)  |
| <hr/> Employment <hr/>                          |                 |                                 |                 |                                          |                 |                 |
| Worked (at all) per 1k                          | -7.5<br>(0.7)   | -9.9<br>(1.8)                   | -6.5<br>(2.2)   | -4.3<br>(0.9)                            | -6.3<br>(0.8)   | -19.1<br>(2.1)  |
| Months Worked                                   | -0.10<br>(0.01) | -0.12<br>(0.02)                 | -0.10<br>(0.03) | -0.06<br>(0.01)                          | -0.09<br>(0.01) | -0.23<br>(0.03) |
| Work Income (NTD)                               | -2321<br>(478)  | -4460<br>(1199)                 | -3373<br>(1467) | -734<br>(490)                            | -1722<br>(554)  | -8020<br>(1423) |
| <hr/> Household Formation (5 years later) <hr/> |                 |                                 |                 |                                          |                 |                 |
| Married per 1k                                  | -1.3<br>(0.5)   | -1.7<br>(1.6)                   | 0.5<br>(1.9)    | -2.8<br>(0.9)                            | -0.8<br>(0.7)   | 0.9<br>(1.0)    |
| Children per 1k                                 | -4.0<br>(0.7)   | -4.6<br>(1.7)                   | -0.4<br>(2.2)   | -8.4<br>(1.3)                            | -1.5<br>(0.8)   | -0.7<br>(0.4)   |
| Treated Units                                   | 1,224,087       | 94,037                          | 66,212          | 453,535                                  | 547,286         | 223,266         |

*Notes:* Estimates are obtained from the event study in Equation 3 pooling all post-event periods. Treated unit counts exclude individuals who failed to match with a single later-treated control unit. This, along with the fact that we focus on adult sons, is why we report fewer treated units than families in Table 2. Columns 4-6 report estimates separately by age in the year before the parent's stroke. Standard errors are reported in parentheses. Standard errors are clustered at the individual level. "Worked (at all) per 1k" indicates if one has earned work income at or above the monthly minimum wage in at least one calendar month of the year; "Months Worked" counts such months, and "Work Income" sums earnings over them. "Earnings" are measured in NTD per year (1 USD is approximately NTD 31). "Depression" is the indicator of whether an individual has been ever diagnosed with depression (a measure of the stock of depression diagnoses). See Appendix E for the depression ICD-9 and ICD-10 codes. Effects for the final two outcomes, Married per Thousand and Children per Thousand, are reported at a specific point in time because these measures are stocks rather than flows.

Table 5: Heterogeneity Analysis: Responses to Strokes and Proxies for the Scope of LTC Needs

|                                                 | (1)             | (2)             | (3)             | (4)               | (5)             |
|-------------------------------------------------|-----------------|-----------------|-----------------|-------------------|-----------------|
|                                                 | By Education    |                 |                 | By Sibling Status |                 |
|                                                 | Below HS        | HS Degree       | Above HS        | Only Child        | Has Siblings    |
| <hr/> Health <hr/>                              |                 |                 |                 |                   |                 |
| Deaths per 1k                                   | 1.68<br>(0.17)  | 0.36<br>(0.12)  | 0.16<br>(0.06)  | 1.12<br>(0.14)    | 0.31<br>(0.07)  |
| Suicides per 1k                                 | 0.13<br>(0.03)  | 0.04<br>(0.02)  | 0.02<br>(0.01)  | 0.09<br>(0.01)    | 0.04<br>(0.01)  |
| Depression per 1k                               | 3.42<br>(0.74)  | 2.62<br>(0.54)  | 2.85<br>(0.44)  | 3.28<br>(0.70)    | 2.72<br>(0.32)  |
| Medical Spending                                | 1254<br>(467)   | 1108<br>(279)   | 786<br>(201)    | 1286<br>(382)     | 869<br>(167)    |
| Hospital Days                                   | 0.25<br>(0.06)  | 0.15<br>(0.04)  | 0.05<br>(0.02)  | 0.14<br>(0.04)    | 0.12<br>(0.02)  |
| <hr/> Employment <hr/>                          |                 |                 |                 |                   |                 |
| Worked (at all) per 1k                          | -12.8<br>(1.5)  | -7.1<br>(1.1)   | -5.1<br>(1.1)   | -13.5<br>(1.4)    | -4.7<br>(0.8)   |
| Months Worked                                   | -0.17<br>(0.02) | -0.09<br>(0.01) | -0.07<br>(0.01) | -0.19<br>(0.02)   | -0.06<br>(0.01) |
| Work Income (NTD)                               | -3838<br>(655)  | -2308<br>(602)  | -1559<br>(901)  | -4349<br>(998)    | -1389<br>(510)  |
| <hr/> Household Formation (5 years later) <hr/> |                 |                 |                 |                   |                 |
| Married per 1k                                  | 1.2<br>(1.0)    | -1.5<br>(0.9)   | -2.4<br>(0.8)   | -1.6<br>(0.9)     | -1.1<br>(0.6)   |
| Children per 1k                                 | -0.9<br>(1.0)   | -3.5<br>(1.1)   | -6.1<br>(1.1)   | -2.6<br>(1.1)     | -4.7<br>(0.8)   |
| Treated Units                                   | 261,047         | 415,227         | 547,813         | 342,361           | 881,726         |

*Notes:* Estimates are obtained from the event study in Equation 3, pooling all post event periods. For other notes, see Table 4. Standard errors are clustered at the individual level.



(a) Marriage responses

(b) Fertility responses

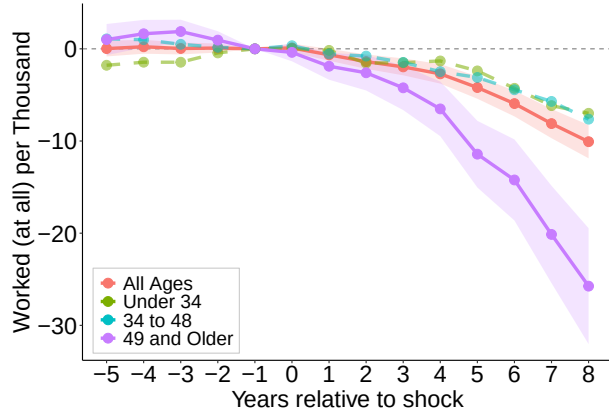
*Notes:* These plots report estimates separately by age of the adult son. The x-axis plots the years relative to the LTC-needer's shock. Estimates reported here correspond to the event study estimates,  $\delta_\ell$ , in Equation 3. 95 percent confidence intervals are plotted as the shaded regions for one subgroup (age 33 and under) and for the pooled estimates. Standard errors are clustered at the individual level.

Figure 3: Event study responses of adult sons to elderly parents' strokes: family formation outcomes

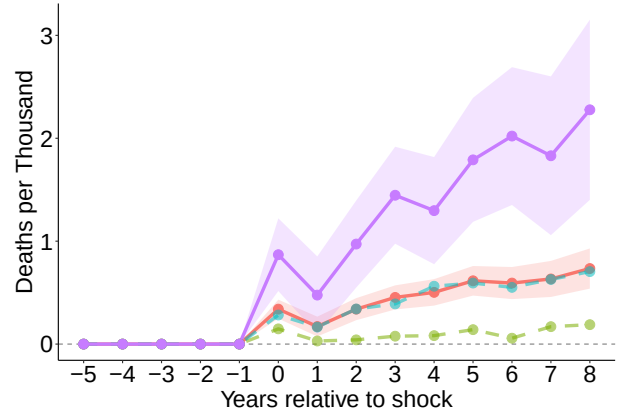
particularly sharp; it is an order of magnitude larger for high school dropouts than for those with schooling beyond a high school degree (1.68 versus 0.16 deaths per thousand per year). Mortality effects are nearly 4 times larger for adult sons with no siblings (1.12 deaths per thousand per year) compared to those who have at least one sibling (0.31 deaths per thousand per year).

This exercise also reveals that the increases in the risk of relative mortality for adult sons coming from vulnerable families are very large. Compared to their similar later-treated peers, mortality among high school dropouts increases by 28 percent and among sons with no identifiable siblings by 38 percent. Suicide rates in these groups increase by 53 percent (0.13 deaths per thousand per year) and 89 percent (0.09 deaths per thousand per year) respectively.

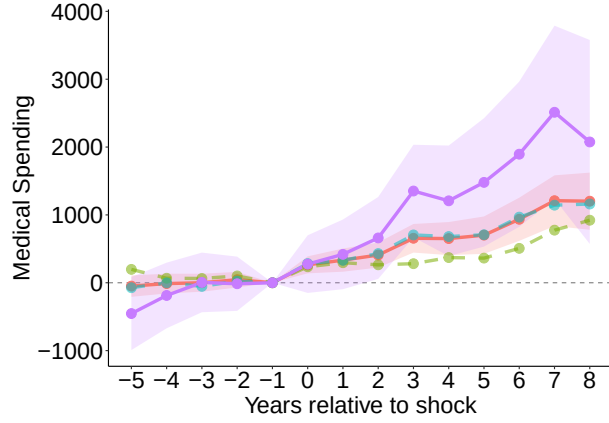
Marriage and fertility effects do not clearly follow this pattern: they are indistinguishable across sons with versus without siblings, and are actually *larger*—not smaller—for sons with more schooling. However, less-educated adult sons tend not to be forming families in the first place. Over the first five post-stroke years for adult sons with less than a high school degree, their comparison later-treated peers have only 78 children per thousand (and *declining* marriage rates). By comparison, the later-treated peers of adult sons with more than a high school degree have 207 children per thousand (and 107 marriages per thousand).



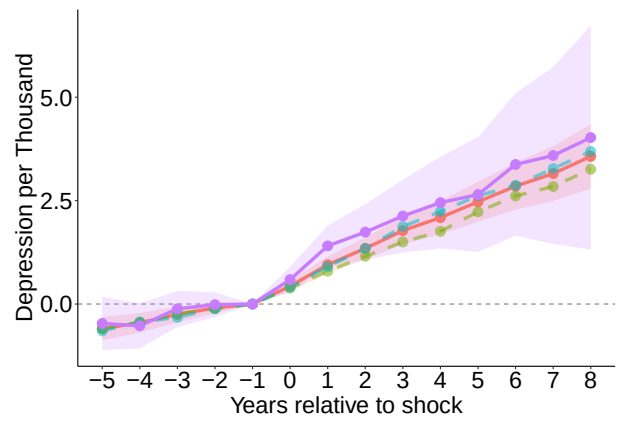
(a) Employment responses



(b) Mortality responses



(c) Medical Cost responses



(d) Depression Diagnosis responses

*Notes:* These plots report estimates separately by age of the adult son. The x-axis plots the years relative to the LTC-needer's shock. Estimates reported here correspond to the event study estimates,  $\delta_{\ell}$ , in Equation 3. 95 percent confidence intervals are plotted as the shaded regions for one subgroup (age 49 and older) and for the pooled estimates. Standard errors are clustered at the individual level.

Figure 4: Event study responses of adult sons to elderly parents' strokes: economic and health outcomes

## 5 Generalizing from Strokes

Strokes, while well-suited to a staggered event-study design, represent only a fraction of LTC-related diseases. Focusing on strokes alone risks providing an incomplete picture of families’ LTC burden, since strokes cause acute and unusually severe care needs relative to other conditions. This motivates us to ask what our design can reveal about other LTC-related shocks as well.

Other prevalent conditions, however, pose a challenge: they often involve a gradual loss of functional ability over time. Neurodegenerative diseases linked to cognitive decline, for example, account for a large share of existing care needs. If these declines materialize or can be anticipated before a formal diagnosis, then the design’s identifying assumptions may not hold. Conversely, if their costs materialize only long after diagnosis, estimates at the horizons our design can credibly support would capture little of them, and any later effects would be difficult to separate from differences between early- and late-diagnosed families.

Nevertheless, we repeat our main empirical exercises for two additional health shocks: hip fractures and dementia diagnoses. We choose these conditions because both are common, serious health problems whose consequences for care needs differ materially from those of strokes. As shown in Table 1, elderly parents who suffer a hip fracture tend to experience less pronounced increases in difficulty bathing, dressing, and eating. Their increases in difficulty with ambulation, walking, and incontinence, however, are more comparable to those of stroke survivors, though still smaller on average. A parent’s hip fracture might therefore be expected to act as a milder shock to adult children than a stroke. Dementia, by contrast, is plausibly associated with a different profile of care needs altogether.

The overall effects of a parent’s hip fracture or dementia diagnosis on adult sons are summarized in the second and third columns of Table 4, alongside our previously-discussed estimates for strokes. Comparing estimates across these events, we find that adverse effects of a parental hip fracture or dementia diagnosis on employment, mortality, and depression are remarkably similar to the adverse effects of a parental stroke. In Appendix Table A.2, we show no substantial systematic differences conditional on the age of the adult son as well. One interpretation of this finding is that the common requirement all three conditions impose on families—that a person be regularly available to support and monitor the LTC-needer—may be a key driver of negative health and family outcomes.

## 5.1 Extrapolating to Broader Functional Limitations

Strokes, hip fractures, and dementia are only three of many potential drivers of parental functional decline. This raises a more ambitious question: how does a decline in a parent’s functional ability *in general* affect adult sons?

While credibly answering this question is difficult, we can make an educated guess based on the responses we have observed to strokes, hip fractures, and dementia diagnoses. If functional limitations (and the care needs that accompany them) are the only channel through which these shocks affect adult sons, then IV allows us to recover the effect of a marginal decline in functional ability itself. Let  $F$  denote a parent’s functional ability,  $Z$  a health shock, and  $Y$  an adult son outcome. The IV estimate of the causal effect of functional limitations on adult sons’ outcomes is:

$$\hat{\beta}^{IV} = \frac{\text{Effect of shock on adult-son outcomes}}{\text{Effect of shock on functional limitations}} = \frac{\hat{\pi}_{YZ}}{\hat{\pi}_{FZ}}, \quad (4)$$

where  $\hat{\pi}_{YZ}$  is the reduced-form effect estimated by our event-study design and  $\hat{\pi}_{FZ}$  is the first-stage effect of the shock on functional ability. In addition to requiring the standard IV assumptions,  $\hat{\beta}^{IV}$  will give a reasonable estimate for the broader average effects of parental functional limitations on adult sons if 1) effects of parental functional ability on outcomes are approximately linear, and 2) the survey-based first stage estimate generalizes to our analytic sample in the administrative data.

Estimating  $\hat{\pi}_{FZ}$  from administrative data alone is not feasible, since individual functional-limitation records are not available at scale in Taiwan. The most readily available source is the TLISA, so we estimate the first stage using the difference-in-differences estimates from Table 1, taking the Barthel score as our measure of functional ability.

Appendix Table A.3 presents the IV estimates for effects of a marginal (1-point in the Barthel index) increase in a parent’s functional ability on adult sons’ outcomes. Panel A reports first stage effects directly from columns 2 and 4 of Table 1, which are approximately 20 points for strokes and 11 points for hip fractures. Panel B then divides the reduced-form estimates from Table 4 by these figures to yield our marginal effect estimates.

This exercise suggests that the costs of parental functional limitations are large for adult sons. Using strokes as the shifter (column 1), a 1 point increase in a parent’s Barthel score reduces annual mortality risks for adult sons by 0.03 per thousand or roughly 1 percent relative to male mortality rates at age 45 (3.37 deaths per thousand in 2020). It increases fertility by 0.2 children per thousand 5 years later, which is a 0.13 percent increase in births over that period. Person-years with a depression diagnosis fall by 0.145 per thousand or

roughly 1 percent of adult sons’ pre-stroke baseline. Annual public medical spending falls by 50 NTD per year, or roughly 0.4 percent of adult sons’ pre-stroke baseline. The decline in hospital bed days, 6 days per thousand per year, amounts to a 0.6 percent decrease in adult sons’ pre-stroke hospital time.

As a benchmark, Table 1 shows that stroke-suffering parents had already lost an average of 6 Barthel points *even before* their stroke occurred, while old-age respondents to the TLSA more generally have lost 10 points on average (Chan and Chen, 2026). If adult sons tend to have *one* parent with the functional ability of the average TLSA respondent, our estimates imply that functional limitations alone account for 0.28 deaths per thousand per year among adult sons. By comparison, age-specific mortality rates for men in their 40s range from 2 to 5 deaths per thousand. Effects on fertility (a decline of 2 children per thousand adult sons) are clearly smaller but not trivial considering the age distribution at which parental limitations occur.

These conclusions are not driven by our choice to focus on stroke events in particular. Using hip fractures instead for this exercise suggests *even larger* effects of functional limitations on adult children, because hip fractures have smaller impacts on functional abilities of parents but similar adverse consequences for adult sons.

## 6 The Expansion of Migrant Caregiving

A natural next question is whether Taiwan’s caregiving-related policies can attenuate these costs of parental health shocks. We exploit large increases over time in migrant caregiver hiring rates, concentrated among families whose parent suffered a health shock at a particularly advanced age. These are exactly the families targeted by the expansionary reforms described in Section 2. Among families that suffer a stroke, Figure 5 shows that post-stroke caregiver hiring rates for LTC-needers age 80 or older rose from 29 to 37 percent between 2007 and 2018.<sup>15</sup> By contrast, hiring rates for LTC-needers aged 50 to 79 have remained stable, just below 10 percent, over the same period.

It is revealing to compare these cross-cohort patterns in caregiver hiring rates to cross-cohort patterns in family responses to a stroke, among those families with LTC-needers over the age of 80. Figure 6 plots caregiver hiring rates in LTC-needing families against near-term estimates of the effect of LTC needs separately by year of the health shock. It shows

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<sup>15</sup>For these cohort-by-cohort effect estimates, we take a 20% random sample of adult sons from our main analytic dataset.

that adverse mortality and health effects of a stroke have attenuated markedly across shock cohorts, coincident with the rise in caregiver hiring rates. By contrast, Figure A.6 shows that the adverse effects for families of LTC-needers between ages 50 and 79 at the time of their stroke—who are not immediately eligible for more lenient hiring criteria under the reforms—have been stable or worsening alongside the slight differences they face in caregiver hiring rates across cohorts. This suggests that the reforms are improving post-stroke outcomes through the actual hiring of caregivers by LTC-needing families, not through changing their expectations over future caregiver hiring or some other indirect mechanism.

Because the expansions in caregiver hiring targeted older LTC-needers, they are less informative about how hiring affects adult sons’ marriage and fertility. As Table 4 shows, those effects are concentrated among younger sons, who tend to have younger parents. Still, because men in Taiwan form families relatively late, there remains some scope for the reforms to have an impact. Appendix Figure A.7 plots cohort-by-cohort near-term effects on marriage and fertility against caregiver hiring rates, as Figure 6 does for other outcomes. The adverse effects of a stroke on these outcomes in families of LTC-needers over the age of 80 have indeed weakened as hiring rates rose.

**The causal effect of caregiver hiring.** If the effects of a stroke on adult sons operate through the burden of care, then hiring a caregiver should attenuate them. To estimate how much, we treat cross-cohort variation in caregiver hiring rates, driven by the expansionary reforms, as a candidate instrument for caregiver hiring.<sup>16</sup> Fitting a line through the cohort-level scatterplots in Figure 6 yields a linear estimate of this attenuation: a 1 percentage point increase in the caregiver hiring rate raises employment by 1.17 workers per thousand per year (panel a), lowers mortality by 0.20 per thousand per year (panel b), reduces medical spending by 71 NTD per person per year (panel c), and reduces depression diagnoses by 0.33 per thousand (panel d). Appendix Figure A.7 implies corresponding increases of 0.2 per thousand in marriage rates and 0.14 per thousand in fertility.

However, caregiver hiring rates are likely not the only relevant factor which has trended over this time. For example, improvements in medical technology and standards of care might also alter the consequences of a parental stroke. One might therefore prefer to more

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<sup>16</sup>The exclusion restriction may fail as an instrument for *individual* hiring: under a more lenient regime, a forward-looking son may adjust his behavior even if his family has not yet hired a caregiver, because he expects that it can do so *in the future*. The IV estimate is then best read as an intent-to-treat effect: the effect on *adult sons as a whole* of living under a marginally more lenient hiring regime, as proxied by the aggregate hiring rate.

narrowly exploit variation in caregiver hiring rates immediately surrounding known expansionary reforms. For this reason, we compare the caregiver hiring rate and average near-term stroke responses across three cohorts preceding the 2012 reform (2007-2009) against the values for three cohorts following that reform (2013-2015). Because the 2015 reform follows so closely after the 2012 reform and targets a similar population, this contrast mixes the effects of both reforms. We exclude the 2010, 2011, and 2012 cohorts because they are partially treated by the 2012 reform. Across these two groups of cohorts, caregiver hiring rates rose by 4.45 percentage points (with a “first-stage” t-statistic of 5.56).

We superimpose these estimates in blue (for the pre-reform cohorts) and red (for the post-reform cohorts) over the scatter plots of Figure 6. They suggest similar—if anything more dramatic—attenuation of the adverse effects of stroke with greater access to migrant caregiver markets.<sup>17</sup> Labor supply responses reversed from a pre-reform decline of 10.56 workers per thousand per year to a slight *increase* of 1.74 workers per thousand per year among post-reform cohorts. Average effects of a stroke on mortality among the pre-reform cohorts amounted to 1.87 additional deaths per thousand per year, but only 0.17 additional deaths per thousand per year among post-reform cohorts. Depression diagnoses reversed across these cohorts as well, from an increase of 0.50 per thousand to a *decrease* of 2.21 per thousand. Effects on marriage and fertility also tend to weaken locally around the reform, in Appendix Figure A.7. On the other hand, medical spending responses are virtually unchanged across these cohorts.

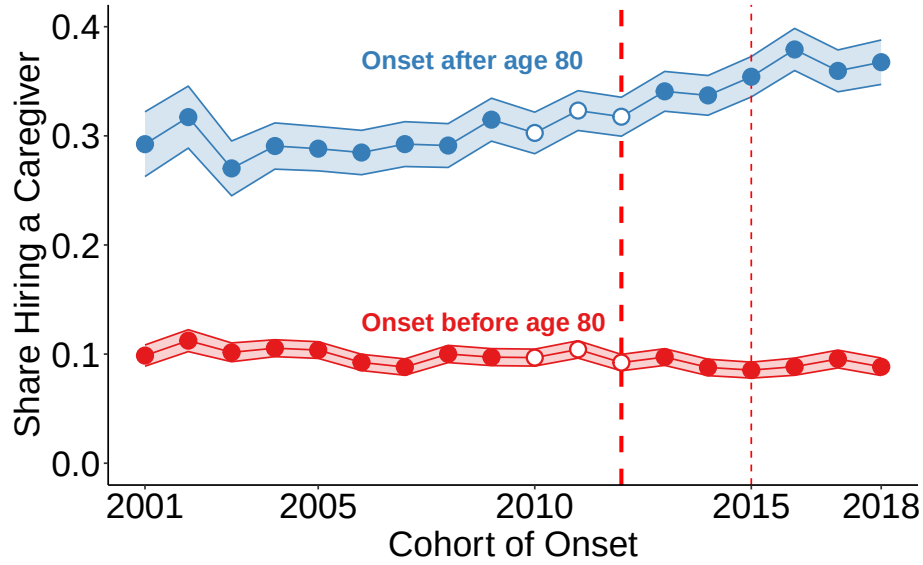
These estimates, though, capture only short-run effects. We can also estimate longer-run effects by comparing families that suffered a stroke in the early 2000s to those that suffered one after the reform; these longer-run effects are qualitatively similar to the short-run effects described above. We verify this in Appendix Figure A.8, comparing effect estimates up to five years after a stroke for pre-reform cohorts (2001–2006) against post-reform cohorts (2013–2015).

## 7 Assessing Willingness to Pay for Caregiver Hiring

For the families that choose to hire a migrant caregiver when given the opportunity to do so, revealed preferences would suggest that caregiver hiring is welfare-improving. In this

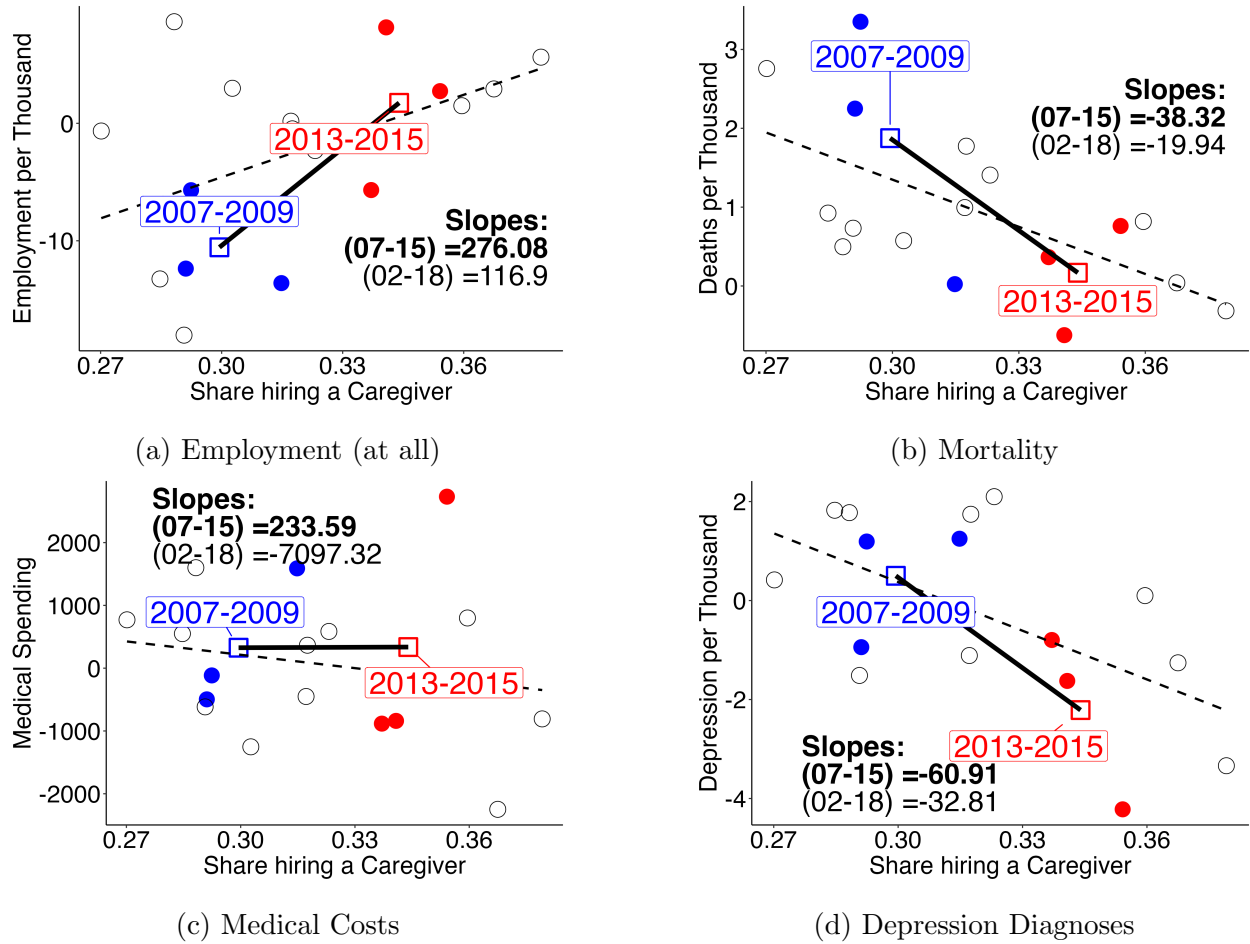
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<sup>17</sup>Recall one concern with these estimates is that our baseline sample is constructed using household registry data, which can only link sons to their parents if they survive to 2011. That may bias *downward* the mortality effects of a parental stroke for pre-reform cohorts, and therefore also the effects of the reform on mortality.



Notes: This figure plots, among adult sons with a stroke-suffering parent, the fraction with a migrant caregiver within two years of the stroke. The hollow dots denote three stroke cohorts (2010, 2011, and 2012) which were partially exposed to the 2012 reform. The vertical dashed lines denote the calendar years in which expansionary reforms occurred, which only targeted families with care-needers age 80 and older. Among families with onset after age 80, caregiver hiring shares averaged 0.344 over 2013-2015 and 0.299 over 2007-2009 (difference: 0.0445,  $t$ -statistic 5.56).

Figure 5: Share of adult sons with stroke-suffering parent under migrant care provision within two years



*Notes:* These graphs plot average effects of a stroke separately by the calendar year in which that stroke occurred, for families with an LTC-needer age 80 or older. Effects are averaged over the stroke year and two subsequent years. The line of best fit between all cohort-specific average effects is plotted in the thin dashed line, and its slope reported as (02-18). Red dots correspond to effects for 2013, 2014, and 2015 cohorts. Blue dots correspond to effects for 2007, 2008, and 2009 cohorts. Squares represent the averages within those groups of cohorts. The line of best fit locally around the expansionary reforms (comparing cohorts 2007-2009 versus 2013-2015) is plotted in solid black and its slope reported as (07-15). The solid points denote the set of cohort-specific estimates entering that local comparison. The sample is based on a 20% draw of sons to families where the health shock occurs at age 80 or older.

Figure 6: Cross-cohort correlations between stroke effects and caregiver hiring propensities among reform-targeted families

section, we quantify the magnitude of welfare gains for the families induced into caregiver hiring by the expansionary reforms, measuring them in terms of private willingness to pay (WTP).

## 7.1 Bounding Ex-Post Willingness to Pay

A natural upper bound on ex-post WTP comes from the cost savings that families enjoy from hiring migrant care services over otherwise-similar domestic ones, which we calculated in Section 2.1 to be roughly 13,000 USD (400,000 NTD) per year. This is an upper bound because many families in practice substitute away from in-home formal care provision when hiring a migrant worker is not possible. It is an *ex-post* measure of WTP because families only pay for a caregiver if and when care needs actually arise.

But ex-post WTP for a migrant caregiver may be substantially lower than ex-ante WTP, owing to the insurance value that households derive from policies that provide support in the event of old-age care needs. This is because LTC insurance markets are plausibly incomplete, meaning families may fail to perfectly smooth marginal utilities across states of the world.

## 7.2 A Framework for Ex-Ante Willingness to Pay

Estimating ex-ante WTP conventionally requires assumed or estimated risk preferences, to recover how the marginal utility of income differs across states of the world. This is true even of sufficient-statistics approaches, which first-order approximate the value of a marginal transfer across states but still require risk preferences to recover the ratio of marginal utilities between them (Hendren, 2021). Our approach instead combines our quasi-experimental estimates with existing valuations of mortality risk and health. It does not require taking a stand on risk preferences.

We sidestep risk preferences by valuing caregiver hiring through its causal effects: on survival and quality of life, which enter utility, and income, which enters the budget. The cost is that we must impose some structure on *how* survival and quality of life enter utility.

Consider an adult son who lives for up to  $L$  years, and let  $W$  denote his ex-ante welfare: his expected payoff over states of parental health and caregiver hiring. We partition these states into two kinds. In the first, a parental shock occurs when the son is age  $a$  and the reform induces the family to hire a caregiver. This happens with probability  $\pi_a$  and yields payoff  $W(a)$ , so it occurs at *some* age with total probability  $\Pi \equiv \sum_a \pi_a$ . In the second, either no parental shock occurs, or one occurs but the family's hiring decision does not depend on

the reform. This happens with probability  $1 - \Pi$  and yields expected payoff  $W_0$ . Ex-ante welfare is then

$$W \equiv (1 - \Pi)W_0 + \sum_a \pi_a W(a)$$

The state-contingent payoff  $W(a)$  consists of a pre-shock payoff  $V^{\text{pre}}(a)$  and post-shock payoff  $V(a)$ :

$$W(a) \equiv V^{\text{pre}}(a) + V(a)$$

We specify nothing else about the pre-shock payoff or pre-shock optimization problem, except that migrant caregiver hiring does not affect them (and that it does not affect the state probabilities,  $\pi_a$  either).<sup>18</sup> This rules out certain dynamic optimization behaviors, for instance that introducing caregiver hiring could crowd out precautionary savings.

We suppose that caregiver hiring can affect three inputs in each year from the age of the shock onward, consistent with our conceptual framework in Section 2.2: survival ( $\lambda_t(a)$ ) and quality of life ( $H_t(a)$ ) which we view as intermediary inputs derived from home production, and disposable income ( $y_t(a)$ ), as faced by a son in year  $t$  having realized an age- $a$  parental shock. Because hiring affects neither pre-shock payoffs nor the states in which the reform does not induce it, it affects welfare only through the post-shock payoffs  $V(a)$ .

**The problem of measuring ex-ante WTP for migrant caregiver hiring.** Converting that effect into a monetary value, however, is difficult. While our reform-based quasi-experimental evidence can reveal how caregiver hiring affects the inputs to welfare, that evidence does not reveal how adult sons value those changes. Estimating a first-order approximation to ex-ante WTP would require knowing the *post-shock* values of survival, quality of life, and income relative to the *ex-ante* marginal utility of income. If sons are not well insured against these shocks (as suggested by the discussion of Section 2.1), these post-shock values can differ substantially from their ex-ante counterparts and can vary with the age at which the shock occurs. No credible estimates of these state-contingent values exist; standard valuations of life and health are not conditioned on such shocks.

**Deriving an empirically tractable measure of WTP under income insurance.** Instead, we make headway by considering the ex-ante WTP for a hypothetical reform in which

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<sup>18</sup>The assumption that caregiver hiring does not change pre-shock circumstances is an innocuous one if caregiver hiring comes from an unannounced policy change. If it is anticipated, abstracting from pre-stroke responses should cause us to understate WTP if it enables the crowd-out of anticipatory savings and other costly self-insurance behaviors.

post-shock adult sons induced by the reforms to hire migrant caregivers *also* receive insurance against effects of the shock on income that operate through margins other than caregiver hiring. Measuring WTP in units of the insured marginal utility of income ( $\mu^*$ , which is not state-contingent), the ex-ante WTP for that hypothetical reform can be decomposed into two parts. The first is the partial WTP for income insurance fixing caregiving arrangements, and the second is the partial WTP for caregiver hiring given income insurance:

$$\underbrace{\frac{W^{\text{insurance+hiring}} - W^{\text{baseline}}}{\mu^*}}_{\text{Full ex-ante WTP}} = \underbrace{\frac{W^{\text{insurance}} - W^{\text{baseline}}}{\mu^*}}_{\text{PWTP for Income Insurance}} + \underbrace{\frac{W^{\text{insurance+hiring}} - W^{\text{insurance}}}{\mu^*}}_{\text{PWTP for Hiring } (\Delta W/\mu^*)}$$

$W^{\text{insurance}}$  denotes ex-ante welfare in an intermediate policy regime where the adult son receives income insurance, but caregiving is fixed at its baseline arrangement (where  $W^{\text{baseline}}$  is the ex-ante welfare under that baseline regime). The complete reform with both income insurance and migrant caregiver hiring delivers ex-ante welfare denoted  $W^{\text{insurance+hiring}}$ . For brevity, we proceed by using stars to denote objects evaluated under income insurance: the son's post-shock payoff  $V^*(a)$  and its derivatives, his survival  $\lambda_t^*(a)$  and quality of life  $H_t^*(a)$ , his consumption  $c_t^*(a)$ , his disposable income  $y_t^*(a)$  and assets at the shock  $A^*(a)$ , and his marginal utility of income  $\mu^*$ . We use the operator  $\Delta$  to denote the causal effect of migrant caregiver hiring under income insurance.

Supposing adult sons are risk averse, the partial WTP (PWTP) for hiring is useful to measure because it is a lower bound on the ex-ante WTP for the full hypothetical reform. A first-order approximation to the PWTP for hiring gives:

$$\begin{aligned} \underbrace{\frac{\Delta W}{\mu^*}}_{\text{PWTP for hiring}} &= \sum_a \pi_a \frac{\Delta V(a)}{\mu^*} \\ &\approx \sum_a \pi_a \left[ \underbrace{\sum_{t=a}^L \frac{V_{\lambda_t}^*(a)}{\mu^*} \times \Delta \lambda_t(a)}_{\text{Survival effect}} + \underbrace{\sum_{t=a}^L \frac{V_{H_t}^*(a)}{\mu^*} \times \Delta H_t(a)}_{\text{Quality of Life effect}} \right. \\ &\quad \left. + \underbrace{\sum_{t=a}^L \frac{V_{y_t}^*(a)}{\mu^*} \times \Delta y_t(a)}_{\text{Income effect}} \right]. \end{aligned} \tag{5}$$

In this expression,  $\Delta \lambda_t(a)$ ,  $\Delta H_t(a)$ , and  $\Delta y_t(a)$  can be informed by our quasi-experimental evidence under the assumption that the effects of caregiver hiring on these inputs (and *who* hires a caregiver) are unchanged by income insurance. Thanks to income insurance, the

state-contingent value of a dollar of year- $t$  disposable income ( $\frac{V_{y_t}^*(a)}{\mu^*}$ ) is simply a discounted survival weight. The existence of income insurance, combined with some additional structure that we impose next on the post-shock problem, will also allow us to identify the value of life ( $V_{\lambda_t}^*(a)$ ) and the value of quality of life ( $V_{H_t}^*(a)$ ) simply from evidence on the value of a statistical life (VSL).

**The post-shock household problem.** For the post-shock optimization problem, we adapt a flexible framework for valuing longevity and quality of life from Murphy and Topel (2006). In that framework, the post-shock payoff to an adult son under income insurance,  $V^*(a)$ , is the result of choosing consumption to maximize the present discounted value of utility flows over the remaining possible years of life:

$$\begin{aligned} V^*(a) \equiv \max_c U(a, c) &= \sum_{t=a}^L H_t^*(a) u(c_t(a)) \frac{1}{(1+\rho)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right] \\ \text{s.t. } \frac{A^*(a)}{(1+r)^a} + \sum_{t=a}^L [y_t^*(a) - c_t(a)] \frac{1}{(1+r)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right] &= 0, \end{aligned} \quad (6)$$

meaning that period utility flows are a multiplicative function of quality of life ( $H_t^*(a)$ ) and a discounted consumption payoff ( $u(c_t(a)) \frac{1}{(1+\rho)^t}$ ), and a survival probability weight ( $\prod_{s=a}^t \lambda_s^*(a)$ ). The son chooses the consumption path to maximize  $U$  subject to a budget constraint, which balances consumption against income net of taxes and fixed expenses (like caregiver hiring) denoted  $y_t^*(a)$ , and initial (age- $a$ ) assets  $A^*(a)$ . We can also allow adult sons to control  $\lambda_t^*(a)$  and  $H_t^*(a)$  through a choice of home production as in Section 2.2, but this is not necessary for the key results from the framework.<sup>19</sup> Note that the budget constraint is deterministic, as implied by assuming a perfect post-shock annuity market. We value all flows in birth-date terms, discounting year- $t$  amounts by  $(1+r)^t$  (and age- $a$  assets accordingly), so that marginal utilities of income are comparable across shock ages. The multiplier on this budget constraint is the marginal utility of income  $\mu^*$ , which is common across post-shock periods because sons face a single deterministic budget constraint, and across shock ages thanks to the hypothetical insurance. The first order condition with respect to consumption implies:

$$\mu^* = H_t^*(a) u_c(c_t^*(a)) \frac{(1+r)^t}{(1+\rho)^t} \quad \text{for all } t \in \{a, a+1, \dots, L\}.$$

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<sup>19</sup>By the envelope theorem, re-optimized investments have no first-order welfare effect, so  $\Delta \lambda_t(a)$  and  $\Delta H_t(a)$  should be read as the direct effects of hiring. Our estimates will overstate them to the extent that responses to reforms also include induced investments whose costs do not appear in  $\Delta y_t(a)$ .

Differentiating the Lagrangian implied by equation 6 with respect to  $y_t^*(a)$  provides the expression for the discounted marginal utility of income:

$$V_{y_t}^*(a) = \frac{\mu^*}{(1+r)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right] \quad (7)$$

**Linking the value of life and quality of life.** A key advantage of this framework is that it links the value of life,  $\frac{V_{\lambda_t}^*(a)}{\mu^*}$ , and the value of proportional changes in quality of life,  $\frac{V_{H_t}^*(a) \times H_t^*(a)}{\mu^*}$ . Assuming the year- $a$  annual survival rate is close to 1 and that extending life does not affect one's budget constraint (because individuals consume their annual income), then we show in Appendix F that this relationship takes on a simple form:

$$\frac{V_{\lambda_a}^*(a)}{\mu^*} \approx \sum_{t=a}^L \frac{V_{H_t}^*(a) \times H_t^*(a)}{\mu^*} = \sum_{t=a}^L \underbrace{\frac{u(c_t^*(a))}{u_c(c_t^*(a))}}_{\text{value of a life-year}} \times \underbrace{\frac{1}{(1+r)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right]}_{\text{discounted survival weight}} \quad (8)$$

This means that the value of a life from year  $a$  equals the value of a permanent proportional increase in quality of life for all years from  $a$  to  $L$ . Both equal the present-discounted value of the utility flows one can expect to enjoy having survived year  $a$ .

Equation 8 would give a straightforward way to express the value of age- $a$  life in terms of the value of quality of life over years  $t \geq a$ . But to go the opposite direction and express the value of quality of life in terms of the value of life, we need two assumptions: that adult sons' post-stroke quality of life is stable over time (which is supported by the data) and that they discount the future at the interest rate ( $\rho = r$ ). These assumptions imply that  $c_t^*(a)$  is constant over post-shock years, and therefore so is  $\frac{u(c_t^*(a))}{u_c(c_t^*(a))}$ . As a result, Equation 8 implies the value of proportional quality-of-life improvements is common across post-shock periods up to a discounted survival weight,  $d$ :

$$\frac{V_{H_t}^*(a) \times H_t^*(a)}{\mu^*} = \frac{V_{H_a}^*(a) \times H_a^*(a)}{\mu^*} \frac{d_t(a)}{d_a(a)}, \quad \text{where} \quad d_t(a) \equiv \frac{1}{(1+r)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right]$$

which means that we can back out the (unweighted) value of quality-of-life improvements in any one year from the value of a life:

$$\frac{V_{H_t}^*(a) \times H_t^*(a)}{\mu^*} = \frac{d_t(a)}{\alpha_a(a)} \times \frac{V_{\lambda_a}^*(a)}{\mu^*}, \quad \text{where} \quad \alpha_t(a) \equiv \sum_{s=t}^L d_s(a) \quad (9)$$

$\alpha_t(a)$  is the annuity factor describing the present-value expected years of life saved when one survives year  $t$ . Applying the same logic to survival gains accruing in later years (using  $\lambda_t^*(a) \approx 1$ ) gives  $\frac{V_{\lambda_t^*}^*(a)}{\mu^*} = \frac{\alpha_t(a)}{\alpha_a(a)} \times \frac{V_{\lambda_a^*}^*(a)}{\mu^*}$ , so the value of both survival and quality of life in *every* post-shock year is pinned down by the single year- $a$  value of life.

**PWTP for hiring among compliers.** We can plug in the expression for the discounted marginal utility of income and the value of quality of life (i.e., plug Equations 7 and 9 into Equation 5). We also normalize by  $\Pi$ , so that our target parameter—which we call *PWTP for hiring per complier*—is WTP *conditional* on a parental shock occurring in conditions where the reform induces caregiver hiring. This provides:

$$\underbrace{\frac{1}{\Pi} \frac{\Delta W}{\mu^*}}_{\text{PWTP for Hiring per complier}} \approx \sum_a \frac{\pi_a}{\Pi} \left[ \underbrace{\frac{V_{\lambda_a^*}^*(a)}{\mu^*} \left( \sum_{t=a}^L \frac{\alpha_t(a)}{\alpha_a(a)} \Delta \lambda_t(a) + \sum_{t=a}^L \frac{d_t(a)}{\alpha_a(a)} \frac{\Delta H_t(a)}{H_t^*(a)} \right)}_{\text{value of survival-equivalent gains}} + \underbrace{\sum_{t=a}^L d_t(a) \times \Delta y_t(a)}_{\text{value of disposable income effects}} \right]. \quad (10)$$

The weights  $\pi_a/\Pi$  sum to one, so this is the gain from caregiver hiring under income insurance among sons whose families are induced to hire by the reform, averaged over the ages at which their parental shocks occur. This is the population to which our quasi-experimental (complier) estimates apply. Unconditional PWTP for hiring (Equation 5) is simply  $\Pi$  times this quantity.

Taking Equation 10 to the data is straightforward. Aside from the interest rate (which enters  $d_t$  and  $\alpha_t$ ), the only key ingredient that cannot come directly from our data is the value of a life under income insurance,  $\frac{V_{\lambda_a^*}^*(a)}{\mu^*}$ . Because the specific circumstances affected by the reform are relatively rare, insurance against income losses in those circumstances should have little effect on the ex-ante value of life. Moreover, a parental shock has—in absolute terms—only a small effect on survival and quality of life. Post-shock survival and quality of life under income insurance ( $\lambda_t^*(a)$  and  $H_t^*(a)$ ) should therefore be close to their no-shock levels, meaning parental shocks should not meaningfully reduce the value of a life through those margins. Together, these imply that we can approximate the conditional value of life under income insurance using a status quo estimate for the value of statistical life (VSL), where we view the VSL as a measure of the ex-ante value of life.<sup>20</sup>

<sup>20</sup>Equation 10 values flows in birth-date terms; we report it in shock-date dollars by multiplying through by  $(1+r)^a$ , so that the VSL enters directly and income effects are discounted from the shock date.

## 8 Estimation of Willingness to Pay

**Deriving estimated effects on survival, quality of life, and disposable income.** Starting from our IV results (the “2007–2015” slope estimates in Figure 6), we extract estimates for the effect of hiring on survival, quality of life, and income. Our mortality effect estimates naturally provide estimates of survival effects. For income, we accumulate effects on employment and net out the costs of caregiver hiring laid out in Section 2.1 (\$32,258 over a three-year contract).<sup>21</sup> As in Section 2.1, we abstract from the costs of accommodation for the caregiver. We assume all reform-induced migrant-hiring families are substituting from informal care provision instead of nursing home care. If some substitute instead from nursing home care, then that will overstate the true cost of migrant caregiver hiring, which would be offset by savings from forgone nursing home care. In practice, the cost of nursing home care is similar to—if not higher than—the cost of migrant caregiver hiring.

Measuring quality of life changes requires more work, because quality of life is not directly observed. We use changes in depression diagnoses as our main proxy, and take the average disability weight of 0.23 from the Global Burden of Disease 2010 to mean that depression proportionally reduces  $H$  by 23 percent (Ferrari et al., 2013).<sup>22</sup> Since we only measure causal effects on depression *diagnoses*, we assume a person diagnosed with depression suffers it for at least the first three years. We report the annual effect estimates for our three key inputs to welfare in Table 6. Column (1) reports the reduced-form effect of reform on the inputs, and Column (2) the 2SLS estimate for the effect on outcomes per complier induced into caregiver hiring by the reform.

**Estimated values of life and quality of life.** We derive an estimate for the value of a life-year using Equation 9. For the following calculations, we assume all sons will live 40 years after the parental health shock with an interest rate ( $r$ ) of 0.03. That means the value of a life and the undiscounted value of quality of life are linked by an annuity factor of  $\alpha(a) = 23$ .

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<sup>21</sup>Recall these calculations are for caregiver hiring in the 2020’s. Wage costs to migrant care workers were roughly 15 percent lower before 2022, meaning the true cost that families paid for a three-year migrant caregiver contract in the reform period was substantially *lower* than \$32,258.

<sup>22</sup>Reductions in medical spending would serve as a more general proxy for good health than depression, but we have no natural way to understand how declines in medical spending relate to a proportional change in health. Conveniently, effects of a stroke on medical spending are virtually flat around the reform. So if one prefers medical spending as a measure of quality of life, one would simply conclude that caregiver hiring has no effect.

We adapt an estimate for the value of a life from a classic hedonic wage regression implemented by Liu et al. (1997) in Taiwan. That study inferred a value of a statistical life (VSL) from mortality-related wage premia of 413,000 USD in the mid-1980s. To account for the high economic growth that Taiwan experienced since then, we borrow the method used in national VSL calculations recommended by the OECD: linearly adjusting a baseline VSL estimate for differences in GDP per capita (i.e., assuming an elasticity with respect to economic output of 1; OECD, 2025a). Taiwan’s per capita GDP grew by a factor of 7.84 in USD terms from 1985 to 2019, implying a 2019 VSL of 3.2 million USD, or 100 million NTD, and a value of a life year of 140,000 USD (4.34 million NTD).<sup>23</sup>

**Estimated PWTP for hiring.** In light of the long-run variation in the value of life—and our limited ability to estimate long-run effects of caregiver hiring—we accumulate costs and benefits only over *the first three years* after the health shock, truncating the sums over  $t$  in Equation 10 at  $t = a + 2$ , which corresponds to the duration of a standard migrant caregiver contract. Over such a short horizon, survival rates are close to one and differences in remaining life expectancy across  $t \in \{a, a + 1, a + 2\}$  are small relative to remaining life:  $\alpha_t(a)/\alpha_a(a) = 1 - \sum_{s=a}^{t-1} d_s(a)/\alpha_a(a) \approx 1$ , meaning a single value of life can approximate survival gains in each of these years. For the same reason, and abstracting from discounting within the horizon,  $d_t(a) \approx 1$  for  $t \in \{a, a + 1, a + 2\}$ . Equation 10 then simplifies to

$$\underbrace{\frac{1}{\Pi} \frac{\Delta W}{\mu^*}}_{\text{PWTP for Hiring per complier}} \approx \sum_a \frac{\pi_a}{\Pi} \left[ \frac{V_{\lambda_a}^*(a)}{\mu^*} \left( \sum_{t=a}^{a+2} \Delta \lambda_t(a) + \frac{1}{\alpha_a(a)} \sum_{t=a}^{a+2} \frac{\Delta H_t(a)}{H_t^*(a)} \right) + \sum_{t=a}^{a+2} \Delta y_t(a) \right].$$

Because three-year survival is concave in the annual mortality reductions, summing the  $\Delta \lambda_t(a)$  slightly overstates the exact change in survival over the contract, so the survival gains we report are a modest upper bound.

We present our calculations for the WTP for caregiver hiring in the remaining rows of Table 6. Our calculations suggest the private welfare effects from hiring are dominated by improvements in survival. For each caregiver hired as a result of the reforms, annual survival rose by 38.3 adult sons per thousand. Accumulating these survival gains over the three years of the hiring contract, our VSL estimate implies a private gain from increased survival alone

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<sup>23</sup>By comparison, the OECD-recommended VSL estimate for Japan and South Korea, based on a meta-analysis of VSL estimates for OECD countries, is 6 million in 2022 USD (OECD, 2025b). VSL benchmarks actually used by the U.S. Department of Transportation (U.S. Department of Transportation, 2021) and the Environmental Protection Agency (U.S. Environmental Protection Agency, 2024) are over 10 million USD.

worth 372,261 USD over that period—an order of magnitude larger than caregiver hiring costs.

Entry into the labor force also partially offsets caregiver hiring costs. For every caregiver hired, 0.276 sons re-enter employment. Assuming average pre-stroke *nominal* earnings conditional on working (400,000 NTD or roughly 13,000 USD, taken from employment and earnings entries of Table 3) are representative of counterfactual earnings for these sons, then 3,600 USD is gained from earnings per caregiver per year which implies 10,800 USD in gains over the hiring contract. These earnings gains would offset one third of the costs of caregiver hiring.

Quality of life effects, measured by changes in depression diagnoses, offset a more modest share of caregiver hiring costs. For each caregiver hired, effects on depression fall by 61 per thousand (panel d of Figure 6). With depression reducing quality of life by 23 percent, this amounts to quality-of-life gains worth 14 life years per thousand. Accumulating this over the three-year hiring interval and multiplying by the value of a life year yields 5,917 USD in gains. These gains offset 18 percent of the costs of caregiver hiring.

Note we make the key assumption that survival gains are “persistent” in the sense that caregiver hiring does not change mortality rates after age  $a + 2$ . To the extent that mortality is driven by acute causes like suicide and accidents, it seems defensible to assume that short-run mortality decreases are not offset by later increases because those causes tend to kill otherwise-healthy people. Truncating at  $a + 2$  arguably understates the total gain from caregiver hiring; recall that Appendix Figure A.8 suggests reform-driven caregiver hiring continues to shut down large adverse effects on mortality and quality of life in later years as well.

**Social costs of caregiving-related policies.** Relating the private welfare gains from reform-induced migrant caregiver hiring to social welfare gains requires accounting for positive and negative spillovers from caregiver hiring. Figure 6 presented evidence of some small *positive* spillovers: higher tax revenue due to higher employment.

Negative externalities, on the other hand, are not obvious. Unlike other forms of immigration, migrant caregivers should not cause direct competition in the labor market with native Taiwanese because in-home caregiving is predominantly done informally by relatives; the costs and benefits of crowding out informal care should be internalized by the hiring family. While substitution from nursing homes to live-in care is one way that migrant caregiver reforms may indirectly impact native-born workers, other research exploiting expansionary reforms in Taiwan has found very few complier families rely on nursing home care in the

Table 6: PWTP for reform-induced migrant caregiver hiring among adult sons with a stroke-suffering parent

|                                             | (1)<br>Annual Effect<br>( $\Delta$ ) | (2)<br>$\Delta$ per<br>Complier | (3)<br>Outcome<br>Value (\$)                | (4)<br>PWTP for Hiring<br>per Complier (\$) |
|---------------------------------------------|--------------------------------------|---------------------------------|---------------------------------------------|---------------------------------------------|
| <b>Disposable Income (<math>y</math>)</b>   |                                      |                                 |                                             | <b>-21,458</b>                              |
| Caregiver Hiring                            | 0.0445                               | 1                               | $\times -10,753 \times 3 \text{ years} =$   | -32,258                                     |
| Work                                        | 0.0123                               | 0.276                           | $\times 13,040 \times 3 \text{ years} =$    | 10,800                                      |
| <b>Survival Rate (<math>\lambda</math>)</b> | 0.00171                              | 0.0383                          | $\times 3,237,920 \times 3 \text{ years} =$ | <b>372,261</b>                              |
| <b>Quality of Life (<math>H</math>)</b>     | 0.000624                             | 0.014                           | $\times 140,779 \times 3 \text{ years} =$   | <b>5,917</b>                                |
| <b>Total Private Gain</b>                   |                                      |                                 |                                             | <b>356,720</b>                              |

*Notes:* Column (1) reports the “reduced form” effect of the 2012 and 2015 reforms on outcomes, comparing effects of a stroke for the 2013-2015 stroke cohorts to effects for the 2007-2009 cohorts (the y-axis difference between the red and blue squares in Figure 6). For “Survival Rate,” we report the negative of the effect for mortality. For “Quality of Life,” we report the negative of the effect on depression diagnoses times 0.23 (the disability weight for depression; Ferrari et al., 2013). Column (2) scales this by the average change in caregiver hiring rates across these two cohorts (4.45 percentage points), and corresponds to the “(07-15)” slope reported in Figure 6. Column (3) reports estimates for the dollar value of changes in each outcome in 2019 USD. For caregiver hiring, this is the cost reported in Section 2.1, and excludes food and accommodation. For work, this is average pre-stroke earnings conditional on doing any work (400k NTD converted to 13k USD, calculated from Table 3). Column (4) converts the outcome effects into PWTP for hiring per complier by multiplying (2) and (3). For the survival rate, this is the value of a life. For the quality of life, this is the value of a life-year. Both are based on the VSL estimate of Liu et al. (1997) adjusted for per-capita GDP growth.

absence of migrant care hiring (Chan and Chen, 2026). In a 2019 survey, 23 percent of families that hired a migrant caregiver reported a nursing home or long-term care facility as their next best alternative mode of care provision, and another 13 percent said they would hire a domestic care worker (Ministry of Labor, 2020, Table 7). That caregiver hiring causes roughly 1 in 4 adult sons to re-enter the labor force is another indirect channel for spillovers into the labor market equilibrium, but that should be second order to the social welfare gains from the re-entry itself.

The direction of Taiwanese policy is also consistent with the idea that social returns to the liberalization of migrant caregiver hiring were high. That the 2012 reform was followed so soon by the 2015 reform is consistent with the idea that policymakers understood the gains from the 2012 reform. Even after that, further liberalization occurred more recently in 2024. To the extent that liberalization in 2024 faced political opposition, that opposition was centered primarily around concerns that further expansions can have distributional consequences by causing supply shortages or increasing the equilibrium cost of caregiver hiring.

**Ex-post gains from survival and quality of life.** Another object our framework allows us to straightforwardly infer is a bound for the *ex-post* welfare gains from survival and quality of life. We report in the first row of Table 6 that the net effects of caregiver hiring on disposable income (offsetting the calculated cost of hiring by increases in earnings) amounts to a cost of 21,458 USD. Adding that to the ex-post WTP of 39,000 USD (13,000 USD per year, from Section 7.1) implies that the ex-post gains from survival and quality of life over a three-year contract must be worth no more than 60,458 USD, or less than one-sixth the value that they contribute to the PWTP for hiring. That substantial gap suggests adult sons are poorly insured against parental health shocks.

## 9 Conclusion

With major demographic changes looming over the horizon for many of the world’s rich countries, the consequences of aging-related health shocks for families of the elderly are increasingly important yet so far poorly understood. In this paper, we contribute to that understanding by examining mortality, well-being, household formation, and labor market responses of adult sons to the old-age health shocks of their parents. We examine these responses in the setting of Taiwan, a rich East Asian society with a limited social safety net,

at the frontier of the demographic transition.

We find that parental health shocks are generally disruptive for the life cycles of adult sons. In the wake of a parental stroke, adult sons face 19 percent higher mortality rates, as well as increases in depression and medical spending, with mortality effects concentrated among sons who are older at the time the stroke occurs. Earlier in the life cycle though, adult sons also face declines in fertility and marriage. Turning towards potential mechanisms underlying these responses, we show that heterogeneity along observable characteristics suggests that the burden of care provision plays an important role—at least for the stroke events that we focus on for this exercise. We also show that adult sons exhibit similar responses to parental hip fractures and even dementia diagnoses. Extrapolating from the effects of a stroke suggests that the total burden of parental functional limitations on adult sons is substantial, especially in terms of excess mortality; it can explain 0.28 deaths per thousand per year among adult sons, relative to annual age-specific mortality rates of 2 to 5 deaths per thousand for men in their 40s.

Finally, we consider whether formal caregiver hiring may help attenuate some of these adverse responses, leveraging variation from national reforms which relaxed eligibility criteria for accessing migrant caregiver markets. Trends over time in how families respond to health shocks suggest that recent increased access to migrant caregiver markets has attenuated the adverse effects of a parental stroke. Using the reform to estimate causal effects of caregiver hiring on the consequences of a stroke, we find that for every thousand migrant caregivers hired, 276 more adult sons work each year, 38 fewer adult sons die each year, and 61 fewer sons are diagnosed with depression. Taking these effect estimates to a simple model of health, longevity, and consumption suggests that the private welfare gains from reform-induced caregiver hiring could be substantial. A VSL estimate points to a WTP for migrant caregiver hiring worth 356,720 USD in net gains per three-year hiring contract—so long as families were first fully insured against the non-hiring-related effects of parental health shocks on their disposable income.

Our results are qualified by two caveats. First, our administrative records are best suited to studying sons, so the consequences for daughters, who in many other societies bear much of the burden of care, remain an open question. Second, the national reforms to caregiver markets which we study only apply to care-needers over the age of 80, so they say little about how caregiver hiring affects adult sons at the ages more critical for family formation. Still, our results suggest that aging-related health conditions have serious consequences for the families of the elderly, but that public policy which reduces the burden of long-term care

can alleviate them.

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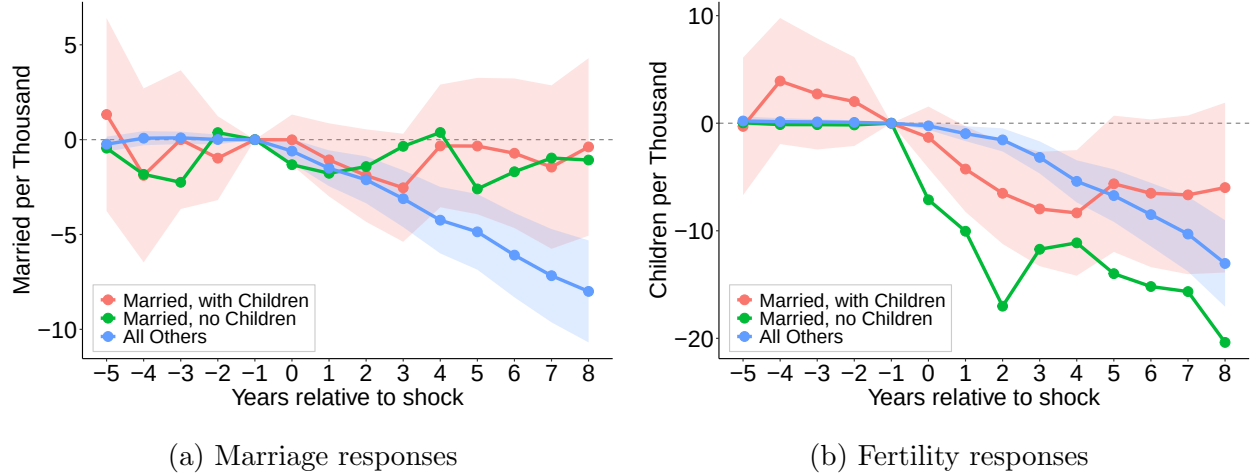
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*Notes:* The x-axis plots the years relative to the LTC-needer's shock. Estimates reported here correspond to the event study estimates,  $\delta_\ell$ , in Equation 3. To construct these estimates, we compare stroke-suffering adult sons to later-treated individuals with the same marital status and number of children in the calendar year before the treated son's stroke, in addition to matching on the observables described in Section 4. These estimates restrict to adult sons younger than age 34 before the stroke. Shaded regions denote 95 percent confidence intervals. We suppress confidence intervals for the "Married, no Children" subgroup for legibility; all point estimates are marginally insignificant for this group in panel (a), and marginally significant only in years 0, 1, 2, and 8 in panel (b). Standard errors are clustered at the individual level.

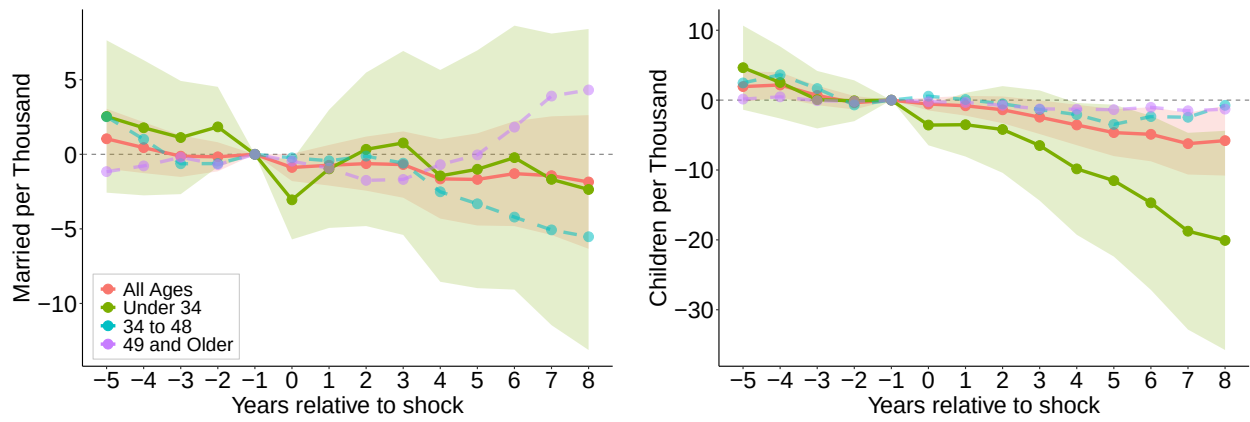
Figure A.1: Family formation responses of younger adult sons, conditional on pre-stroke family structure

## Appendix A Additional Figures and Tables

Table A.1: Economic and Health Effects on Adult Sons of Elderly Parents' Strokes, Using Only Health Insurance Links

|                                                 | (1)            | (2)     | (3)      | (4)     | (5)                          | (6)     | (7)      | (8)     |
|-------------------------------------------------|----------------|---------|----------|---------|------------------------------|---------|----------|---------|
|                                                 | Main Estimates |         |          |         | Not using Household Registry |         |          |         |
|                                                 | All Ages       | ≤ 33    | 34 to 48 | ≥ 49    | All Ages                     | ≤ 33    | 34 to 48 | ≥ 49    |
| <hr/> Health <hr/>                              |                |         |          |         |                              |         |          |         |
| Deaths per 1k                                   | 0.56           | 0.13    | 0.53     | 1.74    | 0.72                         | 0.14    | 0.83     | 2.05    |
|                                                 | (0.06)         | (0.03)  | (0.07)   | (0.24)  | (0.06)                       | (0.04)  | (0.09)   | (0.29)  |
| Suicides per 1k                                 | 0.05           | 0.03    | 0.05     | 0.12    | 0.06                         | 0.03    | 0.07     | 0.13    |
|                                                 | (0.01)         | (0.01)  | (0.01)   | (0.02)  | (0.01)                       | (0.01)  | (0.02)   | (0.02)  |
| Depression per 1k                               | 2.90           | 2.92    | 2.99     | 2.56    | 2.79                         | 2.80    | 2.78     | 2.80    |
|                                                 | (0.31)         | (0.36)  | (0.42)   | (1.04)  | (0.39)                       | (0.43)  | (0.61)   | (1.47)  |
| Medical Spending                                | 1000           | 755     | 927      | 1823    | 1211                         | 661     | 1278     | 2553    |
|                                                 | (166)          | (202)   | (213)    | (476)   | (178)                        | (203)   | (278)    | (664)   |
| Hospital Days                                   | 0.13           | 0.12    | 0.14     | 0.11    | 0.08                         | 0.08    | 0.10     | 0.03    |
|                                                 | (0.02)         | (0.02)  | (0.03)   | (0.05)  | (0.03)                       | (0.03)  | (0.05)   | (0.10)  |
| <hr/> Employment <hr/>                          |                |         |          |         |                              |         |          |         |
| Worked (at all) per 1k                          | -7.5           | -4.3    | -6.3     | -19.1   | -10.7                        | -6.2    | -10.7    | -23.0   |
|                                                 | (0.7)          | (0.9)   | (0.8)    | (2.1)   | (0.8)                        | (1.0)   | (1.1)    | (3.3)   |
| Months Worked                                   | -0.10          | -0.06   | -0.09    | -0.23   | -0.14                        | -0.08   | -0.14    | -0.29   |
|                                                 | (0.01)         | (0.01)  | (0.01)   | (0.03)  | (0.01)                       | (0.01)  | (0.01)   | (0.04)  |
| Work Income (NTD)                               | -2321          | -734    | -1722    | -8020   | -3831                        | -1355   | -4470    | -8982   |
|                                                 | (478)          | (490)   | (554)    | (1423)  | (546)                        | (586)   | (829)    | (2190)  |
| <hr/> Household Formation (5 years later) <hr/> |                |         |          |         |                              |         |          |         |
| Married per 1k                                  | -1.3           | -2.8    | -0.8     | 0.9     | -2.0                         | -3.6    | -1.8     | 1.9     |
|                                                 | (0.5)          | (0.9)   | (0.7)    | (1.0)   | (0.6)                        | (1.0)   | (0.9)    | (1.2)   |
| Children per 1k                                 | -4.0           | -8.4    | -1.5     | -0.7    | -5.1                         | -10.1   | -1.4     | -1.1    |
|                                                 | (0.7)          | (1.3)   | (0.8)    | (0.4)   | (0.8)                        | (1.5)   | (1.1)    | (0.5)   |
| Treated Units                                   | 1,224,087      | 453,535 | 547,286  | 223,266 | 771,954                      | 332,985 | 308,251  | 130,718 |

*Notes:* This table repeats the estimates of Table 4, but identifies the parents of adult sons using only health insurance records, which link dependents to the employment accounts of a close relative (see Appendix D). Unlike links from the household registry, these do not require family members to be alive in 2011. For other notes, see Table 4.

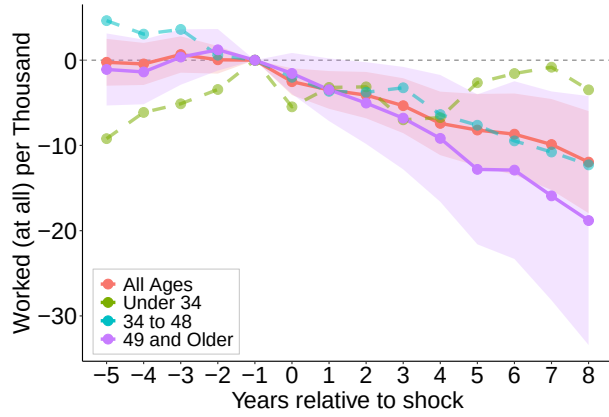


(a) Marriage responses

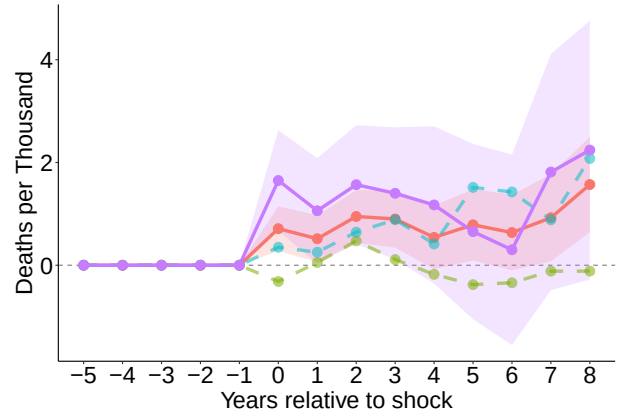
(b) Fertility responses

*Notes:* The x-axis plots the years relative to the LTC-needer's shock. Estimates reported here correspond to the event study estimates,  $\delta_\ell$ , in Equation 3. 95 percent confidence intervals are plotted as the shaded regions for one subgroup (age 33 and under) and for the pooled estimates. Standard errors are clustered at the individual level.

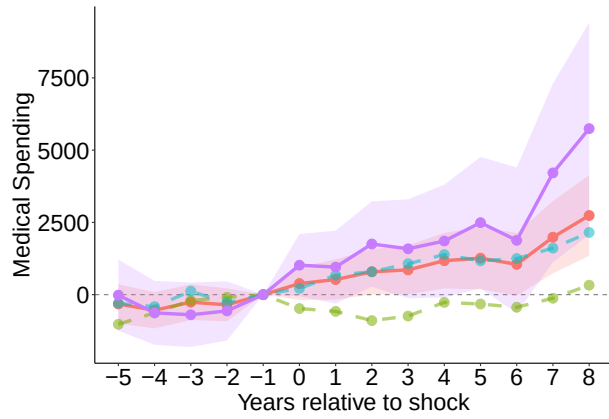
Figure A.2: Event study responses of adult sons to elderly parents' hip fracture: family formation outcomes



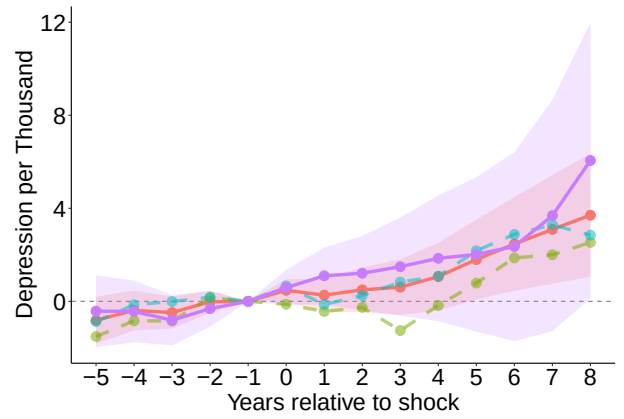
(a) Employment responses



(b) Mortality responses



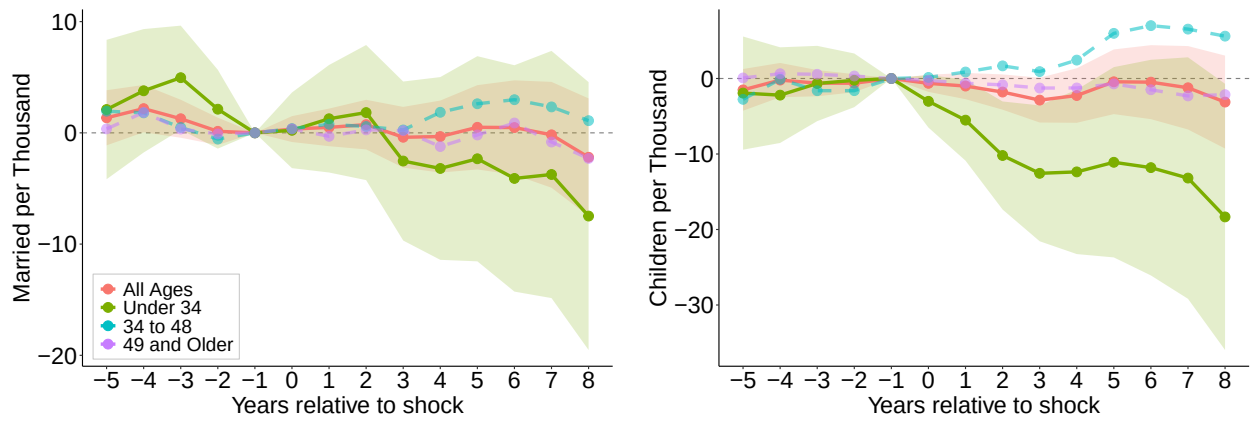
(c) Medical Cost responses



(d) Depression Diagnosis responses

*Notes:* The x-axis plots the years relative to the LTC-needer's shock. Estimates reported here correspond to the event study estimates,  $\delta_{\ell}$ , in Equation 3. 95 percent confidence intervals are plotted as the shaded regions for one subgroup (49 and older) and for the pooled estimates. Standard errors are clustered at the individual level.

Figure A.3: Event study responses of adult sons to elderly parents' hip fracture: economic and health outcomes

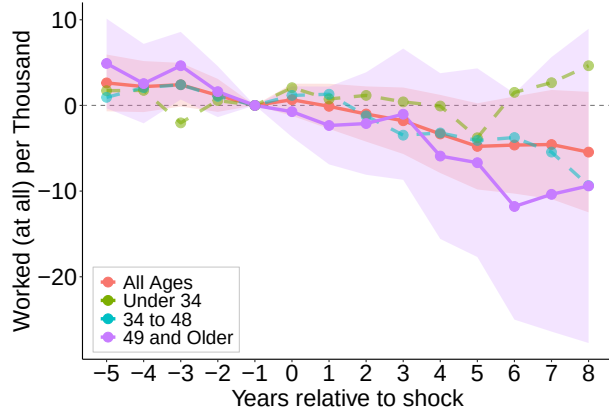


(a) Marriage responses

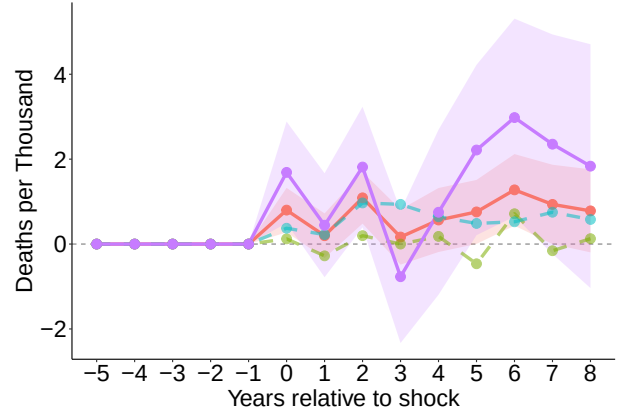
(b) Fertility responses

*Notes:* The x-axis plots the years relative to the LTC-needer's shock. Estimates reported here correspond to the event study estimates,  $\delta_\ell$ , in Equation 3. 95 percent confidence intervals are plotted as the shaded regions for one subgroup (age 33 and under) and for the pooled estimates. Standard errors are clustered at the individual level.

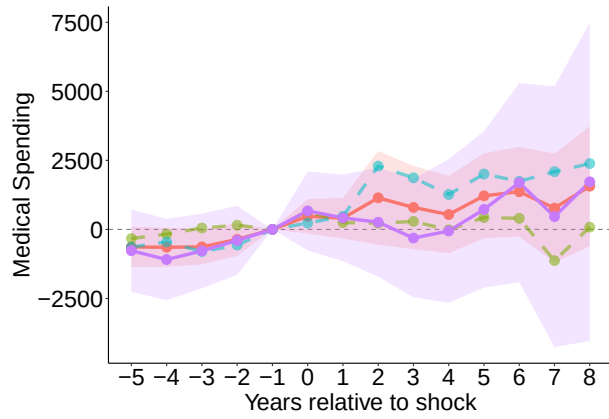
Figure A.4: Event study responses of adult sons to elderly parents' dementia: family formation outcomes



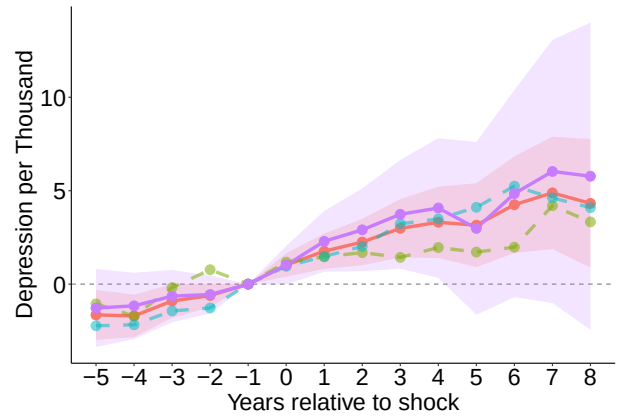
(a) Employment responses



(b) Mortality responses



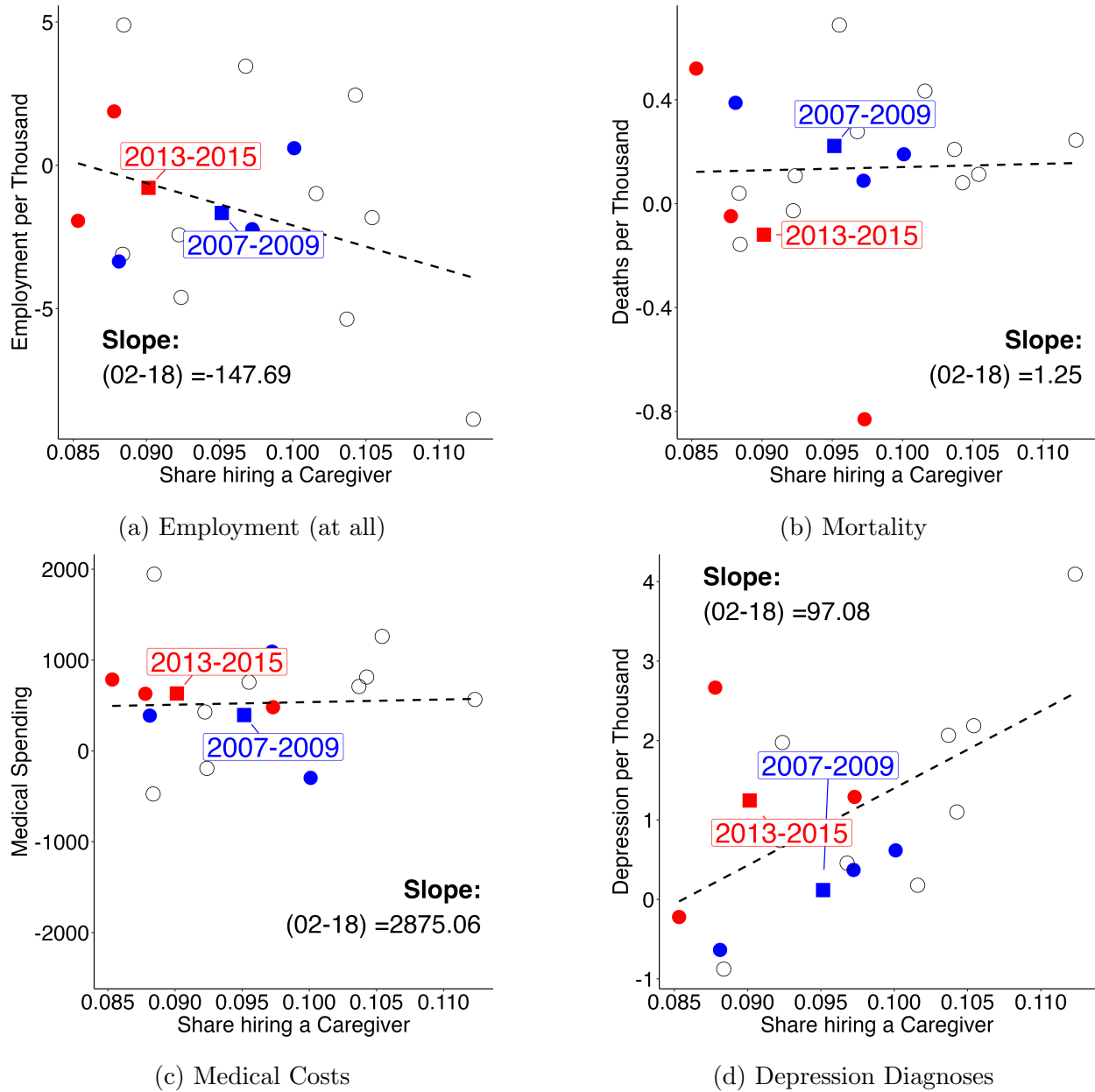
(c) Medical Cost responses



(d) Depression Diagnosis responses

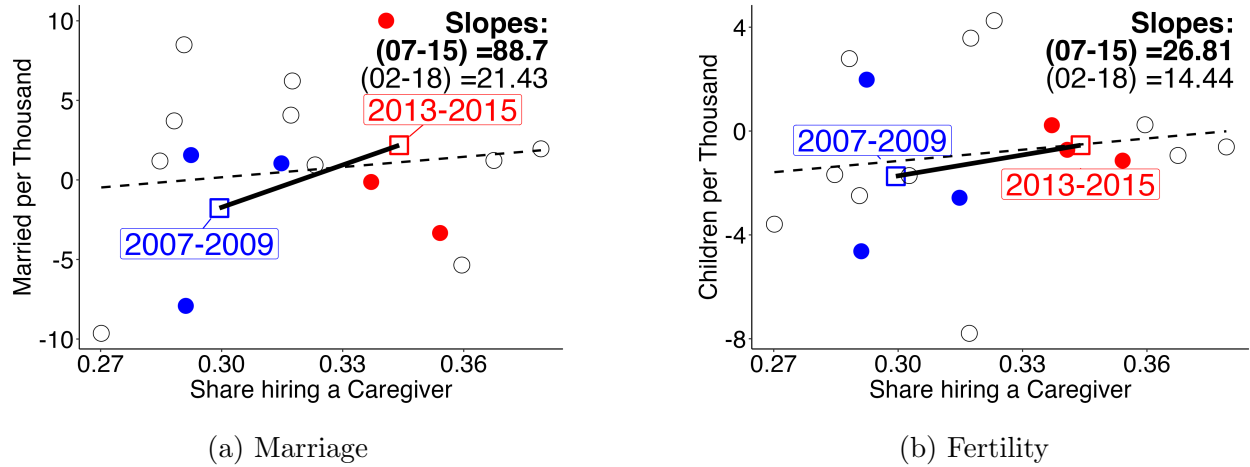
*Notes:* The x-axis plots the years relative to the LTC-needer's shock. Estimates reported here correspond to the event study estimates,  $\delta_{\ell}$ , in Equation 3. 95 percent confidence intervals are plotted as the shaded regions for one subgroup (age 49 and older) and for the pooled estimates. Standard errors are clustered at the individual level.

Figure A.5: Event study responses of adult sons to elderly parents' dementia diagnosis: economic and health outcomes



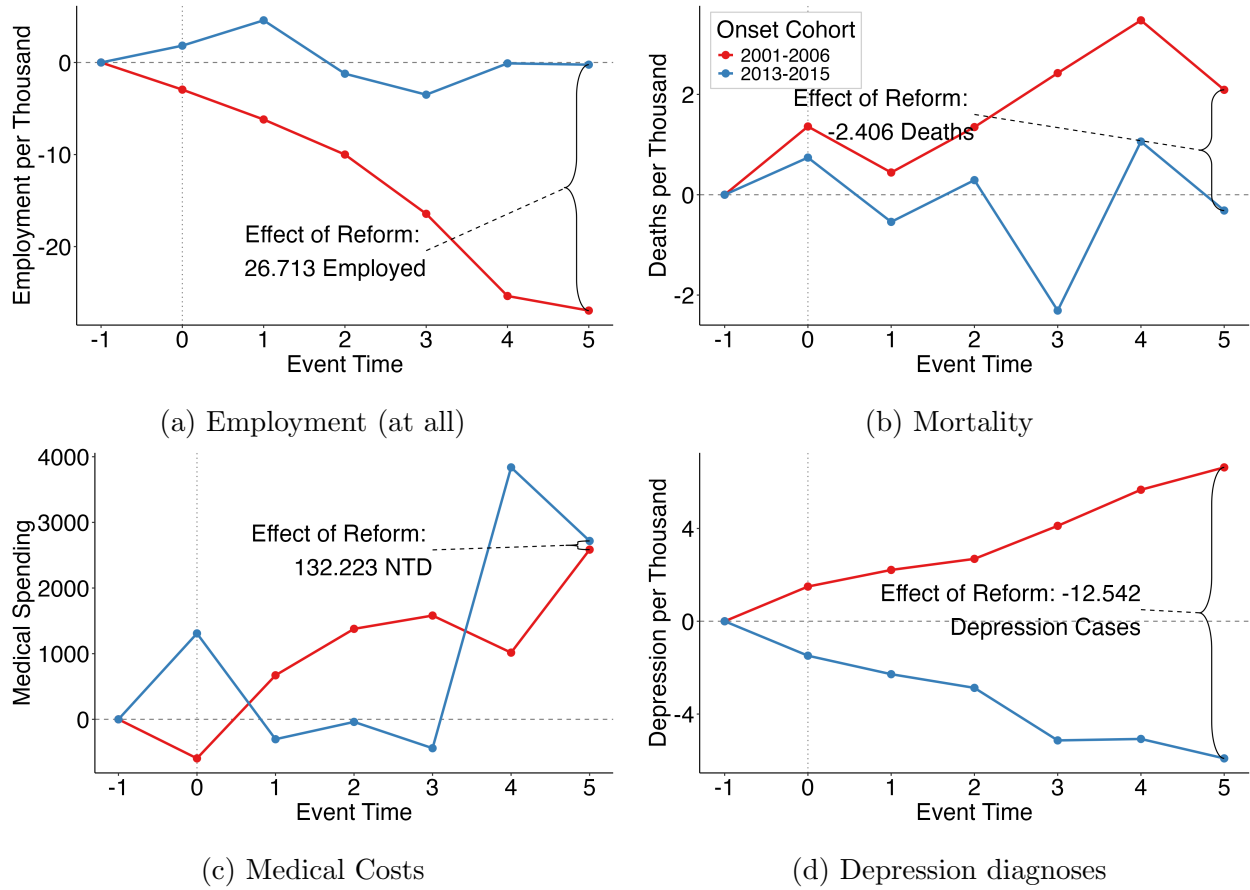
*Notes:* These graphs plot average effects of a stroke separately by the calendar year in which that stroke occurred, for families with an LTC-needer younger than age 80. Effects are averaged over the year of the stroke and the two following years. The line of best fit between all cohort-specific average effects is plotted as the dashed line. Red dots correspond to effects for 2013, 2014, and 2015 cohorts. Blue dots correspond to effects for 2007, 2008, and 2009 cohorts. Squares represent the averages within those groups of cohorts. We omit the line of best fit local to the 2012 reform (plotted in Figure 6) because caregiver hiring rates hardly changed over these years. The sample is based on a 10% draw of sons to families where the health shock occurs at age 79 or younger.

Figure A.6: Cross-cohort correlations between stroke effects and caregiver hiring propensities among non-targeted families



*Notes:* These graphs plot average effects of a stroke separately by the calendar year in which that stroke occurred, for families with an LTC-needer age 80 or older. Effects are averaged over the stroke year and two subsequent years. The line of best fit between all cohort-specific average effects is plotted as the dashed line, and its slope reported as (02-18). Red dots correspond to effects for 2013, 2014, and 2015 cohorts. Blue dots correspond to effects for 2007, 2008, and 2009 cohorts. Squares represent the averages within those groups of cohorts. The line of best fit locally around the reforms (comparing cohorts 2007-2009 versus 2013-2015) is plotted in solid black, with its slope reported as (07-15). The solid points denote the set of cohort-specific estimates entering that local comparison. The sample is based on a 20% draw of sons to families where the health shock occurs at age 80 or older.

Figure A.7: Cross-cohort correlations between stroke effects and caregiver hiring propensities among reform-targeted families



Notes: These plots report effect estimates by year from one's stroke, averaging over the given groups of cohorts.

Figure A.8: Longer-run effects of a stroke on pre-reform versus post-reform cohorts

Table A.2: Average Annual Responses of Adult Sons, by Type of Shock and Age Just Before Shock

|                                     | (1)                | (2)                | (3)                | (4)              | (5)                      | (6)                | (7)             | (8)                  | (9)               |
|-------------------------------------|--------------------|--------------------|--------------------|------------------|--------------------------|--------------------|-----------------|----------------------|-------------------|
|                                     | Under 34           | Stroke<br>34 to 48 | 49 and Older       | Under 34         | Hip Fracture<br>34 to 48 | 49 and Older       | Under 34        | Dementia<br>34 to 48 | 49 and Older      |
| Deaths per 1k                       | 0.13***<br>(0.03)  | 0.53***<br>(0.07)  | 1.74***<br>(0.24)  | 0.02<br>(0.12)   | 0.92***<br>(0.19)        | 1.23**<br>(0.54)   | 0.08<br>(0.21)  | 0.54**<br>(0.24)     | 1.34**<br>(0.66)  |
| Suicides per 1k                     | 0.03***<br>(0.01)  | 0.05***<br>(0.01)  | 0.12***<br>(0.02)  | 0.03<br>(0.04)   | 0.10***<br>(0.04)        | 0.14***<br>(0.05)  | -0.01<br>(0.05) | 0.05<br>(0.06)       | 0.03<br>(0.12)    |
| Depression per 1k                   | 2.92***<br>(0.36)  | 2.99***<br>(0.42)  | 2.56**<br>(1.04)   | 1.13<br>(1.62)   | 2.40**<br>(1.18)         | 2.03<br>(1.77)     | 1.57<br>(2.11)  | 2.60*<br>(1.55)      | 4.35*<br>(2.33)   |
| Medical Spending                    | 755***<br>(202)    | 927***<br>(213)    | 1823***<br>(476)   | -122<br>(426)    | 1271**<br>(527)          | 2388***<br>(847)   | 386<br>(505)    | 1538<br>(1058)       | -219<br>(1365)    |
| Hospital Days                       | 0.12***<br>(0.02)  | 0.14***<br>(0.03)  | 0.11**<br>(0.05)   | -0.03<br>(0.10)  | 0.25**<br>(0.10)         | 0.42***<br>(0.14)  | 0.19<br>(0.16)  | 0.21<br>(0.13)       | -0.22<br>(0.16)   |
| Worked (at all) per 1k              | -4.3***<br>(0.9)   | -6.3***<br>(0.8)   | -19.1***<br>(2.1)  | -3.3<br>(3.9)    | -10.4***<br>(2.2)        | -14.0***<br>(3.8)  | -1.2<br>(4.7)   | -8.2***<br>(2.7)     | -8.6*<br>(5.1)    |
| Months Worked                       | -0.06***<br>(0.01) | -0.09***<br>(0.01) | -0.23***<br>(0.03) | -0.04<br>(0.04)  | -0.12***<br>(0.03)       | -0.18***<br>(0.05) | -0.07<br>(0.05) | -0.10***<br>(0.03)   | -0.12**<br>(0.06) |
| Work Income (NTD)                   | -734<br>(490)      | -1722***<br>(554)  | -8020***<br>(1423) | 158<br>(2104)    | -5064***<br>(1523)       | -7013***<br>(2658) | -2231<br>(2502) | -3698*<br>(1920)     | -3873<br>(3405)   |
| Household Formation (5 years later) |                    |                    |                    |                  |                          |                    |                 |                      |                   |
| Married per Thousand                | -2.8***<br>(0.9)   | -0.8<br>(0.7)      | 0.9<br>(1.0)       | -1.0<br>(4.1)    | -3.3<br>(2.3)            | -0.0<br>(2.4)      | -2.3<br>(4.7)   | 2.6<br>(2.7)         | -0.2<br>(3.0)     |
| Children per Thousand               | -8.4***<br>(1.3)   | -1.5*<br>(0.8)     | -0.7*<br>(0.4)     | -11.5**<br>(5.5) | -3.5<br>(2.5)            | -1.4<br>(1.0)      | -11.1*<br>(6.4) | 6.0*<br>(3.1)        | -0.7<br>(1.1)     |
| Treated Units                       | 453,535            | 547,286            | 223,266            | 18,699           | 39,530                   | 35,808             | 13,305          | 28,903               | 24,004            |

*Notes:* The estimates are obtained from the event study in Equation 3 pooling all post-event periods. For each type of health shock, columns condition on the age of the adult son in the year prior to the health shock. Standard errors are clustered at the individual level. For other notes, see Table 4.

Table A.3: IV Estimates for Effects of Functional Ability on Adult Sons

|                                          | (1)<br>Stroke              | (2)<br>Hip Fracture       |
|------------------------------------------|----------------------------|---------------------------|
| <i>Panel A: First Stage Estimate</i>     |                            |                           |
| Barthel Index                            | -19.95<br>(2.79)           | -10.89<br>(2.55)          |
| <i>Panel B: IV Estimate</i>              |                            |                           |
| Deaths per Thousand                      | -0.02831***<br>(0.004719)  | -0.07436***<br>(0.02515)  |
| Depression per Thousand                  | -0.1453***<br>(0.02491)    | -0.1815**<br>(0.08981)    |
| Medical Spending                         | -50.13***<br>(10.1)        | -119.7***<br>(44.26)      |
| Hospital Days                            | -0.006354***<br>(0.001279) | -0.02188***<br>(0.008128) |
| Worked (at all) per Thousand             | 0.375***<br>(0.06007)      | 0.9054***<br>(0.2702)     |
| Months Worked                            | 0.004956***<br>(0.0007758) | 0.01115***<br>(0.003274)  |
| Work Income (NTD)                        | 116.5***<br>(25.27)        | 409.4***<br>(146)         |
| Children per Thousand<br>(5 Years Later) | 0.2024***<br>(0.04122)     | 0.4267**<br>(0.1866)      |

*Notes:* IV estimates are constructed by dividing effects of health shocks on outcomes (Table 4) by their effects on functional ability (columns 2 and 4 of Table 1 and copied to panel A), measured by the Barthel Index. For these estimates, we have selected the most conservative (largest) first stage estimate for each shock. Standard errors on IV estimates abstract from dependence between the TLSA sample and the population-level administrative data.

## Appendix B The General Safety Net of Taiwan

The Taiwanese general safety net consists primarily of two means-tested programs: the “low income program” and the “middle-low income program.” Both programs provide beneficiaries with a variety of in-kind transfers, including coverage of medical expenses, school tuition, and small housing subsidies.<sup>24</sup> The low income program also provides cash transfers which depend on family size and place of residence: a family in New Taipei City (Taiwan’s largest county) with no income could receive a monthly transfer of 12,393 NTD (approx 413 USD) per person in 2023. By comparison, the 2023 minimum wage was 26,400 NTD (880 USD) per month and the average worker in 2022 earned 57,718 NTD (1,923 USD) per month.<sup>25</sup> Cash transfers through the middle-low income program are smaller and provided only to elderly and disabled individuals.

However, eligibility for both programs is subject to income and asset tests at the *family* level. These means tests are especially aggressive because they can apply even to one’s parents or one’s children *who live in a separate household*, reflecting the social expectation that supporting the needy is first and foremost a responsibility of the family unit. For example, in Taipei, a family must have per-person income no higher than 26,689 NTD (approx. 890 USD) in order to receive cash transfers through the low-income program, and total family savings including investments cannot exceed 150,000 NTD (approx. 5,000 USD). Slightly more lenient means tests apply for the middle-low income program, though they are augmented with a restriction on the value of the family’s total property holdings.<sup>26</sup> In accordance with the stringency of these restrictions, participation in both programs is limited, amounting combined to only 2.6 percent of the Taiwanese population (Ministry of Health and Welfare, 2018).

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<sup>24</sup>Housing subsidies vary across counties. National law imposes a cap on this subsidy of 3,600 NTD (120 USD) per month for a family. The largest county, New Taipei City, has no housing subsidy.

<sup>25</sup>It is common for even minimum wage employers to provide an annual bonus in addition to a monthly salary. In 2022, the average retail service worker received a bonus worth 1.48 months of pay.

<sup>26</sup>Thresholds for per-person income and family savings are roughly 50 percent higher for the middle-low income program. A separate test is applied to property holdings for the middle-low income program; a beneficiary will receive no cash transfers if his or her family owns property worth more than 7.5 million NTD.

## **Appendix C Private Long-Term Care Insurance in Taiwan**

Although commercial LTC policies have been sold since 1995, cumulative sales reached only about 560,000 policies by 2012, of which roughly 416,000 remained in force. By comparison, there were 41.7 million life and 53.8 million health policies outstanding in the same year (Taiwan Insurance Institute, 2015; Tseng and Miao, 2013). The policies that do exist rarely pay out benefits and claims are small when they do. In 2012, insurers paid out 838 initial LTC claims in total, and typical monthly benefit payments of 20,000–30,000 NTD covered only a fraction of prevailing care costs (Taiwan Insurance Institute, 2015). Media reports often highlight cases where claimants face denials due to, for instance, contractual technicalities and disagreements over the intensity of care needs (Lan, 2025).

Beyond commitment and verification problems, another reason for this market failure may lie in regulation itself. Under Article 144 of the Insurance Act, LTC policies are subject to regulatory review. The reviewing agency, the Financial Supervisory Commission, requires prior approval for the prices and terms of LTC insurance policies, and requires policies to use a standardized contract model unless deviations favor the insuree (Legislative Yuan, 2015; Financial Supervisory Commission, 2004, 2006, 2015). Prior approval means products—or changes in their prices—can be delayed or denied, and the standardized contract model leaves insurers little room to experiment with benefit triggers or other contractual adjustments in response to demand.

## **Appendix D Building Family Linkages in Administrative Records**

We identify family links by combining three information channels. The first is health insurance records; access to Taiwan’s universal health care system is nominally tied to employment, but young children and unemployed adults are linked to the employment accounts of a close relative. Observing the adult from whom a child derives her health insurance coverage provides a reliable way of identifying the child’s parent, at least if she is observed as an adolescent at some point in our data. The second source comes from the Household Registry, which is an annually renewed administrative record of individuals’ addresses and co-residing family members. With the Household Registry, we can link any relatives or spouses who were cohabiting at some point between 2000 and 2020. The final source is the birth record,

which contains comprehensive information on parent-child links after 1995.

The weakness of our approach is that we can fail to link an elderly person to adult children if, by 2000, those children were old enough to continuously access health insurance through their own employment account (or that of a spouse) and were old enough to no longer co-reside with the elderly parent. Figure D.1 reports parental match rates by sex and birth cohort, focusing on the cohorts most relevant for our analysis (1950 to 2000). Overall, we manage to identify at least one parent for 55 percent of women and 68 percent of men. Our method is somewhat less successful at matching daughters to their parents because daughters tend to leave the household upon marriage. By contrast, sons (especially the eldest son) tend to cohabit with parents more consistently throughout adulthood, and they may link to their parents through health insurance records even as an adult. Beyond cultural norms that would place the burden of care on the eldest son, this is another reason we choose to focus our analyses mainly on adult sons. However, it is worth noting that our measure of family size will be downward biased. We do have an accurate accounting of individuals who cohabit with elderly family members, who presumably are most responsible for care in the event that a negative health shock arises.<sup>27</sup>

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<sup>27</sup>However, to appear in the household registry, an individual had to be alive in 2011. The links in the household registry are therefore selected on survival of both the adult child and older parent in a way which may bias some of our results—especially when examining death as an outcome. However, our main estimates are similar when we build families exclusively from health insurance records, which do not require survival to 2011 (Appendix Table A.1).

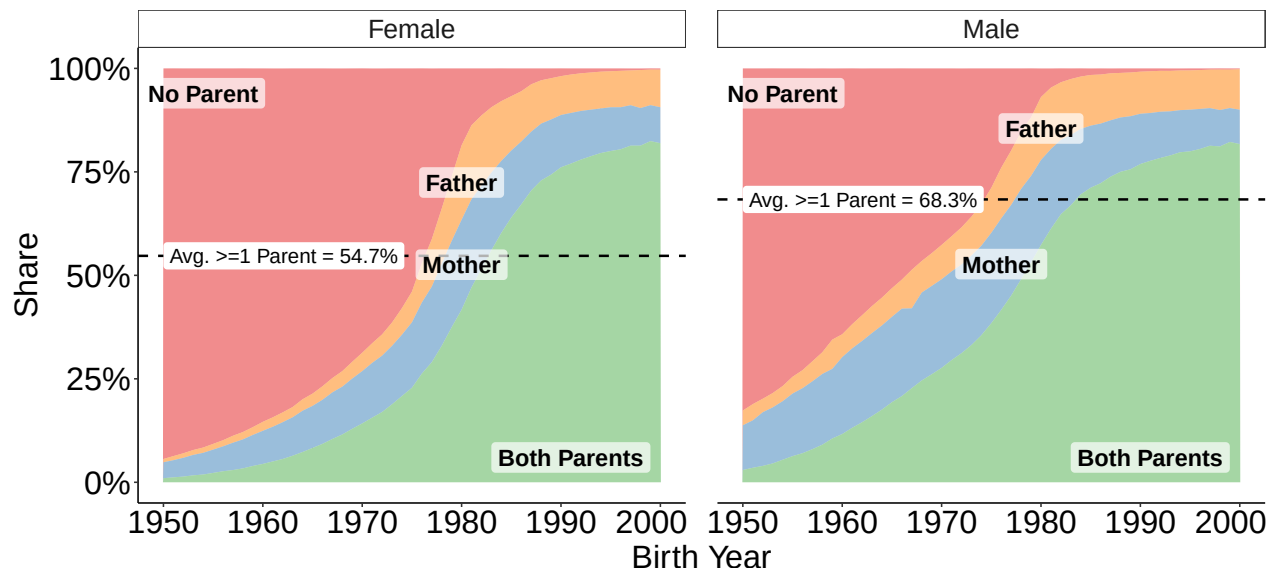


Figure D.1: Match rate of children to their parents

*Notes:* Parent links may be missing because parents were already dead before the periods of our data sources begin.

## Appendix E ICD Codes for Diseases

Here we list the ICD-9 and ICD-10 codes for various diseases discussed in this paper.

### 1. Stroke:

- ICD-9: 430, 431, 4321, 4320, 4329, 43391, 43321, 43301, 43311, 43381, 43401, 43411, 43491, 43331, 43320, 43300, 43310, 43330, 43380, 43390, 43400, 43410, 43490, 4373, 4370, 463, 4372, 4375, 4376, 4374, 4371, 3483, 4359, 436, 4378, 4379
- ICD-10: I6000, I6001, I6002, I6010, I6011, I6012, I6020, I6021, I6022, I6030, I6031, I6032, I604, I6050, I6051, I6052, I606, I607, I608, I609, I610, I611, I612, I613, I614, I615, I616, I618, I619, I6200, I6201, I6202, I6203, I621, I629, I6300, I63011, I63012, I63019, I6302, I63031, I63032, I63039, I6309, I6310, I63111, I63112, I63119, I6312, I63131, I63132, I63139, I6319, I6320, I63211, I63212, I63219, I6322, I63231, I63232, I63239, I6329, I6330, I63311, I63312, I63319, I63321, I63322, I63329, I63331, I63332, I63339, I63341, I63342, I63349, I6339, I6340, I63411, I63412, I63419, I63421, I63422, I63429, I63431, I63432, I63439, I63441, I63442, I63449, I6349, I6350, I63511, I63512, I63519, I63521, I63522, I63529, I63531, I63532, I63539, I63541, I63542, I63549, I6359, I636, I638, I639, I6501, I6502,

I6503, I6509, I651, I6521, I6522, I6523, I6529, I658, I659, I6601, I6602, I6603, I6609, I6611, I6612, I6613, I6619, I6621, I6622, I6623, I6629, I663, I668, I669, I670, I671, I672, I673, I674, I675, I676, I677, I6781, I6782, I6783, I67841, I67848, I6789, I679

## 2. Dementia:

- ICD-9: 29040, 29041, 29042, 29043, 2941, 29410, 29411, 2900, 29010, 29011, 29012, 29013, 29020, 29021, 2903, 2908, 2909, 2948, 2949, 3101, 2940, 29382, 29381, 2939, 3100, 3102, 3108, 3109, 3308, 3309, 3310, 3311, 3312, 3317, 33189, 3319
- ICD-10: F0150, F0151, F0280, F0281, F0390, F0391, F04, F060, F062, F068, F070, F0781, F0789, F079, F09, G3181, G3182, G319, G300, G301, G308, G309, G3101, G3109, G311, G312, G3183, G3185, G3189

## 3. Depression:

- ICD-9: 296, 2960 ~ 2968, 3004, 309, 3090 ~ 3091, 311, 3110 ~ 3115
- ICD-10: F320, F321, F322, F323, F324, F325, F328, F329, F330, F331, F332, F333, F334, F339, F301, F302, F303, F304, F308, F309, F310, F311, F312, F313, F314, F315, F316, F317, F318, F319, F341

# Appendix F Linking the Value of Life and the Value of Quality of Life

In this section, we derive the relationship between the value of life and the value of quality of life, used in Section 7. The post-shock problem of Equation 6 has Lagrangian

$$\begin{aligned} \mathcal{L}(a) = & \sum_{t=a}^L H_t^*(a) u(c_t(a)) \frac{1}{(1+\rho)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right] \\ & + \mu^* \frac{A^*(a)}{(1+r)^a} + \mu^* \sum_{t=a}^L [y_t^*(a) - c_t(a)] \frac{1}{(1+r)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right] \end{aligned}$$

whose first order condition for consumption in year  $t$  pins down the marginal utility of income:

$$H_t^*(a) u_c(c_t^*(a)) \frac{1}{(1+\rho)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right] = \mu^* \frac{1}{(1+r)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right]$$

$$\iff \mu^* = H_t^*(a) u_c(c_t^*(a)) \frac{(1+r)^t}{(1+\rho)^t}$$

where the survival weights cancel because payoffs and the budget constraint are weighted by the same probability of reaching year  $t$ .

Notice by differentiating the Lagrangian that the value of increases in quality of life in year  $j$  is:

$$V_{H_j}^*(a) = \frac{\partial \mathcal{L}(a)}{\partial H_j^*(a)} = u(c_j^*(a)) \frac{1}{(1+\rho)^j} \left[ \prod_{s=a}^j \lambda_s^*(a) \right]$$

where only the payoff sum contributes, because quality of life does not enter the budget constraint. Multiplying by  $H_j^*(a)$  to put the gain in proportional terms, scaling by the marginal utility of income  $\mu^*$ , and substituting for  $\mu^*$  above leaves:

$$\frac{V_{H_j}^*(a) \times H_j^*(a)}{\mu^*} = \underbrace{\frac{u(c_j^*(a))}{u_c(c_j^*(a))}}_{\text{value of living in year } j} \underbrace{\frac{1}{(1+r)^j} \left[ \prod_{s=a}^j \lambda_s^*(a) \right]}_{P(\text{survive to } j)} \quad (11)$$

where  $c_j^*(a)$  denotes optimal consumption in year  $j$  under income insurance. Meanwhile, consider the value of life from year  $a$ . Because  $\lambda_a^*(a)$  appears in the survival weight of every later year, it factors out of the entire sum:

$$V_{\lambda_a}^*(a) = \frac{\partial \mathcal{L}(a)}{\partial \lambda_a^*(a)} = \frac{1}{\lambda_a^*(a)} \left\{ \sum_{t=a}^L H_t^*(a) u(c_t^*(a)) \frac{1}{(1+\rho)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right] \right. \\ \left. + \underbrace{\mu^* \sum_{t=a}^L [y_t^*(a) - c_t^*(a)] \frac{1}{(1+r)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right]}_{= 0 \text{ if } y_t^*(a) = c_t^*(a)} \right\}$$

Supposing that extending life does not affect one's budget constraint (which is implied if individuals do not save, so that  $y_t^*(a) = c_t^*(a)$ ), the second term drops out. Scaling by the marginal utility of income as before gives:

$$\frac{V_{\lambda_a}^*(a)}{\mu^*} = \frac{1}{\lambda_a^*(a)} \sum_{t=a}^L \frac{u(c_t^*(a))}{u_c(c_t^*(a))} \frac{1}{(1+r)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right] \\ \approx \sum_{t=a}^L \frac{u(c_t^*(a))}{u_c(c_t^*(a))} \frac{1}{(1+r)^t} \left[ \prod_{s=a}^t \lambda_s^*(a) \right] \quad (12)$$

where the approximation uses that  $\lambda_a^*(a) \approx 1$  (annual survival rates among prime-age adult men are high). Combining (11) and (12) allows us to relate the value of a life from year  $a$  to the value of quality of life adjustments:

$$\frac{V_{\lambda_a^*}^*(a)}{\mu^*} \approx \sum_{t=a}^L \frac{V_{H_t}^*(a) \times H_t^*(a)}{\mu^*} \quad (13)$$

Equation 13 allows us to relate an estimate for the value of life to the year-specific values of a proportional change in quality of life. Combined with some structure on how those year-specific values of quality of life can differ (as we use in Section 7), we can back out estimates for them.