

# Longevity Risk in Reinsurance and Equity Markets

Laursen, Nicolai Søgaard

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LONGEVITY RISK IN REINSURANCE AND EQUITY MARKETS

Nicolai Sogaard Laursen

# LONGEVITY RISK IN REINSURANCE AND EQUITY MARKETS

Department of Finance

PhD Series 15.2023

**CBS**  COPENHAGEN BUSINESS SCHOOL  
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# Longevity risk in reinsurance and equity markets

Nicolai Søgaaard Laursen

A thesis presented for the degree of

Doctor of Philosophy

Supervisor: Jesper Rangvid

Ph.D. School in Economics and Management

Copenhagen Business School

Nicolai Søgaaard Laursen  
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# Preface

This dissertation represents the tangible output of my enrollment as a Ph.D. student at the Department of Finance, Copenhagen Business School (CBS), between September 2019 and January 2023. I am grateful to the Pension Research Center (PeRCent) and the Department of Finance for providing financial support and an outstanding research environment.

I would like to thank Jesper Rangvid and Svend E. Hougaard Jensen for encouraging me to pursue a Ph.D. at PeRCent. Furthermore, I am grateful to Jesper for inspiring me to study the subject of longevity and, subsequently, for providing helpful feedback and numerous useful comments on every chapter of my thesis. I am also thankful to my secondary supervisor, Paul Whelan, for turning my novice third chapter into a more comprehensive contribution to the financial literature, and to my other secondary supervisor, Peter Feldhütter, for carefully reading and making suggestions on the first chapter. Lastly, I would like to express my gratitude towards the faculty at the Department of Finance for providing helpful suggestions for each chapter at numerous department presentations.

An especially big thanks is reserved for Malene Kallestrup-Lamb, the co-author of my first chapter. Malene profoundly and positively influenced my writing skills, ability to disseminate research, and the depth of my understanding of longevity risk. Due to COVID-19, my research stay at Heriot-Watt University, Scotland, was substituted by a stay in Denmark at the Department of Economics and Business Economics, Aarhus University, where I have had numerous beneficial discussions about my thesis. I am highly appreciative of my stay due to the department's exceptional research environment.

Finally, I would like to thank my family, especially my parents for supporting me and my three siblings in all our endeavors. Our positive traits, accomplishments and joys can largely be attributed to the groundwork you laid. Arguably, the greatest achievement humanly possible is to bring children into the world. My partner, Mette, and I have been blessed twice with the arrival of our children, Lise and Erik. All of you remind me of what is most important in life.



# Abstract

This PhD dissertation centers on systematic longevity risk, which refers to the uncertainty of future survival rates for a group of individuals. Each chapter provides independent contributions to the challenge of how to alternatively redistribute longevity risk.

The first chapter, *Longevity hedge effectiveness using socioeconomic indices* coauthored with Malene Kallestrup-Lamb, uses Danish mortality data stratified into socioeconomic groups to evaluate basis risk in longevity hedging. The study addresses the question of how annuity providers, exposed to different socioeconomic mortality rates, can hedge the variability of a life annuity most efficiently. We assume that the annuity provider can engage in a mortality-linked security via two alternative hedging strategies, with and without basis risk, and evaluate the costs and benefits. The cost is represented by the notional amount of hedging contracts optimally bought times the actuarial risk premium, whereas the benefit is denoted as the risk reduction in the variability of a life annuity. We find that eliminating basis risk is more cost-effective for the annuity provider, as it allows a higher degree of hedge effectiveness at a cost equivalent to a hedge with basis risk. Lastly, the yearly expenses of hedging longevity risk require, at most, an extra added rate of return of no more than 0.2%.

The current assumption in the actuarial literature is that the evolution of mortality rates is independent of financial risks. The second chapter, *Unsystematic mortality and time-varying returns*, however, shows that an increase in unsystematic mortality raises one-year ahead real returns heterogeneously for portfolios sorted by industry and dividend-price ratios in the U.S stock market. For half of the industry portfolios, the magnitude is equivalent to a comparable increase in the dividend-price ratio, whereas the remaining industry portfolios are not influenced. In addition, standardized mortality shocks account for 80% of the average real return gap between value and growth portfolios. I assume that higher unsystematic mortality rates pose temporary adverse demand shocks. Therefore, I analyze unsystematic mortality shocks in a two-good consumption-based asset pricing model and infer economic point estimates of the loss in consumption of various goods for unsystematic mortality's adjacent ages. This validates the heterogeneous impact on real

returns across industry returns; however, the equivalent impact on the dividend-price sorted portfolios cannot be attributed to industry composition alone.

The value premium has previously been suggested as capturing an omitted risk variable (Ball, 1978; Fama and French, 1992), which conforms with longevity risk as the underlying uncertainty, mortality rates, are published with a two-year lag. In the third chapter, *Longevity risk and the value premium*, I introduce stochastic survival rates in the intertemporal budget constraint for a representative agent with standard recursive preferences. Thus, higher longevity (mortality) changes increase (decrease) the representative agent's intertemporal budget constraint. The assumption of a representative agent is justified as changes in conditional survival rates are perfectly positively correlated across ages. Longevity risk also evolves as a random walk, which suggests that any impact on asset prices would be unrelated to hedging demands and, instead, could influence returns comparably to market cash-flow risk. The model suggests that the risk premium linked to longevity risk is slightly higher than the equivalent related to news regarding the market portfolio's cash flows. I find empirical support for the model and show that the monthly Sharpe ratio associated with longevity risk is around 10%.

# Resumé

Denne Ph.d.-afhandling sætter fokus på systematisk levetidsrisiko, hvilket er defineret som usikkerheden omhandlende en gruppe individers fremtidige overlevelsessandsynligheder. Alle kapitler bidrager uafhængigt til udfordringen omkring, hvordan levetidsrisiko alternativt kan allokeres.

Det første kapitel, *Longevity hedge effectiveness using socioeconomic indices*, udarbejdet med Malene Kallestrup-Lamb, bruger dansk dødelighedsdata stratificeret i socioøkonomiske grupper til at evaluere basisrisiko i afdækningen af levetidsrisiko. Studiet adresserer spørgsmålet om hvordan annuitetsudbydere, som er eksponeret for forskellige socioøkonomiske dødelighedsrater, mest efficient kan afdække variationer i en livsannuitet. Vi antager, at annuitetsudbyderen kan indgå i en dødelighedsrelateret kontrakt ved to alternative afdækningsstrategier, med og uden basisrisiko, og evaluerer omkostningerne og fordelene. Omkostningerne er defineret som det nominelle antal af afdækningskontrakter optimalt set købt ganget med den forsikringsmatematiske risikopræmie, hvorimod fordelene er angivet som risikoreduktionen i variationerne af en livsannuitet. Vi finder, at det mest omkostningseffektive for en annuitetsudbyder er at eliminere basisrisiko, da det medfører en højere risikoreduktion til en omkostning, der er ækvivalent med én afdækning med basisrisiko. Slutteligt kræver de årlige udgifter relateret til afdækning af levetidsrisiko højest et ekstraafkast på 0,2 %.

Den nuværende antagelse indenfor litteraturen for forsikringsmatematik er, at udviklingen af dødelighedsrater og finansielle risici er uafhængig. Det andet kapitel, *Unsystematic mortality and time-varying returns*, finder i modsætning hertil, at en stigning i usystematisk dødelighedsrisiko hæver realafkastet et år forud heterogent for porteføljer sorteret på henholdsvis industrier samt dividende og pris-ratioer på det amerikanske aktiemarked. For halvdelen af industriporteføljerne er størrelsesordenen ækvivalent med en tilsvarende stigning i deres dividende-pris-ratio, hvorimod de resterende industriporteføljer er ikke influeret. Ydermere kan 80 % af forskellen på de gennemsnitlige realafkast mellem værdi- og vækstporteføljer tilskrives standardiserede dødelighedsshocks. Jeg formoder, at højere usystematiske dødelighedsrater udgør midlertidige negative efterspørgselsshocks. Derfor analyserer jeg usystematiske dødelighedsshocks i en to-godes

forbrugsbaseret aktiv-prissætningsmodel og udleder økonomiske punkttestimater af forbrugstabet for forskellige goder for usystematisk dødeligheds tilstødende aldre. Dette validerer den heterogene påvirkning af realafkastet på tværs af industriporteføljerne. Dog kan den samme påvirkning af porteføljerne sorteret på dividende og pris-ratioer ikke baseres udelukkende på sammensætningen af industrier.

Værdipræmien er tidligere blevet foreslået at repræsentere en uobserverbar risikopræmie (Ball, 1978; Fama and French, 1992), hvilket stemmer overens med levetidsrisiko, da den underliggende usikkerhed, dødelighedsrater, bliver publiceret med to års forsinkelse. I det tredje kapitel, *Longevity risk and the value premium*, introducerer jeg stokastiske overlevelseshaster i den intertemporale budgetbegrænsning for en repræsentativ agent med standard rekursive præferencer. Dette betyder, at ved højere levetidsændringer (dødelighedsændringer) stiger (falder) den repræsentative agents intertemporale budgetbegrænsning. Antagelsen om en repræsentativ agent er retfærdiggjort, da ændringer i betingede overlevelseshaster er perfekt positivt korreleret på tværs af aldre. Da levetidsrisiko også udvikler sig tilfældigt (random walk), kunne det indikere, at enhver indflydelse fra levetidsrisiko på aktivpriser er urelateret til en tidsvarierende efterspørgsel efter aktiver. Derimod kan levetidsrisiko potentielt påvirke afkast sammenligneligt med dividende-nyheder for markedsporteføljen. Modellen antyder, at risikopræmien forbundet til levetidsrisiko er lidt højere end den tilsvarende relateret til dividende-nyheder for markedsporteføljen. Jeg finder empirisk underbyggelse af modellen og viser, at den månedlige Sharpe ratio forbundet med levetidsrisiko er omkring 10 %.

# Contents

<b>Preface</b>	<b>iii</b>
<b>Abstract</b>	<b>v</b>
<b>Resumé</b>	<b>vii</b>
<b>1 Longevity hedge effectiveness using socioeconomic indices</b>	<b>1</b>
1.1 Introduction . . . . .	2
1.2 Methodology . . . . .	5
1.2.1 Longevity hedging . . . . .	7
1.3 Affluence index and data . . . . .	12
1.4 Results . . . . .	15
1.4.1 Benefits of hedging longevity risk . . . . .	15
1.4.2 Cost of hedging longevity risk . . . . .	17
1.5 Concluding remarks . . . . .	23
1.6 Appendix . . . . .	24
<b>2 Unsystematic mortality and time-varying returns</b>	<b>49</b>
2.1 Introduction . . . . .	50
2.2 Methodology . . . . .	53
2.2.1 Financial magnitude . . . . .	55
2.2.2 Economic mechanism . . . . .	56
2.2.3 Predictions following exogenous shock . . . . .	58
2.3 Data and descriptive statistics . . . . .	59
2.3.1 Mortality data description . . . . .	59
2.3.2 Estimating unsystematic mortality . . . . .	60
2.3.3 Descriptive statistics - financial data . . . . .	64
2.4 Empirical results . . . . .	66
2.4.1 Mortality impact on industry time-varying returns . . . . .	66

2.4.2	Economic magnitude . . . . .	69
2.4.3	Mortality impact on value and growth's time-varying returns . . . . .	71
2.5	Conclusion . . . . .	74
2.6	Appendix . . . . .	76
<b>3</b>	<b>Longevity risk and the value premium</b>	<b>92</b>
3.1	Introduction . . . . .	93
3.2	Methodology . . . . .	96
3.2.1	The stochastic discount factor . . . . .	97
3.2.2	Mutually exclusive factor models . . . . .	99
3.3	Data and descriptive statistics . . . . .	101
3.3.1	Estimating mimicking portfolios . . . . .	102
3.3.2	Time-series relations . . . . .	106
3.4	Empirical estimates of longevity risk premia . . . . .	108
3.4.1	Theory motivated cross-sectional regressions . . . . .	108
3.4.2	Empirical motivated cross-sectional regressions . . . . .	112
3.5	Conclusion . . . . .	116
3.6	Appendix . . . . .	117





# Chapter 1

## 1 Longevity hedge effectiveness using socioeconomic indices

with Malene Kallestrup-Lamb.

### Abstract

This paper evaluates socioeconomic basis risk in longevity hedging. Using data for a full population stratified into socioeconomic groups, we explore the benefits and costs of two alternative hedging strategies, with and without basis risk, in the capital market. The benefit of the longevity hedge is represented by the risk reduction in the variability of a life annuity, whereas the cost is the notional amount of hedging contracts times the actuarial risk premium. We find that hedging is more cost-effective for the annuity provider when basis risk is eliminated. Moreover, it allows for a higher degree of hedge effectiveness at a cost that is equivalent to a hedge where basis risk is present. Finally, the yearly expenses related to hedging longevity risk requires, at most, an extra added rate of return of no more than 0.2 %.

## 1.1 Introduction

Pension systems are heavily exposed to longevity risk (see, e.g., Antolin (2007); Blake et al. (2006)) and globally, the longevity risk exposure in private-sector corporations is estimated to be around USD 25 trillion (Biffis and Blake, 2014). In addition, each year of added life expectancy increases the present value of pension liabilities by 3-4% (Kisser et al., 2012). The literature distinguishes between two types of longevity risk: idiosyncratic and systematic. The former is related to the individual's uncertainty about their life expectancy and can be diversified by pooling enough participants in a pension fund. The latter, in contrast, relates to uncertainty about future mortality rates for a *group* of individuals and cannot be diversified by increasing the number of participants (Hari et al., 2008). When addressing longevity risk, our study focuses only on systematic longevity risk.

Various options are available for pension funds aiming at reducing their longevity risk exposure. Either they can manage the risk internally by, for example, including life insurance in their portfolios (Wang et al., 2010), smoothing the payout profile of a life annuity via bonus payments (Norberg, 1999), or by investing more aggressively under dire circumstances with low returns and longevity improvements (Chaudhry et al., 2017). Alternatively, they can offload their risk in the Life Market that allows pension funds to transfer longevity and mortality risks (Blake et al., 2013). Here, the systematic longevity risk can be reduced by engaging in mortality- or longevity-linked securities either via a capital market solution or an insurance-based solution.

Insurance-based contracts have traditionally been the preferred hedging solution (Barrieu et al., 2012; Blake et al., 2019), but pension funds can also manage their longevity risk through the global capital market using financial instruments with payoffs linked to a predefined mortality index. Capital market solutions differ from insurance-based solutions by favoring investors in terms of standardized indices, whereas insurance-based ones benefit hedge effectiveness as securities are indexed to the specific underlying population. Consequently, insurance-based solutions are, a priori, believed to be more expensive to set up, less liquid, and less attractive to investors, whereas the opposite is true for capital market solutions (Blake et al., 2019).

The low demand for the capital market solution has primarily been attributed to basis risk

represented by the residual risk between the reference group and the annuity providers longevity experience (Coughlan et al., 2011). After taking age and gender into account, the most important determinant of basis risk is the variation between socioeconomic groups (Coughlan et al., 2011). Due to the volume of longevity risk globally, it is clear that the risk capacity of the global insurance industry is insufficient (Michaelson and Mulholland, 2014) and that only capital markets can provide a sustainable risk capacity (Barrieu et al., 2012). Thus, the primary concern is how to structure capital market solutions such that they are still attractive to capital market investors, though without compromising the hedging purpose to relieve longevity risk of annuity providers. An obvious option to consider is a capital market solution that is indexed to the mortality experience of socioeconomic groups instead of the full population to reduce the concern of basis risk.

This motivates an analysis of two alternative hedging strategies. One strategy using an index representing the full population (FP), i.e., *with* basis risk, and one using indices representing specific socioeconomic groups (SEG), i.e., *without* basis risk. Based on this setup, our contribution is twofold. As the first in the literature, we analyse the hedge effectiveness for five different socioeconomic groups within a population for the two alternative hedging strategies. Secondly, we estimate the total costs from hedging longevity risk, allowing an evaluation of costs and benefits simultaneously across the socioeconomic dimension. Moreover, this enable us to relate socioeconomic risk premiums on mortality-linked securities to the aggregate risk premium for the corresponding security.

Benefits are quantified by the relative risk adjustment (RRA) as in Li and Hardy (2011). The percentage point difference in terms of RRA between the two hedging strategies, with and without basis risk, is thus defined as basis risk. Haberman et al. (2014), Li and Hardy (2011), and Ngai and Sherris (2011) all investigate the influence of basis risk but disregard the heterogeneity within a population. In a related paper to our study, Cairns et al. (2014) analyze basis risk involving the use of England & Wales male mortality to hedge liabilities linked to the Continuous Mortality Investigation (CMI) male assured lives. However, the hedge effectiveness is only evaluated for a selected subpopulation.

In the literature, the cost of hedging longevity risk has either focused on finding the optimal notional amount of contracts or pricing various mortality- and longevity-linked securities; see Bravo and Nunes (2021) for an overview on pricing of longevity-linked securities. To quantify costs, we suggest a new measure that multiply actuarial risk premiums using the Wang transform (Wang, 2000) with the optimal notional amount of hedging securities using Cairns (2013)’s framework. The risk measure arising from the Wang transform has an attractive financial application, namely that all prices are conditional on the systematic risk price. The risk premium from engaging in the two alternative hedging strategies can thus be evaluated in a unified framework dependent on the market price for the aggregate risk. Moreover, the risk premiums yield closed-form solutions as we consider q-forward contracts as the hedging instrument, as introduced by Coughlan et al. (2007). In contrast, Boyer and Stentoft (2013) and Hunt and Blake (2020, 2021) price various mortality- and longevity-linked securities, which generally have cumbersome or no closed-form solutions.

The underlying indices follow mortality rates for five different socioeconomic groups, to which we apply the Li and Lee (2005) model. To measure adherence to a socioeconomic group, we use the newly developed affluence measure in Cairns et al. (2019) based on data from Kjærgaard et al. (2020). The affluence index, which allocates individuals based on data on income and wealth, is an attractive feature as both covariates are typically available to annuity providers, enabling them to quantify their own basis risk contingent on their annuity liabilities. Generally, the relationship between affluence and life expectancy has been heavily documented and found to be positively related; see Blane et al. (1997); Brown et al. (2012); Saurel-Cubizolles et al. (2009).

We find that basis risk is a key metric in evaluating how large the advantage of the subgroup-specific hedge is. The benefit is heterogeneous but larger the further away the group is from average life expectancy. Hence, socioeconomic groups exposed to more basis risk have the highest potential benefit. On average, we find that socioeconomic groups obtain RRAs that are 20 and 65 percentage points higher for males and females, respectively. This corresponds well with the findings in Cairns et al. (2014) that shows a reduction in hedge effectiveness for males of 30 percentage points. Moreover, given the same level of RRA, we find that the costs of a hedge

referenced to socioeconomic-specific mortality rates is at least 50 % lower for females compared to a hedge based on the full population. For males, the same result holds for the highest and lowest affluence groups. Given similar total cost, a higher level of RRA is attainable for the subgroup-specific mortality alternative, which is robust to other mortality model specifications. Thus, we find that a hedge referenced to subgroup-specific mortality rates is preferred to the full population equivalent. In addition, the yearly expenses related to hedging longevity risk amounts to an extra annual required rate of return between 0.18 % and 0.10 %, which is decreasing for groups with higher average life expectancy. Generally, this is comparable to the costs of investing in passive equity funds. These findings conclude that hedging longevity risk via securities indexed to the mortality experience of socioeconomic groups may be an attractive alternative to pursue in the future to increase the demand for capital market solutions.

The remainder of this paper is structured as follows. Section 1.2 presents the framework of our analysis, including life annuity, mortality model, and actuarial risk premium. Section 1.3 introduces the data and the affluence index. Section 1.4 summarizes the results of hedging longevity risk, while Section 1.5 concludes and discusses implications.

## 1.2 Methodology

Pension funds issue life-long annuities that pay a series of cash flows, irrespective of how long the annuitant lives. Consequently, pension funds are exposed to longevity risk. We denote the value of a life-long annuity as

$$a(x, i) = \sum_{\tau=1}^{x_u-x} (1+r)^{-\tau} p(x, \tau, i). \quad (1.1)$$

The subscript  $i$  indicates that the value of the annuity differs across the underlying population, i.e., in our case, the socioeconomic groups. The value of the annuity in equation (1.1) is an ex-post relationship, meaning that it can deviate from its ex-ante set terms. We assume that the terms of the annuity contract are set ex-ante at time 0, for an annuitant aged  $x$  whom have a maximum attainable age  $x_u$ , and pays a *fixed* stream of USD 1 per annum, at the end of each  $\tau$  years that the

annuity contract persists. Furthermore, we assume that the discount rate,  $r$ , is constant to focus our attention on uncertainty in the annuity value stemming from longevity risk. Thus, the ex-post value of the annuity unfolds at future time points,  $\tau$ , due to variations in survival rates,  $p(x, \tau, i)$ , which is the best estimate at time 0 that the annuitant, aged  $x$  from group  $i$ , survives another  $\tau$  years. We denote the conditional survival probability as

$$p(x, \tau, i) = \prod_{\tau=0}^{x_u-x-1} [1 - q(x, \tau, i)], \quad (1.2)$$

where  $q(x, \tau, i)$  is the probability that the annuitant, aged  $x$  at time 0, dies in the age interval  $x + \tau$  to  $x + \tau + 1$ . Uncertainty in the annuity value stems from the  $q(x, \tau, i)$  component as it entails longevity risk. Sampling risk or concentration risk, as defined in Blake et al. (2019), are other important risk components which are omitted in our analysis. If we instead considered a *variable* annuity stream with expected payments of USD 1 each year, the issue of systematic longevity risk persists, likely shifted to the annuitant but not eliminated; see Ngai and Sherris (2011). For a variable annuity contract, the annuity provider has the flexibility to scale the stream of payments up or down, depending on future survival probabilities. Thus, the issue becomes whether the annuity provider or the annuitants should bear the systematic longevity risk; see Balter et al. (2020).

We assume that deaths are uniformly distributed within each year across all ages, which allows a simple transformation between end-of-year mortality rates,  $q(x, \tau, i)$ , and the central mortality rate,  $m(x, \tau, i)$ , i.e.,  $q(x, \tau, i) = \frac{m(x, \tau, i)}{1 + 0.5m(x, \tau, i)}$ ; see Pitacco et al. (2009). To model the uncertain mortality process across subgroups,  $m(x, \tau, i)$ , a large variety of multi-population models have been proposed (Enchev et al., 2017; Haberman et al., 2014; Li and Lee, 2005). The multi-population dimension is justified by the correlation in mortality improvements across and within subgroups, e.g., countries or socioeconomic groups, which stems from spillover effects, such as medical advances, improvements in public health, and other best practices (Enchev et al., 2017). The Li and Lee (2005) model captures these considerations by assuming that closely related populations share common factors, which is particularly attractive for our socioeconomic data.

Hence, we apply the Li and Lee model on observed mortality rates,  $m(x, t, i)$ , for  $N$  different subgroups as

$$\log [m(x, t, i)] = \alpha(x, i) + B(x)K(t) + \beta(x, i)\kappa(t, i) + \varepsilon(x, t, i) \text{ for } i = 1, \dots, N, \quad (1.3)$$

where  $\alpha(x, i)$  describes the general age pattern of mortality for group  $i$ , the common factor parameters  $K(t)$  and  $B(x)$  represent an index of the general level of mortality for the average population, and the age-specific response to changes in the index, respectively. The group-specific parameters,  $\beta(x, i)$  and  $\kappa(t, i)$ , are interpreted similarly to the common factor parameters above but relative to changes in them. That is, the model allows group  $i$  to deviate from the general trend in the short run. The age-dependent parameters, i.e.,  $B(x)$  and  $\beta(x, i)$ , are assumed constant over time, as is usually assumed in the actuarial literature (see Lee and Carter, 1992; Li and Lee, 2005). Thus, forecasts of future mortality rates are determined by forecasting and simulating future values of the time-dependent parameters  $K(\tau)$  and  $\kappa(\tau, i)$ . The dynamics of the common time-dependent parameter,  $K(t)$ , is assumed to be a random walk with drift, whereas the time-dependent subgroup-specific parameter,  $\kappa(t, i)$ , follows an  $AR(1)$  process; see Li and Lee (2005). The projected simulated paths of the time-dependent parameters give rise to likely future patterns of end-of-year mortality rates,  $q(x, \tau, i)$  that generate ex-post annuity values in equation (1.1).

To validate our choice of the Li and Lee (2005) model, we consider two other multi-population models: Enchev et al. (2017)'s model 0 and Cairns et al. (2019)'s CBDX model; see Appendix (1.6.1) and (1.6.2). The former poses an alternatively assumed process of  $\kappa(\tau, i)$ , while the latter assumes that mortality rates are log-linear related in the age dimension and relaxes the assumption of constant parameter values. We formally validate our model choice of the Li and Lee model via the model confidence set procedure proposed by Hansen et al. (2011). Appendix (1.6.3) shows that the Li and Lee model fits our mortality data best in- and out-of-sample.

### 1.2.1 Longevity hedging

We consider an annuity provider with the objective of hedging the value of the annuity in the decumulation phase represented by equation (1.1). The hedge is either referenced to the full population or the specific socioeconomic groups, which represents a hedge with basis risk or



without basis risk.<sup>1</sup> Irrespective of the hedging strategy chosen, the value of the hedging portfolio (see Li and Luo, 2012) is defined as

$$H(q) = w(\tau) \cdot F[q], \quad (1.4)$$

where  $w(\tau)$  is the notional amount of the hedging instruments that the annuity provider engages in with maturity  $\tau$ .  $F[q]$  is the payoff from the hedging instruments, which depends on the ex-post realizations of a vector of end-of-year mortality rates,  $q$ , for ages  $x + \tau$ . Intuitively, the hedging portfolio,  $H(q)$ , provides payoffs at future time points,  $\tau$ , conditional on realizations of  $q$ . Depending on whether the hedging instrument is referenced to the FP or SEG indices, the value of the hedging portfolio,  $H(q)$ , will differ as realizations of  $q$  across the socioeconomic dimension vary. Consequently, the benefits and costs will vary between the two hedging alternatives.

### Benefits of longevity hedging

We quantify the benefits as RRA as defined in Li and Hardy (2011)

$$RRA = 1 - \frac{\text{Var}[a(q) - H(q)]}{\text{Var}[a(q)]}, \quad (1.5)$$

where  $a(q)$  is the abbreviated form of the annuity value in equation (1.1). The RRA quantifies the percentage change in the variance of the annuity value when engaging in a hedge (the numerator) relative to not engaging in a hedge (the denominator).

Cairns (2013) provides a framework to estimate how many hedging instruments an annuity provider should optimally enter to hedge the value of the annuity. The hedger's objective is to minimize equation (1.5) using the choice variable  $w(\tau)$ . The optimal notional amount of contracts,  $w(\tau)^*$ , is given by

$$w(\tau)^* = \frac{\text{Cov}[a(q), F(q)]}{\text{Var}[F(q)]}. \quad (1.6)$$

---

<sup>1</sup>Naturally, there would be residual basis risk in a hedge referenced to socioeconomic groups, though, by definition, less population basis risk, which this paper primarily considers.

Thus, we note that the optimal weights in equation (1.6) are found by the linear regression coefficient,  $w(\tau)^*$ , by regressing the simulated annuity values,  $a(q)$ , on the simulated payoffs from the hedging instruments,  $F(q)$ . As noted in Cairns (2013), inserting the optimal weights,  $w^*(\tau)$ , into equation (1.5) gives the squared correlation

$$RRA[w(\tau)^*] = \frac{[\text{Cov}[a(q), F(q)]]^2}{\text{Var}[a(q)] \text{Var}[F(q)]}. \quad (1.7)$$

Equation (1.7) indicates that the linear dependency between the simulated annuity values and the payoffs from the hedging instrument determines the highest achievable hedge effectiveness.

### Costs of longevity hedging

In financial markets, where the underlying asset is traded, no-arbitrage theory determines that the expected value of the derivative is zero from the starting date. Under the assumption of no frictions, e.g., transaction costs and liquidity risk premiums, the payoffs from derivatives can be replicated by continuously trading the underlying asset and the risk-free rate (Black and Scholes, 1973; Merton, 1973). However, Cairns et al. (2006) argue that the underlying uncertain mortality process is currently not traded in a liquid financial market and, consequently, the payoffs of mortality- and longevity-linked securities cannot be replicated. Thus, the expected total cost of the hedging portfolio, defined in equation (1.4), will be non-zero.

As the first in the literature, we suggest measuring the total costs as

$$C = \sum_{\tau} w(\tau)^* \cdot \text{Prem}[\lambda(\tau), \tau, i] \cdot (1+r)^{-\tau}, \quad (1.8)$$

where  $w(\tau)^*$  is the optimal notional amount of the contracts with maturity  $\tau$  defined in equation (1.6) and the latter term is the present value of the risk premium for the hedging instrument,  $\text{Prem}[\lambda(\tau), \tau, i]$ . Thus, total cost can vary due to either the optimal notional amount of the hedging contracts,  $w(\tau)^*$ , or due to the risk premium,  $\text{Prem}[\lambda(\tau), \tau, i]$ .

For a given annuity provider with a specific  $a(q)$ ,  $w(\tau)^*$  is lower if the contract is referenced to higher mortality populations than lower mortality populations. To visualize this, multiply the

reference mortality rate,  $q$ , with a constant,  $c$ , in the hedging instrument,  $F(q)$ . Then, the optimal notional amount of contracts become  $\tilde{w}(\tau)^* = \frac{\text{Cov}[a(q), F(cq)]}{\text{Var}[F(cq)]} = \frac{1}{c} \cdot w(\tau)^*$ . If  $c > 1$ ,  $\tilde{w}(\tau)^* < w(\tau)^*$  and vice versa. Thus, if the mortality index is referenced to the FP mortality index, low (high) mortality groups buy fewer (more) contracts relative to when the mortality index is referenced to the subgroup-specific mortality index. The  $Prem[\lambda(\tau), \tau, i]$ , which we will explore in the next section, depends on the underlying population,  $i$ , the chosen hedging instrument, and the valuation method.

For comparability, we assume that total costs are distributed as an extra yearly required rate of return in the annuity contract. Thus, total costs is an expected fixed cost that is financed by a higher discount rate, which is equal to the alternative of not investing in the hedging contract. The higher rate of return required is denoted  $r^*$  and solves the equation

$$\sum_{\tau=1}^{x_u-x} (1+r^*)^{-\tau} p(x, \tau, i) + \sum_{\tau} w(\tau)^* \cdot Prem[\lambda(\tau), \tau, i] \cdot (1+r)^{-\tau} = \sum_{\tau=1}^{x_u-x} (1+r)^{-\tau} p(x, \tau, i). \quad (1.9)$$

The left hand side of equation (1.9) denotes the investment costs and the right hand side the alternative of not investing. The extra yearly required rate of return is represented by  $er = r^* - r$ .

## The hedging instrument and risk premiums

Several hedging instruments have been proposed in the literature to hedge longevity risk; see Blake et al. (2006) for a review. Due to its simplicity, we apply the q-forward contract introduced by Coughlan et al. (2007) as it can be used to form building blocks for other longevity-linked derivatives (Blake et al., 2019). Moreover, the actuarial risk premium for a q-forward yields a closed-form solution, as mortality risk is assumed to be log normally distributed. On the contrary, longevity-related securities cover risks that are products of log normally distributed variables, which is considerably more cumbersome; see Boyer and Stentoft (2013); Hunt and Blake (2020, 2021). The fixed rate receiver's present value of the q-forward,  $F[q]$ , is defined in Li and Luo (2012) as

$$F[q] = (1+r)^{-\tau} [q^f - q], \quad (1.10)$$

where  $q^f$  is the vector of fixed mortality rates for ages  $x + \tau$  agreed at the inception of the contract with maturity  $\tau$ , and  $q$  is the corresponding vector of realized end-of-year mortality rates at time  $\tau$  for age  $x + \tau$ . The q-forward has a positive pay-off when ex-post mortality rates are lower than expected. That is, if the realized  $q$  for age  $x + \tau$  at maturity  $\tau$  is lower than the corresponding  $q^f$ , the pension fund receives payments to compensate higher annuity liabilities and vice versa. Hence, the q-forward hedges the duration risk of the liability in equation (1.1) with respect to mortality; see Li and Luo (2012).

In contrast to other derivatives, the underlying mortality risk is not continuously traded. Consequently, the real-life and risk-neutral probability of mortality risk will differ, i.e.,  $q^f \neq \mathbb{E}_0[q]$ . Wang (2000) proposes a valuation method that can be applied to a risk or asset with any distribution and can be related to other actuarial risk-neutral transforms; see Kijima (2006); Labuschagne and Offwood (2010). The Wang (2000) transform is denoted as

$$\tilde{G}(z) = \Phi[\Phi^{-1}[G(z)] - \lambda], \quad (1.11)$$

where  $\tilde{G}(z)$  is the risk-neutral cumulative distribution function (CDF) of the asset  $Z$ ,  $\Phi$  is the standard CDF,  $G(z)$  is the real CDF, and  $\lambda$  is a distortion parameter. Taking the expected value of  $\tilde{G}(z)$ , yields the risk-neutral price of the asset, i.e., in our case  $q^f$ . Wang (2000) shows that if the asset is normally distributed, the distortion parameter,  $\lambda$ , can be interpreted as the risk-adjusted return in the capital asset pricing model (Lintner, 1965a; Sharpe, 1964), which makes it possible to price any risk depending on its correlation with aggregate risk. Following the same approach, we apply the Wang transform to the log normally distributed aggregate mortality risk and assume that it is the only relevant systematic factor for pricing individual q-forwards. In Appendices (1.6.4) and (1.6.4), we outline the aggregate mortality risk premium and the socioeconomic q-forward premiums respectively, which largely follows the results presented in Wang (2000). Thus, the risk

premium of the q-forward for the fixed-rate payer is given as

$$Prem[\lambda(\tau), \tau, i] = \mathbb{E}[q(\tau, i)] \left[ 1 - \exp \left[ -\rho_{i,M} \cdot \lambda(\tau) \cdot \sqrt{\text{Var}\{\ln(q(\tau, i))\}} \right] \right], \quad (1.12)$$

where  $\mathbb{E}[q(\tau, i)]$  is the real-life expected mortality rate at maturity  $\tau$  for age  $x + \tau$  from group  $i$ ;  $\rho_{i,M}$  is the correlation coefficient between the aggregate mortality rate; and the group  $i$ 's specific mortality rate,  $\lambda(\tau)$ , is the Wang distortion parameter, and  $\sqrt{\text{Var}\{\ln(q(\tau, i))\}}$  is the real-life standard deviation of the natural logarithm of the mortality rate at maturity  $\tau$ . Generally, the exponentiated term can be interpreted as an enlargement of the risk premium. As the last term,  $\sqrt{\text{Var}\{\ln(q(\tau, i))\}}$ , is increasing for higher mortality groups, the premium on the individual q-forwards is generally increasing for higher mortality groups. Cairns et al. (2014) suggest that an illiquidity premium might imply a higher risk premium for subgroup-specific hedges. More generally, Bates (2003) finds that frictions under hedging can distort the risk-neutral pricing of derivatives, e.g., transaction costs, discrete time hedging, and liquidity. We refrain from these considerations as the impact on mortality- and longevity-linked securities is still unobserved.

### 1.3 Affluence index and data

To measure adherence to a socioeconomic group, we use the newly developed affluence measure introduced by Cairns et al. (2019). Denoted  $A(i, t, x)$ , the index is defined as a linear combination of lagged wealth and income for individual  $i$  at time  $t$  for age  $x$

$$A(i, t, x) = W(i, t - 1, x - 1) + K \cdot Y(i, t - 1, x - 1), \quad (1.13)$$

where  $W(\cdot)$  is wealth,  $Y(\cdot)$  is income, and  $K$  is a constant, serving as a weighting factor in balancing income (flow variable) and wealth (stock variable). Intuitively,  $K$  can be thought of as the capitalization factor for retirement income in a low interest-rate environment. Thus, it should approximately equal average life expectancy at the retirement age. We follow Cairns et al. (2019) in assuming  $K = 15$  and is constant across affluence groups. The affluence measure is robust to different choices of  $K$  in the interval from 10 to 20. The gender-specific allocation procedure

assigns each individual in each year and at each age into one of  $G$  socioeconomic groups based on the individual's ranking in terms of the affluence measure. The main advantage is the ability to obtain equally sized groups for each year and age. We deliberately avoid using educational attainment as a socioeconomic indicator since it represents a suboptimal choice in the time dimension. Generally, the proportion of highly educated individuals has been higher in recent years (Bound et al., 2015), forcing education fractiles to vary over time. This creates a bias in the relationship between life expectancy and education; see Brønnum-Hansen and Baadsgaard (2012); Hendi (2015); Olshansky et al. (2012).

We use data from Kjærgaard et al. (2020), containing five affluence groups for both genders in the age interval 50 to 100 from 1985 to 2016. The five different groups are denoted group 1 (G1) to group 5 (G5), with G1 being the 20% least affluent and G5 the most. In accordance with Kallestrup-Lamb et al. (2020), we top-code the data at age 95 due to the small number of deaths at an advanced age. Furthermore, we smooth mortality rates for all ages using splines, in accordance with Wood (1994), and apply Hyndman et al. (2012)'s unconstrained regression splines, which smooths the mortality rates so they increase monotonically with age. The unconstrained regression splines are equivalent to the weighted penalized regression splines with a monotonicity constraint when setting the lower age of increasing monotonicity on the mortality rates equal to the lowest age in the data sample.

Fig. (1.1) shows the period life expectancy at age 67 for males and females across socioeconomic groups.<sup>2</sup> Age 67 is a natural choice as it represents the current statutory retirement age in Denmark, which will remain the same until 2030, at which point it will be linked to developments in life expectancy (Danish Ministry of Economic Affairs and Interior, 2018). The target of this indexation is an expected retirement period of 14.5 years for the average individual; see Andersen (2015).

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<sup>2</sup>Life expectancy is approximated as the partial period life expectancy, as in Cairns et al. (2019), i.e.,

$$LE(t, x) \approx \frac{1}{2} + \sum_{y=x+1}^{x_u-1} S_p(t, x, y) + \frac{1}{2} S_p(t, x, x_u)$$

where  $S_p(t, x, y) = \exp \left[ - \sum_{s=x}^{y-1} \hat{m}(t, s) \right]$  is the observed period survival probabilities and  $x_u = 95$  is the maximum age.

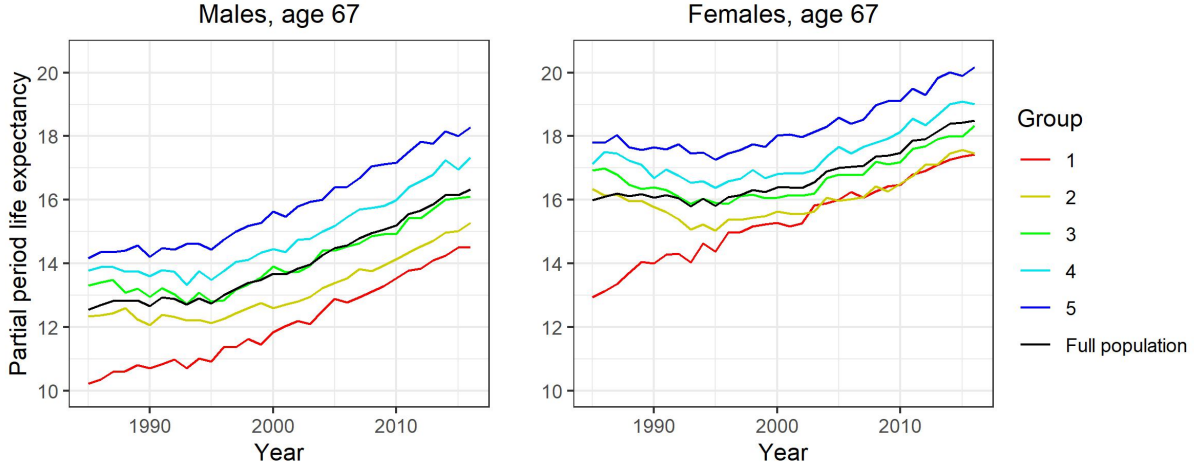


Fig. 1.1: Partial period life expectancy for all groups, the full population, and both genders.

The affluence measure obtains a clear and detailed partitioning between the five subgroups in terms of life expectancy, from the lowest to the highest affluence group. A consistent gap of four years is found across time for 67-year-old men. For women of the same age, the gap has been decreasing over time from five to three years due to different improvement rates across socioeconomic groups. We observe a stagnation in female life expectancy for groups G2 to G5 until 1995, which is related to cancer and smoking-related lung and bronchial issues (Kallestrup-Lamb et al., 2020). The least affluent group, G1, experiences continued improvements throughout the entire period. After 1995, life expectancy has been increasing for all groups at a similar pace. Fig. (1.1) clearly shows that disaggregating by socioeconomic status reveals heterogeneity with different mortality patterns compared to the aggregated level.

In the modeling framework, we assume that G3 can serve as a proxy for the mortality experience of the FP, which seems reasonable based on the overlap observed in Fig. (1.1). This assumption is based on the fact that when modeling G3 in deviation from the FP using the Li and Lee model, the subgroup-specific parameters,  $b(x, 3)$  and  $k(t, 3)$ , become difficult to estimate as there is very limited variation left to explain. The same argument holds for the Enchev model examined in Appendix (1.6.2).

## 1.4 Results

### 1.4.1 Benefits of hedging longevity risk

Benefits are quantified by the RRA using equation (1.7), under the assumption that the annuity provider has a liability as defined in equation (1.1). Li and Luo (2012) show that the optimal risk reduction can be achieved if the annuity provider engages in five evenly spaced q-forwards. We follow this approach and assume that the distance between maturities is five years, i.e.,  $\tau = \{5, 10, 15, 20, 25\}$ ; thus, the indexed ages for the q-forwards are  $x + \tau$ , where  $x = 67$ . The interest rate is assumed fixed and equal to  $r = 0.02$  per annum, while the mortality projections are based on the Li and Lee (2005) model from equation (1.3). The fixed forward rate,  $q^f$ , is indexed to either the FP or the individual SEGs. Table (1.1) shows RRAs for all five q-forwards for each gender. For each socioeconomic group, the RRA depends on the reference index, thus allowing for a comparison between the two alternatives.

Maturity		5	10	15	20	25	5	10	15	20	25
Policyholder	Index	Males					Females				
G1	FP	0.29	0.55	0.67	0.69	0.70	0.07	0.14	0.18	0.20	0.20
	SEG	0.38	0.70	0.89	0.96	0.98	0.17	0.45	0.75	0.94	0.98
G2	FP	0.40	0.76	0.92	0.95	0.96	0.17	0.30	0.40	0.43	0.43
	SEG	0.41	0.78	0.95	0.98	0.98	0.22	0.48	0.73	0.92	0.98
G3	FP	0.39	0.76	0.94	0.98	0.98	0.35	0.68	0.88	0.97	0.98
	SEG										
G4	FP	0.30	0.65	0.86	0.90	0.90	0.10	0.21	0.28	0.32	0.32
	SEG	0.37	0.74	0.93	0.98	0.98	0.16	0.38	0.64	0.90	0.98
G5	FP	0.18	0.41	0.56	0.59	0.59	0.08	0.17	0.25	0.30	0.31
	SEG	0.34	0.70	0.91	0.97	0.98	0.13	0.32	0.59	0.88	0.99

Table 1.1: Relative risk adjustments (RRA) for all subgroups using five q-forwards. Column 1 lists the five groups and column 2 the reference group. Each gender has five columns, which represent the incremental cumulative increase in RRA using the Li and Lee (2005) model for the five q-forward contracts. Each group has two rows. The top row in each group quantifies the RRA with basis risk, where the reference group is the full population (FP), while the bottom row shows the RRA without basis risk, where the reference group is the specific socioeconomic group (SEG).

Hedging longevity risk with the q-forward referenced to the FP shows a general pattern consistent for both genders, namely that the RRA more or less decreases the further away the group



is from average life expectancy. However, there is substantial variation in RRA across different groups and between the genders. For males, RRA is higher than the equivalent female groups. For instance, the two groups closest to the median, G2 and G4, have an RRA of 96 and 90% for the males and 43 and 32% for the females, respectively. This is a rather large discrepancy and illustrates the challenges involved in longevity hedging when using the mortality experience of the FP as the underlying index. The reason for the difference between males and females can be explained by the differences in mortality improvements experienced in the socioeconomic groups. As Fig. (1.1) shows, life expectancy for males comoves across time, whereas females have experienced different mortality improvements at the start of the sample period. This is explained by the stagnation in mortality improvements experienced until 1995 for females above 60 years in all groups except G1 (Kallestrup-Lamb et al., 2020). As the projected mortality rates mirror the experienced mortality rates in terms of variation across the groups, the discrepancy experienced for the female groups leads to lower hedge effectiveness when using FP as the underlying index. More specifically, the magnitude of basis risk is determined by  $\beta(x, i)$ , the estimated relative slope of the mortality curve across the age dimension. The  $\beta(x, i)$  estimates are around zero in the retirement stage for males, whereas it increases in the age dimension for females; see Fig. (1.7) and (1.8) in Appendix (1.6.1).

In contrast, hedging directly on the SEG-specific index diminishes the variability, which amounts to an RRA of 98-99% using all five q-forward contracts, irrespective of the group. Thus, all groups have higher RRAs for this strategy, e.g., males in G1 achieve an RRA that is 28 percentage points higher (98 versus 70%) when using all five q-forwards indexed to a specific SEG. On average across all socioeconomic groups, we obtain RRAs that are 20 and 65 percentage points higher for this alternative for males and females, respectively. Cairns et al. (2014) find equivalent differences in RRAs of approximately 30 percentage points between an insured and full male population from England and Wales. Moreover, Table (1.1) clearly shows that there is a substantial gain to be had from hedging with a q-forward indexed to the specific SEG relative to being indexed to FP. For instance, males in G1 attain the same cumulative RRA of 70% using only the first two q-forwards indexed to the specific SEG compared to using all five q-forwards indexed to

FP.

If we consider alternative mortality models, RRA indexed to FP differ somewhat from the results in Table (1.1). With the Enchev model, RRAs are slightly higher than in Table (1.1), whereas for the CBDX model, they are approximately 25-30% across groups. The RRAs indexed to the specific SEGs, in contrast, are almost identical to Table (1.1). Appendix (1.6.5) contains the results.

#### **1.4.2 Cost of hedging longevity risk**

To illustrate the financial magnitude of hedging longevity risk, we calculate the expected costs for a hedge with and without basis risk. We focus specifically on risk premiums arising from exposure to the systematic risk factor of aggregate mortality risk and disregard other probable factors, e.g., illiquidity and transaction costs. Additionally, we explore the total costs, defined in equation (1.8), which is the risk premium multiplied by the optimal notional amount of the contracts,  $w(\tau)^*$ .

For all q-forwards, we apply the actuarial risk premium in equation (1.12), assuming that the Wang distortion parameter is  $\lambda_\tau = \sqrt{\tau} \cdot 0.18$  for males and  $\lambda_\tau = \sqrt{\tau} \cdot 0.23$  for females (Lin and Cox, 2005). The parameters in Lin and Cox (2005) are estimated on the U.S. life annuity market and applied to a 30-year survivor bond. As mentioned in Section (1.2.1), the Wang distortion parameter can also be interpreted as the risk-adjusted return demanded from the investor.

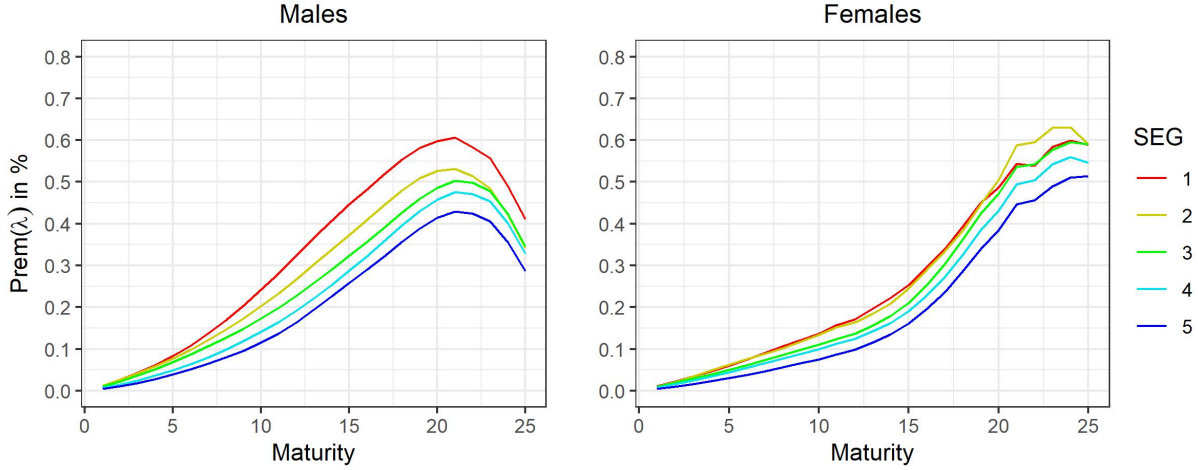


Fig. 1.2:  $q$ -forward premiums for all socioeconomic groups (SEG) by maturity. On the vertical axis, the premiums are measured as the percentage-point difference between the  $q$ -forward rate,  $q^f$ , and the expected mortality,  $\mathbb{E}[q]$ . The premiums vary on the horizontal axis depending on maturity  $\tau$  for ages  $x + \tau$ .

In Fig. (1.2), we see that the risk premiums covary positively with mortality rates, i.e., the high mortality groups have an added risk premium in excess of the aggregate mortality risk premium, whereas the opposite is true for low mortality groups. The clear ranking is explained by the sensitivity of the price movements in the  $q$ -forwards. The price sensitivity is larger for high mortality groups compared to low mortality groups, which equation (1.12) clearly shows. The aggregate mortality risk premium is represented by the green line, which approximates the mortality experience of FP. Barrieu and Veraart (2016) show a pattern similar to the one seen for the FP proxy for females, G3, in Fig. (1.2). That is, the aggregate risk premium increases until 20 years of maturity, after which the growth in premiums dampens.<sup>3</sup>

Next, we consider the optimal notional amount,  $w(\tau)^*$ , as the last component of total costs. To simplify, we show the total notional amounts for all five  $q$ -forward contracts for the two hedging alternatives in Fig. (1.3).

<sup>3</sup>Barrieu and Veraart (2016) estimate the  $q$ -forward risk-neutral price,  $q^f$ , using four different pricing rules on the England and Wales' male population. Our risk premium is the absolute difference between  $q^f$  and the expected mortality rate at maturity and can be estimated after replicating  $q^f$  in Barrieu and Veraart (2016). We do not have the replication in this paper but the claim can easily be verified using either our deducted risk premium in equation (1.12) with  $\rho_{i,M} = 1$  or their standard deviation pricing rule (PR2) which is approximately the same pricing rule as ours for aggregate mortality risk.

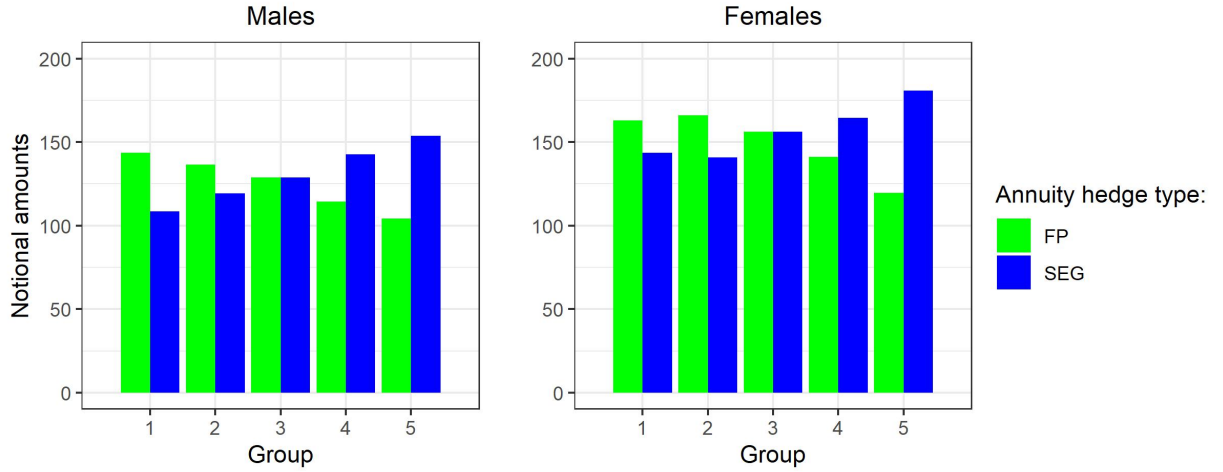


Fig. 1.3: Total notional amounts for the five q-forward contracts. For each socioeconomic group, there are two categories of bar plots (full population (FP) and socioeconomic group (SEG)). The green bar (FP) represents the total notional amount for all five q-forward contracts when the hedge is referenced to the median G3, i.e., with basis risk. The blue bar (SEG) is the total notional amount for all five q-forward contracts when the hedge is referenced to the specific SEGs for all five q-forwards, i.e., without basis risk.

The optimal total notional amount of contracts depends on the underlying reference index. When the index is referenced to a specific SEG, represented by the blue bars in Fig. (1.3), the notional amount increases with higher affluence groups. This is intuitive since the notional amount is determined by the sensitivity between the payoff of the q-forward and the value of the life annuity based on the same underlying mortality rates for a specific group. As a life annuity for higher affluence groups pays out for a longer horizon, small deviations in the mortality rates for earlier maturities affect the value of the annuity to a greater extent. This is analogous to duration matching in the fixed-income literature (see Li and Luo, 2012; Ho, 1992). Thus, if a bond has a higher duration, a larger notional amount in an interest-rate swap is required. When the index is referenced to FP green bars in Fig. (1.3), the notional amount of contracts decreases with higher affluence groups. Under this alternative, the notional amount is based on the mortality rates represented by FP and the mortality rates underlying the life annuity of the specific group. Consistent with Fig. (1.2), higher affluence groups are less sensitive to changes in the mortality rates in FP than lower affluence groups. Thus, the optimal notional amount is inversely related to the SEG-specific risk premium. For example, if G1 hedges with the q-forward referenced to FP, the notional amount required is higher but at a lower price, cf. Fig. (1.2). If, in contrast, the

q-forward is referenced to SEG, the notional amount required is lower but at a higher price. The exact relationship can be seen by comparing the risk premium equation (1.33) in Appendix (1.6.4) with the optimal notional amount in equation (1.6) in Section (1.2.1), i.e., the high  $\beta_{i,M}$  exposure for high mortality groups is offset by lower notional amounts of  $w(\tau)^*$  by the same magnitude.

As shown in Table (1.1), hedging with a q-forward referenced to the specific SEG leads to a higher RRA than when referenced to FP. Thus, Fig. (1.4) presents the yearly expense ratio, as defined in equation (1.9), for three different cases: 1) the green bar shows the expense ratio when the q-forward is referenced to FP using all five q-forwards, 2) the light blue bar (SEG: RRA Min) represents the expense ratio for attaining a RRA at least as high as with the FP hedge, e.g., two q-forwards for G1 males, and 3) the dark blue bar (SEG: RRA Max) represents the extra expense ratio to attain the maximum RRA for this hedging alternative using all five q-forward contracts.

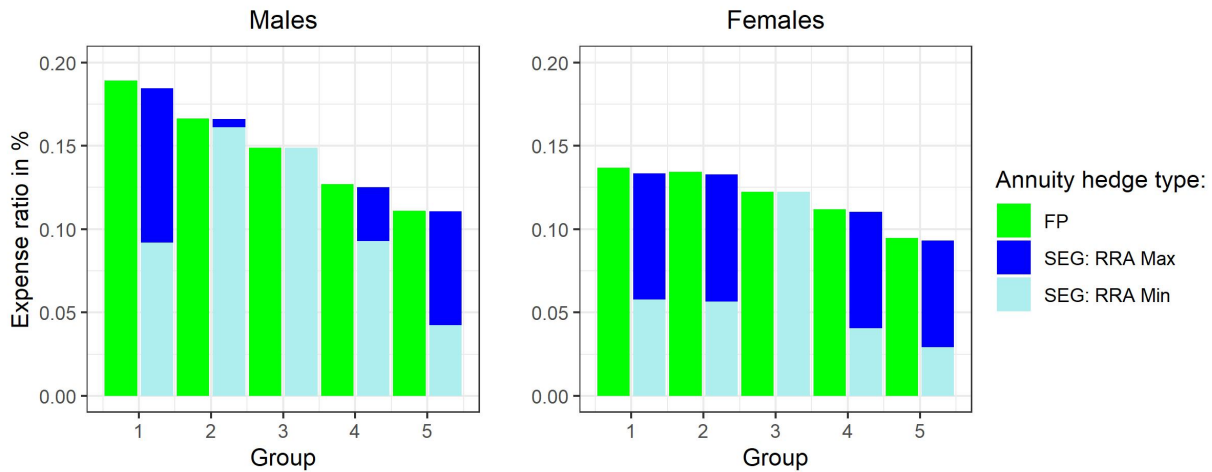


Fig. 1.4: Annual expense ratios for hedging longevity risk. For each group, there are three categories of bar plots: FP, SEG: RRA Min, and SEG: RRA Max. The green bar (full population (FP)) is the expense ratio by hedging referenced to the median group 3 for all five q-forwards, i.e., with basis risk. The light blue bar (socioeconomic group (SEG), SEG: RRA Min) is the expense ratio for attaining an relative risk adjustment (RRA) at least as high as with the FP hedge. The dark blue bar (SEG: RRA Max) is the expense ratio when referenced to the specific SEGs for all five q-forwards, i.e., without basis risk.

When hedging with basis risk (FP), the green bars indicate that the expense ratio is generally lower for higher affluence groups, with an expense ratio of approximately 0.1% in G5. The expense ratio for the lowest affluence group is considerable higher, with roughly 0.18% for males and 0.13% for females. As the risk premium is constant across socioeconomic groups, it follows

that the optimal notional amount of q-forwards bought is decreasing for higher affluence groups when the hedge is referenced to FP; see Fig. (1.3). The magnitude of the expense ratios, imply that hedging longevity risk would be moderately higher than investing in passively managed bond and equity mutual funds, which have on average annual expense ratios of 0.12% and 0.06%, respectively (Duvall and Johnson, 2022).

When the RRA is held constant across the hedging strategies, hedging without basis risk is clearly cheaper, cf. Fig. (1.4), light blue bar (SEG: RRA Min). For females, hedging without basis risk leads to an expense ratio less than half the size of hedging with basis risk. For males in G1 and G5 the same pattern applies. For males, the groups closest to the median, i.e., G2 and G4, the gain of hedging without basis risk is less prevalent. The hedge referenced to specific SEGs using all five q-forwards allows for a significantly higher hedge effectiveness across all groups and genders, cf. Table (1.1). Surprisingly, this is without increasing the cost of the hedge (SEG: RRA Max) compared to a hedge referenced to the FP. Thus, irrespectively of whether or not the hedge is referenced to FP or SEG when using all five q-forwards, the total costs remain the same but with lower RRAs for the FP hedge. The reason for this is that the optimal notional amount,  $w(\tau)^*$ , and the risk premium operate in opposite directions. Informally, comparing Fig. (1.2) and (1.3) shows this. Appendix (1.6.4) contains a more detailed analysis.

Considering alternative mortality models, we confirm the results in Fig. (1.4) related to the green bar (FP) and dark blue bar (SEG: RRA Max); see Appendix (1.6.7). In contrast, the light blue bar (SEG: RRA Min) is, to some extent, dependent on the assumed mortality model. This follows from the differences in RRAs when indexed to FP across the alternative mortality models, as shown in Appendix (1.6.5), as well as the differences in socioeconomic risk premiums, as shown in Appendix (1.6.6).

When engaging in all five q-forwards, the expense ratio is the same for the assumed mortality models. Consequently, any deviation from our expense ratios ought to arise from other components constituting the risk premium. As seen in equation (1.12), or equivalently, in equations (1.30) and (1.33) in Appendix (1.6.4), the risk premium is increasing in the Wang distortion parameter,  $\lambda(\tau)$ . However, this has generally been utilized in the literature (cf. Barrieu et al.,

2012; Boyer and Stentoft, 2013). Thus, other components not considered in this paper may distort the expense ratio to be higher. This could include compensation for holding illiquid assets and for transaction costs, e.g., by higher frequency dynamic hedging, or alternatively, by integrating parameter uncertainty into the risk premium. We leave those concerns for future research.

## 1.5 Concluding remarks

Life annuity providers are exposed to longevity risk; however, measures have been initiated to establish a financial market to mitigate this risk. At present, capital market solutions remain relatively underused compared to insurance-based solutions. The main impediment for capital market solutions is believed to be subpopulation basis risk. This is the main motivation for examining basis risk across different socioeconomic groups in this paper.

We find that the estimated basis risk varies between genders and subgroups. Generally, basis risk is higher the further away the group is from the median group, assumed to proxy the full population. Thus, mitigating longevity risk through hedges indexed to the full population is suboptimal. Referencing the mortality-linked security to the specific socioeconomic group, in contrast, yields high hedge effectiveness irrespective of the group. Quantifying basis risk in terms of longevity risk reduction within a population only focuses on the benefits of longevity risk reduction. Another important aspect is at what cost. Under the assumption that capital or reinsurance markets are efficient, i.e., buyers of  $q$ -forwards can diversify the residual risk after taking aggregate mortality risk into account, the hedges indexed to specific socioeconomic groups are more cost-effective. Moreover, higher hedge effectiveness can be achieved without any extra costs compared to a hedge referenced to the full population. The reason is that risk premiums for socioeconomic groups operate in the opposite direction of the optimal notional amount of hedging contracts. Thus, hedges indexed to a specific subgroup that has high (low) risk premiums optimally buy fewer (more) hedging contracts.

Our study suggests that hedging longevity risk via capital market securities, indexed to the mortality experience of socioeconomic groups, significantly reduce basis risk. Moreover, expenses related to implementing such a hedging solution is comparable to investing in passive equity funds. Disregarding capital market solutions and assuming status quo, exposes pension funds to aggregate longevity risk in a global insurance industry with insufficient risk capacity.



## 1.6 Appendix

This appendix outlines the actuarial risk premium, the fitted parameters, and the forecasted mortality rates for all of the mortality models examined, the risk premiums, and the hedge effectiveness using Enchev et al. (2017)'s model 0 and Cairns et al. (2019)'s CBDX model, as well as the result from the Hansen et al. (2011) model confidence set (MCS) procedure.

### 1.6.1 Fitting and forecasting Li and Lee- and Enchev models

The Li and Lee (2005) model is denoted as

$$\ln [m(x, t, i)] = a(x, i) + B(x) K(t) + b(x, i) k(t, i) + \varepsilon(x, t, i) \quad \forall i = 1, \dots, N \quad (1.14)$$

$$K(t) = \mu_K + K(t-1) + \xi(t), \quad \xi \sim N(0, \sigma_{LC}^2) \quad (1.15)$$

$$k(t, i) = \mu_k(i) + \phi(i) k(t-1, i) + \zeta(t, i), \quad \zeta(i) \sim N(0, \sigma_{LL}^2) \quad (1.16)$$

$$\forall k(t, i) = \Phi k(t-1, i) + Z(t), \quad Z \sim MVN(0, \Sigma), \quad (1.17)$$

where equation (1.14) is the fitted Li and Lee model and equation (1.15) is the assumed dynamics for the FP time-specific parameter,  $K(t)$ . Equations (1.16) and (1.17) are the assumed dynamics for respectively, the originally Li and Lee (2005) model and the Enchev model (Enchev et al., 2017, model 0). The specification in equation (1.17) was introduced in Enchev et al. (2017) since the Li and Lee assumption might produce non-coherent and diverging mortality rate forecasts.

$m(x, t, i)$  is the crude mortality rates, which are equal to  $\frac{D(x, t, i)}{E(x, t, i)}$ , where  $D(x, t, i)$  is the number of deaths at age  $x$ , at time  $t$ , for population  $i$ . We assume that  $m(x, t, i)$  is approximately equal to the underlying force of mortality,  $\mu(x, t, i)$ , i.e., the force of mortality is approximately constant over each age  $x$  and time  $t$ .

Since we follow the method put forward in Currie (2016) for fitting, we treat  $D(x, t, i)$  as a poisson-distributed random variable with the rate parameter  $m(x, t, i) \cdot E(x, t, i)$ , making it possible to infer equation (1.14) as a generalized non-linear model. Following Enchev et al. (2017), we use

the identifiability constraints

$$\sum_x b(x, i) = 1$$

$$\sum_t k(t, i) = 0.$$

We also follow Li and Lee (2005) in setting

$$a(x, i) = \frac{\sum_t \ln[m(x, t, i)]}{T + 1}.$$

Fig. (1.5) shows the base mortality rate parameter  $a(x, i)$ .

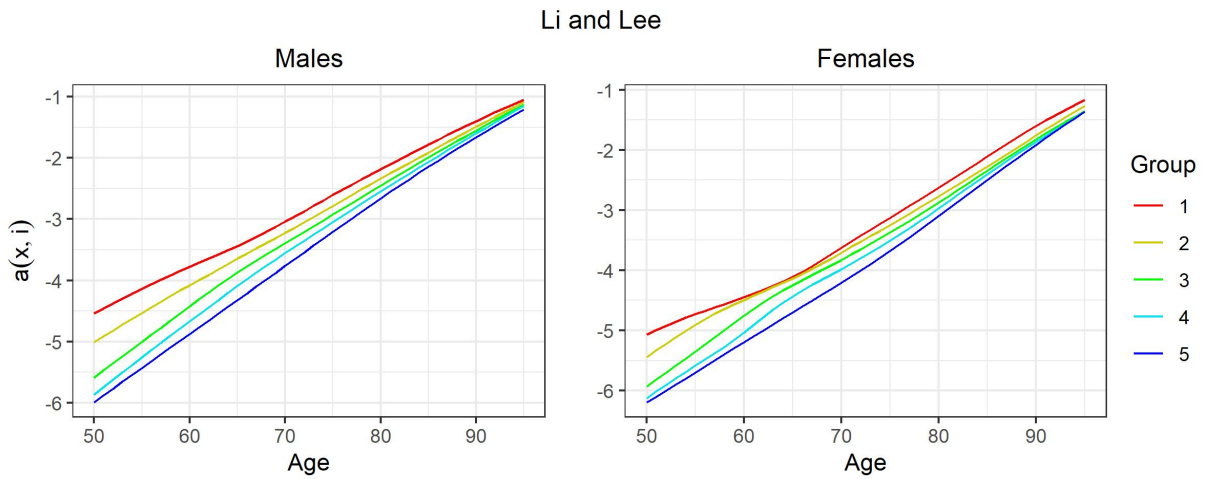


Fig. 1.5: Base-rate mortality rates,  $a(x, i)$ , for all groups and both genders. The horizontal axis illustrates the age for the groups and the vertical axis displays  $a(x, i)$ .

As usual, the base mortality rate increases log linearly with age. The  $a(x, i)$  is increasing for a given age in terms of the affluence index. The difference in log-mortality shrink with age and leads to late-life convergent mortality rates.

Fig. (1.6) illustrates FP's parameter estimates, i.e.,  $B(x)$  and  $K(t)$ .

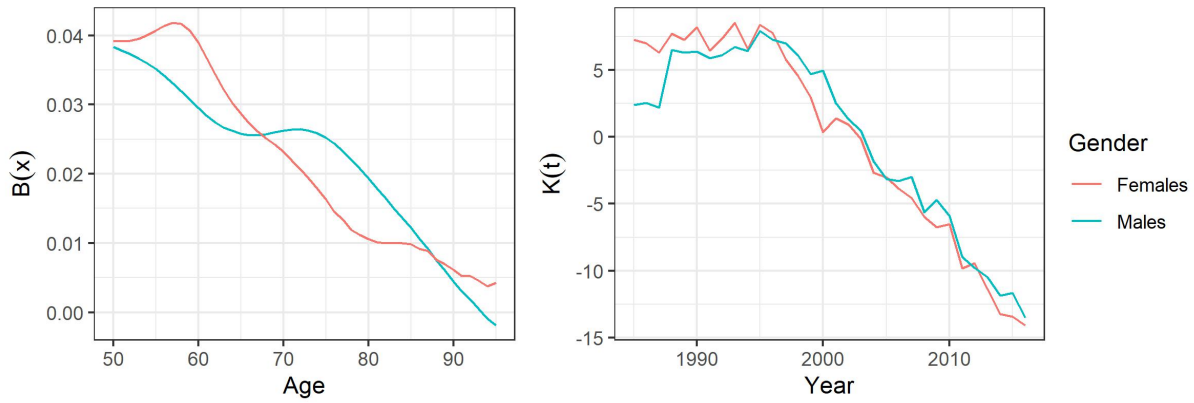


Fig. 1.6: Full population (FP) estimates for the relative slope parameter,  $B(x)$ , and the time dependent parameter  $K(t)$ . The vertical axis displays  $B(x)$  and  $K(t)$ , and the horizontal axis displays the age and year for the two FP parameters, respectively.

The parameter estimates for both genders follow each other closely. However, the male FP estimate of the relative slope parameter,  $B(x)$ , converges close to zero for the elderly part of the population. Since the time-dependent parameter  $K_t$  is downward sloping, which reflects the mortality improvements experienced during the data period, the male parameter estimate close to zero for  $B(x)$  reflects that the general mortality improvements have been less prevalent in this part of the population.

Fig. (1.7) illustrates the relative slope of the mortality curve and the time-dependent parameter for the males that is not explained by the trends in the FP.

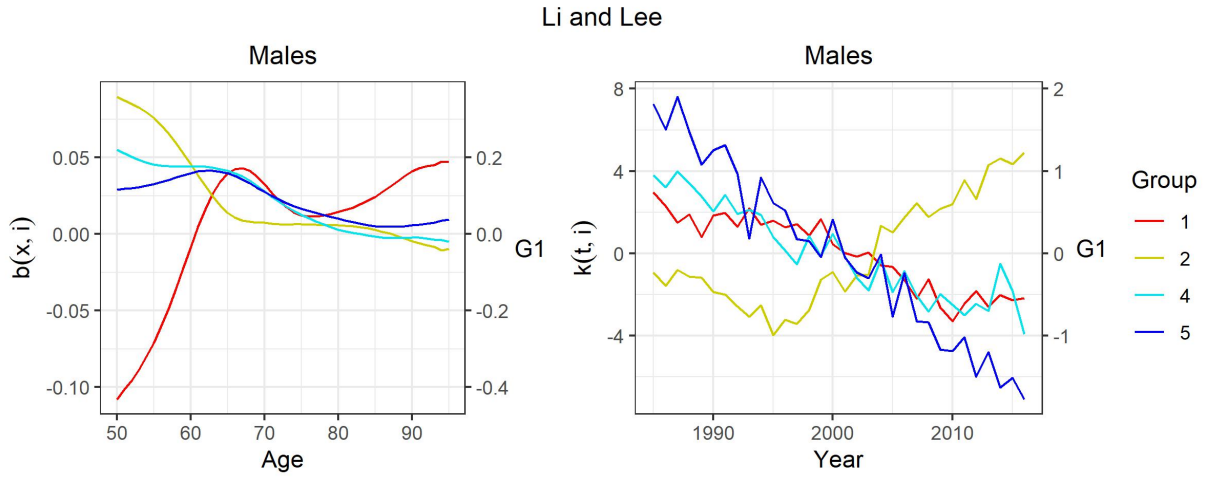


Fig. 1.7: Relative slope of the mortality curve,  $b(x, i)$ , and subgroup-specific, time-dependent parameter,  $k(t, i)$  for males and all subgroups. The vertical axis displays  $b(x, i)$  and  $k(t, i)$ , and the horizontal axis displays the age and year for the subgroup-specific parameters, respectively.

The lowest socioeconomic group, G1, has larger variation in the relative slope of the mortality curve,  $b(x, i)$ , than the rest of the groups. Thus, the relative slope of the mortality curve,  $b(x, 1)$ , is depicted on a separate, secondary axis in the panel to the right in Fig. (1.7). Coupled with the downward slope of the subgroup-specific time-dependent parameter  $k(t, 1)$ , the ages with a positive  $b(x, 1)$  indicate the ones that have experienced mortality improvements in excess of G3. In similar terms, the affluent groups, i.e., G4 and G5, have experienced mortality improvements as depicted by the downward sloping  $k(t, i)$ . However,  $b(x, i)$  is largest for younger ages, which indicates that mortality improvements in excess of G3 have happened for the younger age groups in the sample.

Fig. (1.8) shows the subgroup-specific parameters for females.

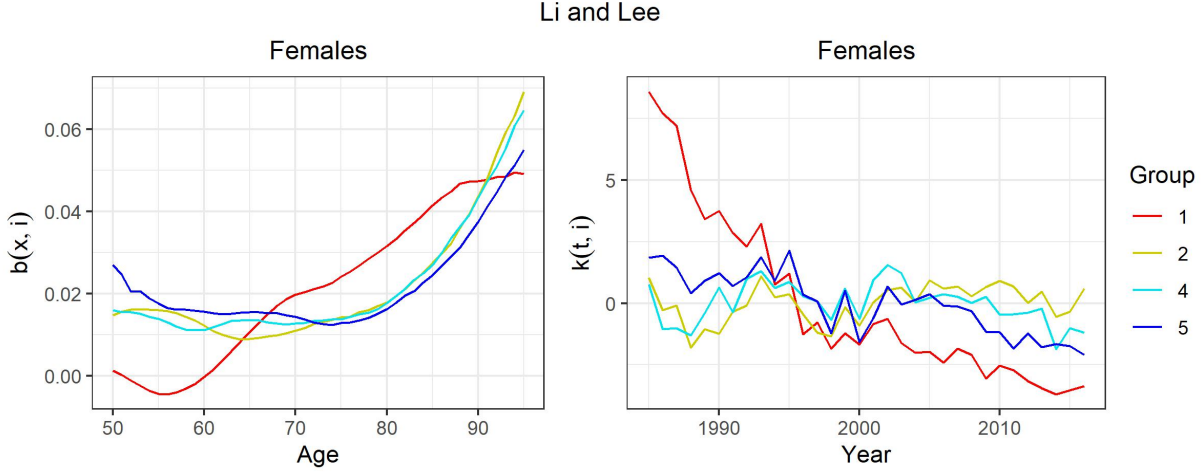


Fig. 1.8: Relative slope of the mortality curve,  $b(x, i)$ , and subgroup-specific, time-dependent parameter,  $k(t, i)$  for females and all subgroups. The vertical axis displays  $b(x, i)$  and  $k(t, i)$ , and the horizontal axis displays the age and year for the subgroup-specific parameters, respectively.

Fig. (1.8) shows that in all groups with the exception of G1, the subgroup-specific time-dependent parameter  $k(t, i)$  is relatively flat. This indicates that all groups, excluding G1, have experienced equivalent mortality improvements compared to G3. The negative slope of  $k(t, i1)$  and the positive slope of  $b(x, 1)$  for G1 indicates that the latter age groups have experienced mortality improvements, as argued earlier (Fig. (1.1)).

Looking at forecasts, we present the two specifications used in this analysis, i.e., varying assumptions regarding the modeling assumption of  $k(t, i)$ . Equation (1.16) shows the original Li and Lee assumption, while equation (1.17) represents the Enchev model assumption. For former, we set  $\phi = 1$ , since for three groups, the parameter is indistinguishably different from 1 and cannot be estimated.

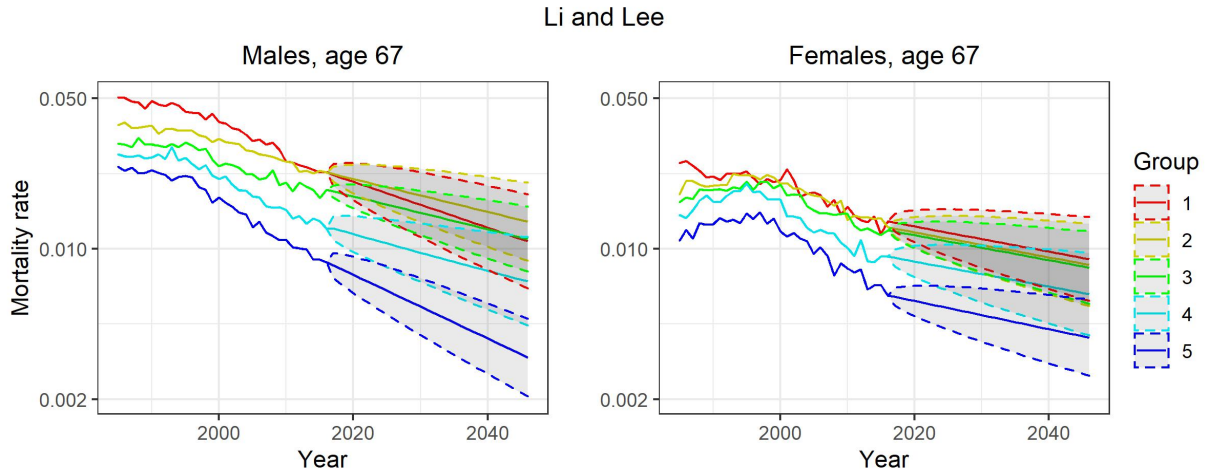


Fig. 1.9: Forecasted mortality rates at age 67 for all groups and both genders using the Li and Lee assumption. The mortality rates, on the vertical axis, are rescaled to a logarithmic scale. The horizontal axis contains the in- and out-of-sample years. The uncertainty span contains the 95% confidence interval.

Fig. (1.9) illustrates the forecasted mortality rates at age 67 using the Li and Lee assumption. The mortality trajectory for G1 and G2 males follows each other closely at the end of the data sample, leading them to cross over in the forecast, which is undesirable. For females, the mortality trajectory for groups 1-3 follow each other closely without crossing over.

Fig. (1.10) forecasts mortality rates using the Enchev model assumption for age 67.

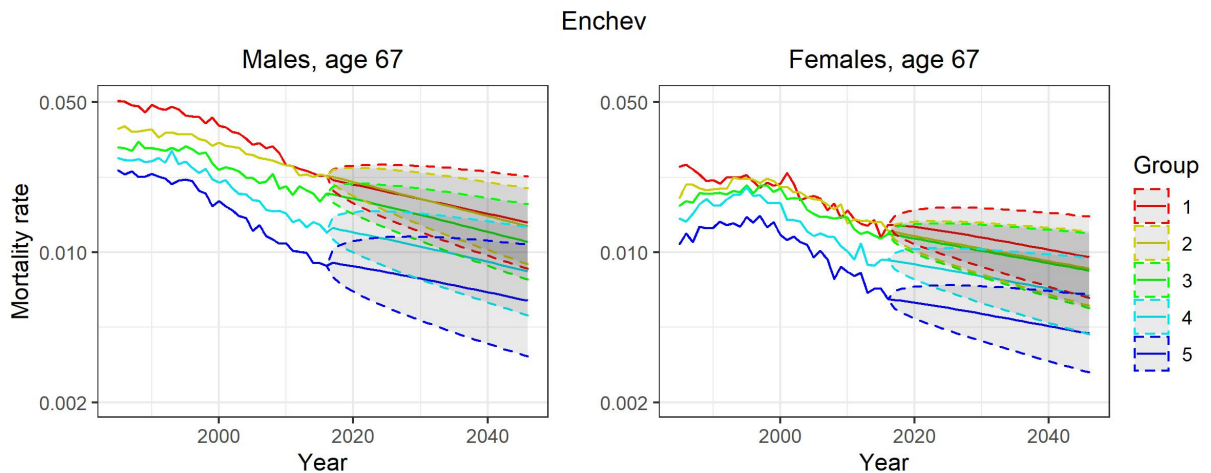


Fig. 1.10: Forecasted mortality rates for all groups and both genders using the Enchev assumption at age 67. The mortality rates, on the vertical axis, are rescaled to a logarithmic scale. The horizontal axis contains the in- and out-of-sample years. The uncertainty span contains the 95% confidence interval.

Now, G1 and G2 males are still close to each other, however without crossing over. Similarly,

G1-G3 females are still close to each other without crossing over.

### 1.6.2 Fitting and forecasting the CBDX Model

To supplement the popular multi-population Li and Lee and Enchev models, we show the fitting and forecast of the CBDX model, as introduced in Cairns et al. (2019). The CBDX model is denoted as

$$\log [m(x, t, i)] = \beta_0(x, i) + \kappa_1(t, i) + \kappa_2(t, i) \cdot [x - \bar{x}] \quad \forall i = 1, 2, \dots, N \quad (1.18)$$

$$\begin{aligned} \kappa_1(t, i) &= \kappa_1(t-1, i) - \psi[\kappa_1(t-1, i) - \bar{\kappa}_1(t-1)] + \mu_1 + Z_{1i}(t) \\ \kappa_2(t, i) &= \kappa_2(t-1, i) - \psi[\kappa_2(t-1, i) - \bar{\kappa}_2(t-1)] + \mu_2 + Z_{2i}(t), \end{aligned} \quad (1.19)$$

where  $\beta_0(x, i)$  is the base-rate mortality for subgroup mortality rates, which according to Cairns et al. (2019) preserves the mortality ranking between groups.  $\kappa_1(t, i)$  captures changes in levels of log mortality, whereas  $\kappa_2(t, i)$  models changes in the slope of the log mortality curve relative to the base-rate mortality. The dynamics from the time-dependent variables,  $\kappa$ , are bivariate random walks with common intercept  $\mu$  and subgroup specific innovations,  $Z_i$ .  $\bar{\kappa}$  is the mean of the group's  $\kappa$  processes at time  $t$ , i.e.,  $\bar{\kappa}_j(t) = \frac{1}{N} \sum_{i=1}^N \kappa_j(t, i) \quad \forall j = 1, 2$ . Thus,  $\bar{\kappa}_1$  and  $\bar{\kappa}_2$  can be interpreted as the mortality improvement on the national population.  $\kappa_1(t, i)$  and  $\kappa_2(t, i)$  mean revert to the national population time-dependent variables  $\bar{\kappa}_1$  and  $\bar{\kappa}_2$  by the parameter  $\psi$ , which is equal for both processes. Hence,  $\psi$  determines how much the individual time-dependent parameters can diverge from the national population's time-dependent parameters,  $\bar{\kappa}_1$  and  $\bar{\kappa}_2$ . Cairns et al. (2019) termed this the gravity effect. The random innovations  $Z_i$  are multivariate normal, with the mean equal to 0 and with a covariance matrix with equal correlation between the two  $\kappa$  processes for all groups but group-specific variances.

Overall, we apply the same hyperpriors on the hyperparameters as in Cairns et al. (2019). An important exception is the hyperprior on  $\psi$ , which in Cairns et al. (2019) is assumed to be  $\psi \sim \text{Beta}(2, 2)$ . The model fit with  $\psi \sim \text{Beta}(2, 2)$  leads to a proper posterior and model fit. However, when examining the mean of the predictive posterior distribution for  $\kappa$ , this leads to

exponentially decreasing and increasing  $\kappa$  for  $\kappa_1$  and  $\kappa_2$ , respectively. Thus, we assume that  $\psi \sim U(0, 1)$ , which keeps the mean of the predictive posterior of  $\kappa$  (almost) linear.

We fit the model and then forecast it using RSTAN (Stan Development Team, 2021). The model features hyperpriors,  $\psi$ ,  $Z_i$ .  $Z_i$ , consisting of the correlation,  $\rho$ , which is assumed equal for all  $\kappa_j(t, i)$  processes, and the subgroup-specific variances for the  $\kappa_j(t, i)$  processes. We experiment with the hyperprior only influences the prior on the  $\kappa_j(t, i)$  processes. Thus, the prior on  $\beta_0(x, i)$  can be freely set. Since the model requires estimating  $N \times X \times T$  parameters, in addition to the hyperpriors, the model is highly susceptible to model misspecification, multi-modality, and collinearity. To alleviate this, we set a strict prior on  $\beta_0(x, i) \sim N(\log \hat{m}(x, 1985, i), 0.005)$ , where  $\hat{m}(x, 1985, i)$  is the crude mortality rates for the ages  $x$  in group  $i$  in year 1985. This deviates from Cairns et al. (2019), where  $\beta_0(x, i)$  has a uniform prior. We impose the strict prior on  $\beta_0(x, i)$  to gain computational efficiency and to get proper model fit. The parameter estimates are close to that of Cairns et al. (2019).

Since the CBDX model is fitted by Bayesian inference, we can construct two different types of forecasts, one with (WPU) and one without parameter uncertainty (NPU). If we forecast WPU, we include the uncertainties arising from the hyperparameters and all parameters. If we forecast NPU, we assume that the estimates of the parameters are known after fitting the model, and the only uncertainty in the forecast arises in the different draws of nature from  $Z_{1i}$  and  $Z_{2i}$  at a future time point.

Below is the fitted base-rate mortality parameter,  $\beta_0(x, i)$ , which has a strict prior centered around the 1985 natural logarithm of crude mortality rates.



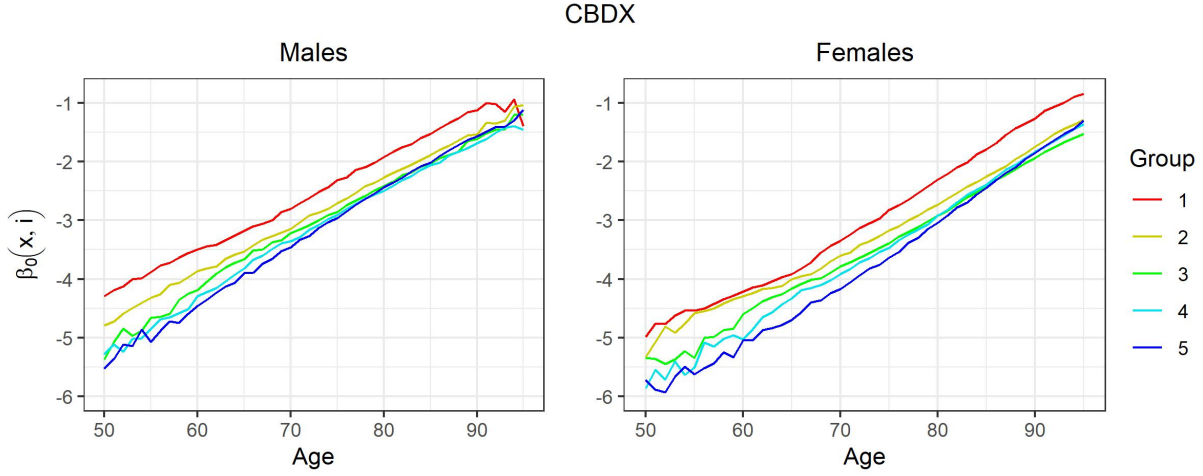


Fig. 1.11: Fitted base-rate mortality for both genders and all socioeconomic groups . For each gender, the five groups are abbreviated by G1-G5. The horizontal axis illustrates the age for the groups and the vertical axis displays  $\beta_0(x, i)$ .

Fig. (1.11) shows that the fitted base-rate parameter,  $\beta_0(x, i)$  increases linearly with age. The fitted parameter values are less smooth than the one from the Li and Lee model in Fig. (1.5), which is probably due to applying the crude and not the smoothed mortality rates. However, the base-rate mortality parameters are ranked using the affluence index, as in the Li and Lee model.

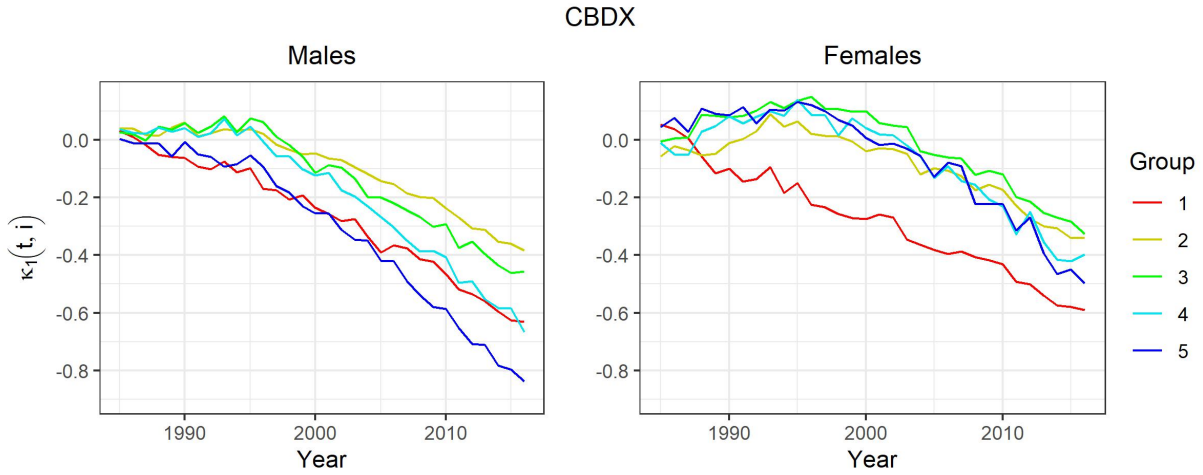


Fig. 1.12: Fitted  $\kappa_1(t, i)$  for both genders and all socioeconomic groups. The vertical axis displays  $\kappa_1(t, i)$ , and the horizontal axis displays the year for the subgroup parameters.

Fig. (1.12) shows fitted  $\kappa_1(t, i)$ . Overall, the fitted males,  $\kappa_1(t, i)$ , look similar in shape, magnitude, and rank as in Cairns et al. (2019), who use 10 affluence indices. The fitted females,  $\kappa_1(t, i)$ , are new in the literature. The lower fitted  $\kappa_1(t, i = 1)$  for G1 is lower than the other groups,

$\kappa_1(t, i)$ , which captures the mortality improvement seen in the data sample for G1 females.

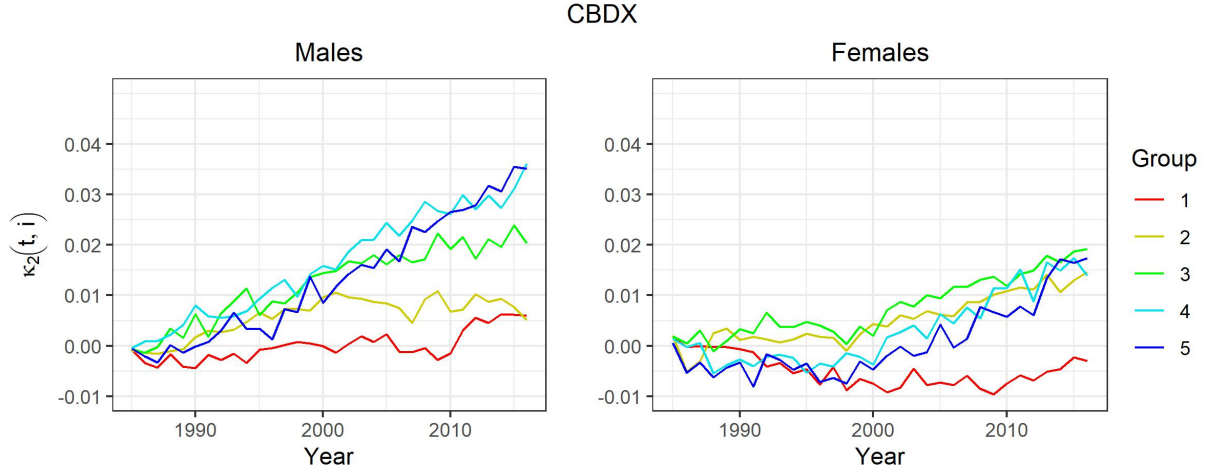


Fig. 1.13: Fitted  $\kappa_2(t, i)$  for both genders and all socioeconomic groups. The vertical axis displays  $\kappa_2(t, i)$ , and the horizontal axis displays the year for the subgroup parameters.

Fig. (1.13) shows fitted  $\kappa_2(t, i)$ , which captures changes in the slope of the mortality curve through time. As with the fitted  $\kappa_1(t, i)$ , the fitted  $\kappa_2(t, i)$  looks similar in shape, magnitude, and rank as in Cairns et al. (2019), e.g., the fitted  $\kappa_2(t, i)$  for the two most affluent groups, i.e., groups 4 and 5 have a sharper increase throughout the sample.

Next, we present the forecasted mortality rates at age 67 using the CBDX model WPU and without parameter uncertainty NPU.

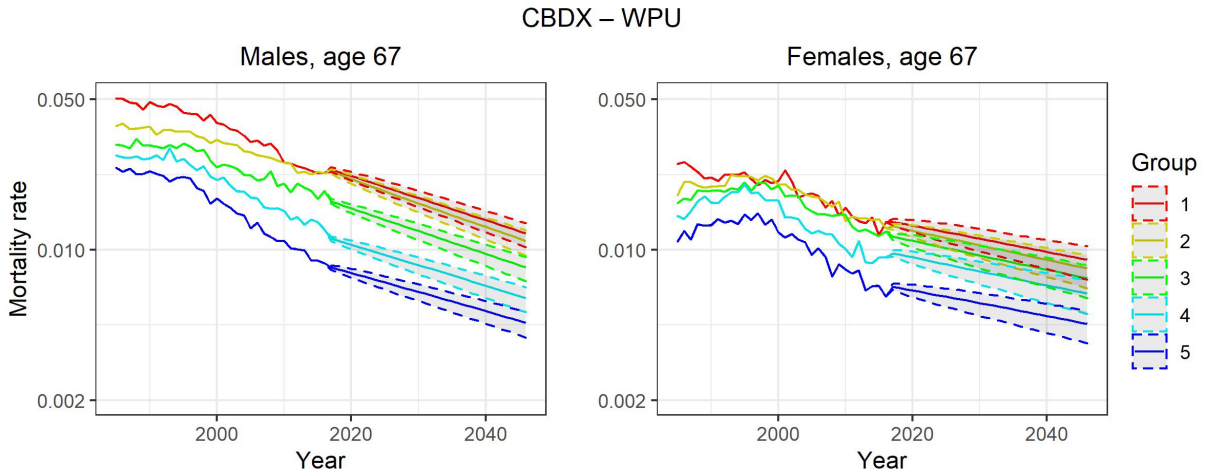


Fig. 1.14: Forecasted mortality rates for both genders and all socioeconomic groups using the CBDX model with parameter uncertainty (WPU) at age 67. The mortality rates, on the vertical axis, are rescaled to a logarithmic scale. The horizontal axis contains the in- and out-of-sample years. The uncertainty span contains the 95% confidence interval.

Fig. (1.14) shows that the forecasted mortality rates using the CBDX model WPU exhibits an expected mortality trajectory somewhat similar to that from the Li and Lee model and Enchev model, however with a narrower confidence interval. Also, the G1 males stay above G2 males mortality trajectory, as with the Enchev model.

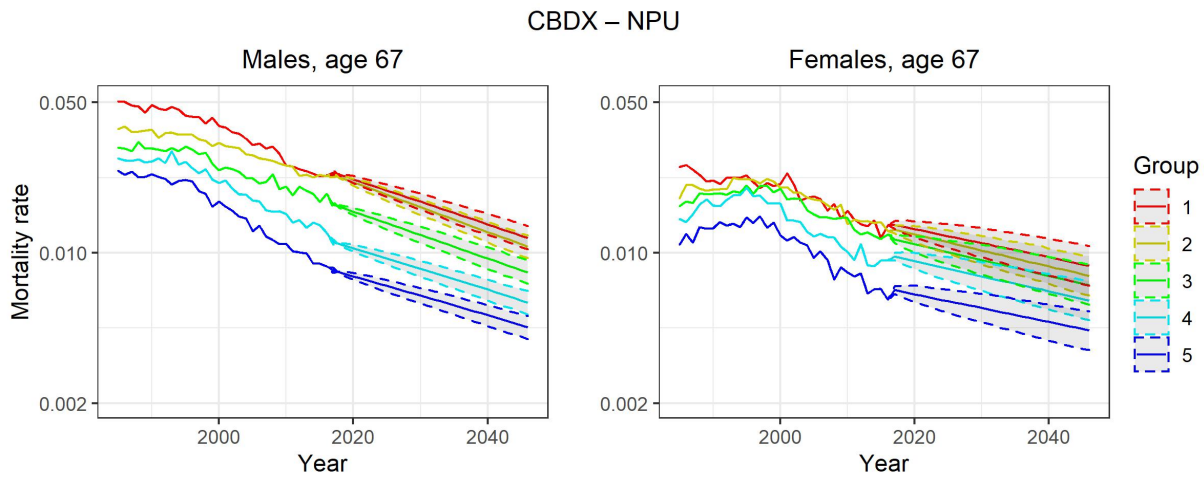


Fig. 1.15: Forecasted mortality rates for both genders and all socioeconomic groups using the CBDX model without parameter uncertainty (NPU) at age 67. The mortality rates, on the vertical axis, are rescaled to a logarithmic scale. The horizontal axis contains the in- and out-of-sample years. The uncertainty span contains the 95% confidence interval.

Fig. (1.15) shows forecasted mortality rates at age 67 using the CBDX model NPU. The expected mortality trajectory looks similar to the forecast using the CBD-X model with parameter uncertainty. Comparing the females' uncertainty confidence interval WPU and NPU, the uncertainty confidence interval is larger NPU than WPU.

### 1.6.3 Model selection procedure

To compare the model fit of the Li and Lee and CBDX models, we use the explanation ratio introduced in Li and Lee (2005) traditionally used in the actuarial literature, i.e.,

$$ER = 1 - \frac{\sum_{x,t,i} L_j(x,t,i)}{\sum_{x,t,i} L_0(x,t,i)} \quad (1.20)$$

$$\begin{aligned} L_j(x,t,i) &= \left[ \ln[m(x,t,i)] - \ln[m_j(\hat{x}, t_0 + T, i)] \right]^2 \\ L_0(x,t,i) &= \left[ \ln[m(x,t,i)] - a(\hat{x}, i) \right]^2. \end{aligned} \quad (1.21)$$

The explanation ratio is the proportion of variance explained in-sample relative to the variance explained by the base-rate mortality. The  $L_j$  is the loss function for the two chosen models in-sample, whereas the  $L_0$  is the difference between actual log-mortality rates and base-rate mortality rates.

To add to the rigorousness of choosing the optimal model, we use the MCS procedure (Hansen et al., 2011) to rank which of the models considered are the best given the data. The MCS procedure leads to a model set,  $\mathcal{M}^*$ , with a certain probability, analogously to a confidence interval for a population parameter that consists of the best models given the data.

We use the MCS procedure for both in-sample and out-of-sample evaluation. Thus, in-sample, the model set  $\mathcal{M}^0$  consists of the Li and Lee and CBDX models, and out-of-sample, the model set  $\mathcal{M}^0$  consists of the two models with the two variations of forecast assumptions.

In this context, the MCS procedure evaluates each mortality model,  $j$  and  $h$ , at the index  $x, t, i$  using the loss difference between model  $j$  and  $h$ ,  $d_{hj}(x, t, i)$ . This relative performance variable is

$$d_{hj}(x, t, i) \equiv L_h(x, t, i) - L_j(x, t, i), \quad \forall h, j \in \mathcal{M}^0, \quad (1.22)$$

where  $L_h(x, t, i)$  and  $L_j(x, t, i)$  is the squared error loss, as defined in equation (1.21). The null and

alternative hypothesis is defined as

$$H_{0,\mathcal{M}} : \mathbb{E} [d_{hj}(x, t, i)] = 0, \forall h, j \in \mathcal{M} \quad (1.23)$$

$$H_{A,\mathcal{M}} : \mathbb{E} [d_{hj}(x, t, i)] \neq 0, \forall h, j \in \mathcal{M}, \quad (1.24)$$

where  $\mathcal{M} \subset \mathcal{M}^0$ . To test the above hypothesis, Hansen et al. (2011) construct two t-statistics, where we will use the first, i.e.,

$$t_{hj} = \frac{\bar{d}_{hj}}{\sqrt{\hat{\mathbb{V}}ar(\bar{d}_{hj})}},$$

where  $\bar{d}_{hj} \equiv \frac{1}{X \times N \times T} \sum_{x,t,i} d_{hj}(x, t, i)$ .  $\bar{d}_{hj}$  is the relative sample loss between the  $h$ 'th and  $j$ 'th models, and  $\hat{\mathbb{V}}ar(\bar{d}_{hj})$  denotes estimates of  $\mathbb{V}ar(\bar{d}_{hj})$ . Hansen et al. (2011) argue that the null hypothesis can be tested by the t-statistics

$$T_{max,\mathcal{M}} = \max_{h,j \in \mathcal{M}} |t_{hj}|.$$

The MCS algorithm follows sequentially by eliminating the worst model, until the equal predictive ability is accepted. This ability is given by

$$e_{R,M} = \arg \max_h \left\{ \sup_{j \in M} \frac{\bar{d}_{hj}}{\sqrt{\hat{\mathbb{V}}ar(\bar{d}_{hj})}} \right\}.$$

Thus, if the equal predictive ability is accepted at the first iteration, every model is in  $\mathcal{M}^*$  with probability  $1 - \alpha$ .

**Result of the procedure** To empirically evaluate which of the mortality models, and thus RRA that are closest to the true underlying mortality evolution, we move forward with the MCS procedure (Hansen et al., 2011) described in Section (1.6.3).

Table (1.2) below shows the in-sample explanation ratios, the sum of the loss function, and the model set containing the best models with confidence level 5%.

Model	Li and Lee	CBDX
Criteria		
ER	0.635	0.6005
$\sum_{x,i,t} L_j(x,t,i)$	313.8	343.5
$\mathcal{M}^*$	Yes	No

Table 1.2: In this table is explanation ratios (ERs), sum of loss function and the best model in-sample. The ER is relative to the Li and Lee base-rate mortality rate,  $a(x,i)$ .  $\sum_{x,i,t} L_j(x,t,i)$  is the squared error loss relative to the fitted log-mortality rates, and  $\mathcal{M}^*$  is evaluated using the Hansen et al. (2011) MCS procedure with a confidence level of 5% and 5,000 bootstrap samples.

Table (1.2) shows that the explanation ratios are relatively close between the Li and Lee and the CBDX models, with the Li and Lee model having a higher explanation ratio than the CBDX model. The explanation ratios are somewhat lower than the ones from Wen et al. (2021), where they test 12 different multi-population models, e.g., the Li and Lee model which has between 0.71-0.79 in their 10 different socioeconomic groups in England. However, their maximum age is 89, whereas ours is 95.<sup>4</sup>

To compare the models out-of-sample, we repeat the analysis and MCS procedure by fitting the model up to year 2011 and evaluate the loss function with forecast up to the end of our sample in 2016. Table (1.3) presents the results.

Model	Li and Lee	Enchev	CBDX-WPU	CBDX-NPU
Criteria				
ER	0.8184	0.8118	0.7966	0.6995
$\sum_{x,i,t} L_j(x,t,i)$	68.1	70.6	76.3	112.7
$\mathcal{M}^*$	Yes	No	No	No

Table 1.3: In this table is explanation ratios (ERs), sum of loss function and the best model out-of-sample. The ER is relative to the Li and Lee base-rate mortality rate,  $\sum_{x,i,t} L_j(x,t,i)$  is the squared error loss relative to out-of-sample log-mortality rates, and  $\mathcal{M}^*$  is evaluated using the Hansen et al. (2011) MCS procedure with a confidence level of 5% and 5,000 bootstrap samples.

Somewhat surprisingly, the explanation ratios have generally increased in this sample for all of the models. The Li and Lee model, which still has the highest explanation ratio, is the only

<sup>4</sup>The result holds for all groups for both genders. The only group where there is marginal evidence of the CBDX model ought to be included in the model set  $\mathcal{M}^*$  is G3 males with a  $p$ -value of 5.48%, i.e., it is marginally included in the model set  $\mathcal{M}^*$  for G3 males.

model in the best model set,  $\hat{\mathcal{M}}^*$ . Thus, we have validated that the Li and Lee model is the best model given our dataset.

#### 1.6.4 Outline of risk premium propositions

This section presents the propositions behind the aggregate mortality risk premium and the individual q-forward prices conditional on the aggregate mortality risk premium.

**The  $q^f$  - price for aggregate mortality risk** Denote the Wang (2000) distortion operator as

$$\tilde{G}(z) = \Phi \left[ \Phi^{-1} [G(z)] - \lambda \right], \quad (1.25)$$

where  $\Phi$  is the standard cumulative distribution function (CDF),  $G(z)$  is the CDF for a risk,  $\lambda$  is a real-valued parameter, and  $\tilde{G}(z)$  is the risk-neutral distribution. Mortality is log-normal distributed, i.e.,  $\ln[q(\tau)] \sim N(\mu, \sigma)$ , where  $\mu$  is the expected log-mortality rate at time  $\tau$  for age  $x$  and the last parameter,  $\sigma$ , is the standard deviation. Applying the Wang transformation for aggregate mortality risk to the q-forward implies that the risk neutral price,  $q^f = \mathbb{E} [\tilde{G}[q(t)]]$ , is

$$q^f = \mathbb{E}[q(\tau)] \cdot \exp \left[ -\lambda(\tau) \cdot \sqrt{\text{Var}\{\ln(q(\tau))\}} \right]$$

*Proof.* If mortality is log-normally distributed then

$$G[q(t)] = \Phi \left( -\frac{\ln[q(t)] - \mu}{\sigma} \right). \quad (1.26)$$

Inserting equation (1.26) into equation (1.25) leads to

$$\begin{aligned}
\tilde{G}[q(t)] &= \Phi \left[ \Phi^{-1}(G[q(t)]) - \lambda \right] \\
&= \Phi \left[ \Phi^{-1} \left[ \Phi \left( -\frac{\ln[q(\tau)] - \mu}{\sigma} \right) \right] - \lambda \right] \\
&= \Phi \left[ \left( -\frac{\ln[q(\tau)] - \mu}{\sigma} \right) - \lambda \right] \\
&= \Phi \left[ \left( -\frac{\ln[q(\tau)] - (\mu - \lambda \sigma)}{\sigma} \right) \right] \\
\log [\tilde{G}[q(t)]] &\sim N(\mu - \lambda \sigma, \sigma). \tag{1.27}
\end{aligned}$$

Thus, equation (1.27) states that the risk-neutral distribution of mortality is another log-normal distributed variable with a distorted mean,  $\mu - \lambda \sigma$ , but with the same uncertainty,  $\sigma$ . Taking the expectation of equation (1.27) leads to the risk-neutral price, denoted by  $q^f$ , i.e.,

$$\begin{aligned}
q^f &\equiv \mathbb{E} [\tilde{G}[q(t)]] \\
&= \exp \left[ \mu - \lambda \sigma + \frac{1}{2} \sigma^2 \right] \\
&= \mathbb{E} [q(\tau)] \exp [-\lambda \sigma(\tau)] \\
&= \mathbb{E} [q(\tau)] \cdot \exp \left[ -\lambda(\tau) \cdot \sqrt{\text{Var} \{ \ln(q(\tau)) \}} \right], \tag{1.28}
\end{aligned}$$

which concludes the proof of the q-forward's price for aggregate mortality risk.  $\square$

**The  $q^f$ -price conditioned on aggregate mortality risk** The net rate of return from holding the q-forward,  $R(q) = \frac{q}{q^f} - 1$  is, per definition, log normally distributed as  $q$  is log normally distributed and  $q^f$  is a constant. Thus, the logarithm of the gross rate of return,  $r(q) = \ln[1 + R(q)]$ , is normally distributed. Furthermore, assume that the net risk-free rate is known and given by  $r_f = 1 + R_f$ , and denote, for simplicity, the risk-neutral distribution from equation (1.25) as the Choquet Integral, cf. Wang (2000), as  $H[Z; \lambda]$ , where  $Z$  is the risk and  $\lambda$  is the distortion parameter. Following Wang (2000), the current risk-neutral price,  $q^f$ , can be deducted as the present value of



end of period price,  $q$ , i.e.,

$$\begin{aligned}
q^f &= H \left[ \frac{q}{1+r_f}; -\lambda_i \right] \\
&= H \left[ \frac{q^f (1+R(q))}{1+R_f}; -\lambda_i \right] \\
r_f &= H[r(q); -\lambda_i] \\
r_f &= \mathbb{E}[r(q)] - \lambda_i \cdot \sigma[r(q)] \\
\lambda_i &= \frac{\mathbb{E}[r(q)] - r_f}{\sigma[r(q)]}.
\end{aligned} \tag{1.29}$$

Now,  $\lambda_i$  has an asset return interpretation, i.e., the risk distortion parameter is equal to the Sharpe ratio, i.e., excess return divided by the standard deviation of the return of holding the  $q$ -forward. Assuming there is a market price of risk for aggregate mortality risk, denoted by  $\lambda_M$ , the market price of risk for aggregate mortality risk is

$$\lambda_M = \frac{\mathbb{E}[r(q)] - r_f}{\sigma[r(q)]}. \tag{1.30}$$

Finally, assume that investors price aggregate mortality risk as in Ross (1976) i.e.,

$$\mathbb{E}[r_i] = r_f + \sum_{j=1}^n \beta_{j,i} \lambda_j, \tag{1.31}$$

where  $\beta_{j,i} = \frac{\text{Cov}[r_i, \lambda_j]}{\text{Var}[\lambda_j]}$  is the covariation between the return of asset,  $i$ , and the  $j$ 'th factor loading,  $\lambda_j$ , divided by the variance of the  $j$ 'th factor loading. Since equation (1.29) holds for every  $q$ -forwards (or every asset that is normal or log normally distributed), combining this with the Ross (1976) asset pricing model and assuming that aggregate mortality risk,  $\lambda_M$ , is uncorrelated with any other factor loading, we have that

$$\lambda_i = \rho_{i,M} \lambda_M, \tag{1.32}$$

where  $\rho_{i,M}$  is the correlation between the return of asset  $i$  and the return of holding the aggregate

mortality risk. Inserting equation (1.32) into the risk-neutral price for the q-forward in equation (1.28) concludes the proof.

Applying the risk-neutral price of the  $i$ th q-forward from equation (1.28), the risk loadings relation with systematic mortality risk in equation (1.32), and denoting the return of the  $i$ th q-forward as  $r(i) = \log[1 + R(i)]$  leads back to Ross (1976) in equation (1.31)

$$\begin{aligned}
q^f(i) &= \mathbb{E}[q(\tau, i)] \cdot \exp \left[ -\lambda(\tau, i) \cdot \sqrt{\text{Var}\{\ln(q(\tau, i))\}} \right] \\
r(i) &= \log \left( \frac{\mathbb{E}[q(\tau, i)]}{q^f} \right) \\
&= \lambda(\tau, i) \cdot \rho_{i,M} \sqrt{\text{Var}\{\ln(q(\tau, i))\}} \\
&= \beta_{i,M} \cdot \mathbb{E}[r_M],
\end{aligned} \tag{1.33}$$

where  $r_M$  is the rate of return holding the q-forward assumed to be constituting systematic mortality risk. Thus, the same arguments as in Ross (1976) apply, i.e., the individual q-forwards cannot deviate from the pricing rule in equation (1.33) as this would imply arbitrage opportunities that cannot hold up in a rational equilibrium.

**Total costs are independent of hedging strategy** Combining the total cost equation in equation (1.8) with the risk premium in equation (1.33) leads to the total cost being independent of whether the hedge is initiated with or without basis risk.

As deducted in equation (1.33), the risk premium is dependent on the sensitivity of the subgroup mortality rates to the FP mortality rate. More formally, marginally increasing the expected return of aggregate mortality risk in equation (1.33) increases the expected return of the  $i$ th q-forward with  $\beta_{i,M} \equiv \frac{\partial q(i)}{\partial q}$ . Likewise, the optimal notional amount of contracts,  $w(\tau)^*$  when hedging with basis risk is the marginal increase of the annuity value following a marginal increase of the aggregate mortality rate, i.e.,  $w(\tau)^* \equiv \frac{\partial a(q)}{\partial q}$ . Similarly, when hedging without basis risk,  $w(\tau)^*$  measures the marginal increase of the annuity value following a marginal increase of the subgroup-specific mortality rate, i.e.,  $w(\tau)^* \equiv \frac{\partial a(q)}{\partial q(i)}$ . Combining the optimal notional amount of

contracts and the risk premium when hedging without basis risk leads to the total cost

$$\begin{aligned}
C &= \sum_{\tau} \frac{\partial a(q)}{\partial q(i)} \cdot \frac{\partial q(i)}{\partial q} \cdot r_M \cdot (1+r)^{-(\tau-t_0)} \\
&= \sum_{\tau} \frac{\partial a(q)}{\partial q} \cdot r_M \cdot (1+r)^{-(\tau-t_0)},
\end{aligned} \tag{1.34}$$

which is exactly the same as when hedging with basis risk.

The reasoning behind this result is that the two mechanisms, i.e., the optimal notional amount of q-forward contracts and the risk premium, operate in exactly opposite directions. When hedging with basis risk, the risk premium is  $r_M$ , whereas the risk premium is  $\frac{\partial q(i)}{\partial q} r_M$  without basis risk. If the hedging population has lower mortality rates than the aggregate mortality rates as in the previous example, then  $\frac{\partial q(i)}{\partial q} < 1$  per definition. The deduction in equation (1.34) shows that the discount in risk premium when hedging without basis risk is exactly offset by the extra notional amount of contracts the low-mortality hedging population optimally buys. A similar interpretation is that both hedging with or without basis risk has benefits, but not simultaneously. A low mortality hedging population optimally buys fewer q-forward contracts when hedging with rather than without basis risk, but at a higher price. The risk premium, in contrast, is cheaper when hedging without basis risk but optimally entails buying more q-forward contracts.

The total cost equation above is novel in the literature of hedging longevity risk. The objective of the cost function is to elicit the differences in actuarial costs between hedging with and without basis risk. We show that, within our framework, the decision between hedging with or without basis risk for each hedging population is irrelevant in terms of costs.

### 1.6.5 Hedge effectiveness

In this section, we evaluate the hedge effectiveness with the Enchev and CBDX models. For the CBDX model, we evaluate both WPU and NPU. Table (1.4) shows the RRA using all five q-forwards and for all the four different models.

	Model	Li and Lee	Enchev	CBDX-WPU	CBDX-NPU
Policyholder	Index				
Males					
G1	SEG	0.98	0.97	0.98	0.98
	FP	0.70	0.91	0.31	0.25
	$\Delta$ RRA	<b>0.28</b>	<b>0.06</b>	<b>0.67</b>	<b>0.73</b>
G2	SEG	0.98	0.98	0.98	0.98
	FP	0.96	0.97	0.33	0.23
	$\Delta$ RRA	<b>0.02</b>	<b>0.01</b>	<b>0.65</b>	<b>0.75</b>
G3	SEG	0.98	0.98	0.98	0.98
	FP			0.31	0.23
	$\Delta$ RRA			<b>0.67</b>	<b>0.75</b>
G4	SEG	0.98	0.98	0.98	0.98
	FP	0.90	0.97	0.32	0.22
	$\Delta$ RRA	<b>0.07</b>	<b>0.01</b>	<b>0.66</b>	<b>0.76</b>
G5	SEG	0.98	0.98	0.98	0.98
	FP	0.59	0.87	0.31	0.24
	$\Delta$ RRA	<b>0.39</b>	<b>0.10</b>	<b>0.67</b>	<b>0.74</b>
Policyholder	Index				
Females					
G1	SEG	0.98	0.96	0.98	0.98
	FP	0.20	0.74	0.30	0.26
	$\Delta$ RRA	<b>0.78</b>	<b>0.22</b>	<b>0.68</b>	<b>0.72</b>
G2	SEG	0.98	0.97	0.98	0.98
	FP	0.43	0.97	0.33	0.24
	$\Delta$ RRA	<b>0.55</b>	<b>0.00</b>	<b>0.65</b>	<b>0.74</b>
G3	SEG	0.98	0.98	0.98	0.98
	FP			0.33	0.25
	$\Delta$ RRA			<b>0.65</b>	<b>0.73</b>
G4	SEG	0.98	0.98	0.98	0.98
	FP	0.32	0.96	0.33	0.23
	$\Delta$ RRA	<b>0.66</b>	<b>0.01</b>	<b>0.65</b>	<b>0.75</b>
G5	SEG	0.99	0.97	0.98	0.98
	FP	0.31	0.95	0.35	0.25
	$\Delta$ RRA	<b>0.68</b>	<b>0.02</b>	<b>0.63</b>	<b>0.73</b>

Table 1.4: Relative risk adjustments (RRA) for all subgroups. The columns list the Li and Lee, Enchev, and CBDX models with (WPU) and without parameter uncertainty (NPU). For the specific subgroup, row 1 quantifies the RRA where the reference population is the specific socio-economic group (SEG), row 2 shows the RRA when hedging with basis risk where the reference population is the full population (FP), and row 3 illustrates the difference in RRA between hedging without and with basis risk.

Across the different mortality models, one thing stands out, hedging directly on the subgroup-specific index diminishes the variability by 98-99% as is the case with the Li and Lee model.

Thus, hedge effectiveness when the q-forward is indexed to the subgroup-specific mortality rate is high irrespective of the assumed model.

However, if the q-forward is indexed to the FP, the hedge effectiveness varies across models. Comparing the Enchev model with the Li and Lee model, the hedge effectiveness is higher for the former than for the latter. Since the modeling assumption for the parameter  $k(t, i)$  in the Enchev model is a vector autoregression model, the different  $k(t, i)$  projections are more highly correlated than in the Li and Lee model. This leads to higher correlation in mortality rates, and possibly annuity values, between the groups. As the optimal RRA in equation (1.7) denotes the linear dependency between the hedging instrument and the annuity value for the different group, the RRA is higher using the Enchev model than when using the Li and Lee model.

For the CBDX model WPU and NPU, the pattern across groups is quite different. The loss in hedge effectiveness, summarized in row 3 for each group, is almost uniform across subgroups and also rather similar between the two genders. In general, the hedge effectiveness attained is higher for the WPU model than NPU.

### **1.6.6 Risk premium for q-forwards - Enchev and CBDX models**

It is not only hedge effectiveness that is exposed to model risk. In this section, we show the actuarial fair risk premiums using equation (1.12) for the Enchev and CBDX NPU models.

Starting with the Enchev model, we expect the SEG-specific actuarial risk premiums to be more highly correlated with the aggregate mortality risk premium as the hedge effectiveness is higher with the Enchev model than the Li and Lee model, as Table (1.1) shows.

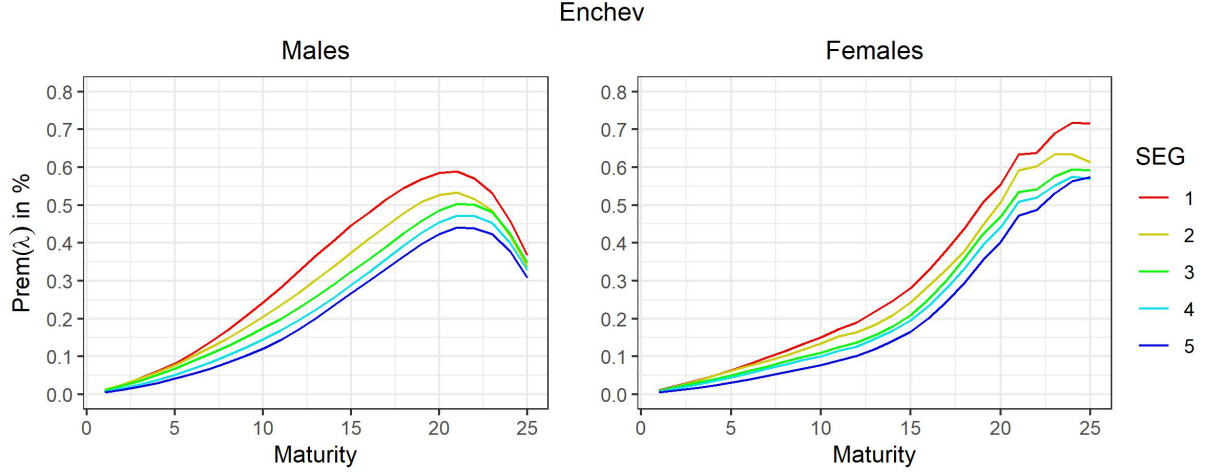


Fig. 1.16: Enchev model:  $q$ -forward premiums for all socioeconomic groups (SEG) by maturity. On the vertical axis, the premiums are measured as the percentage-point difference between the  $q$ -forward rate,  $q^f$ , and the expected mortality,  $\mathbb{E}[q]$ . The premiums vary on the horizontal axis depending on maturity  $\tau$  for ages  $x + \tau$ .

Fig. (1.16) illustrates the same pattern of mortality risk premiums as in the Li and Lee model in Fig. (1.2). However, the risk premiums using the Enchev model are a bit closer to the aggregate mortality risk premium. Furthermore, the actuarial risk premium for the G1 females is the highest risk premium.

With the CBDX model, we model aggregate mortality risk as the risk arising from the mean of the individual time-dependent and age-dependent parameters, i.e.,  $\bar{\kappa}_j(t) = \frac{1}{N} \sum_{i=1}^N \kappa_j(t, i) \quad \forall j = 1, 2$  and  $\bar{\beta}_0(x) = \frac{1}{N} \sum_{i=1}^N \beta_0(x, i)$ .

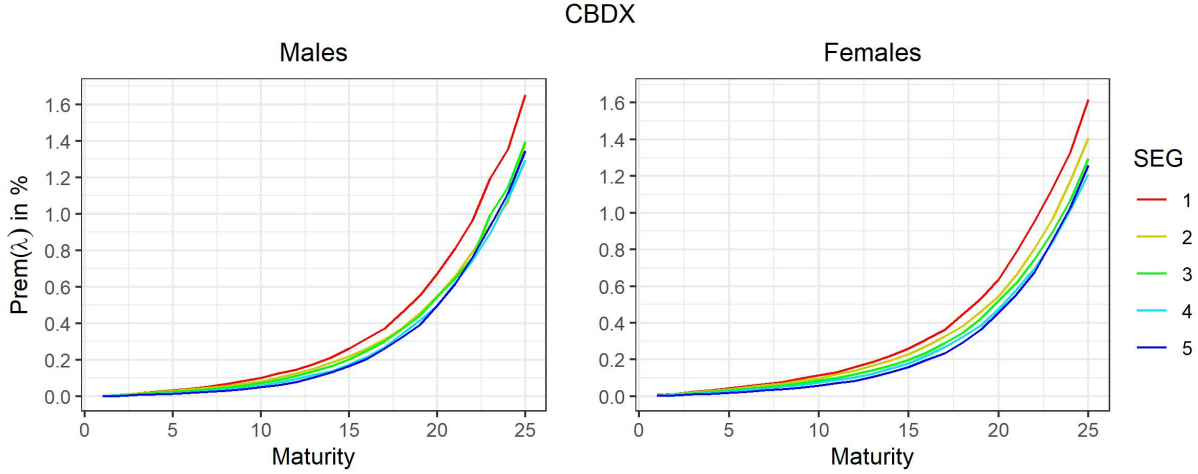


Fig. 1.17: CBDX model:  $q$ -forward premiums for all socioeconomic groups (SEG) by maturity. On the vertical axis, the premiums are measured as the percentage-point difference between the  $q$ -forward rate,  $q^f$ , and the expected mortality,  $\mathbb{E}[q]$ . The premiums vary on the horizontal axis depending on maturity  $\tau$  for ages  $x + \tau$ .

Fig. (1.17) shows a rather different pattern for the aggregate risk premium, which increases with a higher exponential rate with maturity for both genders. As a result, the risk premiums in the longer maturities, i.e., in excess of 20 years, do not decline but rise with a faster rate than in the Li and Lee and Enchev models.

In summary, the actuarial fair risk premium is also exposed to model risk, especially at longer maturities in excess of 20 years.

### 1.6.7 Cost of hedging longevity risk - Enchev and CBDX models

Applying the present value of the risk premiums in the previous section, along with the optimal notional amount of  $q$ -forward contracts,  $w(\tau)^*$ , we arrive at the total costs of engaging in the  $q$ -forwards for the two alternative mortality models, i.e., the Enchev and CBDX models.

Fig. (1.18) shows the total costs as a percentage of the annuity value using the Enchev model by hedging with and without basis risk.

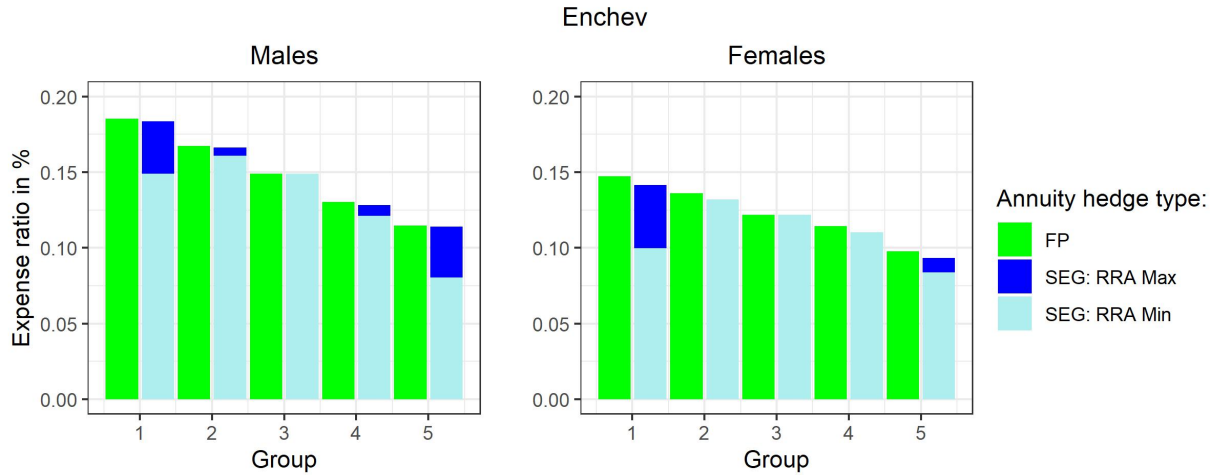


Fig. 1.18: Annual expense ratios for hedging longevity risk. Enchev model: For each group, there are three categories of bar plots: FP, SEG: RRA Min, and SEG: RRA Max. The green bar (full population (FP)) is the expense ratio by hedging referenced to the median group 3 for all five q-forwards, i.e., with basis risk. The light blue bar (socioeconomic group (SEG), SEG: RRA Min) is the expense ratio for attaining an relative risk adjustment (RRA) at least as high as with the FP hedge. The dark blue bar (SEG: RRA Max) is the expense ratio when referenced to the specific SEGs for all five q-forwards, i.e., without basis risk.

Overall, the total costs per unit of annuity value is the same when applying the Enchev model as it is with the Li and Lee model, cf. Fig. (1.18) and Fig. (1.4). However, the gain in terms of costs by hedging without basis risk is lower for the groups closest to group 3.

On contrary to the Li and Lee and Enchev models, the total costs with the CBDX model have slightly higher total costs when hedging with basis risk than without basis risk, cf. Fig. (1.19). Furthermore, the costs is approximately uniform across groups for the males, whereas the total costs are decreasing for higher annuity-value groups for the females.



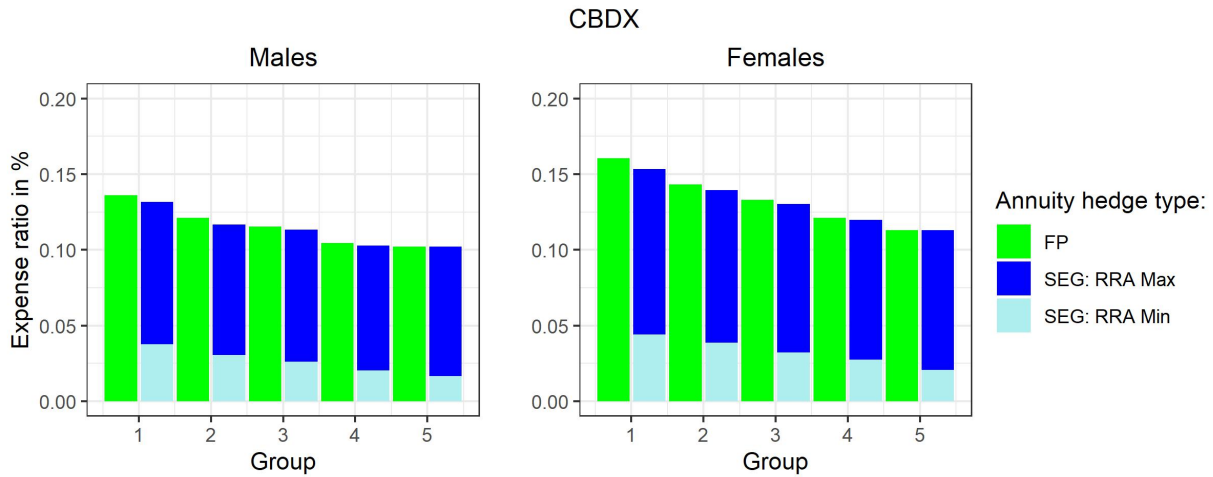


Fig. 1.19: Annual expense ratios for hedging longevity risk. CBDX model: For each group, there are three categories of bar plots: FP, SEG: RRA Min, and SEG: RRA Max. The green bar (full population (FP)) is the expense ratio by hedging referenced to the median group 3 for all five q-forwards, i.e., with basis risk. The light blue bar (socioeconomic group (SEG), SEG: RRA Min) is the expense ratio for attaining an relative risk adjustment (RRA) at least as high as with the FP hedge. The dark blue bar (SEG: RRA Max) is the expense ratio when referenced to the specific SEGs for all five q-forwards, i.e., without basis risk.

When measuring the total costs per unit of annuity value the total costs are slightly decreasing for higher annuity-value groups. Furthermore, the magnitude of costs is within the same range as in the Li and Lee and Enchev models.

# Chapter 2

## 2 Unsystematic mortality and time-varying returns

### Abstract

This paper explores the impact of unsystematic mortality risk on the US stock market. Using portfolios sorted on industry and dividend-prices, I study the link between unsystematic mortality and real returns one year ahead. A higher unsystematic mortality rate represents a temporary adverse demand shock that increases next periods' returns. I find that an increase in unsystematic mortality raises one-year ahead real returns heterogeneously. The magnitude for half of industry portfolios is equivalent to a comparable increase in the dividend-price ratio, whereas the other half is not influenced. In addition, the real returns of value firms' have increased by 3.26 % on average following a standardized mortality shock, while growths' equivalents, in contrast, remain unaltered. The difference accounts for 80% of the average real return gap between value and growth portfolios.

## 2.1 Introduction

Does mortality risk affect financial assets? The current assumption in the actuarial literature is that the evolution of mortality rates is independent of financial risks; see Cairns et al. (2006) and Hunt and Blake (2021). Mortality, which varies by year and age, is an omitted variable as it is publicly available with a two-year lag. Moreover, mortality risk can be divided into two components: unsystematic and systematic (Cairns et al., 2008). The latter refers to permanent changes in mortality that affect all individuals in the same way, whereas the former relates to transitory fluctuations in the number of deaths. Naturally, unsystematic mortality risk is the largest for smaller populations; thus, the volatility of mortality rates is higher for the older ages due to smaller populations.

Theoretical financial models are generally preoccupied with systematic mortality risk, which adds to the uncertainty of intertemporal substitution decisions; see, e.g., Cocco and Gomes (2012) and Heimer et al. (2019). In contrast, unsystematic mortality specifies abrupt losses in demand arising from fluctuations in the number of deaths and is unrelated to intertemporal choices due to the limited time horizon for the aging part of a population.

Empirically, mortality shocks do not alter aggregate predictive asset returns or consumption growth significantly; see Bisetti et al. (2017) and Friedberg and Webb (2007). Nevertheless, approximately 50% of stock price variations are temporary (Lee, 1995; Poterba and Summers, 1988), which corresponds to changes in current dividends. Furthermore, the temporary variations are either due to irrational bubbles or time-varying expected returns; see Fama (1991). Accordingly, Cochrane (2011) emphasizes that temporary price deviations need to have common variations across various characteristics, to be identified as rational time-varying returns. For example, Fama and French (1992) show that various pricing ratios summarize price information concerning risk and expected return. This corresponds to Ball (1978)'s notion that earnings-price ratios, and other pricing ratios, capture various omitted risk variables.

In addition, Ball (1978) argues that the experimental design must be predictive to assess the impact on returns of a potential omitted variable. That is, the impact of the omitted variable on the pricing ratio is already in the information set at time  $t$ ; thus, a predictive setting avoids

any simultaneous price adjustment in returns and corresponds to identifying time-varying stock returns.

Motivated by this, I examine the time-varying real returns following yearly variations in unsystematic mortality risk in two dimensions. First, I analyze differences in time-varying returns across industry portfolios. Moreover, I infer economic point estimates of the loss in consumption of various goods in older age groups, which validates the heterogeneous impact on real returns. Second, I explore differences in time-varying returns across portfolios sorted on dividend-price ratios, which proxy for the value factor, due to variations in unsystematic mortality risk.

I examine the predictability of returns by including the dividend-price ratio for the specific portfolio in conjunction with the unsystematic mortality variable. I also standardize all variables and multiply with the portfolio's sample standard deviation of return. This allows a comparison between the two predictor variables on an equal scale, in accordance with Cochrane (2008). Changes in dividend-price ratios signal *temporary* and *permanent* changes in expected returns and/or expected dividend growth due to the Campbell and Shiller (1988) approximation. Most of the variation in the dividend-price ratio for the aggregate market portfolio accounts for changes in long-term expected returns; see Campbell and Ammer (1993). Moreover, Lettau and Van Nieuwerburgh (2008) show that changes in the next period dividend-price ratio signals changes in expected return less expected dividend growth and that next period returns, i.e., contemporaneous returns, are negatively related to the next period dividend-price ratio. Thus, an exogenous variable that increases (decreases) expected returns leads to higher (lower) discounting, which generates a lower (higher) contemporaneous return.

To quantify the economic magnitude of changes in unsystematic mortality on time-varying returns, I examine the two-good model of Piazzesi et al. (2007), which involves a numéraire good and a composite good. The latter can be extended to several goods, which allows comparison of the economic magnitude between distinct goods following demand shocks. Consumers have a generalized constant elasticity of substitution for the composite good relative to the numéraire good. In contrast to Piazzesi et al. (2007)'s model, the interest is in how the composite good is affected, relative to the numéraire good, following an exogenous demand shock, and how much

these disintegrated demand shocks affect the stochastic discount factor and the expected dividend growth. Following a larger negative (positive) demand shock for the composite good in period  $t$ , relative to the numéraire good, the model predicts that expected returns and expected dividend growth increase (decrease). This is consistent with Piazzesi et al. (2007)'s two-good model. Moreover, the model's predictions are comparable with the model Menzly et al. (2004) introduced, where a low (high) relative share of the dividend of good  $i$  in relation to aggregate consumption positively (negatively) predicts the next-period dividend growth and real returns. Similarly, for the two-good model, the interpretation is that relative consumption of the composite good affects the impact of an exogenous shock on the economy. However, in the model presented here, a uniformly positive or negative shock can affect the returns and dividend growth of several goods with equal sign, which is unattainable in Menzly et al. (2004)'s model as all relative dividends must add up to one, which leads to counteracting impacts.

I study the real returns of 10 portfolios sorted on dividend-price ratios and industry that follow Fama and French (1997)'s definition, whereas the dividend-price sorted portfolios adhere to the formalization in Fama and French (1988). In accordance with most of the literature, I use the portfolios available on Kenneth French's website and compute dividend-price ratios as in Cochrane (2008), assuming that dividends are reinvested. Moreover, I use the aggregate market portfolio and the consumer price index from Shiller (2015). All financial variables are yearly as mortality rates from the Human Mortality Database have yearly frequencies. Last, I use the Lee and Carter (1992) model to infer the unsystematic mortality variable.

I find that an increase in unsystematic mortality increases contemporaneous returns for all industries on average, exempt the durable industry portfolio; however, the impact is highly heterogeneous. For the five industries most exposed to unsystematic mortality, a standard deviation increase of unsystematic mortality raises contemporaneous returns by 4.3% on average. For the five industries least exposed to unsystematic mortality, in contrast, a standard deviation shock to unsystematic mortality increases contemporaneous returns by 0.6% on average. A standard deviation increase of the dividend-price ratio, in comparison, leads to a 4.5% increase in one-year-ahead returns for the five industries most exposed to unsystematic mortality, which is an equivalent mag-

nitude to a comparable increase in unsystematic mortality. In addition, unsystematic mortality raises the contemporaneous return of the aggregate market portfolio by 1% on average. Thus, unsystematic mortality affects roughly half of the disintegrated industry portfolios by a significant amount.

Using the intertemporal elasticity of substitution of demand as a free parameter to equate the economic impact on the aggregate market demand leads to an estimate of around 0.35 and 0.47, which is within the interval of estimates microeconomic studies propose. The economic magnitude is relatively modest and equals, at most, a temporary negative demand shock of 1% for the most exposed industries and a negative impact of 0.24% for aggregate demand. The demand shocks correspond well with observed consumption patterns in recent consumer expenditure surveys for the older population, as well as the expected relative total consumption. In addition, the point estimates for dividend-price sorted portfolios show that unsystematic mortality accounts for 80% of the differences in real returns between the highest and lowest dividend-price sorted portfolios. This matches Ferson and Harvey (1991)'s observation that 80% of the time variation in returns is due to time-varying returns. These findings show that unsystematic mortality affects time-varying stock returns one year ahead. In addition, the findings suggest that mortality shocks account for a significant amount of the value anomaly.

In the remainder of this study, section (2.2) introduces the methodology, including the predictive framework and the two-good model. Next, section (2.3) describes the data and mortality model, while section (2.4) contains the empirical results of unsystematic mortality's predictability of real returns. Finally, section (2.5) presents the conclusions.

## 2.2 Methodology

Dividend-price ratios vary over time. Consequently, the variations in the time-series of dividend-price ratios affect either expected returns,  $r_{t+1}$ , and/or dividend growth,  $\Delta d_{t+1}$ , which follows from Campbell and Shiller (1988)'s log-linear approximation

$$\begin{aligned}
r_{t+1} &= k + \rho (pd_{t+1}) + \Delta d_{t+1} - (pd_t) \\
dp_t &= r_{t+1} - \Delta d_{t+1} + \rho (dp_{t+1}) - k,
\end{aligned} \tag{2.1}$$

where  $p_t \equiv \log(P_t)$ ,  $d_t \equiv \log(D_t)$ , and  $r_{t+1} \equiv \log\left(\frac{P_{t+1} + D_{t+1}}{P_t}\right)$ . The constants  $k$  and  $\rho = \frac{\bar{P}\bar{D}}{1 + \bar{P}\bar{D}}$  are functions of the long-run average log dividend-price ratio  $\bar{d}p = \log\left(\frac{\bar{r} - \bar{d}}{1 + \bar{d}}\right)$ . The log-linear approximate identity in equation (2.1) is a *transitory* relation, signifying that temporary deviation in the current dividend-price ratio leads to momentary changes in expected return and/or dividend growth. Iterating equation (2.1) forward, imposing the transversality condition  $\lim_{j \rightarrow \infty} \rho^j [dp_{t+j}] \rightarrow 0$ , and taking the conditional expectation at time  $t$  amounts to:

$$dp_t = \bar{d}p + \mathbb{E}_t \sum_{j=1}^{\infty} \rho^{j-1} [(r_{t+j} - \bar{r}) - (\Delta d_{t+j} - \bar{d})]. \tag{2.2}$$

Equation (2.2) relates *permanent* variations in the dividend-price ratio to changes in expected returns and dividend growth. A rise (fall) in the dividend-price ratio must be due to a rise (fall) in expected returns and/or fall (rise) in expected dividend growth.

The dividend-price ratio on the aggregate market has mainly predicted future long-term returns and not future long-term dividend growth in the post-Second World War era; see Campbell and Ammer (1993) and Chen (2009), which is represented by the first term in equation (2.2), whereas the opposite pattern has been observed for firm-level stocks; see Vuolteenaho (2002). In addition, dividend-price ratios account for temporary changes in returns and dividend growth; cf. equation (2.1). Thus, dividend-price ratios play a dual role as a predictor of permanent and transitory changes for long-term returns. A variable can alter expected returns momentarily, cf. equation (2.1), without changing long-run expected returns. According to Cochrane (2011), such a variable conveys changes in the term structure of risk premia.

### 2.2.1 Financial magnitude

The dividend-price ratio is also a persistent predictor, which entails that revisions in dividend-price ratios are related to persistent alterations of expected returns and/or expected dividends growth, which justifies large changes in stock prices. Lettau and Van Nieuwerburgh (2008) propose a structural model that envisions how changes in the dividend-price ratio carry information from changes in expected returns and dividend growth. The structural model assumes that expected dividend growth and expected returns follows a first-order autoregressive (AR(1)) process with an autoregressive coefficient  $\phi$

$$\Delta d_{t+1} - \bar{d} = \mathbb{E}_t [\Delta d_{t+1}] + \varepsilon_{t+1}, \quad \mathbb{E}_{t+1} [\Delta d_{t+2}] = \phi \mathbb{E}_t [\Delta d_{t+1}] + \zeta_{t+1} \quad (2.3)$$

$$r_{t+1} - \bar{r} = \mathbb{E}_t [r_{t+1}] + \eta_{t+1}, \quad \mathbb{E}_{t+1} [r_{t+2}] = \phi \mathbb{E}_t [r_{t+1}] + \xi_{t+1}, \quad (2.4)$$

where  $\varepsilon_{t+1}$  and  $\eta_{t+1}$  are unexpected dividend and return shocks, respectively, and innovations to expected dividend growth and expected returns are denoted as  $\zeta_{t+1}$  and  $\xi_{t+1}$ . Combining equation (2.2) with equations (2.3)-(2.4) reveals that the dividend-price ratio is a function of the variables: expected returns and dividend growth, and the autoregressive coefficient,  $\phi$ , and  $\rho$

$$dp_t - \bar{d}p = \frac{\mathbb{E}_t [r_{t+1}] - \mathbb{E}_t [\Delta d_{t+1}]}{1 - \rho\phi}. \quad (2.5)$$

Similar to equation (2.2), equation (2.5) shows that a larger dividend-price ratio indicates higher expected returns that, however, can be attenuated by expected dividend growth. In addition, even minor changes in expected returns or expected dividend growth can lead to significant changes in the dividend-price ratio, provided that  $\rho$  and  $\phi$  are positive and  $\rho\phi < 1$ . Combining equation (2.1) with equations (2.3)-(2.5) leads to a structural decomposition of innovations to unexpected returns:

$$\eta_{t+1} = \frac{-\rho}{1 - \rho\phi} [\xi_{t+1} - \zeta_{t+1}] + \varepsilon_{t+1}. \quad (2.6)$$

Higher (lower) expected returns (dividend growth) lead to lower contemporaneous returns, which is denoted as  $\eta_{t+1}$ . The economic reasoning for the former is that a higher expected return



leads to a higher discounting of future dividends, resulting in a lower price and thus return. In addition, the term in square brackets is defined as the part of the next period's dividend-price ratio not explained by the current one, cf. equations (2.3)-(2.5). Hence, if a variable predicts the next period's dividend-price ratio with the current dividend-price ratio in the information set, it ought to predict the next period's return with the opposite sign.

Moreover, since the price-dividend ratios of equities are high and very persistent, even small shocks to the next period's expected return or dividend growth lead to large shocks in contemporaneous returns. In addition, the next period's dividend-price ratio includes temporary deviations that fade in the long term. Indeed, temporary deviations in dividend shocks explain 50% of the response of stock prices on the aggregate market; see Lee (1995) and Poterba and Summers (1988). Empirically, this means that a positive (negative) shock to the next period's dividend is followed by an even higher price appreciation (depreciation). This occurs even if the dividend shock is not large enough to alter expected dividends. As dividend-price ratios mostly convey information about expected returns, a temporary dividend shock would generally lead to a contemporaneous return of the opposite sign.

### **2.2.2 Economic mechanism**

As a result, temporary dividend shocks may affect contemporaneous returns through dividend-price ratios. This could be effectuated by sudden changes in consumption patterns that change expected returns and expected dividend growth. However, consumers substitute more consumption between goods and within a time interval than across time. That is, the intertemporal elasticity of substitution of consumers is normally lower than 1 (Attanasio and Weber, 1995; Hall, 1988), whereas intratemporal elasticities of goods are generally larger than one due to Hick's second law of demand; cf. Hicks (1946) and Pollak (1969).

Thus, temporary changes in consumption will affect assets heterogeneously, depending on the preferences for different goods. Piazzesi et al. (2007)'s two-good model, which can be extended to include multiple goods, illustrates this economic intuition; see Appendix (2.6.2). Based on Piazzesi et al. (2007), I assume that two goods are present for a representative consumer, where  $c$

is the numéraire good,  $h$  denotes the composite good, and the consumer has a generalized constant elasticity of substitution utility function:

$$C_t = \left[ c_t^{(\varepsilon-1)/\varepsilon} + \omega h_t^{(\varepsilon-1)/\varepsilon} \right]^{\varepsilon/(\varepsilon-1)},$$

where  $\omega$  represents the consumption shares of good  $h$ , and  $\varepsilon$  denotes the constant intratemporal elasticity of substitution between the goods. Then the stochastic discount factor,  $M_{t+1}$ , resulting from the two-good's economy can be represented as:

$$M_{t+1} = \beta \left[ \frac{c_{t+1}}{c_t} \right]^{-\frac{1}{\sigma}} \left( \frac{c_t}{c_{t+1}} \frac{p_t^c}{p_{t+1}^c} \frac{D_{c,t+1} + D_{h,t+1}}{D_{c,t} + D_{h,t}} \right)^{-\frac{(\varepsilon-\sigma)}{\sigma[\varepsilon-1]}}, \quad (2.7)$$

where  $\beta$  is the subjective discount rate,  $\sigma$  denotes the intertemporal elasticity of substitution,  $p^c$  represents the consumer price for the numéraire good,  $c$ , and  $D_i = p_t^i i$  constitutes the real expenditure of good  $i$  as all expenditures are paid out as dividends. Equation (2.7) designates two different mechanisms.

The first component in the square bracket represents the consumption capital asset pricing model and states the consumer's preference to allocate consumption intertemporally, which is not volatile enough to explain the variations in stock prices; see Grossman and Shiller (1981). In contrast, the latter component in parentheses represents aggregated real dividend growth between time  $t$  and  $t+1$  relative to numéraire consumption. Furthermore, it is raised to a negative exponent indicating that current low relative dividend levels lead to a high expected dividend growth and return. Thus, the dynamics of the latter expression is analogous to the consumption capital asset pricing model. In addition, the latter dynamic is partially comparable to Menzly et al. (2004)'s. In their model, a relatively low dividend level of good  $i$  to aggregate consumption leads to higher expected returns due to mean-reverting dividends but their model implies that an exogenous shock that affect dividends uniformly negatively or positively leads to offsetting impacts on the goods' respective returns as the sum of all relative dividends have to equal one. In contrast, the two-good model solely requires that an impact on the composite good's expenditure is larger than the numéraire good's equivalent.

### 2.2.3 Predictions following exogenous shock

Suppose consumption expenditures are exposed to a temporary exogenous shock,  $g_t$ , that affects only the composite good,  $h_t$ . Moreover, consider  $g_t$  as an age-dependent shock affecting a decline in the older population, e.g., a severe flu season. As agents of different ages consume different goods (see DellaVigna and Pollet, 2007), such a shock would leave an imprint that is independent of the first component in equation (2.7). In addition, consumers in the late part of their life cycle use endowments to invest in the risk-free asset or for consumption (Cocco et al., 2005), leaving the intertemporal influence on stocks unaltered; see Vissing-Jørgensen (2002). Subsequently, the stochastic discount factor is affected by the fall in real expenditures of the composite good, denoted by  $\log\left(\frac{\tilde{C}_{t+1}}{\tilde{C}_t}\right) = \log\left(\frac{p_t^c}{p_{t+1}^c} \frac{D_{h,t+1}}{D_{h,t}}\right)$ , which is assumed to be log normally distributed.

If dividend growth of the aggregate market follows an AR(1) process, as in equation (2.3), and the joint log-normal distribution of expected returns implied by equation (2.7) follows the AR(1) process in equation (2.4), a negative  $g_t$ , ceteris paribus, leads to:

$$\frac{\partial \mathbb{E}_t(r_{t+1})}{\partial g_t} = (\varepsilon - \sigma) / [\sigma(\varepsilon - 1)] \quad (2.8)$$

$$\frac{\partial \mathbb{E}_t(\Delta d_{t+1})}{\partial g_t} = 1. \quad (2.9)$$

Thus, expected returns and dividend growth at time  $t$  increases proportionally with the exogenous shock,  $g_t$ . However, as  $g_t$  is temporary, the consumption level of the composite good,  $h_{t+1}$ , is unaltered, leading to a decrease in the innovations of the next period's expected returns and dividend growth, which are denoted as  $\xi_{t+1} \equiv \mathbb{E}_{t+1}[r_{t+2}] - \phi \mathbb{E}_t[r_{t+1}]$  and  $\zeta_{t+1} \equiv \mathbb{E}_{t+1}[\Delta d_{t+2}] - \phi \mathbb{E}_t[\Delta d_{t+1}]$  in equations (2.3) and (2.4). Hence, the  $g_t$  temporary shocks at time  $t$  lead to time-varying expected returns and dividend growth:

$$\frac{\partial [\xi_{t+1} - \zeta_{t+1}]}{\partial g_t} = -[\varepsilon - \sigma] / \sigma[\varepsilon - 1] - 1. \quad (2.10)$$

The first expression in equation (2.10), which is related to innovations in expected returns in equation (2.8), is higher than the latter, expected dividend growth in equation (2.9), as intratem-

poral substitution ( $\epsilon$ ) is higher than intertemporal substitution ( $\sigma$ ); cf. section (2.2.2). Moreover, innovations in expected returns and expected dividend growth comove. Thus, following a  $g_t$  shock, as considered above, composite consumption expenditure declines, expected returns and expected dividend growth in period  $t$  increases, innovations in expected returns and expected dividend growth decline and contemporaneous returns are high, cf. equation (2.6).

In addition, a consumption-related impact would result in heterogeneous time-varying returns due to differing preferences for goods. Since the aggregated impact leads to equation (2.10), the relative impact on different consumption goods can be inferred by dividing with the aggregate loss in demand following a  $g_t$  shock.

## 2.3 Data and descriptive statistics

The previous section describes how sudden changes in consumption expenditures can affect expected returns, expected dividends growth, and contemporaneous returns. In this section, I introduce an exogenous variable that may affect consumption expenditures abruptly, namely unsystematic mortality.

### 2.3.1 Mortality data description

In this regard, I use annual mortality data from the Human Mortality Database, which contains death and population counts from various countries and has been used extensively in actuarial research and industry. To keep the mortality analysis simple, I focus on the US population. Reliable data on annual death counts are available every year for ages 85 and older from 1951 to present time in the US. Moreover, population counts have been adjusted for 1940 to 1969 to exclude overseas Armed Forces since they are excluded from death counts. Since the variability of deaths in older ages is much larger due to small population sizes (see Cairns et al., 2008), the age span I consider for unsystematic mortality is that above 90 years of age where reliable death counts are available. In addition, I use mortality rates from 1953 up to 2020, the most recent year. When fitting the mortality model to extract unsystematic mortality, I use the age span above 65 years of age to provide more robust parameter estimates.

### 2.3.2 Estimating unsystematic mortality

Widely used in actuarial research, Lee and Carter (1992)'s mortality model is recognized as the leading mortality model (Deaton and Paxson, 2004). Log mortality,  $\log[m(x, t)]$ , is modeled as:

$$\log[m(x, t)] = a(x) + b(x)k(t) + \vartheta(x, t) \quad (2.11)$$

$$k(t) = \mu + k(t-1) + \omega(t), \quad (2.12)$$

where  $\log[m(x, t)] \equiv \log\left[\frac{D(x, t)}{E(x, t)}\right]$ . At time  $t$  for ages  $x$ , death counts are denoted as  $D(x, t)$ , and  $E(x, t)$  represents the population size.  $a(x)$  tracks the general mortality level in the population, and  $b(x)$  represents the relative mortality level that depends on  $k(t)$ , the stochastic time trend, which captures general mortality improvements in the population.  $k(t)$  is assumed to follow a random walk with drift  $\mu$ . The Lee-Carter model is estimated using a generalized non-linear model with deaths modeled as a Poisson-distributed variable and by imposing the identifiable restrictions  $\sum_x b(x) = 1$  and  $\sum_t k(t) = 0$ ; see Currie (2016). The fitted Lee-Carter parameters follow Cairns et al. (2009)'s estimates, which cover the mortality of the US population for a similar time and age group. Appendix (2.6.1) contains the estimates.

The three parameters on the right side of equation (2.11) describe the evolution of mortality rates in the age and time dimension. The Lee-Carter model also contains two mortality risk variables: unsystematic,  $\vartheta(x, t)$ , and systematic,  $\omega(t)$ . The latter measures the mortality evolution that affects the entire population. The former quantifies the risk of the number of deaths. Even if future population estimates and mortality rates were correct,  $\vartheta(x, t)$  would still be random due to stochastic deaths; see Cairns et al. (2008). As a result,  $\vartheta(x, t)$  is more variable in the older ages due to sparser populations. Bisetti et al. (2017) and Friedberg and Webb (2007) define excess mortality,  $EM_t$ , as the observed average mortality rates in excess of the fitted mortality rates across

ages in the population

$$\begin{aligned}
 \log (EM_t) &= \frac{1}{X} \sum_x \left\{ \log [m(x, t)] - \log [\hat{m}(x, t)] \right\} \\
 \log (EM_t) &= \underbrace{\frac{1}{X} \omega(t)}_{SM_t} + \underbrace{\frac{1}{X} \sum_x \vartheta(x, t)}_{UM_t},
 \end{aligned} \tag{2.13}$$

where  $X$  is the total number of ages in the population,  $\log [\hat{m}(x, t)]$  is fitted mortality rates, and I have used the property  $\sum_x b_x = 1$ . The latter term for unsystematic mortality risk ( $UM_t$ ), dominates the former ( $SM_t$ ) as observed by Bisetti et al. (2017). Moreover, the part of unsystematic mortality ( $UM_t$ ) arising from ages above 90 years of age overshadows the equivalent emerging from ages below 90 as shown in the top left panel in Fig. (2.1). Thus, I only use unsystematic mortality from ages above 90 years of age in this analysis, as it is sufficient to describe excess mortality in equation (2.13).

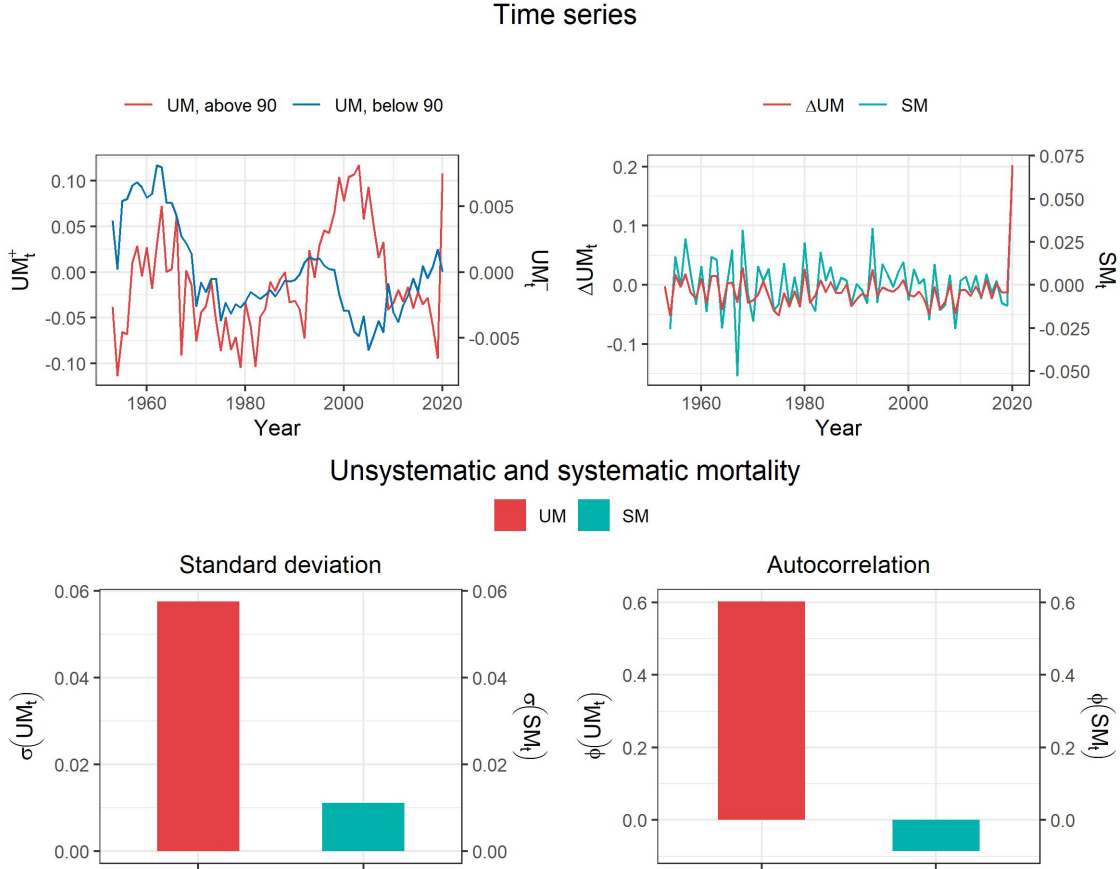


Fig. 2.1: The plots depict descriptive statistics and the evolution of unsystematic mortality (UM) and systematic mortality (SM) across time. The top-left plot depicts unsystematic mortality differentiated between above 90 years of age (red), and between 65 and 89 years of age (blue), and displays the time series of the two variables. For the remaining three plots, unsystematic mortality above 90 years of age is denoted as UM. The top-right plot illustrates systematic mortality and the first-difference of unsystematic mortality across time. The two plots at the bottom illustrate standard deviations,  $\sigma$ , and autocorrelations,  $\phi$ , of UM and SM. The sample covers 1953 to 2020.

The top left plot in Fig. (2.1) shows the evolution of unsystematic mortality for ages less than or greater than 90 years of age. As the different scales indicate, unsystematic mortality risk is largest at the older ages. For above 90 years of age, unsystematic mortality varies between 10% lower or higher than the Lee-Carter average, represented by 0, whereas the equivalent of below 90 years of age ranges between 0.5% higher or lower. The same applies to systematic mortality risk, which generally fluctuates between -2 and 2%. The bottom of Fig. (2.1) shows the differences in fluctuations. Unsystematic mortality has a standard deviation five times higher than systematic mortality. Moreover, it is more persistent, cf. the positive autocorrelation of approximately 60%

for unsystematic mortality.

However, unsystematic and systematic mortality rates are connected, as indicated by the top right plot in Fig. (2.1). In this sample, changes in unsystematic mortality are 83% correlated with systematic mortality. Thus, growth in unsystematic mortality between adjacent time periods approximates a systematic mortality shock, which heavily affects the entire mortality distribution's mortality expectations and life expectancy; see Denuit and Dhaene (2007) and Denuit (2008). Moreover, the relationship between unsystematic mortality and systematic mortality holds well for COVID 19 in 2020, where unsystematic mortality was 20% higher than expected. Systematic mortality increased by approximately 7.5%, leading to a decrease in life expectancy of 1.26 years for a US resident aged 65, according to the Human Mortality Database. This is in contrast to the conjecture Cairns et al. (2020) put forward, i.e., that the impact of COVID 19, which affected mainly the elderly, would be modest for the surviving population. The connection between unsystematic and systematic mortality has not been examined in the actuarial literature. The two mortality risks could be related due to time autocorrelation of consecutive deaths, which is assumed Poisson distributed. For example, the autocorrelation ( $\phi$ ) of two Poisson-distributed random variables can be denoted as  $\phi [\vartheta(x, t+1), \vartheta(x, t)] = a_x^2 + a_x$ ; see Miller and Childers (2012, Eq. 8.51), where  $a_x$  represents the average mortality rates for ages  $x$  across time. As the mortality rates in the older population are high, they may convey information on systematic mortality risk in the time dimension.

The interpretation of the unsystematic mortality shock is as follows: a 1% positive unsystematic mortality in period  $t$  means that the average mortality rate was 1% higher than expected from the Lee-Carter fit. A higher unsystematic mortality rate can be related to two factors. First, transitory components, e.g., a severe flu season can impact the older ages severely; see Dushoff et al. (2006). Second, differences in cohort-specific mortality influence the persistent pattern of the unsystematic mortality variable; see Cairns et al. (2009). Thus, peaks and troughs in unsystematic mortality can be interpreted as cohorts with ages above 90 that have experienced higher mortality rates than expected. For instance, in 2000 the peak suggests that cohorts born around 1910 had generally higher mortality rates than the other cohorts in the sample. As economic prosperity is



an important factor for mortality improvements (Brenner, 2005), the average high mortality rate of the cohort around 1910 is less surprising.

### **2.3.3 Descriptive statistics - financial data**

I use portfolios sorted by industry and dividend-prices, as defined in Fama and French (1997, 1988). Both datasets are extracted from Kenneth French' website (French, 2020). The analysis focuses on real returns and dividend-price ratios as one can infer the impact of a predicting variable on dividend growth by comparing it to the predictive ability for the dividend-price ratio and return; see Engsted et al. (2012). Moreover, the return information contained in dividend announcements reflects permanent earnings changes, which supports the use of dividend-price ratios and not other pricing ratios; see Ham et al. (2020).

For dividend-price ratios, denoted  $dp$ , I assume that dividends are reinvested; see Cochrane (2008). For the market portfolio, represented by "Agg", I use value-weighted returns from Shiller (2015). All returns are deflated by the consumer price index based on Shiller (2015)'s dataset, and both financial variables are log transformed. Fig. (2.2) present the industry means, standard deviations, autocorrelation, and correlations between unsystematic mortality and the financial variables: log real return and log real dividend growth. Appendix (2.6.3) contains the equivalent statistics for the portfolios sorted by dividend-price ratios.

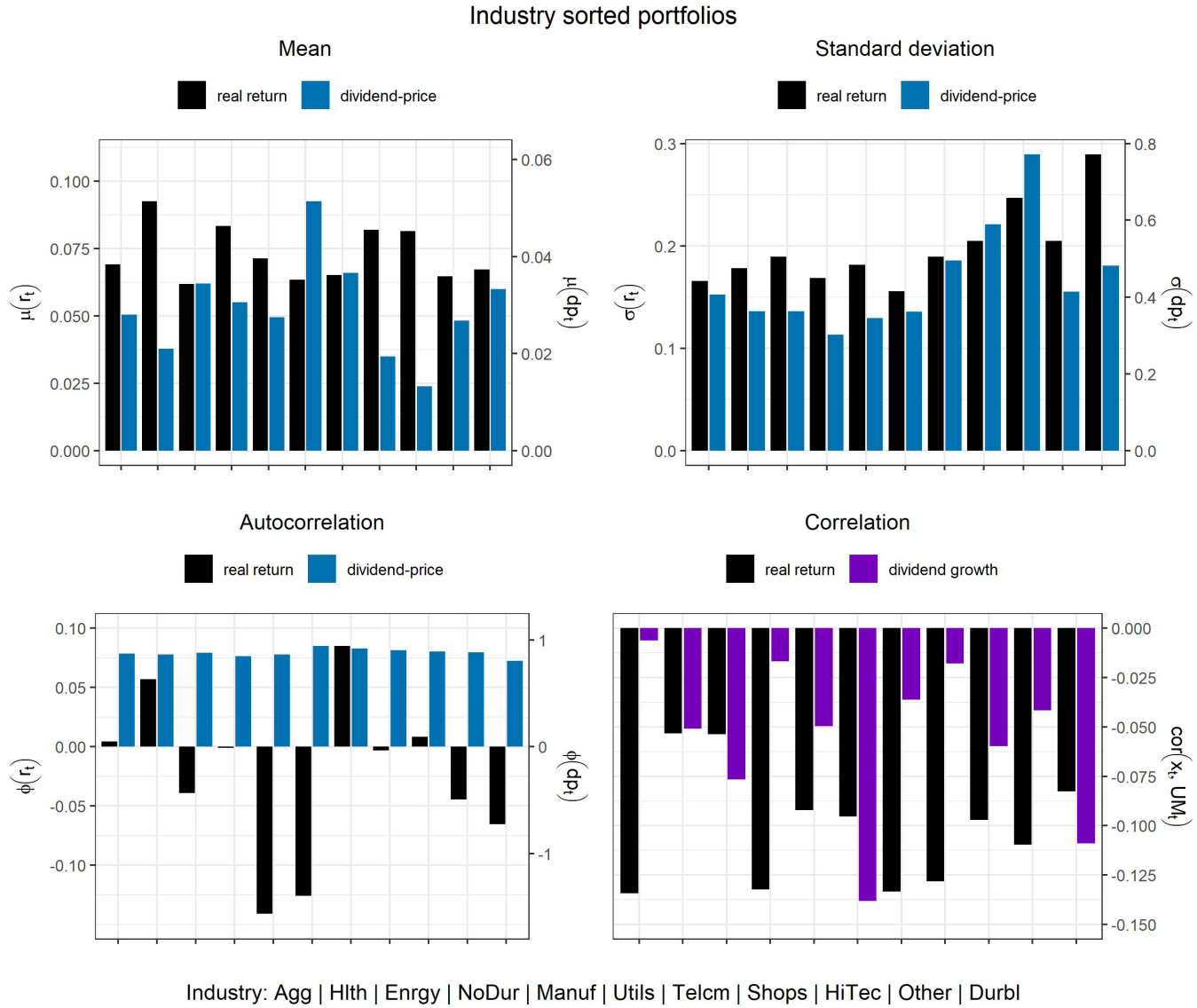


Fig. 2.2: Descriptive statistics for industry portfolios. The black bars show real returns, blue the dividend-price ratios, and purple dividend growth. All variables are log transformed, except the average dividend-price ratios in the top left graph, which are denoted in levels.  $\mu$  denote the sample average,  $\sigma$  the standard deviation, and  $\phi$  the autocorrelation coefficient. The bottom right graph contains sample correlations between unsystematic mortality (UM) and the variables real return and dividend growth. The sample covers 1953 to 2020.

Fig. (2.2) illustrates stylized facts regarding the evolution of aggregate returns. The average real return is approximately 7%, the standard deviation of the real return is just below 20%, and the autocorrelation coefficient is slightly positive. Based on the annual dataset, Fig. (2.2) also shows large industry differences in terms of the two first moments of real returns. For example, the health industry portfolio (Hlth), non-durable portfolio (NoDur), manufacturing portfolio (Manuf), shops

portfolio, and utilities portfolio (Utils) have experienced relatively large average returns compared to their volatility.

The bottom right graph also shows that the evolution of unsystematic mortality correlates negatively with real returns and dividend growth, which indicates that during times with high unsystematic mortality, expected returns and expected dividend growth are low.

## 2.4 Empirical results

This section begins by examining the return predictability of real returns when including dividend-price ratios and the unsystematic mortality variable as predictors across the 10 different industries as defined in Fama and French (1997).

### 2.4.1 Mortality impact on industry time-varying returns

I regress industry real returns of year  $t + 1$  on to their dividend-price ratios from year  $t$  and the unsystematic mortality variable in year  $t$ . Moreover, I subtract the sample mean of each variable and divide by its standard deviation:

$$\frac{r_{i,t+1} - \bar{r}_i}{\sigma(r_{i,t})} = a_i + b_{i,1} \frac{dp_{i,t} - \bar{dp}_i}{\sigma(dp_{i,t})} + b_2 \frac{UM_t - \bar{UM}}{\sigma(UM_t)} + v_{i,t+1}. \quad (2.14)$$

Due to the large difference in standard deviations between unsystematic mortality and dividend-price ratios, this allows a comparison of the magnitude of the predictability from the two predictor variables. In addition, I multiply the regressor's coefficients with  $\sigma(r_{i,t})$ . The interpretation from the rescaling is how much next year's return varies due to a standard deviation event from the predictors. This is in accordance with Cochrane (2008), who stresses the economic significance of variation in the dividend-price ratio indicating time-varying returns. Fig. (2.3) illustrates the regression estimates and confidence intervals.

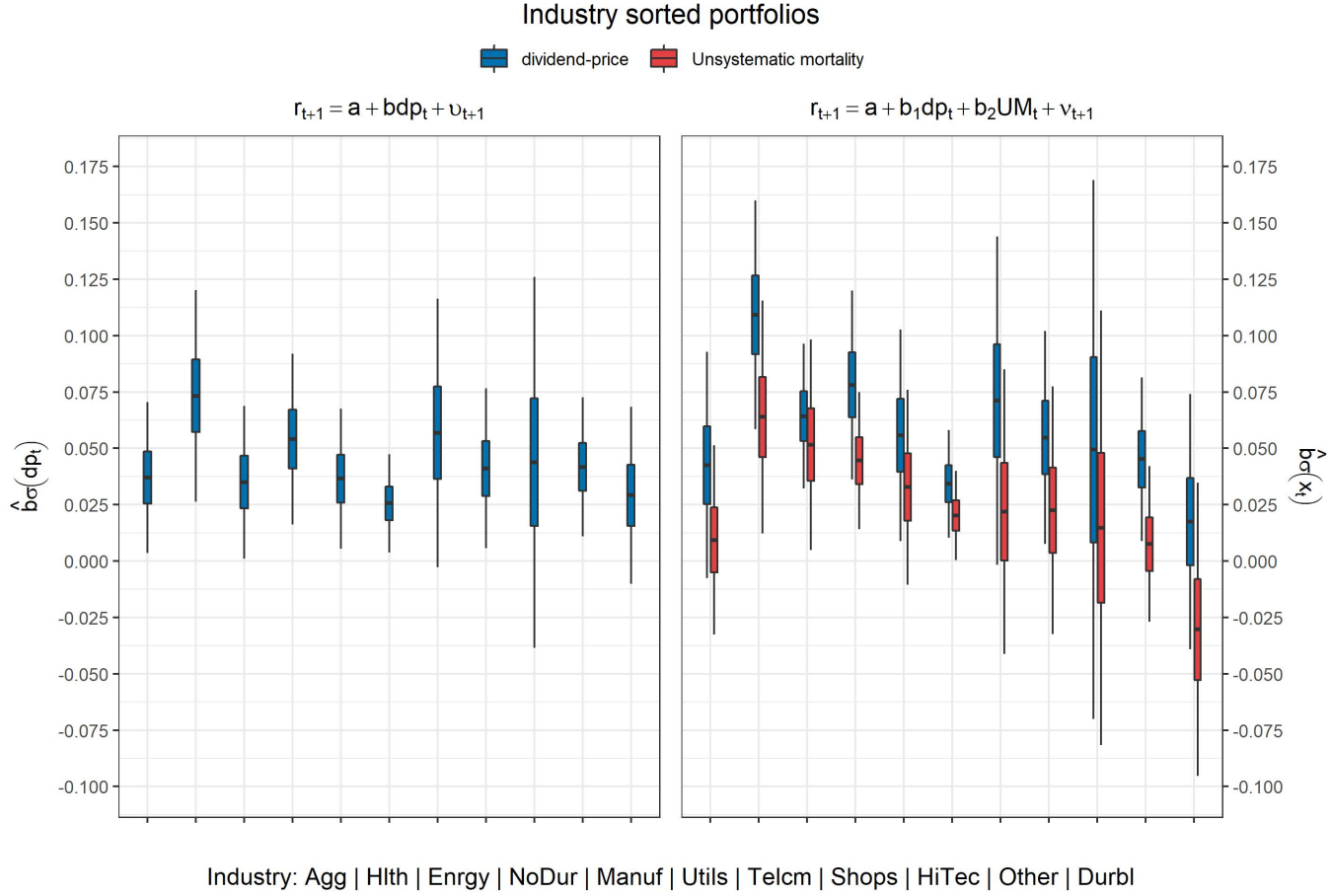


Fig. 2.3: Predictive real return regression with dividend-price ratios and unsystematic mortality as predictors. The predictive estimates from dividend-price ratios are blue and the unsystematic mortality equivalents are red. The variables are standardized to be standard normally distributed, and the estimate is multiplied with the standard deviation of the log real return. The interpretation based on the standardized estimates is how much real returns at time  $t + 1$  increase following a standard deviation event from the predictor variable. The lines parallel to the x-axis are the point estimates, the red and blue boxes cover the 50% confidence interval, and the whiskers represent the 95% confidence interval. Standard errors are the heteroscedasticity and autocorrelation consistent estimators using automatic lag selection (Newey and West, 1987, 1994). The sample covers 1953 to 2020.

A standard deviation increase in the dividend-price ratio leads to an increase in the contemporaneous return of the aggregate market of approximately 4%, as in Cochrane (2008). In addition, an increase in the dividend-price ratio increases contemporaneous returns for all industries. The dividend-price ratio primarily conveys consistent information on expected returns and not expected dividend growth (Appendix (2.6.3)). However, for the durable (Durbl) and hi-tech (HiTec) portfolios, the point estimates of the proportion of return variance explained by dividend news are the largest. Moreover, the point estimates of the proportion of return variance explained by

return news are the largest for the health industry portfolio and lowest for the durable sector. This suggests that for most industry portfolios, a fall in expected returns leads to an increase in contemporaneous returns.

Thus, following a standard deviation increase of unsystematic mortality, the contemporaneous returns of all industries' increase, except the durable portfolio. Furthermore, all estimates of the dividend-price ratio increase. Correspondingly, Lettau and Van Nieuwerburgh (2008) argue that the point estimate of regressing dividend-price ratios on future returns is downward biased, arguing that the downward bias arises due to averaging dividend-price ratios over structural breaks. In contrast, the downward bias arises in this setting due to unsystematic mortality constituting an omitted variable not captured by investors. As unsystematic mortality predicts contemporaneous returns positively, and if dividend-price ratios are predicted negatively, using dividend-price ratios in a univariate regression would lead to a downward bias, which supports the use of predictive regressions to infer the impact of unsystematic mortality on returns; see Ball (1978). However, for the aggregate value-weighted portfolio, unsystematic mortality predicts a 1% higher contemporaneous return, leading to a negligible downward bias.

Nonetheless, the 1% higher contemporaneous return is equivalent to the magnitudes estimated in studies examining the return predictability from macro variables. For example, Maio and Philip (2015) extract six principal components from 124 macro variables and show that an increase from the principal components alter annual contemporaneous returns between -3.6 and 1.6%, where one of the six principal components is statistically significant. This signifies that the impact of unsystematic mortality through changes in expenditures is equivalent to other macro variables in the aggregate.

In addition, Fig. (2.3) clearly shows that an increase in unsystematic mortality increases industry contemporaneous returns heterogeneously. For example, the five industries exposed most to unsystematic mortality, from health to utilities, show an increase in contemporaneous returns of approximately 4.3%, on average, following a standard deviation increase in unsystematic mortality. In comparison, using only the dividend-price ratio in the predictive regression increases contemporaneous returns with approximately 4.5%, on average, for the same industries, cf. left

plot in Fig. (2.3). In contrast, the five industries least exposed show an increase in contemporaneous returns of approximately 0.6%, on average, following an increase in unsystematic mortality and an equivalent dividend-price ratio prediction as the aforementioned five industries.

The results in Fig. (2.3) follow the economic reason suggested in relation to equation (2.10). That is, an unsystematic mortality shock alters contemporaneous returns through changes in conditional expected returns and expected dividends growth. The heterogeneous impact arises from the differing consumption composition in the older ages. For example, the industry most exposed to mortality, the health portfolio, depends heavily on consumption from older ages. In particular, French et al. (2017) find that medical spending in the last three years of life accounts for approximately 20% of total medical expenditures in the US.

#### 2.4.2 Economic magnitude

To substantiate the validity of the economic magnitude, I regress industry dividend-price ratios in year  $t + 1$  to their dividend-price ratio and the unsystematic mortality variable at time  $t$ :

$$dp_{i,t+1} = a_i + c_{i,1}dp_{i,t} + c_{i,2}UM_t + v_{t+1}. \quad (2.15)$$

The equation relates to two measures of interest for this analysis. First, the equation is associated with the financial magnitude to changes in expected returns minus expected dividends, denoted as  $\xi_{t+1}$  and  $\zeta_{t+1}$ , respectively, in equation (2.6). Moreover, the economic impact for the aggregate market can be inferred from equation (2.10). As the dividend-price ratio is a function of expected returns and expected dividends, the contribution of the two cannot be disentangled. Concurrently, unsystematic mortality affects discount rates through the impact on aggregate consumption as well as heterogeneous dividend growth expectation for the specific industry. As a result,  $c_{Agg,2}$  conveys information on the aggregate loss in consumption on the aggregate market, while  $c_{i,2}$  reports information on the change in expected return minus expected dividends for industry  $i$ . Fig. (2.4) depicts the transformed estimates from equation (2.15) to reflect the two measures of interest.

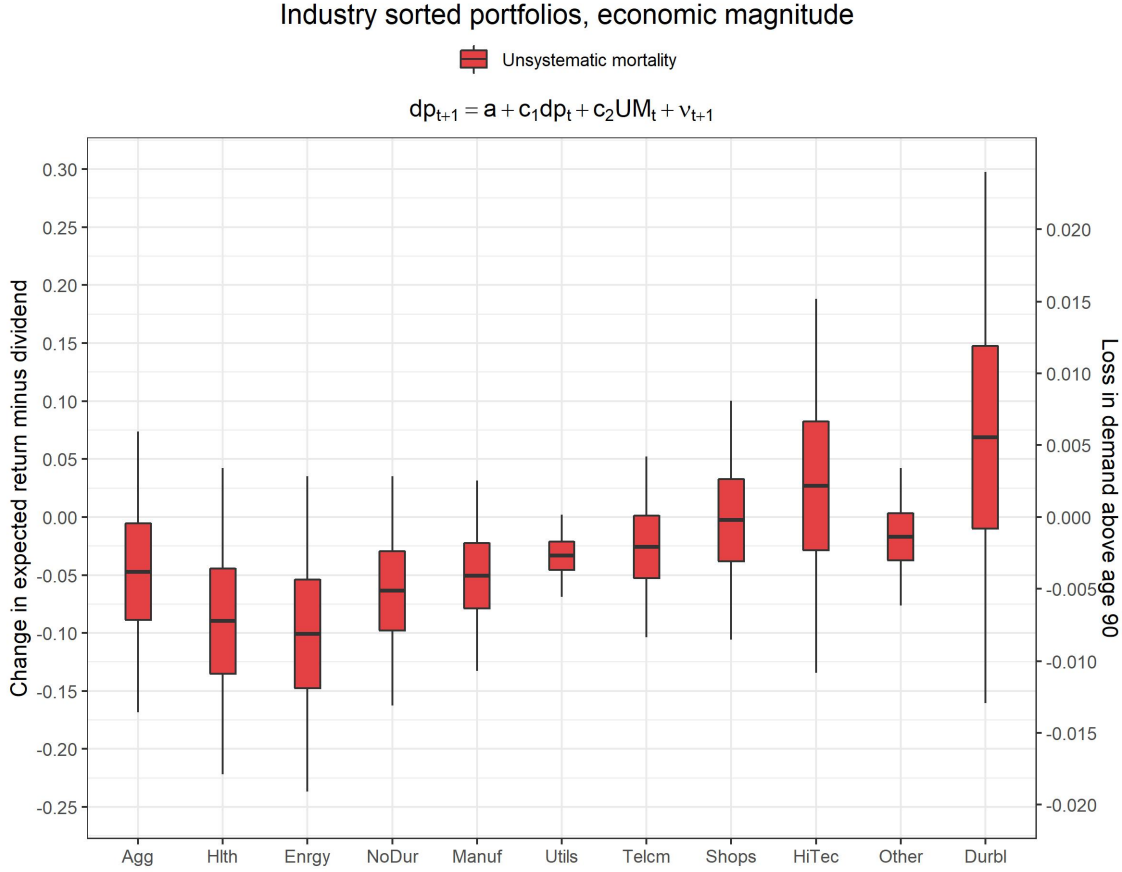


Fig. 2.4: Changes in expected returns with expected dividends growth subtracted and absolute economic loss in consumption following a marginal increase in unsystematic mortality. The left y-axis denotes the marginal contribution to changes in  $\xi_{i,t+1} - \zeta_{i,t+1}$  by multiplying each  $c_{i,2}$  by  $(1 - \rho_i \phi_i)$ , where  $\rho_i = \frac{\exp(\bar{p}d_i)}{1 + \exp(\bar{p}d_i)}$  and  $\phi_i$  denotes the autocorrelation of the dividend-price ratio for industry  $i$ . The right y-axis represents the absolute economic loss in consumption assuming that the intertemporal and intratemporal elasticity of substitution parameters are  $\sigma = 0.47$  and  $\varepsilon = 1.1$ , respectively, where the  $\hat{\xi}_{i,t+1} - \hat{\zeta}_{i,t+1}$  estimates are divided by  $(\varepsilon - \sigma) / [\sigma(\varepsilon - 1)] - 1$ . The lines parallel to the x-axis are the point estimates, the red boxes cover the 50% confidence interval, and the whiskers represent the 95% confidence interval. Standard errors are the heteroscedasticity and autocorrelation consistent estimators using automatic lag selection (Newey and West, 1987, 1994). The sample covers 1953 to 2020.

The average drop in expected return with expected dividend growth for the value-weighted portfolio subtracted following a marginal increase in unsystematic mortality amounts to  $-4.7\%$ . In comparison, the average across the 10 industry portfolios is  $-2.8\%$ . According to equation (2.10), the estimate is amplified by  $(\varepsilon - \sigma) / [\sigma(\varepsilon - 1)] - 1$ . The average estimate of  $\varepsilon$  across industries in Hall (2018) is 1.10; see Appendix (2.6.2). Thus, the remainder input to compute the loss in aggregate consumption is the elasticity of intertemporal substitution parameter,  $\sigma$ .

Hall (1988) estimates this to be close to zero and unlikely to be much above 0.1 using aggregate consumption data. In contrast, using micro data from consumer expenditure surveys, Attanasio and Weber (1995) report estimates in the interval 0.34 to 0.48. The average proportion of people over 90 years of age relative to the full population in the data sample is 0.48%. However, Banerjee (2014) reports that average expenditures of ages above 90 years old drops about 50% relative to 50-64-year-olds. Assuming that average expenditures match 50% of the average relative population rates, an impact of  $-2.5\%$  to  $-4.7\%$  amounts to an estimate of  $\sigma$  around 0.35 to 0.47, which is within the interval Attanasio and Weber (1995) report.<sup>5</sup>

If  $\sigma$  is assumed to be 0.47, the right y-axis in Fig. (2.4) indicates the absolute loss in demand following a marginal increase in unsystematic mortality. Banerjee (2014)'s survey shows that relative health and food expenditures account for 19.4% and 11.4% of the total expenditure for consumers aged 85 or older between 2003 and 2011. This is in accordance with the parameter estimates that show a loss in demand of approximately 18% and 12% on health and non-durable consumption relative to the absolute sum of losses in Fig. (2.4). Moreover, the right y-axis estimates indicate that the loss in demand is relatively modest, accruing to on average less than 1%. However, due to the persistence of dividend-price ratios, even small deviations in demand shocks alter expected returns.

### 2.4.3 Mortality impact on value and growth's time-varying returns

The previous two sections show that unsystematic mortality affects expected returns and dividends growth negatively across industry portfolios. As unsystematic mortality constitutes latent short term and temporary deviations in consumption demand, mortality could carry information on time-varying returns across the value dimension. In this section, I repeat the analysis in section (2.4.1) using the dividend-price sorted portfolios (Fama and French, 1988) to proxy for the value anomaly.

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<sup>5</sup>The impact is inferred by solving for  $\sigma$  in the equation  $\left[ \frac{\varepsilon - \sigma}{\sigma[\varepsilon - 1]} - 1 \right] \partial D = \hat{c}_{Agg,2} \cdot (1 - \rho_{Agg} \phi_{Agg})$ , where  $\partial D = 0.48\% \cdot 0.5$ ,  $\varepsilon = 1.1$ , and the right-hand side contains empirical estimates.



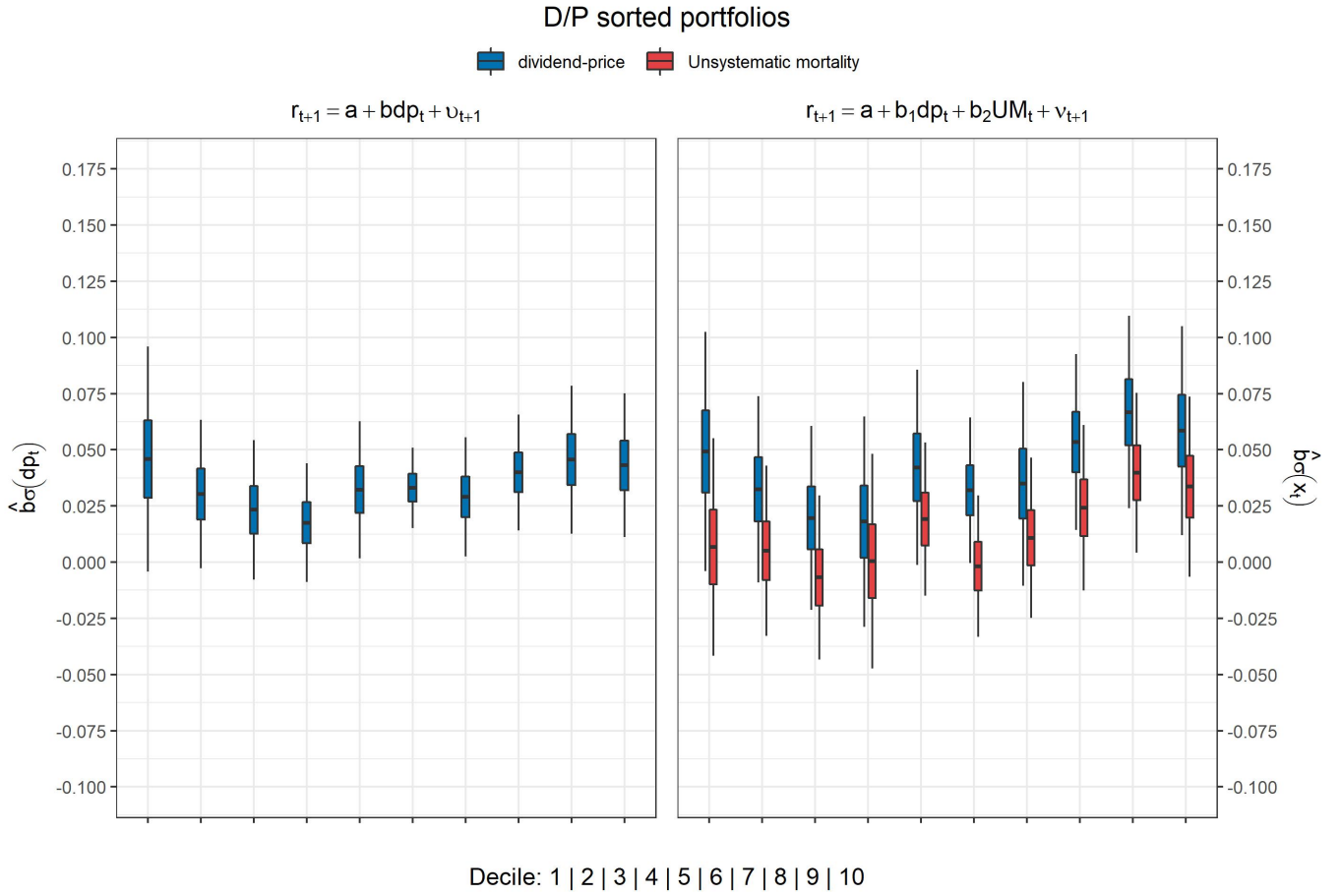


Fig. 2.5: Predictive real return regression with dividend-price ratios and unsystematic mortality as predictors. The predictive estimates from dividend-price ratios are blue and the unsystematic mortality equivalents are red. The variables are standardized to be standard normally distributed, and the estimate is multiplied with the standard deviation of the log real return. The interpretation from the standardized estimates is how much expected returns increase following a standard deviation event from the predictor variable. The lines parallel to the x-axis are the point estimates, the red and blue boxes cover the 50% confidence interval, and the whiskers represent the 95% confidence interval. Standard errors are the heteroscedasticity and autocorrelation consistent estimators using automatic lag selection (Newey and West, 1987, 1994). The sample covers 1953 to 2020.

Fig. (2.5) shows that unsystematic mortality primarily raises contemporaneous returns of the value portfolios. For example, the three highest portfolios sorted by dividend-price ratios, deciles 8-10, increase contemporaneous returns with 3.26%, on average. In contrast, the three lowest portfolios sorted by dividend-price ratios, deciles 1-3, increase contemporaneous returns with 0.15%, on average, following a standard deviation shock of unsystematic mortality. Thus, the difference between the three highest and lowest portfolios sorted by dividend-price ratios is 3.10%. In comparison, the difference in average real returns between the highest and lowest portfolios

sorted by dividend-price ratios is 3.8%. Hence, the time-varying returns account for approximately 80% of the difference in real returns between value and growth stocks. This is equivalent in magnitude to Ferson and Harvey (1991)'s finding that time-varying expected returns account for about 80% of the time-series variation in returns.

Industry composition is unlikely to explain the difference in exposure to unsystematic mortality between value and growth stocks. If the gap were explained by the formation of industries across the value dimension, the five industries most exposed to unsystematic mortality, cf. Fig. (2.3), must constitute a negative proportion for the growth portfolios and 71.9% for the value portfolios.<sup>6</sup> Based on my replication of the portfolios sorted by dividend-price ratios, I find that the five industries most exposed to unsystematic mortality accounts for 44.2% and 50.5% of the value weighted stocks in the growth and value portfolios, respectively. Alternatively, the differential exposures may be related to the observation that value stocks have higher sensitivities with respect to aggregate cash flow rather than discount rate news, as suggested by, e.g., Campbell and Vuolteenaho (2004) and Lettau and Van Nieuwerburgh (2008).

However, this would require that mortality changes impact cash flows or discount rates permanently; see Campbell (1991). Since unsystematic mortality risk is temporary, it is improbable that the unsystematic component of mortality alone can explain mortality's influence on the value anomaly. Instead, two consecutive shocks of unsystematic mortality proxy a systematic mortality shock; cf. Fig. (2.1). This would suggest that systematic mortality risk may illuminate the observed differences in risk-adjusted returns across the value dimension. I leave those concerns for future research.

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<sup>6</sup>The calculation can be summarized by solving the equation:  $\bar{r}_{i,UM} = \pi \cdot 4.3 + (1 - \pi) \cdot 0.6$ , for the proportion  $\pi$  with the point estimates of exposure to unsystematic mortality for portfolios  $\bar{r}_{i,UM}$ . That is, for the conditional returns:  $\bar{r}_{G,UM} = 0.15$  and  $\bar{r}_{V,UM} = 3.26$ , where  $G$  represents growth portfolios and  $V$  constitutes value portfolios.

## 2.5 Conclusion

The current belief in the actuarial and financial literature is that mortality rates and asset returns move independently. However, no studies have examined the influence of the most unpredictable part of mortality risk, unsystematic mortality, on equity returns. Financial research suggests that the time-series variation in equity real returns is largely attributed to variation in expected returns. Most studies have concluded that the dividend-price ratio is a strong predictor of long horizon returns. However, short horizon real returns contain fluctuations associated with temporary changes that are related to both changes in expected returns and dividend growth, where the former masks the latter. This motivates an analysis of the impact of unsystematic mortality on disaggregated returns, as consumers have heterogeneous preferences, which affects short term expected returns and dividend growth diversely.

I find that the influence of unsystematic mortality on contemporaneous returns varies between industries. Approximately half of the industries realize contemporaneous returns following an increase in unsystematic mortality that are of equivalent magnitudes to an increase in their dividend-price ratio. The remaining industries experience an insignificant exposure to unsystematic mortality. To assure that the variations in unsystematic mortality contain reliable information on abrupt changes in disaggregated consumption growth, I examine a two-good model, where consumers have heterogeneous preferences. The empirical changes in the next period dividend-price ratio are largely consistent with observed consumption behavior at the older ages. Furthermore, the free parameter of intertemporal elasticity of substitution that matches the consumption patterns is within the interval estimated in micro-economics studies.

In addition, I introduce a possible link between mortality risk and the value anomaly. Including unsystematic mortality in a predictive regression across portfolios sorted dividend-price ratios shows that contemporaneous returns increase heterogeneously. More specifically, value portfolio returns increase, on average, more than the growth portfolios' equivalent, which accounts for a large part of the differences in average real returns. This suggests that changes in mortality are related to the value anomaly. However, industry compositions alone cannot account for the difference; thus, future research may reveal if the differential impact is due to differences in exposure

to cash flow and discount rates.

## 2.6 Appendix

### 2.6.1 Fit of Lee-Carter Model

This section will validate the fit of the Lee-Carter model for the population of the US. (Fig. (2.6)).

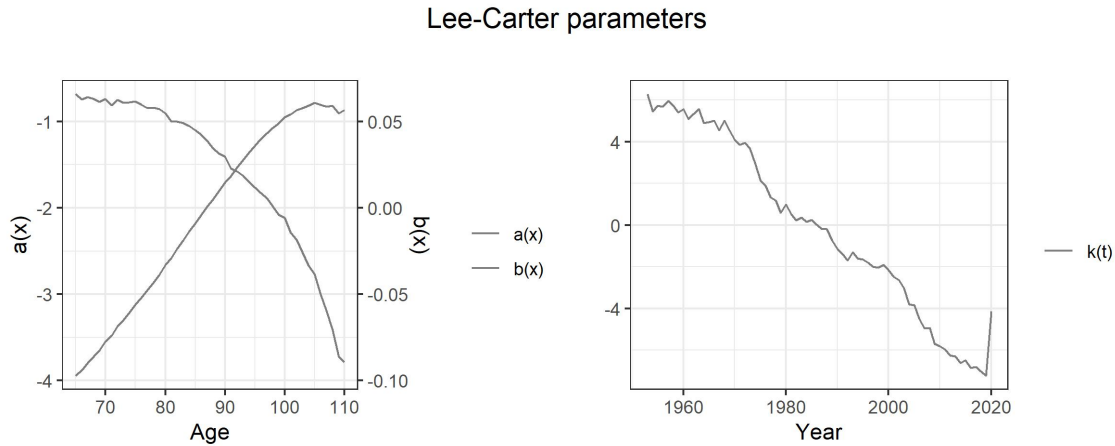


Fig. 2.6: Model fit of parameters from the Lee-Carter model. The x-axis comprises age and year for the age-dependent and time-dependent parameters. The y-axis contains the parameter estimates. The model is based on the US population for ages over 65 and between 1953 and 2020.

The base-rate mortality rate,  $a(x)$ , increases with age. The upward slope of the base-rate mortality rate implies that mortality rates are higher for more advanced ages. The notion of  $a(x)$  declining at old ages reflects late-life deceleration, i.e., the observation that mortality rate growth dampens more than expected after roughly 90 years of age; see review in Bebbington et al. (2014). The parameters  $b(x)$  and  $k(t)$  can be interpreted jointly. The time-dependent parameter,  $k(t)$ , decreases almost throughout the sample years, which reflects the mortality improvements experienced in the sample period. The exception is COVID 19, which decreased the mortality improvements. A positive  $b(x)$  indicates that these ages experienced the most mortality improvements, while a negative  $b(x)$  means that mortality for these ages tends to increase when mortality in other ages declines. The point estimates are approximately equal to that estimated in Cairns et al. (2009, Fig. 12), where they estimate the Lee-Carter model (M1) on the US population for ages 60-90 from 1968-2003.

An underlying modeling assumption when fitting the Lee-Carter model is that death counts are independently Poisson distributed. One way to evaluate this assumption is to test if the stan-

standardized residuals given by:

$$\varepsilon(t, x) = \frac{D_{x,t} - E_{x,t} \cdot \hat{m}_{x,t}}{\sqrt{E_{x,t} \hat{m}_{x,t}}} \sim N(0, 1)$$

are approximately independently identically distributed standard normal variables; see Cairns et al. (2009). A frequently encountered feature of mortality models is that they generally are overdispersed, which means that the variance of the standardized residuals are greater than one, i.e., the mortality model's fitted values have less variance than the data. This is also the case with this fit. The mean is  $-0.006$ , whereas the variance is  $27.4$ . Thus, the standardized residuals of this Lee-Carter model have a variance approximately six times larger than in Cairns et al. (2009). Overdispersion is not believed to have a significant impact on the validity of the estimates of the parameter. Rather, the main issue is in underestimating the future variability of mortality rates. Since this study focuses on variations in sample, overdispersion is not a concern for the inference of the Lee-Carter parameters; see Cairns et al. (2009).

To illustrate any systematic differences in the standardized residuals, I plotted them conditioned on the year and age in the first row in the graph to the left, in Fig. (2.7).

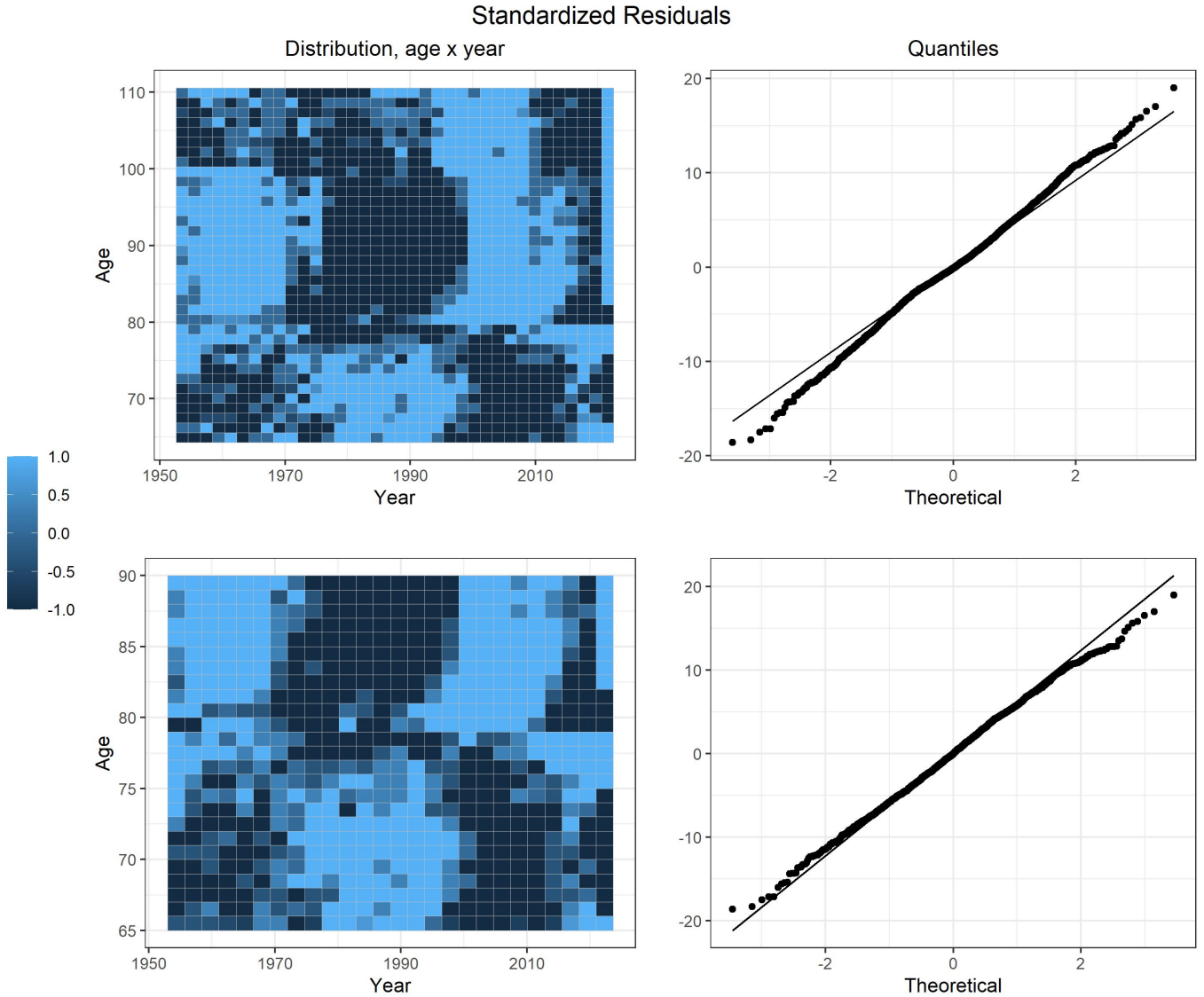


Fig. 2.7: Standardized residuals for the model fit of US population (top row) and when excluding residuals above age of 90 (bottom row). The graphs on the left-hand side illustrate standardized residuals conditioned on year and age. The graphs on the other side compare the empirical distribution of the standardized residuals with a standardized normal distribution.

Light blue indicates that there were more deaths than expected at a given age and time, whereas dark blue indicates the opposite. The shaded areas do not seem to have any pattern for a specific age group across years, except for the age groups above 90, i.e., the late-life mortality deceleration is in effect. The bias in the age dimension related to excess mortality above age 90 has implications for the standardized residuals across age and time, as depicted in the first row, right-hand side of Fig. (2.7). Roughly around the mean, the standardized residuals are reasonable distributed, whereas the variations in the tails are underestimated, i.e., the standardized residuals have heavy

tails measured by a kurtosis of 3.65, which ought to be 3.

If the ages above age of 90 are excluded, the standardized residuals change significantly. Now the standardized residuals conditioned on year and age do not have systematic differences in years nor in ages, as seen in the second row on the left-hand side of Fig. (2.7). Furthermore, the standardized residuals across ages and years are closer aligned to the shape of a normal distribution, as depicted in the second row on the right-hand side of Fig. (2.7). This is stressed by the kurtosis of 3.02, i.e., excluding excess mortality above the age of 90 means that there are not more cases of excess mortality than to be expected in the tails of the distribution. The model is still overdispersed with a variance of 34 for the standardized residuals. However, as argued above overdispersion is not important for the reliability of the parameter estimates.

## 2.6.2 The N-good consumption economy

**Deduction of stochastic discount factor** The deduction of the stochastic discount factor in the  $N$ -good consumption economy follows the same sequence of steps as Piazzesi et al. (2007)'s two-good model. That is, the representative consumer has a time additive utility function:

$$u(C_t) = \frac{C_t^{1-\frac{1}{\sigma}}}{1-\frac{1}{\sigma}},$$

where  $C_t$  denotes aggregate consumption,  $\sigma$  represents the intertemporal elasticity of substitution, and the utility of consumption is maximized over  $u(C_t) = \mathbb{E}[\sum_{t=0}^{\infty} \beta^t U(C_t)]$ , where  $\beta$  is the subjective discount factor.

In addition, assume that the aggregate consumption is aggregated over  $N$  goods via the generalized constant elasticity of substitution utility function:

$$C_t = \left[ \sum_{i=1}^N \omega_i c_{i,t}^{(\varepsilon-1)/\varepsilon} \right]^{\varepsilon/(\varepsilon-1)},$$

where  $\varepsilon$  is the elasticity of intratemporal substitution and  $\omega_i$  is the share parameter for good  $i$ , i.e.,  $\sum_i \omega_i = 1$ . Without loss of generality, suppose there are three goods, i.e.,  $c_t$ ,  $h_{1,t}$ , and  $h_{2,t}$ , where  $c_t$  is the numéraire good and  $h_{i,t}$  denotes the composite goods. Then the stochastic discount factor is



denoted as:

$$\begin{aligned}
M_{t+1} &= \beta \frac{u'(C_{t+1}) C'_{c,t+1}}{u'(C_t) C'_{c,t}} \\
&= \beta \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{1}{\sigma}} \left[ \frac{1 + \omega_{h_1} \left( \frac{h_{1,t+1}}{c_{t+1}} \right)^{(\varepsilon-1)/\varepsilon} + \omega_{h_2} \left( \frac{h_{2,t+1}}{c_{t+1}} \right)^{(\varepsilon-1)/\varepsilon}}{1 + \omega_{h_1} \left( \frac{h_{1,t}}{c_t} \right)^{(\varepsilon-1)/\varepsilon} + \omega_{h_2} \left( \frac{h_{2,t}}{c_t} \right)^{(\varepsilon-1)/\varepsilon}} \right]^{\frac{\sigma-\varepsilon}{\sigma[\varepsilon-1]}}. \quad (2.16)
\end{aligned}$$

The expression in equation (2.16) follows readily from the two-good model. Moreover, as the representative consumer equates all marginal rates of substitution, the relevant relative expenditure shares are given as:

$$\begin{aligned}
\frac{p_t^c c_t}{p_t^{h_1} h_{1,t}} &= \omega_{h_1}^{-1} \left( \frac{c_t}{h_{1,t}} \right)^{(\varepsilon-1)/\varepsilon} \\
\frac{p_t^c c_t}{p_t^{h_2} h_{2,t}} &= \omega_{h_2}^{-1} \left( \frac{c_t}{h_{2,t}} \right)^{(\varepsilon-1)/\varepsilon}.
\end{aligned}$$

Thus, equation (2.16) can be rewritten as

$$\begin{aligned}
M_{t+1} &= \beta \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{1}{\sigma}} \left[ \frac{1 + \frac{p_{t+1}^{h_1} h_{1,t+1}}{p_{t+1}^c c_{t+1}} + \frac{p_{t+1}^{h_2} h_{2,t+1}}{p_{t+1}^c c_{t+1}}}{1 + \frac{p_t^{h_1} h_{1,t}}{p_t^c c_t} + \frac{p_t^{h_2} h_{2,t}}{p_t^c c_t}} \right]^{\frac{\sigma-\varepsilon}{\sigma[\varepsilon-1]}} \\
&= \beta \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{1}{\sigma}} \left[ \frac{c_t}{c_{t+1}} \frac{p_t^c}{p_{t+1}^c} \cdot \frac{p_{t+1}^c c_{t+1} + p_{t+1}^{h_1} h_{1,t+1} + p_{t+1}^{h_2} h_{2,t+1}}{p_t^c c_t + p_t^{h_1} h_{1,t} + p_t^{h_2} h_{2,t}} \right]^{\frac{\sigma-\varepsilon}{\sigma[\varepsilon-1]}} \\
&= \beta \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{1}{\sigma}} \left[ \frac{c_t}{c_{t+1}} \frac{p_t^c}{p_{t+1}^c} \cdot \frac{p_{t+1}^c c_{t+1} + p_{t+1}^{h_1} h_{1,t+1} + p_{t+1}^{h_2} h_{2,t+1}}{p_t^c c_t + p_t^{h_1} h_{1,t} + p_t^{h_2} h_{2,t}} \right]^{-\frac{\varepsilon-\sigma}{\sigma[\varepsilon-1]}}. \quad (2.17)
\end{aligned}$$

As all expenditures spent on the consumption goods are assumed paid out as dividends, the term in the square brackets is equivalent to the aggregate dividend growth adjusted for changes in prices of the numéraire good, i.e.,  $\frac{p_t^c}{p_{t+1}^c}$ , and changes in the consumption in the numéraire good,  $\left( \frac{c_{t+1}}{c_t} \right)^{\frac{\varepsilon-\sigma}{\sigma[\varepsilon-1]}}$ . The composite goods,  $h_{1,t}$  and  $h_{2,t}$ , can be extended to include  $N - 1$  goods. Then

the stochastic discount factor in equation (2.17) takes the form:

$$\begin{aligned} M_{t+1} &= \beta \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{1}{\sigma}} \left[ \frac{c_t}{c_{t+1}} \frac{p_t^c}{p_{t+1}^c} \cdot \frac{p_{t+1}^c c_{t+1} + \sum_{i=1}^{N-1} p_{t+1}^{h_i} h_{i,t+1}}{p_t^c c_t + \sum_{i=1}^{N-1} p_t^{h_i} h_{i,t}} \right]^{-\frac{\varepsilon-\sigma}{\sigma[\varepsilon-1]}} \\ &= \beta \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{1}{\sigma}} \left[ \frac{c_t}{c_{t+1}} \frac{p_t^c}{p_{t+1}^c} \cdot \frac{D_{c,t+1} + \sum_{i=1}^{N-1} D_{i,t+1}}{D_{c,t} + \sum_{i=1}^{N-1} D_{i,t}} \right]^{-\frac{\varepsilon-\sigma}{\sigma[\varepsilon-1]}}. \end{aligned}$$

**Empirical estimate of intratemporal elasticity** Hall (2018) estimates the own-price elasticity,  $\varepsilon_{ii}$ , of 19 North American Industry Classification System sectors, which are generally equal to the intratemporal elasticity; cf. Pollak (1969). The own-price elasticity enters the maximization problem for a price-setting firm as:

$$\max_{\{p\}} \pi = p(h) \cdot h - c(h),$$

where  $p(h)$  is the price set by the firms to maximize profits,  $h$  is the consumption level conditional on prices, and  $c(h)$  is total costs. Maximizing profits with respect to  $p$  and inserting back into the profit equation leads to the profit given the optimal price:

$$\pi^* = c'(h) \cdot h \cdot \left( -\frac{1}{1 + \varepsilon_{h,h}} \right).$$

The optimal profit,  $\pi^*$ , is equal to marginal costs,  $c'(h)$ , times consumption  $h$ , and  $\left( -\frac{1}{1 + \varepsilon_{h,h}} \right)$ , where  $\varepsilon_{h,h}$  is the own-price elasticity of demand. Using Hicks (1946)' second law of demand, the intratemporal elasticity of substitution between good  $i$  and the numéraire good is equal to  $\varepsilon_{c,h} = -\varepsilon_{hh}$ . That is, assuming that cross-price elasticities between all other goods is zero, the intratemporal elasticity of substitution is inferred from the transformation:

$$\varepsilon_{c,h} = \frac{1+x}{x}, \tag{2.18}$$

where  $x$  is the estimated quantity,  $x = -\frac{1}{1+\varepsilon_{h,h}}$ , in Hall (2018). Averaging over all estimates from Hall (2018) using the transformation in equation (2.18) leads to an average intratemporal elasticity of substitution of 1.10.

**The health industry case** To exemplify, Hall (2018) estimates the own-price elasticity,  $\varepsilon_{ii}$ , for the health care and social assistance sector, equivalent to the health services portfolio, and the manufacturing sector, which contains the Standard Industrial Classification codes for the pharmaceutical drug and medical equipment industry portfolios. Hall (2018) estimates the own-price elasticity of health care and social assistance sector to be positive and, thus, the intratemporal elasticity to be negative, whereas the manufacturing sector, which includes firms producing pharmaceutical drugs and medical equipment, the elasticity is estimated to be  $\varepsilon_{c,h} = 3.43$ . This implies that the pharmaceutical drug and medical equipment industry portfolios are substitute goods with residual consumption and that the health service portfolio contains complement goods with respect to residual consumption.

Following Piazzesi et al. (2007), I estimate the intratemporal elasticity as:

$$\ln \frac{h_t}{c_t} = \text{constant} - \eta_{c,h} \cdot \ln \frac{p_t^h}{p_t^c} + \text{error}, \quad (2.19)$$

where baseline consumption data relates to nondurable goods and services expenditure less shoes and clothing and is attained from the national income and product accounts, based on Piazzesi et al. (2007). Moreover, I augment equation (2.19) with a categorical variable for these three periods: 1965-1971, 1972-2002, and 2003 to the end of the data sample due to changes in institutional factors; see Acemoglu et al. (2013).

Nondurable goods covers pharmaceutical and other medical products, whereas health care is included in service expenditures. Hence, non-health consumption refers to baseline consumption with either health care or pharmaceutical and other medical products subtracted. The result of this exercise can be seen in Table (2.1)

	Health Services	Drugs and Medical Equipment
LR	35.92	27.97
CV	[20.20]	[20.20]
$\varepsilon_{c,h}$	-1.56 (0.82)	3.07 (0.94)

Table 2.1: Intratemporal elasticity estimate between the two health consumption categories: health care services goods and drugs and medical equipment equivalents, relative to non-health consumption. The first and second row represent the Johansen test value for cointegration's likelihood ratio (LR) and the 1% critical value (CV) in square brackets. The third row shows the  $\varepsilon_{c,h}$  estimates and standard errors in parentheses.

The Johansen test for no cointegration is rejected at the 1% critical value when comparing the two top rows, using two lags and allowing for a constant term. Moreover, consistent with Hall (2018)'s evidence, the intratemporal elasticity is negative for the health service sector and within the range of value for the drugs and medical equipment sector, with a parameter estimate of 3.07.

Thus, the consumption of health services is unaffected by changes in relative prices, which yields no profits. As a result, changes in demand do not alter expected returns of the health service industry portfolio, but which is the case for the pharmaceutical drugs and medical equipment industries. This is consistent with the estimates in Fig. (2.8).

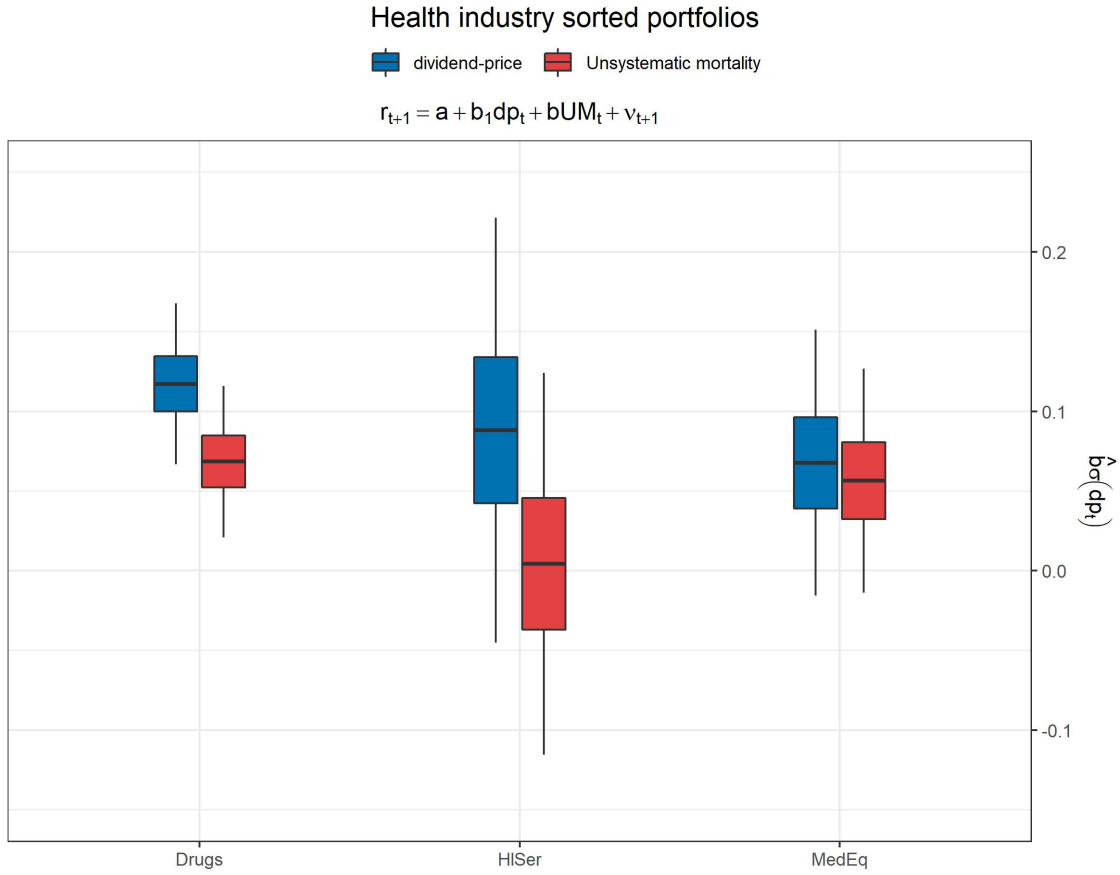


Fig. 2.8: Return predictability from dividend-price ratio and unsystematic mortality. Dividend-price ratios are blue and unsystematic mortality estimates are red. The variables are standardized to be standard normally distributed and the estimate is multiplied with the standard deviation of the log real return. The interpretation based on the standardized estimates is how much expected returns increase following a standard deviation event from the predictor variable. The lines parallel to the x-axis in the colored boxes are the coefficient estimates, the red and blue boxes are the 50% confidence interval, and the whiskers are the 95% confidence interval. Standard errors are HAC-adjusted. The sample covers 1953 to 2020.

Fig. (2.8) shows that a standard deviation increase of unsystematic mortality increases contemporaneous returns by approximately 8%, which is equivalent to the results in Fig. (2.3) for the health industry portfolio. The contemporaneous return in the health service industry portfolio, in contrast, is unaffected following changes in unsystematic mortality, which suggests that a positive intratemporal elasticity is necessary for a sector to be exposed to temporary demand changes.

### 2.6.3 Return predictability

**Return and cash-flow news, industry** This section will show the return innovation variance attributed to news about future returns ( $DR_{t+1}$ ) or dividends ( $CF_{t+1}$ ), also referred to as discount-rate and cash-flow news. The estimation relates to equation (2.2), which states that changes in the current dividend-price ratio must be due to changes in expected future returns or dividends. The news components can be extracted from:

$$Tot_{t+1} = e1^T \varepsilon_{t+1}$$

$$DR_{t+1} = e1^T \rho A (I - \rho A)^{-1} \varepsilon_{t+1}$$

$$CF_{t+1} = Tot_{t+1} + DR_{t+1},$$

where  $Tot_{t+1}$  is the total return innovation,  $e1$  denotes an indicator variable selecting the first entry from the return innovations,  $\varepsilon_{t+1}$ , and  $A$  represents the coefficient matrix from a first-order vector autoregressive regression model; see Campbell (1991).

Table (2.2) contains the results for the portfolios sorted by industry.

	$\frac{\text{Var}[DR_{t+1}]}{\text{Var}[Tot_{t+1}]}$	$\frac{\text{Var}[CF_{t+1}]}{\text{Var}[Tot_{t+1}]}$	$2 \cdot \frac{\text{Cov}[DR_{t+1}, CF_{t+1}]}{\text{Var}[Tot_{t+1}]}$
Hlth	1.21 (0.04)	0.28 (0.26)	−0.49 (0.28)
Enrgy	0.31 (0.04)	0.40 (0.25)	0.29 (0.27)
NoDur	0.60 (0.04)	0.29 (0.26)	0.11 (0.29)
Manuf	0.31 (0.03)	0.36 (0.25)	0.33 (0.26)
Utils	0.28 (0.04)	0.30 (0.26)	0.43 (0.29)
Telcm	0.79 (0.04)	0.35 (0.26)	−0.14 (0.28)
Shops	0.40 (0.04)	0.47 (0.26)	0.14 (0.28)
Agg	0.59 (0.03)	0.16 (0.25)	0.25 (0.27)
HiTec	0.33 (0.03)	0.71 (0.26)	−0.04 (0.28)
Other	0.34 (0.04)	0.48 (0.27)	0.18 (0.29)
Durbl	0.11 (0.03)	0.62 (0.25)	0.27 (0.26)

Table 2.2: Portfolios sorted by industry: The table reports the proportion of return innovation variance contributed by discount rate,  $DR_{t+1}$ , and cash-flow news,  $CF_{t+1}$ , relative to the total variance,  $Tot_{t+1}$ . It also outlines the correlation between the two news components. The top row for each industry denotes the sample estimate, whereas the bottom one shows the nonparametric bootstrap standard errors using 1,000 simulations; see e.g., Davison and Hinkley (1997).

The variance decomposition shows that the proportion of variance explained by return innovations constitutes the largest part for the aggregate market portfolio, as shown by Campbell and Ammer (1993). The point estimates suggest that most of the industry portfolio return innovations are driven by cash-flow news. However, the standard errors of the estimates for the proportion of return innovation variance contributed by returns, in the bottom row for each portfolio, have less variability than the dividend equivalents. This suggests that the dividend-price ratios identify discount-rate news more consistently.

Moreover, the portfolios with the highest proportion of return variance explained by changes in dividends growth are hi-tech (HiTec) and durable (Durbl). This suggests that for these portfolios,

the dividend-price ratio (contemporaneous return) may increase (decrease) following a decrease in expected dividends growth; cf. equation (2.6).

**Industry, tables** Tables (2.3) contains estimates for the predictive regression of dividend-price ratios and unsystematic mortality on contemporaneous returns, while Table (2.4) shows the equivalent predictive estimates on dividend-price ratios.

$r_{t+1}$	$dp_t$	$\bar{R}^2$	$dp_t$	$UM_t$	$\bar{R}^2$
Hlth	0.20 (3.06)	0.16	0.30 (4.12)	1.11 (2.56)	0.23
Enrgy	0.10 (2.11)	0.02	0.18 (4.09)	0.89 (2.24)	0.05
NoDur	0.18 (2.76)	0.09	0.26 (3.63)	0.77 (2.94)	0.12
Manuf	0.11 (2.41)	0.03	0.16 (2.33)	0.57 (1.57)	0.03
Utils	0.07 (1.85)	0.01	0.09 (2.53)	0.35 (1.45)	0.01
Telcm	0.11 (1.89)	0.08	0.15 (1.89)	0.38 (0.68)	0.07
Shops	0.07 (3.02)	0.03	0.09 (2.24)	0.39 (0.77)	0.02
Agg	0.09 (1.95)	0.03	0.10 (1.65)	0.16 (0.42)	0.02
HiTec	0.06 (1.04)	0.02	0.07 (0.78)	0.23 (0.26)	0.00
Other	0.10 (2.67)	0.03	0.11 (2.34)	0.13 (0.42)	0.01
Durbl	0.06 (1.36)	-0.01	0.04 (0.79)	-0.58 (-0.88)	-0.01

Table 2.3: Portfolios sorted by industry: Return predictability from dividend-price ratios and unsystematic mortality. Parentheses indicate t-statistics using heteroskedasticity- and autocorrelation-consistent standard errors. The data sample covers 1953 to 2020.

Similar to Fig. (2.3), Table (2.3) shows that the contemporaneous return for three industries, health, non-durable, and manufacturing, increases following an increase in unsystematic mortality. The parameter estimates entail that following a 1% increase in unsystematic mortality, contemporaneous returns increase between 0.77% and 1.11%. Multiplying the estimates with the standard deviation of unsystematic mortality of approximately 6%, cf. Fig. (2.1) allows an equivalent



interpretation like the estimates in Fig. (2.3).

$dp_{t+1}$	$dp_t$	$\bar{R}^2$	$dp_t$	$UM_t$	$\bar{R}^2$
Hlth	0.87 (15.19)	0.81	0.82 (12.76)	-0.60 (-1.32)	0.82
Enrgy	0.90 (21.61)	0.81	0.83 (18.15)	-0.72 (-1.46)	0.81
NoDur	0.85 (15.09)	0.77	0.81 (13.39)	-0.36 (-1.27)	0.77
Manuf	0.88 (19.37)	0.81	0.85 (19.97)	-0.32 (-1.20)	0.81
Utils	0.96 (25.18)	0.90	0.93 (25.97)	-0.34 (-1.84)	0.90
Telcm	0.93 (12.65)	0.86	0.91 (13.86)	-0.24 (-0.65)	0.86
Shops	0.91 (20.05)	0.87	0.91 (17.14)	-0.03 (-0.06)	0.86
Agg	0.90 (19.05)	0.82	0.87 (12.74)	-0.31 (-0.75)	0.82
HiTec	0.90 (8.68)	0.84	0.91 (7.82)	0.25 (0.34)	0.83
Other	0.90 (35.04)	0.84	0.89 (37.20)	-0.12 (-0.54)	0.84
Durbl	0.89 (15.36)	0.74	0.91 (13.42)	0.38 (0.62)	0.73

Table 2.4: Portfolios sorted by industry: Predictive dividend-price regression using current dividend-price ratios and unsystematic mortality. Parentheses indicate t-statistics using heteroskedasticity- and autocorrelation-consistent standard errors. The data sample covers 1953 to 2020.

Table (2.4) presents the predictive regression estimates for the dividend-price ratio. Similar to Fig. (2.4), all industry dividend-price ratios are predicted to fall following an increase in unsystematic mortality, though the hi-tech and durables industry portfolios are exempt. Multiplying the estimates in Table (2.4) with the industry portfolios' respective  $(1 - \rho_i \phi_i)$ , presented in Fig. (2.2), leads to the estimates in Fig. (2.4).

**Dividend-price sorted portfolios, descriptive statistics** Fig. (2.9) shows descriptive statistics for the portfolios sorted on dividend-price ratios, where low and high deciles represent growth and value portfolios, respectively.

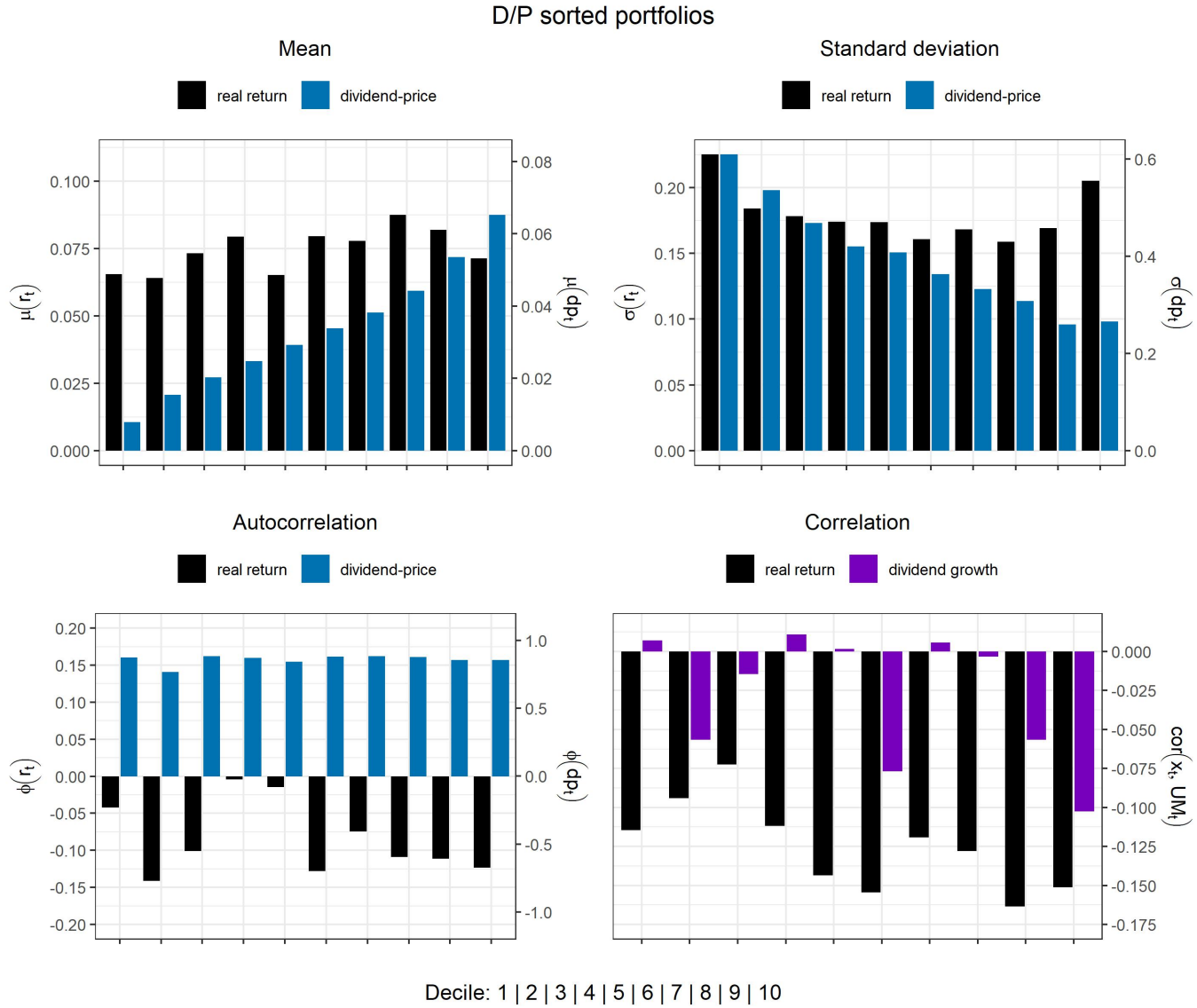


Figure 2.9: Descriptive statistics for portfolios sorted by dividend-price ratios. Real returns are black, dividend-price ratios are blue, and dividend growth is purple. All variables are log transformed, except the average dividend-price ratios, which are denoted in levels.  $\mu$  denote the sample average,  $\sigma$  the standard deviation, and  $\phi$  the autocorrelation coefficient.

The graph in the lower right-hand corner presents sample correlations between the variables' real return and dividend growth and the exogenous shock unsystematic mortality (UM). The sample covers 1953 to 2020.

The real returns (volatility) are higher (lower), on average, for value portfolios than for growth portfolios. Moreover, current unsystematic mortality is more correlated with current returns.

$r_{t+1}$	$dp_t$	$\bar{R}^2$	$dp_t$	$UM_t$	$\bar{R}^2$
Dec.1	0.07 (1.76)	0.03	0.08 (1.82)	0.11 (0.25)	0.01
Dec.2	0.06 (1.64)	0.01	0.06 (1.45)	0.10 (0.30)	0.00
Dec.3	0.05 (1.54)	0.00	0.04 (1.00)	-0.11 (-0.36)	-0.01
Dec.4	0.05 (1.38)	0.00	0.05 (0.85)	0.04 (0.09)	-0.02
Dec.5	0.08 (1.74)	0.02	0.10 (1.79)	0.32 (1.04)	0.01
Dec.6	0.09 (2.86)	0.03	0.09 (1.71)	-0.03 (-0.09)	0.01
Dec.7	0.09 (1.74)	0.01	0.11 (1.58)	0.20 (0.61)	0.00
Dec.8	0.13 (2.77)	0.05	0.17 (2.73)	0.41 (1.25)	0.05
Dec.9	0.18 (2.69)	0.06	0.26 (3.15)	0.71 (2.30)	0.09
Dec.10	0.16 (2.89)	0.03	0.21 (2.52)	0.56 (1.41)	0.03

Table 2.5: Portfolios sorted by dividend-price ratios: Return predictability from dividend-price ratios and unsystematic mortality. Parentheses indicate t-statistics using heteroskedasticity- and autocorrelation-consistent standard errors. The data sample is for 1953 to 2020.

**Portfolios sorted by dividend-price ratios, tables** Table (2.5) shows that the contemporaneous return of the three highest portfolios sorted on dividend-price ratios increases more than the three lowest portfolios sorted on dividend-price ratios. Furthermore, unsystematic mortality predicts a negative dividend-price ratio in the following period for all portfolios, cf. Table (2.6).

$dp_{t+1}$	$dp_t$	$\bar{R}^2$	$dp_t$	$UM_t$	$\bar{R}^2$
Dec.1	0.88 (22.31)	0.84	0.88 (21.29)	-0.01 (-0.02)	0.84
Dec.2	0.78 (5.96)	0.63	0.76 (5.84)	-0.28 (-0.73)	0.63
Dec.3	0.89 (26.07)	0.84	0.89 (30.11)	-0.05 (-0.25)	0.84
Dec.4	0.88 (23.87)	0.82	0.85 (24.04)	-0.50 (-1.84)	0.82
Dec.5	0.87 (19.62)	0.80	0.85 (16.82)	-0.34 (-1.42)	0.80
Dec.6	0.89 (21.20)	0.83	0.86 (17.56)	-0.29 (-1.04)	0.83
Dec.7	0.89 (26.05)	0.83	0.87 (21.06)	-0.14 (-0.57)	0.83
Dec.8	0.88 (22.39)	0.82	0.85 (16.06)	-0.28 (-1.04)	0.82
Dec.9	0.86 (17.22)	0.76	0.81 (11.82)	-0.42 (-1.42)	0.76
Dec.10	0.85 (14.18)	0.74	0.79 (13.02)	-0.66 (-4.01)	0.76

Table 2.6: Portfolios sorted by dividend-price ratios: Predictive dividend-price regression using current dividend-price ratios and unsystematic mortality. Parentheses indicate t-statistics using heteroskedasticity- and autocorrelation-consistent standard errors. The data sample covers 1953 to 2020.

# Chapter 3

## 3 Longevity risk and the value premium

### Abstract

This paper proposes that longevity risk is a key factor in understanding the value premium. Higher longevity (mortality) changes increase (decrease) the representative agent's intertemporal budget constraint. Moreover, longevity changes follow a random walk, which suggests that any impact on asset prices would be unrelated to hedging demands and, instead, could influence returns comparably to market cash-flow risk. Empirically, I find risk premiums consistent with the model and that the monthly Sharpe ratio associated with longevity risk is around 10%. In addition, my findings indicate that Fama and French (1992)'s size factor is an inferior proxy for longevity risk and that including size can lead to double counting of the longevity risk premium.

### 3.1 Introduction

Deviations from the representative agent's consumption-wealth ratio predicts and prices risk in stock returns; see Lettau and Ludvigson (2001a,b). Comparably, deviations in the equivalent ratio due to uncertain survival rates entail life insurance premiums (Pissarides, 1980; Pliska and Ye, 2007), which relates to life annuities; see Scott (2015). Individuals' stochastic survival rates can be diversified by pooling enough policyholders. However, the diversification benefits reach a saturation point where aggregate survival rates are undiversifiable, which defines systematic longevity risk; see Hari et al. (2008). The focus of this research is solely on systematic longevity risk, which is identical to negative systematic mortality risk, when analyzing longevity risk. Time-variations in longevity risk evolve as a random walk (Lee and Carter, 1992) and is perfectly positive correlated across ages; see Denuit (2008). Thus, longevity risk can be considered as an instant shock to intertemporal budget constraints, which is undiversifiable and unpredictable.

In addition, current aggregate survival rates are unobservable as the underlying uncertainty, mortality rates, are published with a two-year lag. Hence, longevity risk pose an omitted risk variable that cannot be accounted for in real time. In the stock market, an omitted risk variable that yields a positive risk premium may restrain prices. According to Ball (1978), various pricing ratios such as the earnings-price and dividend-price ratios exhibit time-variations that is conjectured to be due to omitted risk variables. Furthermore, Fama and French (1993)'s size and value factors, small-minus-big (SMB) and high-minus-low (HML) are sorted on price-augmented variables, i.e., market-equity and market-to-book equity, and asserted to potentially capture rational variations in expected returns; however, the underlying rationale is left unexplained; see Fama and French (1992).

The Fama and French (1993) three-factor model's success in cross-sectional asset pricing tests has prompted several attempts to explain their empirical findings. Among others, Ghosh et al. (2017) find that unobserved consumption risk correlates with the Fama and French (1993) factors, whereas Lettau and Ludvigson (2001a)'s model suggests that positive deviations in consumption-wealth ratios is associated with a positive risk premium. Moreover, Campbell and Vuolteenaho (2004)'s model suggests that stocks' risk premia for immediate changes in the market portfolio

are higher than the equivalent regarding hedging future market news. An obvious candidate that could reconcile the above findings, would be longevity risk due to its short-term and lagged data characteristics.

Accordingly, I explore the cross-sectional pricing implications from instant longevity shocks in the stock market. To begin with, I develop a consumption-based asset pricing model, where the representative agent's intertemporal budget constraint depends on stochastic survival rates and the agent's utility function is recursive. Furthermore, I substitute consumption out to show how longevity risk affects stock returns via instantaneous consumption shocks. Next, I test the models asset pricing ability in the cross-section for multiple portfolios.

In the financial economics literature, longevity shocks are often modeled as shocks to heterogeneous consumers' time-preference; see Chen and Yang (2019); Heimer et al. (2019). Thus, their models relate to behavioral models which argue that fluctuations in market sentiments explain asset prices through excessive volatility; see Albuquerque et al. (2016); Barberis et al. (1998); Dumas et al. (2009). In contrast, my proposed model suggests that positive (negative) longevity changes result in higher (lower) intertemporal budget constraints and can be regarded as instantaneous shocks to current consumption, which the empirical findings in Chen and Yang (2019) supports. Consequently, I substitute out consumption from the resulting stochastic discount factor, following Campbell (1993), and present a model where longevity changes' risk premium is equivalent to Campbell and Vuolteenaho (2004)'s cash-flow beta with an additional risk premium attributed for the agent's preference for early resolution of uncertainty; see Epstein and Zin (1989). These proposed dynamics suggest that investor's are more risk averse to value stocks due to their covariation with cash-flow risk, which includes short-term variations in intertemporal budget constraints via longevity shocks.

I construct mimicking longevity portfolios using six market equity and book-to-market sorted portfolios using prices up until portfolio formation dates in June, following Asness and Frazzini (2013). Moreover, I construct equivalent mimicking portfolios using separately three portfolios sorted on market equity and book-to-market portfolios, respectively, which I address as exposition mimicking portfolios. Since both the size and value factors are sorted on characteristics dependent

on prices, they are both potentially exposed to the same omitted risk variable. This allows a comparison of the risk premium between the two exposition mimicking portfolios. Furthermore, the exposition mimicking portfolios may illuminate why the value, in contrast to the size factor, has been more persistent observed as a risk premium for systematic risk; see Jensen et al. (2022); van Dijk (2011).

I test the model by Fama and MacBeth (1973) regressions, assuming unconditional risk premiums using monthly data between June 1963 to December 2020, for two different specifications. The theory motivated model includes Campbell and Vuolteenaho (2004)'s model in conjunction with the longevity mimicking portfolios, whereas the empirical motivated address the Fama and French (1993) three-factor model. The test assets include 25 portfolios sorted on size and book-to-market equity. In addition, I include 25 portfolios sorted on size and investment, 25 portfolios sorted on size and operating profitability, and five industry-sorted portfolios, following Kan et al. (2013). Moreover, I report the r-squared ( $R^2$ ) statistics using both ordinary least squares (OLS) and generalized least squares (GLS) estimates. The latter refers to the proposed factors mean-variance efficiency (Kandel and Stambaugh, 1995) and relates to other asset pricing tests; see Lewellen et al. (2010).

Empirically, I find that longevity changes is highly correlated with the size and value factors. The mimicking portfolio extracted from market-equity and book-to-market sorted portfolios is 80% correlated with the value factor and 9% with the equivalent size factor. The exposition mimicking portfolios are 81% and 52% correlated with the value and size factor, respectively. Furthermore, the mimicking portfolios are 50%, 45%, and 24% correlated with longevity changes suggesting that the portfolio extracted from size and value sorted portfolios capture best the uncertainty from longevity changes. Consistent with the proposed theory, the high correlation occurs due to higher sensitivities to cash-flow and longevity risk for value and small portfolios. Regarding the theory-motivated model, I find that the longevity mimicking portfolio yields a marginally higher risk premium than Campbell and Vuolteenaho (2004)'s cash-flow beta, which can be attributed to the agent's preference for early resolution of future uncertainty. In contrast, including the exposition mimicking portfolios yield negative cash-flow and discount-rate betas. Moreover,



the pricing error becomes significantly positive and both  $R^2$ -metrics doubles. A similar pattern emerge, when testing using the mimicking portfolios in conjunction with the market portfolio. The market portfolio's risk premium increases when including the longevity mimicking portfolio and decreases when using the exposition mimicking portfolios. Furthermore, the pricing error is insignificant for the single mimicking portfolio and significant positive for the exposition mimicking portfolio. Thus, these findings conclude that the value premium proxy for longevity risk, whereas the size premium is an inferior proxy for the equivalent risk. Consequently, including both the size and value factor leads to double-counting the risk premium associated with longevity risk.

### 3.2 Methodology

Consider a representative agent endowment economy, where all wealth is tradable; see Campbell (1993). Denote all wealth including human capital at the beginning of period  $t + 1$  as  $W_{t+1}$ ,  $C_t$  is consumption at time  $t$ , and  $R_{m,t+1}$  represents the gross rate of return on aggregate wealth, i.e., the market portfolio. In addition, assume that the probability of survival is stochastic and as a consequence savings in period  $t$  may be lost in period  $t + 1$ . Then the intertemporal budget constraint can be written as

$$W_{t+1} = \frac{L_{t+1}}{L_t} R_{m,t+1} (W_t - C_t), \quad (3.1)$$

where  $L_{t+1}$  denotes the survival probability from time  $t$  to  $t + 1$  and  $L_{t+1}/L_t$  represents the conditional survival equivalent. In the life insurance literature, the wealth accumulation equation in discrete time is commonly denoted as  $W_S = \sum_{i=0}^S R_{t+i} [W_{t+i} - C_{t+i}]$ , where  $s \leq T$  is the stochastic time of death; see Pissarides (1980); Pliska and Ye (2007). In insurance markets, individual agents are willing to pay a life insurance premium to hedge against stochastic survival rates. Thus, the augmented intertemporal budget constraint in equation (3.1) introduces the equivalent possibility of agents willing to hedge aggregate survival rates.

In addition, equation (3.1) follows readily the intertemporal budget constraint examined in

Campbell (1993). Thus, revisions in current consumption rates follow as

$$\begin{aligned}
c_{t+1} - \mathbb{E}_t[c_{t+1}] &= (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=0}^{\infty} \rho^j r_{m,t+1+j} \\
&\quad - (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=1}^{\infty} \rho^j \Delta c_{t+1+j} \\
&\quad + (\mathbb{E}_{t+1} - \mathbb{E}_t) \Delta l_{t+1}.
\end{aligned} \tag{3.2}$$

In accordance with Campbell (1993), this equation states that an upward surprise in current consumption is due to an unexpected return on wealth today,  $(\mathbb{E}_{t+1} - \mathbb{E}_t) r_{m,t+1}$ , higher returns in the future,  $(\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=1}^{\infty} \rho^j r_{m,t+1+j}$ , or because future consumption growth rates are expected to decline. In addition, a positive current consumption shock can be due to unexpected higher longevity expectations,  $(\mathbb{E}_{t+1} - \mathbb{E}_t) \Delta l_{t+1}$ . Note that higher expected changes in longevity,  $(\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=1}^{\infty} \rho^j \Delta l_{t+1+j}$ , is omitted and that probabilities of survival can be defined as longevity changes. In equation (3.1), the probability of survival varies across ages, which stipulates potential heterogeneous intertemporal budget constraints. However, aggregate changes in survival probabilities across time are conventionally captured by a single time-varying parameter that follows a random walk, which indicates that longevity changes are permanent and fully summarized by current changes; see Lee and Carter (1992). In addition, these time-varying changes in longevity affect different ages with a perfect positive dependence, which is termed comonotonicity; see Denuit (2008).

### 3.2.1 The stochastic discount factor

I examine the asset pricing implications stemming from the approximate budget constraint in equation (3.2) combined with a loglinear Euler equation. I assume that the representative agent has standard recursive preferences (Epstein and Zin, 1991) as

$$\mathcal{U}_t = \left\{ (1 - \beta) C_t^{\frac{1-\gamma}{\theta}} + \beta^* \left[ \mathbb{E}_t \mathcal{U}_{t+1}^{1-\gamma} \right]^{\frac{1}{\theta}} \right\}^{\frac{\theta}{1-\gamma}}, \tag{3.3}$$

where  $\mathcal{U}_t$  and  $\mathcal{U}_{t+1}$  is current and future utility, respectively,  $C_t$  is consumption at time  $t$ ,  $\gamma$

denotes relative risk aversion,  $\theta = \frac{1-\gamma}{1-\frac{1}{\psi}}$ ,  $\psi$  is the intertemporal elasticity of substitution, and  $\beta^*$  represents the subjective discount rate. The  $\beta^*$  could include the conditional survival probability,  $\frac{L_{t+1}}{L_t}$  following most heterogeneous agents' models, e.g., Gomes and Michaelides (2008). This introduces the possibility of endogenous behaviour to increase life expectancy as considered in Bommier et al. (2020), which intuitively suggests that the agent's longevity increases is due to investments in mortality risk reduction; see Hall and Jones (2007). However, Acemoglu et al. (2013) show that US consumers' life expectancies have not increased due to wealth effects. Rather, Acemoglu et al. (2013) propose that exogenous factors, such as the introduction of Medicare and Medicaid and pharmaceutical innovations have increased life expectancies in the US, which this paper's theory reflects.

As the intertemporal budget constraint in equation (3.2) has been augmented with longevity changes,  $L_{t+1}/L_t$ , the stochastic discount factor will differ from the standard case examined in Epstein and Zin (1991). Now the return on the aggregate wealth portfolio,  $R_{w,t+1}$ , includes the conditional survival probability in addition to consumption changes and the future utility certainty equivalent; cf. equation (3.13) in Appendix (3.6.1). The stochastic discount factor is expressed as

$$M_{t,t+1} = \beta^\theta \left( \frac{L_{t+1}}{L_t} \right)^{\theta-1} \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{\theta}{\psi}} (R_{w,t+1})^{\theta-1}; \quad (3.4)$$

see Appendix (3.6.1). If  $L_{t+1}/L_t$  is ignored in equation (3.1), the stochastic discount factor reduces to the standard case with Epstein and Zin (1991)'s recursive preferences. Moreover, if  $\gamma = \frac{1}{\psi}$  survival probabilities become obsolete and the stochastic discount factor reduces to the standard power utility case.

The stochastic discount factor in equation (3.4) is equivalent to the longevity-derived expression considered in Chen and Yang (2019) if one multiply the equation with  $L_{t+1}/L_t$ . Chen and Yang (2019) assumes the usual intertemporal budget constraint,  $W_{t+1} = R_{m,t+1}(W_t - c_t)$ , and that changes in longevity affects the time-preference parameter,  $\beta$ , in equation (3.3), which represents how agents trade off current and future utility; see Albuquerque et al. (2016). Intuitively, this suggests that agents changes their subjective discount rate in response to changes in survival expectations. In contrast, the specification examined here presumes that changes in longevity affect the

economy through exogenous and instantaneous shocks by the intertemporal budget constraint.

The Euler equation corresponding to equation (3.4) appears as

$$1 = \mathbb{E}_t \left[ \beta^\theta \left( \frac{L_{t+1}}{L_t} \right)^{\theta-1} \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{\theta}{\psi}} (R_{m,t+1})^{\theta-1} R_{i,t+1} \right], \quad (3.5)$$

assuming that the market portfolio constitutes the aggregate wealth portfolio.

### 3.2.2 Mutually exclusive factor models

In the following, I proceed as in Campbell (1993) and assume that life expectancy changes, consumption, and asset returns are jointly log-normal distributed and homoskedastic. Calculating the expected value of the logarithm of the Euler equation, I obtain the expected return equation

$$\begin{aligned} \mathbb{E}_t r_{i,t+1} - r_{f,t+1} + \frac{1}{2} \text{Var}_t(r_{i,t+1}) &= \frac{\theta}{\psi} \text{Cov}_t(\Delta c_{t+1}, r_{i,t+1}) + (1 - \theta) \text{Cov}_t(r_{m,t+1}, r_{i,t+1}) \\ &+ (1 - \theta) \text{Cov}_t(\Delta L_{t+1}, r_{i,t+1}); \end{aligned} \quad (3.6)$$

see Appendix (3.6.1). The asset pricing model in equation (3.6) contains the Giovannini and Weil (1989) model and in addition a risk premium related to the covariance of changes in longevity with asset return  $i$ . Furthermore, the model considered here suggests that the size of the risk premiums related to the market portfolio and longevity is of equal size. If the agent has preference for early resolution of future uncertainty (Epstein and Zin, 1989), i.e.,  $\gamma < \frac{1}{\psi}$ , and  $\psi > 1$  (Vissing-Jørgensen and Attanasio, 2003), the risk premiums will be positive. Alternatively, if the agent prefer late resolution or is indifferent, the risk premiums will be negative or zero, respectively.

In contrast, Chen and Yang (2019)'s model suggests that the risk premium related to longevity should be higher than the market portfolios equivalent. In addition, Chen and Yang (2019) finds that their consumption and longevity factors yields multicollinear factor estimates. This conform with the intuition regarding the intertemporal budget constraint in equation (3.1). That is, aggregate wealth at time  $t + 1$  can be high due to low consumption rates or low survival rates,  $L_t$ , at time  $t$ .

This motivates substituting out consumption in equation (3.6) to consider the isolated impact

of the pricing ability of the market portfolio and longevity. The risk premium arising from this exercise leads to the risk premium of longevity to be a factor of  $(1 - \theta)$  higher than the risk premium of the instant return on the invested market portfolio, i.e.,

$$\begin{aligned} \mathbb{E}_t r_{i,t+1} - r_{f,t+1} + \frac{1}{2} \mathbb{V}ar_t(r_{i,t+1}) &= \gamma \mathbb{C}ov_t(r_{m,t+1}, r_{i,t+1}) + (\gamma - 1) V_{ih} \\ &+ (\gamma + 1 - \theta) \mathbb{C}ov_t(\Delta l_{t+1}, r_{i,t+1}), \end{aligned} \quad (3.7)$$

where  $V_{ih} \equiv \mathbb{C}ov_t \left[ (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=1}^{\infty} \rho^j r_{m,t+1+j}, r_{i,t+1} \right]$  represents the covariation of asset  $i$ 's return with expected future market returns. The first line in equation (3.7) is equivalent to Campbell (1993)'s model, whereas the second line is the risk premium related to changes in longevity. In Campbell (1993)'s terminology, the model is a discrete time version of Merton (1973)'s continuous-time model where the right-hand side factors are hedging portfolios for changes in the investment opportunity set. Likewise, the longevity factor hedges changes in the investment opportunity set regarding longevity risk, equivalent to the discussion after equation (3.1).

In addition, the first line regarding the market portfolio in equation (3.7) has been examined by Campbell and Vuolteenaho (2004). In their paper, they derive a two-factor model with respect to the market portfolio. Correspondingly, Campbell and Vuolteenaho (2004)'s model can be augmented to contain

$$\mathbb{E}_t r_{i,t+1} - r_{f,t+1} + \frac{1}{2} \mathbb{V}ar_t(r_{i,t+1}) = \gamma \sigma_m^2 \beta_{i,CF} + \sigma_m^2 \beta_{i,DR} + (\gamma + 1 - \theta) \sigma_l^2 \beta_{i,l}, \quad (3.8)$$

where  $\beta_{i,CF}$  and  $\beta_{i,DR}$  represent asset  $i$ 's sensitivity to market news regarding instantaneous and future changes in market returns. Furthermore,  $\beta_{i,l}$  is the marginal sensitivity of asset  $i$  with respect to longevity risk; see Appendix (3.6.1) for further clarification.

Comparing equation (3.6) with equation (3.8) highlights several properties. From equation (3.8) it is clear that the longevity factor contains an equivalent risk premium as the risk premium regarding cash-flow news after adjusting for differences in variability, which reflects the direct correspondence between surprises in consumption and longevity changes in equation (3.2). In

addition, the longevity factor obtains an extra risk premium, if the agent has preference for early resolution of future uncertainty and the intertemporal elasticity of substitution,  $\psi$ , is larger than one.

Moreover, as the sum of the risk premia from the cash-flow and discount-rate betas is equivalent to the market portfolio risk premium, it is possible to compare the risk premiums from equation (3.6) when excluding the consumption factor. If consumption, after taking longevity into account, is an important factor for cross-sectional asset pricing, then the risk premiums for longevity and the market portfolio will differ between equation (3.6) and equation (3.8) due to omitted variable bias. That is, the risk premium for the market portfolio and the longevity factor ought to be lower in equation (3.6) than in equation (3.8), since consumption is positively related to market returns and longevity changes and obtain a negative risk premium, cf. equation (3.2) and (3.6),.

### 3.3 Data and descriptive statistics

Annual time-series data for changes in longevity is obtained from the Human Mortality Database (HMD). The database contains observed population counts and deaths for each year and age for 41 industrialized countries. Moreover, the database includes the remaining life expectancy at age  $x$ . A feature characterizing life expectancies is that across ages, they share a common systematic longevity component. In the popular Lee and Carter (1992)-mortality model, time-variation in mortality rates is represented by a single parameter that follows a random walk with drift. These time-variations affect all ages with a perfect positive dependence, which Denuit (2008) denotes as comonotonicity. Thus, to extract systematic longevity risk amounts to choosing an arbitrary age  $x$  and compute the life expectancy first-order time difference. In this paper, I consider the age of 65 in the US population. The equivalent changes in life expectancies for ages 20, 40, and 80 are correlated with between 97% and 98% with changes in life expectancy for age 65. The only difference between changes in life expectancies across ages is that younger ages' life expectancy differences are more volatile than older due to a longer time horizon. However, as will become apparent later, I will adjust for that.

For constructing market cash-flow and discount-rate news, I follow Campbell and Vuolteenaho (2004)’s approach except that I exclude the term yield spread in the vector autoregressive regression as it diminish the correlation with Campbell and Vuolteenaho (2004)’s equivalent measures. Data on test assets, the market portfolio, and the Fama and French (1992) three-factor model examined later are obtained from Kenneth French’ website. I consider monthly samples between June 1963 to the latest observation in December 2020; however, I follow Campbell and Vuolteenaho (2004) in estimating mimicking portfolios as well as market cash-flow and discount-rate news for the full available data period.

### **3.3.1 Estimating mimicking portfolios**

To relate changes in longevity to the asset pricing models in Section (3.2.2), I rely on the mimicking portfolios approach, as longevity-related assets are still not traded in public markets; see Blake et al. (2019). The method has been utilized since Breeden (1979), to project non-traded factors on a set of base assets. According to equation (3.7), longevity changes ought to have close resemblance to Campbell and Vuolteenaho (2004)’s cash-flow beta, which prices risk within the size and book-to-market sorted portfolios. Thus, to motivate the construction of the mimicking portfolio, I regress time-series excess returns of size and book-to-market portfolios on to Campbell and Vuolteenaho (2004)’s two factors and additionally changes in longevity. The result is contained in Table (3.1).

$\hat{\beta}_{i,CF}$	Growth	2	3	4	Value	Diff.
Small	0.15	0.19	0.29	0.28	0.38	0.23
2	0.07	0.19	0.17	0.30	0.28	0.21
3	0.07	0.16	0.24	0.26	0.34	0.27
4	0.04	0.18	0.21	0.32	0.24	0.20
Large	0.10	0.10	0.17	0.19	0.23	0.13
Diff.	-0.05	-0.09	-0.13	-0.09	-0.15	

$\hat{\beta}_{i,DR}$	Growth	2	3	4	Value	Diff.
Small	1.10	0.82	0.70	0.46	0.64	-0.47
2	0.94	0.78	0.60	0.48	0.58	-0.36
3	0.96	0.66	0.55	0.59	0.51	-0.44
4	0.91	0.70	0.58	0.64	0.70	-0.20
Large	0.97	0.72	0.54	0.57	0.77	-0.20
Diff.	-0.13	-0.10	-0.16	0.11	0.13	

$\hat{\beta}_{i,dLE}$	Growth	2	3	4	Value	Diff.
Small	-0.06	0.06	0.11	0.17	0.18	0.23
2	0.02	0.08	0.17	0.23	0.16	0.14
3	-0.02	0.06	0.22	0.17	0.29	0.30
4	-0.01	0.11	0.17	0.29	0.32	0.33
Large	0.02	0.10	0.13	0.26	0.22	0.20
Diff.	0.08	0.04	0.02	0.09	0.04	

Table 3.1: Annual joint time-series regressions of the sensitivity of the returns of 25 portfolios sorted on size and book-to-market equity to cash-flow beta,  $\hat{\beta}_{i,CF}$ , discount-rate beta,  $\hat{\beta}_{i,DR}$ , and changes in longevity,  $\hat{\beta}_{i,dLE}$ . The differences (diff.) is derived by subtracting the most extreme cells across the two dimensions. The estimates are obtained for the period 1963 to 2020.

Consistent with Campbell and Vuolteenaho (2004), cash-flow (discount-rate) betas are clearly increasing (decreasing) for higher book-to-market portfolios, i.e., from growth to value stocks. In contrast, the difference in cash-flow and discount-rate betas are less apparent across the size dimension, where the two measures are modestly decreasing for higher market-equity portfolios. In comparison, changes in longevity ( $dLE$ ) is increasing with higher book-to-market portfolios and on average increasing for higher market-equity portfolios. This comovement pattern between cash-flow risk and changes in longevity is consistent with equation (3.8) and the a priori motivation that changes in longevity affects current consumption rates. Furthermore, it is in accordance with Chen and Yang (2019)'s empirical findings that longevity risk is related to short-term consumption



risk and stocks with shorter equity durations, which characterizes value stocks; see Lettau and Wachter (2007).

Due to the patterns observed in Table (3.1), I construct a mimicking portfolio of longevity by projecting longevity risk onto six portfolios sorted on size and book-to-market. Following Asness and Frazzini (2013), I update the return factors with prices up until portfolio formation dates in June for each year. Thus, the mimicking portfolio for longevity risk is inferred by the regression

$$dLE_t = \kappa_{R,SV}^T R_t + u_t, \quad (3.9)$$

where  $R_t$  represents the observed returns for the base assets and  $\kappa_R^T$  denotes the coefficient estimates including an intercept, following Breeden et al. (1989). The coefficients are estimated to be  $\kappa_{R,SV}^T = \begin{bmatrix} -0.17 & 0.53 & -0.42 & -0.53 & -0.22 & 0.86 \end{bmatrix}$ , where the first three portfolios represent increasing book-to-market portfolios with the lowest 50% market equity and the latter three portfolios denote the equivalent book-to-market portfolios with the higher 50% market equity. Thus, in essence the longevity mimicking portfolio is long in the portfolios in the bottom right corner and top middle of Table (3.1) and short in the remainder. The positive exposure to the bottom right loads positive on value portfolios, whereas the positive exposure to the top middle loads positively on small stocks, which are more exposed to cash-flow risk than large stocks.

To emphasize the possible connection between longevity risk and the size and value factors, I repeat the steps in equation (3.9) now using three portfolios sorted on size and book-to-market, respectively. The equivalent coefficient estimates are  $\kappa_{R,SMB}^T = \begin{bmatrix} -0.44 & 1.16 & -0.77 \end{bmatrix}$  and  $\kappa_{R,HML}^T = \begin{bmatrix} -0.48 & -0.38 & 0.83 \end{bmatrix}$ . Thus, the mimicking portfolio extracted from book-to-market portfolios are increasing in higher book-to-market portfolios, whereas the equivalent from market equity portfolios are long in the average market equity portfolio and short in both small and large market equity portfolios. These latter mimicking portfolios allow a comparison with the Fama and French (1993) three-factor model and are only for expositional use. Figure (3.1) presents the mimicking portfolios graphically.

### Time series, annual

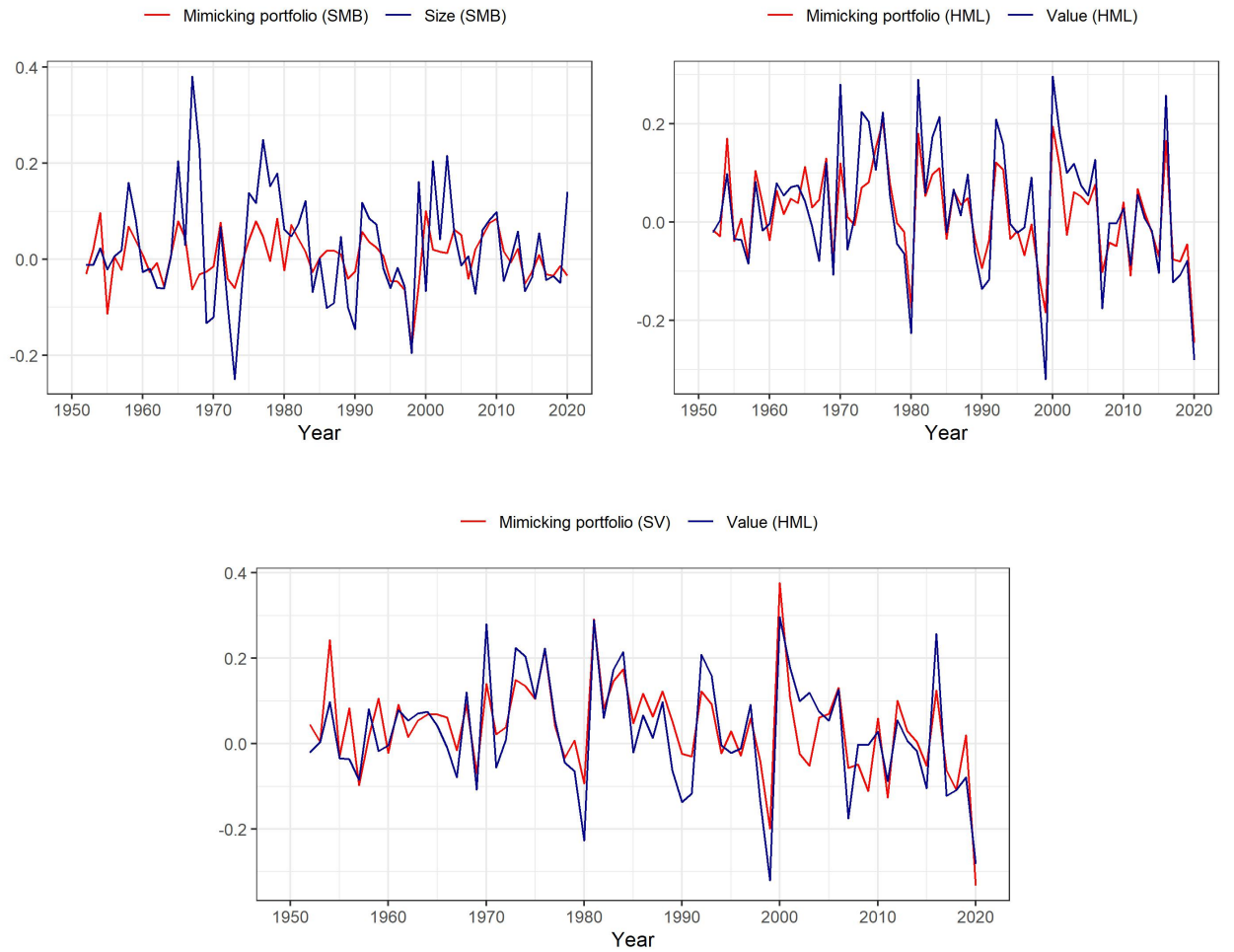


Fig. 3.1: Annual small-minus-big (SMB), high-minus-low (SMB) empirical factor returns and the mimicking factors of longevity risk. The mimicking factors constitute returns extracted from: six portfolios sorted on market equity and book-to-market equity (SV), three portfolios sorted on market equity (SMB) and book-to-market equity (HML), respectively. All return factors are constructed using prices up until portfolio formation dates in June; see Asness and Frazzini (2013). The sample period is for the full sample from 1952 to 2020.

Figure (3.1) clearly shows that both the mimicking portfolios extracted from the six portfolios sorted on size and book-to-market equity ( $FMP_{SV}$ ) and from portfolios sorted on book-to-market equity ( $FMP_{HML}$ ) are highly correlated. The  $FMP_{SV}$  is 85% correlated with the  $HML$  factor and the  $FMP_{HML}$  is 88% correlated with the equivalent factor. In addition, the mimicking portfolio extracted from portfolios sorted on market equity ( $FMP_{SMB}$ ) is 52% correlated with the  $SMB$  factor. These correlations could indicate that longevity risk proxy for a common risk for both the  $HML$  and  $SMB$  factor. This is further supported by the fact that the mimicking portfolios are

correlated with changes in longevity of 50%, 45%, and 24% for the portfolios  $FMP_{SV}$ ,  $FMP_{HML}$ , and  $FMP_{SMB}$ , respectively. Thus,  $FMP_{SV}$  track longevity risk marginally better than  $FMP_{HML}$ , whereas  $FMP_{SMB}$  is an inferior proxy for longevity risk.

### 3.3.2 Time-series relations

Using the estimated weights in the former section on annual data, I apply the same set of weights on monthly base assets' returns. As the weights from the former section approximately sum to 0, normalizing the weights to 1 leads to excessive mimicking portfolio weights. Alternatively, Balduzzi and Robotti (2010) highlights that dividing a mimicking portfolio with its standard deviation allows an interpretation of the mimicking portfolio as its Sharpe ratio on portfolios tracking the slope of the factor risk. This allows a comparison between the different longevity mimicking portfolios and additionally it avoids the concern of using different ages to proxy longevity risk, cf. the discussion in the beginning of Section (3.3). Table (3.2) contains descriptive statistics for the factors used in this paper.

	Panel A: Annual descriptive statistics							
	Mean	Median	Std. dev.	Min.	Max.	AR(1)		
$dLE$	0.11	0.12	0.21	−1.25	0.42	0.08		
$CF$	−0.00	−0.01	0.07	−0.12	0.20	0.03		
$DR$	0.01	0.02	0.14	−0.44	0.30	−0.13		
$FMP_{SV}$	0.04	0.04	0.11	−0.33	0.38	0.10		
$FMP_{SMB}$	0.00	0.01	0.05	−0.18	0.10	−0.06		
$FMP_{HML}$	0.02	0.02	0.09	−0.25	0.20	0.03		
	Panel B: Monthly means, standard deviations, and correlations							
(in %)	Mean	Std. dev.	Sharpe ratio	$SMB$	$HML$	$FMP_{SV}$	$FMP_{SMB}$	$FMP_{HML}$
$R_m$	0.63	4.33	14.64	27.16	−19.13	−9.96	10.31	−11.93
$SMB$	0.16	2.87	5.49		−15.83	−9.47	51.82	14.67
$HML$	0.25	2.71	9.31			80.00	9.15	81.23
$FMP_{SV}$	0.11	2.21	4.84				17.91	90.24
$FMP_{SMB}$	0.07	1.44	4.63					23.19
$FMP_{HML}$	0.13	2.06	6.39					

Table 3.2: Longevity risk and factors' descriptive statistics. Panel A contains the annual statistics concerning changes in longevity ( $dLE$ ), market news regarding cash-flow ( $CF$ ) and discount-rate changes ( $DR$ ), and the factor-mimicking portfolios ( $FMP$ ).  $AR(1)$  shows the first-order autocorrelation coefficient. Panel B summarize monthly means, standard deviations (Std. dev.), Sharpe ratios, and correlations between the factors. The factor-mimicking portfolios constitute returns extracted from: six portfolios sorted on market equity and book-to-market equity ( $SV$ ), three portfolios sorted on market equity ( $SMB$ ) and book-to-market equity ( $HML$ ), respectively. All return factors are constructed using prices up until portfolio formation dates in June; see Asness and Frazzini (2013). The sample period is for the full sample from 1952 to 2020.

Panel A in Table (3.2) shows that changes in life expectancy for an individual aged 65 have increased on average 0.11 years pro anno. Furthermore, the autoregressive coefficients point estimate is 0.08. Moreover, market cash-flow and discount-rate news are approximately zero as they are extracted from the errors of a vector autoregressive model; see Campbell and Vuolteenaho (2004). Panel B in Table (3.2) contains monthly descriptive statistics. The market portfolio is positively correlated with both Fama and French (1993)'s small-minus-big ( $SMB$ ) portfolio and the longevity mimicking portfolio equivalent. As the market portfolio beta is the sum of the  $\beta_{CF}$  and  $\beta_{DR}$  in Table (3.1) this reflect the higher market-betas towards the small market-equity portfolios. The correlation between the market portfolio and Fama and French (1993)'s high-minus-low ( $HML$ ) portfolio and the longevity mimicking portfolio equivalent follows a similar rationale. In addition, the longevity mimicking portfolios are highly correlated with Fama and French (1993)'s two factors as illustrated in Figure (3.1).

### 3.4 Empirical estimates of longevity risk premia

I examine unconditional cross-sectional regression models using Fama and MacBeth (1973)'s procedure

$$\mathbb{E} [R_{it} - R_f] = \alpha + \lambda \hat{\beta} + \varepsilon_{it}, \quad (3.10)$$

where  $\mathbb{E} [R_{it} - R_f]$  is the excess return of asset  $i$  in month  $t$ ,  $\alpha$  is the zero-beta rate in excess of the risk-free rate,  $\hat{\beta}$  represents the time-series regression of returns on factors,  $\lambda$  denotes cross-sectional prices of risks, and  $\varepsilon_{it}$  are residuals. I consider 25 size and book-to-market sorted portfolios in addition to 25 portfolios sorted by size and investment, 25 portfolios sorted on size and operating profitability, and five industry-sorted portfolios. The five industry portfolios are included to add to the robustness of the risk price estimates, following Kan et al. (2013).

#### 3.4.1 Theory motivated cross-sectional regressions

I test various versions of the model

$$R_{it} - R_f = a_i + \beta_{CFi} CF_t + \beta_{DRi} DR_t + \beta_{ij}^{FMP} FMP_{tj} + \varepsilon_{it}, \quad (3.11)$$

where  $R_{it} - R_f$  denotes excess returns, the variables,  $CF_t$  and  $DR_t$ , represent Campbell and Vuolteenaho (2004)'s model and  $FMP_{tj}$  is the return from one or a combination of the longevity mimicking portfolios presented in Section (3.3). In accordance with equation (3.8),  $\alpha$  should be zero. Furthermore, the ratio of risk prices from aggregate cash-flow and discount-rate news identifies the relative risk aversion parameter,  $\gamma$ ; see Campbell and Vuolteenaho (2004). Moreover, equation (3.8) states that the risk price from the longevity beta should be a factor of  $(1 - \theta)$  larger than the cash-flow beta after adjusting for their respective variances; see Table (3.2). Table (3.3) presents the results from the Fama-MacBeth regressions.

	25 portfolios			80 portfolios		
	(1.1)	(2.1)	(3.1)	(1.2)	(2.2)	(3.2)
$\alpha$	-0.18 (-0.15)	0.24 (0.75)	1.10 (3.28)	-0.04 (-0.04)	0.34 (1.60)	0.93 (5.19)
$\lambda_{CF}$	2.93 (2.34)	0.58 (0.45)	-0.88 (-0.75)	2.86 (2.29)	0.45 (0.34)	-0.16 (-0.15)
$\lambda_{DR}$	0.37 (0.90)	0.31 (0.90)	-0.59 (-1.84)	0.25 (0.60)	0.24 (0.69)	-0.48 (-1.58)
$\lambda_{SV}^{FMP}$		0.22 (1.49)			0.25 (1.66)	
$\lambda_{SMB}^{FMP}$			0.27 (2.68)			0.21 (2.36)
$\lambda_{HML}^{FMP}$			0.22 (1.31)			0.15 (0.92)
$OLS : R^2$	0.30	0.34	0.65	0.24	0.31	0.58
$GLS : R^2$	0.07	0.07	0.19	0.06	0.06	0.11

Table 3.3: Cross-sectional regression results. This table shows estimated risk premia for the longevity augmented asset pricing model presented in equation (3.8) using two different set of test assets and two specifications of the longevity factor. The first set of test assets constitute 25 portfolios sorted by size and book-to-market equity. The second set of test assets consist of the former and; in addition, 25 portfolios sorted by size and investment, 25 portfolios sorted on size and operating profitability, and five industry-sorted portfolios, following Kan et al. (2013). Campbell and Vuolteenaho (2004)'s factors are represented by news to aggregate cash-flows ( $CF$ ) and discount-rates ( $DR$ ). In addition, the mimicking portfolios of longevity across the size-value dimension, ( $FMP_{SV}$ ), and the portfolios extracted from size ( $FMP_{SMB}$ ) and value ( $FMP_{HML}$ ) portfolios, respectively, are explored. The mimicking portfolios are constructed using prices up until formation dates in June; see Asness and Frazzini (2013). All coefficients are multiplied by 100. In parenthesis are test statistics adjusted for errors-in-variables, following Shanken (1992). Moreover,  $OLS : R^2$  denotes standard r-squared statistics, whereas  $GLS : R^2$  represents the equivalent r-squared statistics using generalized least squares estimates; see Kandel and Stambaugh (1995).

The risk price from market cash-flow news is higher than the equivalent for market discount-rate news for the standard Campbell and Vuolteenaho (2004) model in column (1.1) and (1.2). Furthermore, the risk aversion parameter identified as  $\gamma \equiv \lambda_{CF}/\lambda_{DR}$  are between approximately 8 and 11.5 depending on the test assets. This is consistent with Campbell and Vuolteenaho (2004), whom find that the implied risk aversion parameter is between 8 and 11. Moreover, the risk aversion interval includes the estimate of 9.5 reported in Bansal and Yaron (2004).

Introducing the factor-mimicking portfolio of longevity extracted from size- and book-to-market sorted portfolios, ( $FMP_{SV}$ ), in columns (2.1) and (2.2) does not alter the risk price con-

cerning the discount-rate beta. In contrast, it transmit a large portion of the risk price associated with the cash-flow beta towards the longevity factor. The risk price for the cash-flow beta turns insignificant; however, still positive and higher than the risk price arising from news of market discount-rate. Nonetheless, using the point estimates of the risk prices for market cash-flow and discount-rate news leads to an implied risk aversion parameter of around 1.9 for the two test assets. The low risk aversion estimates are inconsistent with risk aversion estimates using aggregate consumption measures as in Bansal and Yaron (2004); Vissing-Jørgensen and Attanasio (2003). However, they are consistent with risk aversion estimates using individual-level consumption measures. For example, Brav et al. (2002) find support of risk aversion estimates between 2 and 4 using individual-level consumption data. This suggests that the risk premium regarding market risk relates to individual-level consumption risk, when controlling for longevity risk.

The longevity factor yields a monthly risk premium of 0.22 and 0.25 for the two test assets, which is significant at the 10% and 5% critical values. According to equation (3.8), the risk price for cash-flow risk ought to be  $\lambda_{CF} = \frac{\sigma_m^2}{\sigma_l^2} \lambda_{FMP}$ , when assuming  $\theta = 1$ , which amount to point estimates of 0.92 to 1.06, which is within the confidence intervals of  $\lambda_{CF}$  in Table (3.3), columns (2.1) and (2.2).<sup>7</sup> Thus,  $\theta$  cannot be rejected to be equal to one; however, a higher  $(1 - \theta)$  would decrease the implied risk premium for cash-flow risk lending support to the notion that the longevity factor obtains a larger risk premium due to preference for early resolution of future uncertainty, i.e.,  $\theta < 1$ .

The difference in the price of risk for market cash-flow and discount-rate news with and without the longevity mimicking portfolio clearly shows that the latter accounts for a large part of the priced risk in this sample. Thus, introducing the two mimicking portfolios used for exposition,  $FMP_{SMB}$  and  $FMP_{HML}$ , is expected to distort the model prediction discussed around equation (3.8). This conjecture is supported by column (3.1) and (3.2) in Table (3.3), where the point estimates of market cash-flow and discount-rate news decrease to  $-0.88$  ( $-0.16$ ) and  $-0.59$  or ( $-0.48$ ), respectively for the 25 (80) test assets. The significantly negative risk premium for market discount rate news indicate that this empirical model may be misspecified.

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<sup>7</sup>The relation between the risk price of market cash-flow news and longevity can be deduced similar to that between market news of cash-flow and discount rate risk premia, i.e.,  $\gamma \equiv \frac{\lambda_{CF}}{\lambda_{DR}} = \gamma \frac{\sigma_m^2}{\sigma_l^2}$  and  $\frac{\lambda_{CF}}{\lambda_{FMP}} = \frac{\gamma \sigma_m^2}{[\gamma + (1 - \theta)] \sigma_l^2}$ .

In addition, the zero-beta rate in excess of the risk-free rate,  $\alpha$ , is significantly different from zero amounting to a monthly estimate of 1.1 and 0.93 for the two different test assets. In contrast, the Campbell and Vuolteenaho (2004) model and the equivalent augmented with the longevity factor obtain an  $\alpha$  insignificantly different from zero measured to  $-0.18$  to  $-0.04$  and  $0.24$  to  $0.34$  for the two models, respectively. The reason is straightforward that  $\alpha$  is empirically denoted as the time-average return not explained by factor betas and risk premia, i.e.,  $\alpha = \bar{R}_T - \hat{\beta}\hat{\lambda}$ , cf. equation (3.10). Thus, including two factors, i.e.,  $FMP_{SMB}$  and  $FMP_{HML}$ , demanding an equivalent risk premium increases the zero-beta rate above zero.

The equivalence of the three different longevity factor mimicking portfolios is clear after dividing the risk premiums in Table (3.3) by their respective standard deviations from Table (3.2). Besides establishing a common scale, the risk premia can be interpreted as the Sharpe ratios on portfolios tracking the slope of the common longevity factor; see Balduzzi and Robotti (2010). The different longevity factors obtain monthly Sharpe ratios of 0.099 (0.114), 0.186 (0.148), and 0.106 (0.071) for  $FMP_{SV}$ ,  $FMP_{SMB}$ , and  $FMP_{HML}$  in the 25 (80) test assets, respectively. The null hypothesis of the difference between the most different longevity factors' Sharpe ratios are equal to zero is not rejected at the 10% critical value. Thus, if the longevity model in equation (3.8) is the true null model, one cannot reject the hypothesis that including both the  $FMP_{SMB}$  and the  $FMP_{HML}$  leads to overfitting risk premia, which results in rejection of the null hypothesis of the zero-beta rate in excess of the risk-free rate being equal to zero.

More formally, Kandel and Stambaugh (1995) consider the generalized least squares  $R^2$  as informative for evaluating a factor model's ex ante mean-variance efficiency. They show that the square root of the GLS  $R^2$  provides a numerical estimation of the relative distance between the proposed factors and the mean-variance boundary. Furthermore, Lewellen et al. (2010) shows that the GLS  $R^2$  is linked to other asset pricing tests such as Shanken (1985)'s test of linearity and the Hansen-Jagannathan distance (Kan and Zhou, 2004). Considering the GLS  $R^2$  above for the Campbell and Vuolteenaho (2004) model and the equivalent augmented with the longevity factor, reveals that these factors constitute about 25% of the mean-variance boundary. The model including the two longevity mimicking portfolios,  $FMP_{SMB}$ , and  $FMP_{HML}$ , could lead one to conclude



that this model explain more of the mean-variance boundary with the factors representing 44% and 33% of the mean-variance boundary. However, Kandel and Stambaugh (1995) emphasize that a bundle of proposed factors are inefficient if  $\alpha > 0$  or  $\lambda < \bar{R}_b - R^f$ , where  $\bar{R}_b$  represents the mean return of portfolio factor  $b$ . Comparing the risk premium estimates in Table (3.3) with the descriptive statistics in Table (3.2) reveal that all risk premiums are close to their respective monthly mean excess returns. In contrast, the  $\alpha > 0$  for the model including two longevity factors suggest that they are mean-variance inefficient. This implies that including separate risk factors for both size and value in asset-pricing tests may overfit the observed returns due to their common variation with longevity risk.

### 3.4.2 Empirical motivated cross-sectional regressions

Motivated by the results in the former section, I examine cross-sectional regressions using the capital asset-pricing model (CAPM) (Sharpe, 1964; Lintner, 1965b), Fama and French (1993)'s three-factor model, and a two-factor version of equation (3.6) omitting the consumption factor but including the market portfolio and the longevity factors. As discussed in Section (3.2), longevity changes is a priori believed to affect current consumption. Thus, including both longevity and consumption factors would result in multicollinear factors as shown in Chen and Yang (2019). Moreover, longevity risk is priced at least equivalent to market cash-flow risk, which Table (3.3) clearly shows. Hence, longevity may be an omitted factor that would add to the pricing ability of the CAPM, which then should equal  $\lambda_{CF} + \lambda_{DR}$ , if consumption is not an important omitted risk variable. Table (3.4) presents the results from the Fama-MacBeth regressions.

	25 portfolios				80 portfolios			
	(1.1)	(2.1)	(3.1)	(4.1)	(1.2)	(2.2)	(3.2)	(4.2)
$\alpha$	0.73 (2.41)	1.15 (4.37)	0.35 (0.97)	1.08 (2.70)	0.73 (2.41)	1.09 (5.39)	0.40 (1.20)	1.07 (3.30)
$\lambda_{RM}$	0.01 (0.04)	-0.55 (-2.24)	0.30 (1.19)	-0.54 (-2.22)	0.01 (0.04)	-0.47 (-1.93)	0.26 (1.04)	-0.48 (-1.98)
$\lambda_{SMB}$		0.16 (0.97)				0.15 (0.87)		
$\lambda_{HML}$		0.27 (1.74)				0.26 (1.64)		
$\lambda_{SV}^{FMP}$			0.23 (1.59)				0.27 (1.92)	
$\lambda_{SMB}^{FMP}$				0.27 (2.62)				0.22 (2.22)
$\lambda_{HML}^{FMP}$				0.25 (1.95)				0.22 (1.74)
$OLS : R^2$	0.00	0.61	0.24	0.65	0.00	0.56	0.24	0.58
$GLS : R^2$	0.01	0.20	0.05	0.23	0.01	0.05	0.01	0.05

Table 3.4: Cross-sectional regression results. This table shows estimated risk premia for the CAPM, Fama and French (1993)'s empirical model and the longevity augmented asset pricing model presented in equation (3.6), omitting the consumption factor, using two different set of test assets and two specifications of the longevity factor. The first set of test assets constitute 25 portfolios sorted by size and book-to-market equity. The second set of test assets consist of the former and; in addition, 25 portfolios sorted by size and investment, 25 portfolios sorted on size and operating profitability, and five industry-sorted portfolios, following Kan et al. (2013). Fama and French (1993)'s factors are represented by: excess market return ( $R_M$ ), small-minus-big ( $SMB$ ), and high-minus low ( $HML$ ). The mimicking portfolios of longevity are represented by the portfolio derived from the size-value dimension, ( $FMP_{SV}$ ), and the portfolios extracted from size ( $FMP_{SMB}$ ) and value ( $FMP_{HML}$ ) portfolios, respectively, using prices up until formation dates in June; see Asness and Frazzini (2013). All coefficients are multiplied by 100. In parenthesis are test statistics adjusted for errors-in-variables, following Shanken (1992). Moreover,  $OLS : R^2$  denotes standard r-squared statistics, whereas  $GLS : R^2$  represents the equivalent r-squared statistics using generalized least squares estimates; see Kandel and Stambaugh (1995).

As a basis for comparison, the unconditional risk prices for the CAPM model and Fama-French's three-factor model is represented in the two left columns for each set of test assets, i.e., columns (1.1),(1.2), (2.1), and (2.2). The risk prices for CAPM amounts to an insignificant monthly risk price of 0.01 for both test assets. Furthermore, the zero-beta rate in excess of the risk-free rate,  $\alpha$ , is significantly positive and both r-squared statistics are close to zero.

In contrast, the Fama-French three-factor model obtains relatively high r-squared statistics amounting to 0.61 (0.56) and 0.20 (0.05) for the 25 (80) test assets in the OLS and GLS r-

squared statistics, respectively. The  $\alpha$ 's for the three-factor model is significantly positive for both test assets. Furthermore, the market risk premium is significantly negative, whereas SMB is insignificantly positive and HML is significantly positive at the 5% and 10% critical values for the two test assets. The literature provides different inference on the risk prices for the market factor, whereas the risk price for SMB is usually insignificant or vaguely significant and the HML factor is significantly positive. For example, Lettau and Ludvigson (2001b) reports an insignificant positive risk price for the market factor, whereas Lewellen et al. (2010) arrive at an insignificant negative risk price, despite using the same data set and time frequency; however, for different periods.

In comparison, utilizing the two-factor model represented by the market portfolio and the longevity mimicking portfolio,  $FMP_{SV}$  leads to a positive, although insignificant, risk premium for the market risk premium. Consistent with the results in Table (3.3), the risk premium for the market portfolio departs slightly with the equivalent model including the Campbell and Vuolteenaho (2004) model in Table (3.3), columns (2.1) and (2.2). The sum of the risk premiums for the Campbell and Vuolteenaho (2004) factors are 0.89 (0.69) for the 25 (80) test assets, which is higher than the market risk premium of 0.30 (0.26) for the equivalent test assets in Table (3.4), columns (3.1) and (3.2). Thus, aggregating cash-flow and discount-rate beta to market returns appear to aggravate the pricing from the market portfolio slightly, which is compatible with an omitted variable from consumption risk. However, the difference between the two market premium methods are not statistically significant with p-values of 0.33 (0.38) for the 25 (80) test assets.<sup>8</sup> Thus, this suggests that including the longevity factor in conjunction with the market portfolio can, to a large extent, account for the pricing ability between cash-flow and discount-rate betas and that omitting the consumption factor is not critical for inference of equation (3.6).

In addition, comparing the longevity factor price of risk,  $\lambda_{SV}^{FMP}$ , between Table (3.3) and Table (3.4) shows that the risk premium is consistent between the different specifications. Comparably, including the two longevity factors,  $FMP_{SMB}$  and  $FMP_{HML}$ , corresponds well with the results in Table (3.3). Moreover, the model obtains an equivalent significant zero-beta rate in excess of the

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<sup>8</sup>The test statistics,  $Z$ , can be calculated using the point estimates and t-statistics in Table (3.3) and (3.4) as:  $Z = \frac{\lambda_M - (\lambda_{CF} + \lambda_{DR})}{\sqrt{\text{Var}(\lambda_M) + \text{Var}(\lambda_{CF}) + \text{Var}(\lambda_{DR})}}$ ; see Gelman and Stern (2006).

risk-free rate and r-squared statistics as the Fama-French three-factor model, cf. columns (2.1), (2.2) with (4.1) and (4.2) in Table (3.4). Thus, this indicate that the Fama-French model assumes longevity risk in both the size and value dimension, which lead to high r-squared statistics and significant  $\alpha$ 's, cf. the discussion in Section (3.4.1).

### 3.5 Conclusion

The value premium has been attributed to deviations in consumption-wealth ratios, differences in risk premiums for market cash-flow and discount rate news, and argued as capturing omitted risk premiums. I introduce stochastic survival rates for the representative agent's intertemporal budget constraint and present a model consistent with these previous findings.

Higher longevity (mortality) changes increase (decrease) the representative agent's intertemporal budget constraint; thus, longevity changes covary positively with consumption-wealth ratios comparable to aggregate wealth. Moreover, since longevity changes follow a random walk with drift, all information of future longevity changes is contained in current changes. Hence, I show that longevity risk convey a risk premium comparable to market cash-flow risk, when substituting out consumption. In addition, the underlying source of uncertainty, mortality, is published with a two-year lag lending support to the notion that longevity risk is a key factor in explaining the value premium.

Empirically, I find that a model including a longevity mimicking portfolio and the market portfolio or its two components, market cash-flow and discount rate news, fits well with the model's prediction. That is, the longevity risk risk premium is slightly higher than the cash-flow equivalent. Furthermore, the zero-beta rate in excess of the risk-free rate is economic small and insignificantly different from zero. In contrast, including the market portfolio along two exposition mimicking portfolios extracted from respectively market equity and book-to-market sorted portfolios, leads to higher r-squared statistics and significant positive pricing errors. This suggests that the size and value factors proxy for the same factor, namely longevity risk, and including both leads to double-counting the risk premium.

## 3.6 Appendix

### 3.6.1 Derivation of economy

**Derivation of the stochastic discount factor** The state-price deflator from the recursive utility function in equation (3.3) follows Epstein and Zin (1991) as

$$M_{t,t+1} = \beta \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{1}{\psi}} \left[ \frac{\mathcal{U}_{t+1}}{q_t(\mathcal{U}_{t+1})} \right]^{\frac{1}{\psi}-\gamma}, \quad (3.12)$$

where  $q_t(\mathcal{U}_{t+1}) \equiv \left[ \mathbb{E}_t \left( \mathcal{U}_{t+1}^{1-\gamma} \right) \right]^{\frac{1}{1-\gamma}}$  for brevity. To introduce return on aggregate wealth, I use the notion that

$$\mathbb{E}_t [M_{t,t+1} W_{t+1}] = \frac{\beta}{1-\beta} c_t^{\frac{1}{\psi}} q_t(\mathcal{U}_{t+1})^{1-\frac{1}{\psi}}.$$

This can be verified by using the standard intertemporal budget constraint,  $W_{t+1} = R_{w,t+1} [W_t - C_t]$ , in the following deductions. Introducing survival probability as in equation (3.1), the intertemporal budget constraint is

$$\begin{aligned} \frac{L_{t+1}}{L_t} [W_t - c_t] &= \mathbb{E}_t [M_{t,t+1} W_{t+1}] \\ &= \frac{\beta}{1-\beta} c_t^{\frac{1}{\psi}} q_t(\mathcal{U}_{t+1})^{1-\frac{1}{\psi}}. \end{aligned}$$

Thus, the return on wealth is

$$\begin{aligned} R_{w,t+1} &= \left( \frac{L_{t+1}}{L_t} \right)^{-1} \frac{W_{t+1}}{[W_t - c_t]} \\ &= \left( \frac{L_{t+1}}{L_t} \right)^{-1} \frac{\frac{1}{1-\beta} c_{t+1}^{\frac{1}{\psi}} \mathcal{U}_{t+1}^{1-\frac{1}{\psi}}}{\frac{\beta}{1-\beta} c_t^{\frac{1}{\psi}} q_t(\mathcal{U}_{t+1})^{1-\frac{1}{\psi}}} \\ &= \left( \frac{L_{t+1}}{L_t} \right)^{-1} \beta^{-1} \left( \frac{c_{t+1}}{c_t} \right)^{\frac{1}{\psi}} \left( \frac{\mathcal{U}_{t+1}}{q_t(\mathcal{U}_{t+1})} \right)^{1-\frac{1}{\psi}}. \end{aligned} \quad (3.13)$$

The above equation can be rewritten as

$$\left[ \frac{\mathcal{U}_{t+1}}{q_t(\mathcal{U}_{t+1})} \right]^{\frac{1}{\psi}-\gamma} = \beta^{\theta-1} \left( \frac{L_{t+1}}{L_t} \right)^{\theta-1} \left( \frac{c_{t+1}}{c_t} \right)^{\frac{1-\theta}{\psi}} (R_{w,t+1})^{\theta-1}, \quad (3.14)$$

when noting that  $\theta = \frac{1-\gamma}{1-\frac{1}{\psi}}$ . Substituting equation (3.14) into equation (3.12), I obtain

$$\begin{aligned} M_{t,t+1} &= \beta \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{1}{\psi}} \beta^{\theta-1} \left( \frac{L_{t+1}}{L_t} \right)^{\theta-1} \left( \frac{c_{t+1}}{c_t} \right)^{\frac{1-\theta}{\psi}} (R_{w,t+1})^{\theta-1} \\ &= \beta^{\theta} \left( \frac{L_{t+1}}{L_t} \right)^{\theta-1} \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{\theta}{\psi}} (R_{w,t+1})^{\theta-1}, \end{aligned}$$

which corresponds to equation (3.4).

**Derivation of factor model** Taking the logarithm of the Euler equation (3.5) leads to

$$\begin{aligned} 0 &= \theta \log \beta + (\theta - 1) \mathbb{E}_t \Delta l_{t+1} - \frac{\theta}{\psi} \mathbb{E}_t \Delta c_{t+1} + (\theta - 1) \mathbb{E}_t r_{m,t+1} + \mathbb{E}_t r_{i,t+1} \\ &+ \frac{1}{2} \left[ (\theta - 1)^2 V_{ll} + \left( \frac{\theta}{\psi} \right)^2 V_{cc} + (\theta - 1)^2 V_{mm} + V_{ii} \right] \\ &+ \frac{1}{2} \left[ -2 \frac{(\theta - 1) \theta}{\psi} V_{lc} + 2 (\theta - 1)^2 V_{lm} + (\theta - 1) V_{li} \right] \\ &+ \frac{1}{2} \left[ -2 \frac{\theta^2}{\psi} V_{cm} - 2 \frac{\theta}{\psi} V_{ci} + 2 (\theta - 1) V_{im} \right], \end{aligned} \tag{3.15}$$

where lower case letters denote the logarithm of the real variables and  $V_{yy}$  or  $V_{yx}$  corresponding to variances and covariances, respectively. The riskfree real return,  $r_{f,t+1}$ , is assumed to have zero variance and covariances with the state variables' changes in longevity, consumption growth, and the aggregate market portfolio. Hence, the expected real return of asset  $i$  in excess of the riskfree return simplifies to

$$\mathbb{E}_t r_{i,t+1} - r_{f,t+1} + \frac{1}{2} V_{ii} = \frac{\theta}{\psi} V_{ci} + (1 - \theta) [V_{im} + V_{li}], \tag{3.16}$$

which conform with equation (3.6).

**Substituting out consumption** The Euler equation (3.5) concerning the aggregate market portfolio itself can be rewritten as

$$1 = \mathbb{E}_t \left[ \beta^\theta \left( \frac{L_{t+1}}{L_t} \right)^{\theta-1} \left( \frac{c_{t+1}}{c_t} \right)^{-\frac{\theta}{\psi}} (R_{m,t+1})^\theta \right]. \quad (3.17)$$

Thus, taking the logarithm of equation (3.17) and assuming the variables are joint log-normal distributed and homoskedastic leads to

$$\begin{aligned} 0 = & \theta \log \beta + (\theta - 1) \mathbb{E}_t \Delta l_{t+1} - \frac{\theta}{\psi} \mathbb{E}_t \Delta c_{t+1} + \theta \mathbb{E}_t r_{m,t+1} \\ & + \frac{1}{2} \left[ (\theta - 1)^2 V_{ll} + \left( \frac{\theta}{\psi} \right)^2 V_{cc} + \theta^2 V_{mm} \right] \\ & + \frac{1}{2} \left[ -2 \frac{(\theta - 1) \theta}{\psi} V_{lc} + 2 (\theta - 1) \theta V_{lm} - 2 \frac{\theta^2}{\psi} V_{cm} \right]. \end{aligned} \quad (3.18)$$

From equation (3.18), I can isolate the expected value of consumption as

$$\mathbb{E}_t \Delta c_{t+1} = \mu + \frac{\psi}{\theta} (\theta - 1) \mathbb{E}_t \Delta l_{t+1} + \psi \mathbb{E}_t r_{m,t+1}. \quad (3.19)$$

Substituting equation (3.19) into equation (3.2) for consumption generates

$$\begin{aligned} c_{t+1} - \mathbb{E}_t [c_{t+1}] = & r_{m,t+1} - \mathbb{E}_t r_{m,t+1} \\ & + (1 - \psi) (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=1}^{\infty} \rho^j r_{m,t+1+j} \\ & + \left( 1 - \frac{\theta - 1}{\theta} \psi \right) [\Delta l_{t+1} - \mathbb{E}_t \Delta l_{t+1}]. \end{aligned} \quad (3.20)$$

Thus, any covariation of asset  $i$  with respect to consumption can be restated as

$$\begin{aligned} \text{Cov}_t (r_{i,t+1}, \Delta c_{t+1}) & \equiv V_{ic} \\ & = V_{im} + (1 - \psi) V_{ih} + \left( 1 - \frac{\theta - 1}{\theta} \psi \right) V_{il}. \end{aligned} \quad (3.21)$$



Inserting equation (3.21) into the expected return formula of equation (3.16) effectuates

$$\mathbb{E}_t r_{i,t+1} - r_{f,t+1} + \frac{1}{2} V_{ii} = \gamma V_{im} + (\gamma - 1) V_{ih} + (\gamma + 1 - \theta) V_{il}, \quad (3.22)$$

which corresponds to equation (3.7).

**Assuming consumer-investors react to news about future longevity expectations** Suppose that equation (3.2) is augmented to include all future longevity expectations, i.e.,

$$\begin{aligned} c_{t+1} - \mathbb{E}_t [c_{t+1}] &= (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=0}^{\infty} \rho^j r_{m,t+1+j} \\ &\quad - (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=1}^{\infty} \rho^j \Delta c_{t+1+j} \\ &\quad + (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=0}^{\infty} \rho^j \Delta l_{t+1}. \end{aligned}$$

Then the corresponding equation (3.20) emerges as

$$\begin{aligned} c_{t+1} - \mathbb{E}_t [c_{t+1}] &= r_{m,t+1} - \mathbb{E}_t r_{m,t+1} \\ &\quad + (1 - \psi) (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=1}^{\infty} \rho^j r_{m,t+1+j} \\ &\quad + \left( 1 - \frac{\theta - 1}{\theta} \psi \right) [\Delta l_{t+1} - \mathbb{E}_t \Delta l_{t+1}] \\ &\quad + \left( 1 - \frac{\theta - 1}{\theta} \psi \right) (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=1}^{\infty} \rho^j \Delta l_{t+1}, \end{aligned}$$

where the last line is the longevity addition to equation (3.20), which corresponds to news about future life expectancies. Now equation (3.21) contains four covariances, i.e.,

$$V_{ic} = V_{im} + (1 - \psi) V_{ih} + \left( 1 - \frac{\theta - 1}{\theta} \psi \right) (V_{il} + V_{il,h}). \quad (3.23)$$

where  $V_{il,h} \equiv \text{Cov}_t \left[ r_{i,t+1}, (\mathbb{E}_{t+1} - \mathbb{E}_t) \sum_{j=1}^{\infty} \rho^j \Delta l_{t+1} \right]$ . Inserting equation (3.23) into equation

(3.16) leads to

$$\mathbb{E}_t r_{i,t+1} - r_{f,t+1} + \frac{1}{2} V_{ii} = \gamma V_{im} + (\gamma - 1) V_{ih} + \gamma (V_{il} + V_{il,h}) + (1 - \theta) V_{il}. \quad (3.24)$$

Following Campbell (1993), assuming changes in longevity follow a univariate stochastic process  $V_{il,h} = [A(\rho) - 1] V_{il}$ , where the lag operator  $A(\rho) = 1 + \phi_1 L + \dots + \phi_n L^n$ , where  $\phi_i$  represents the  $i$ 'th autocorrelation coefficient with respect to the lagged variable,  $L$ . Since longevity changes follow a random walk,  $A(\rho) = 1$  and  $\phi_i = 0 \forall i = 1, \dots, n$ ,  $V_{il,h} = 0$ . Thus, equation (3.24) collapses to equation (3.7).



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|--|---|
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|---|---|

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*Access Decisions in a Partly-Digital World*  
*Comparing Digital Piracy and Legal Modes for Film and Music*
9. Marie Mathiesen  
*Making Strategy Work*  
*An Organizational Ethnography*
10. Arisa Shollo  
*The role of business intelligence in organizational decision-making*
11. Mia Kaspersen  
*The construction of social and environmental reporting*
12. Marcus Møller Larsen  
*The organizational design of offshoring*
13. Mette Ohm Rørdam  
*EU Law on Food Naming*  
*The prohibition against misleading names in an internal market context*
14. Hans Peter Rasmussen  
*GIV EN GED!*  
*Kan giver-idealstyper forklare støtte til velgørenhed og understøtte relationsopbygning?*
15. Ruben Schachtenhaufen  
*Fonetisk reduktion i dansk*
16. Peter Koerver Schmidt  
*Dansk CFC-beskatning*  
*I et internationalt og komparativt perspektiv*
17. Morten Froholdt  
*Strategi i den offentlige sektor*  
*En kortlægning af styringsmæssig kontekst, strategisk tilgang, samt anvendte redskaber og teknologier for udvalgte danske statslige styrelser*
18. Annette Camilla Sjørup  
*Cognitive effort in metaphor translation*  
*An eye-tracking and key-logging study*
19. Tamara Stucchi  
*The Internationalization of Emerging Market Firms: A Context-Specific Study*
20. Thomas Lopdrup-Hjorth  
*"Let's Go Outside": The Value of Co-Creation*
21. Ana Alačovska  
*Genre and Autonomy in Cultural Production*  
*The case of travel guidebook production*
22. Marius Gudmand-Høyer  
*Stemningssindssygdommenes historie i det 19. århundrede*  
*Omtydningen af melankolien og manien som bipolære stemningslidelser i dansk sammenhæng under hensyn til dannelsen af det moderne følelseslivs relative autonomi.*  
*En problematiserings- og erfarings-analytisk undersøgelse*
23. Lichen Alex Yu  
*Fabricating an S&OP Process*  
*Circulating References and Matters of Concern*
24. Esben Alfort  
*The Expression of a Need*  
*Understanding search*
25. Trine Pallesen  
*Assembling Markets for Wind Power*  
*An Inquiry into the Making of Market Devices*
26. Anders Koed Madsen  
*Web-Visions*  
*Repurposing digital traces to organize social attention*
27. Lærke Højgaard Christiansen  
*BREWING ORGANIZATIONAL RESPONSES TO INSTITUTIONAL LOGICS*
28. Tommy Kjær Lassen  
*EGENTLIG SELVLEDELSE*  
*En ledelsesfilosofisk afhandling om selvledelsens paradoksale dynamik og eksistentielle engagement*

- |  |   |
|--|---|
| <p>29. Morten Rossing<br/><i>Local Adaption and Meaning Creation in Performance Appraisal</i></p> <p>30. Søren Obed Madsen<br/><i>Lederen som oversætter<br/>Et oversættelsesteoretisk perspektiv på strategisk arbejde</i></p> <p>31. Thomas Høgenhaven<br/><i>Open Government Communities<br/>Does Design Affect Participation?</i></p> <p>32. Kirstine Zinck Pedersen<br/><i>Failsafe Organizing?<br/>A Pragmatic Stance on Patient Safety</i></p> <p>33. Anne Petersen<br/><i>Hverdagslogikker i psykiatrisk arbejde<br/>En institutionsetnografisk undersøgelse af hverdagen i psykiatriske organisationer</i></p> <p>34. Didde Maria Humle<br/><i>Fortællinger om arbejde</i></p> <p>35. Mark Holst-Mikkelsen<br/><i>Strategieksekvering i praksis – barrierer og muligheder!</i></p> <p>36. Malek Maalouf<br/><i>Sustaining lean<br/>Strategies for dealing with organizational paradoxes</i></p> <p>37. Nicolaj Tofte Brenneche<br/><i>Systemic Innovation In The Making<br/>The Social Productivity of Cartographic Crisis and Transitions in the Case of SEEIT</i></p> <p>38. Morten Gylling<br/><i>The Structure of Discourse<br/>A Corpus-Based Cross-Linguistic Study</i></p> <p>39. Binzhang YANG<br/><i>Urban Green Spaces for Quality Life - Case Study: the landscape architecture for people in Copenhagen</i></p> | <p>40. Michael Friis Pedersen<br/><i>Finance and Organization:<br/>The Implications for Whole Farm Risk Management</i></p> <p>41. Even Fallan<br/><i>Issues on supply and demand for environmental accounting information</i></p> <p>42. Ather Nawaz<br/><i>Website user experience<br/>A cross-cultural study of the relation between users' cognitive style, context of use, and information architecture of local websites</i></p> <p>43. Karin Beukel<br/><i>The Determinants for Creating Valuable Inventions</i></p> <p>44. Arjan Markus<br/><i>External Knowledge Sourcing and Firm Innovation<br/>Essays on the Micro-Foundations of Firms' Search for Innovation</i></p> <p><b>2014</b></p> <p>1. Solon Moreira<br/><i>Four Essays on Technology Licensing and Firm Innovation</i></p> <p>2. Karin Strzeletz Ivertsen<br/><i>Partnership Drift in Innovation Processes<br/>A study of the Think City electric car development</i></p> <p>3. Kathrine Hoffmann Pii<br/><i>Responsibility Flows in Patient-centred Prevention</i></p> <p>4. Jane Bjørn Vedel<br/><i>Managing Strategic Research<br/>An empirical analysis of science-industry collaboration in a pharmaceutical company</i></p> <p>5. Martin Gylling<br/><i>Processuel strategi i organisationer<br/>Monografi om dobbeltheden i tænkning af strategi, dels som vidensfelt i organisationsteori, dels som kunstnerisk tilgang til at skabe i erhvervsmæssig innovation</i></p> |
|--|---|

6. Linne Marie Lauesen  
*Corporate Social Responsibility in the Water Sector: How Material Practices and their Symbolic and Physical Meanings Form a Colonising Logic*
7. Maggie Qiuzhu Mei  
*LEARNING TO INNOVATE: The role of ambidexterity, standard, and decision process*
8. Inger Høedt-Rasmussen  
*Developing Identity for Lawyers Towards Sustainable Lawyering*
9. Sebastian Fux  
*Essays on Return Predictability and Term Structure Modelling*
10. Thorbjørn N. M. Lund-Poulsen  
*Essays on Value Based Management*
11. Oana Brindusa Albu  
*Transparency in Organizing: A Performative Approach*
12. Lena Olaison  
*Entrepreneurship at the limits*
13. Hanne Sørum  
*DRESSED FOR WEB SUCCESS? An Empirical Study of Website Quality in the Public Sector*
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*Knowing networks How experts shape transnational governance*
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*Entrepreneurial Individuals Empirical Investigations into Entrepreneurial Activities of Hackers and Makers*
16. Robert Spliid  
*Kapitalfondenes metoder og kompetencer*
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*Public-private partnerships & the need, development and management of trusting A processual and embedded exploration*
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*Management of design as a translation process*
19. Kåre Moberg  
*Assessing the Impact of Entrepreneurship Education From ABC to PhD*
20. Alexander Cole  
*Distant neighbors Collective learning beyond the cluster*
21. Martin Møller Boje Rasmussen  
*Is Competitiveness a Question of Being Alike? How the United Kingdom, Germany and Denmark Came to Compete through their Knowledge Regimes from 1993 to 2007*
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*Can Language be Managed in International Business? Insights into Language Choice from a Case Study of Danish and Austrian Multinational Corporations (MNCs)*
24. Rikke Kristine Nielsen  
*Global Mindset as Managerial Meta-competence and Organizational Capability: Boundary-crossing Leadership Cooperation in the MNC The Case of 'Group Mindset' in Solar A/S.*
25. Rasmus Koss Hartmann  
*User Innovation inside government Towards a critically performative foundation for inquiry*

- |   |   |
|---|---|
| <p>26. Kristian Gylling Olesen<br/><i>Flertydig og emergerende ledelse i folkeskolen</i><br/><i>Et aktør-netværksteoretisk ledelsesstudie af politiske evalueringsreformers betydning for ledelse i den danske folkeskole</i></p> <p>27. Troels Riis Larsen<br/><i>Kampen om Danmarks omdømme 1945-2010</i><br/><i>Omdømmearbejde og omdømmepolitik</i></p> <p>28. Klaus Majgaard<br/><i>Jagten på autenticitet i offentlig styring</i></p> <p>29. Ming Hua Li<br/><i>Institutional Transition and Organizational Diversity: Differentiated internationalization strategies of emerging market state-owned enterprises</i></p> <p>30. Sofie Blinkenberg Federspiel<br/><i>IT, organisation og digitalisering: Institutionelt arbejde i den kommunale digitaliseringsproces</i></p> <p>31. Elvi Weinreich<br/><i>Hvilke offentlige ledere er der brug for når velfærdstænkningen flytter sig – er Diplomuddannelsens lederprofil svaret?</i></p> <p>32. Ellen Mølgaard Korsager<br/><i>Self-conception and image of context in the growth of the firm</i><br/><i>– A Penrosian History of Fiberline Composites</i></p> <p>33. Else Skjold<br/><i>The Daily Selection</i></p> <p>34. Marie Louise Conradsen<br/><i>The Cancer Centre That Never Was</i><br/><i>The Organisation of Danish Cancer Research 1949-1992</i></p> <p>35. Virgilio Failla<br/><i>Three Essays on the Dynamics of Entrepreneurs in the Labor Market</i></p> | <p>36. Nicky Nedergaard<br/><i>Brand-Based Innovation</i><br/><i>Relational Perspectives on Brand Logics and Design Innovation Strategies and Implementation</i></p> <p>37. Mads Gjedsted Nielsen<br/><i>Essays in Real Estate Finance</i></p> <p>38. Kristin Martina Brandl<br/><i>Process Perspectives on Service Offshoring</i></p> <p>39. Mia Rosa Koss Hartmann<br/><i>In the gray zone</i><br/><i>With police in making space for creativity</i></p> <p>40. Karen Ingerslev<br/><i>Healthcare Innovation under The Microscope</i><br/><i>Framing Boundaries of Wicked Problems</i></p> <p>41. Tim Neerup Thomsen<br/><i>Risk Management in large Danish public capital investment programmes</i></p> <p><b>2015</b></p> <p>1. Jakob Ion Wille<br/><i>Film som design</i><br/><i>Design af levende billeder i film og tv-serier</i></p> <p>2. Christiane Mossin<br/><i>Interzones of Law and Metaphysics</i><br/><i>Hierarchies, Logics and Foundations of Social Order seen through the Prism of EU Social Rights</i></p> <p>3. Thomas Tøth<br/><i>TRUSTWORTHINESS: ENABLING GLOBAL COLLABORATION</i><br/><i>An Ethnographic Study of Trust, Distance, Control, Culture and Boundary Spanning within Offshore Outsourcing of IT Services</i></p> <p>4. Steven Højlund<br/><i>Evaluation Use in Evaluation Systems – The Case of the European Commission</i></p> |
|---|---|

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*AMBIGUOUS WINDS OF CHANGE – OR FIGHTING AGAINST WINDMILLS IN CHINESE WIND POWER*  
*A CONSTRUCTIVIST INQUIRY INTO CHINA'S PRAGMATICS OF GREEN MARKETISATION MAPPING*  
*CONTROVERSIES OVER A POTENTIAL TURN TO QUALITY IN CHINESE WIND POWER*
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*A Multi-case Analysis of the Development of Enterprise Resource Planning Systems (ERP) Business Practices*  
  
Morten Friis-Olivarius  
*The Associative Nature of Creativity*
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*New Cooperativism: A study of emerging producer organisations in India*
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*Implementering af ITIL® IT-governance - når best practice konflikt med kulturen Løsning af implementerings-problemer gennem anvendelse af kendte CSF i et aktionsforskningsforløb.*
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*A Real Options Approach to Determining Power Prices*
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*Toward a Digital Strategy for Omnichannel Retailing*
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*A relational processual investigation moving beyond the conceptual horizon of welfare management*
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*In Search of Opportunities: Three Essays on Global Linkages for Innovation*
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*Embedded Diversity: A critical ethnographic study of the structural tensions of organizing diversity*
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*Learning through Scenario Planning*
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*Knowledge dissemination based on terminological ontologies. Using eye tracking to further user interface design.*
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*PUBLIC-PRIVATE PARTNERSHIPS FOR INNOVATION AND SUSTAINABILITY TRANSFORMATION*  
*An embedded, comparative case study of municipal waste management in England and Denmark*
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*Follwoing the Content of Reported Risk Across the Organization*
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*Language strategies in multinational corporations. A cross-sector study of financial service companies and manufacturing companies.*
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*Designing performance management for operational level*  
*- A closer look on the role of design choices in framing coordination and motivation*



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*Objects and Social Actions  
– on Second-hand Valuation Practices*
24. Thorhildur Hansdottir Jetzek  
*The Sustainable Value of Open  
Government Data  
Uncovering the Generative Mechanisms  
of Open Data through a Mixed  
Methods Approach*
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*Innovation-based M&A  
– Technological-Integration  
Challenges – The Case of  
Digital-Technology Companies*
26. Mie Plotnikof  
*Challenges of Collaborative  
Governance  
An Organizational Discourse Study  
of Public Managers' Struggles  
with Collaboration across the  
Daycare Area*
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*Who Are the Post-Bureaucrats?  
A Philosophical Examination of the  
Creative Manager, the Authentic Leader  
and the Entrepreneur*
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*Constituting Performance Management  
A field study of a pharmaceutical  
company*
29. Rasmus Ploug Jenle  
*Engineering Markets for Control:  
Integrating Wind Power into the Danish  
Electricity System*
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*Complex Business Negotiation:  
Understanding Preparation and  
Planning*
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*TRUST AND TRANSPARENCY FROM AN  
ALIGNMENT PERSPECTIVE*
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*Byregimer og styringsevne: Politisk  
lederskab af store byudviklingsprojekter*
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*Essays on Competition, Innovation and  
Firm Strategy in Digital Markets*
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*The Value of Talent Management  
Rethinking practice, problems and  
possibilities*
35. Johanna Sax  
*Strategic Risk Management  
– Analyzing Antecedents and  
Contingencies for Value Creation*
36. Pernille Rydén  
*Strategic Cognition of Social Media*
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*The Measurable Me  
- The Influence of Self-tracking on the  
User Experience*
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*Towards a Fair Global Economic  
Regime? A critical assessment of Fair  
Trade through the examination of the  
Argentinean wine industry*
39. Marie Henriette Madsen  
*Emerging and temporary connections  
in Quality work*
40. Yangfeng CAO  
*Toward a Process Framework of  
Business Model Innovation in the  
Global Context  
Entrepreneurship-Enabled Dynamic  
Capability of Medium-Sized  
Multinational Enterprises*
41. Carsten Scheibye  
*Enactment of the Organizational Cost  
Structure in Value Chain Configuration  
A Contribution to Strategic Cost  
Management*

## 2016

1. Signe Sofie Dyrby  
*Enterprise Social Media at Work*
2. Dorte Boesby Dahl  
*The making of the public parking attendant  
Dirt, aesthetics and inclusion in public service work*
3. Verena Girschik  
*Realizing Corporate Responsibility  
Positioning and Framing in Nascent Institutional Change*
4. Anders Ørding Olsen  
*IN SEARCH OF SOLUTIONS  
Inertia, Knowledge Sources and Diversity in Collaborative Problem-solving*
5. Pernille Steen Pedersen  
*Udkast til et nyt copingbegreb  
En kvalifikation af ledelsesmuligheder for at forebygge sygefravær ved psykiske problemer.*
6. Kerli Kant Hvass  
*Weaving a Path from Waste to Value:  
Exploring fashion industry business models and the circular economy*
7. Kasper Lindskow  
*Exploring Digital News Publishing  
Business Models – a production network approach*
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*The chameleon workforce:  
Assembling and negotiating the content of a workforce*
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*Aesthetic encounters  
Rethinking autonomy, space & time in today's world of art*
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*EU PERSPECTIVES ON INTERNATIONAL COMMERCIAL ARBITRATION*
11. Abid Hussain  
*On the Design, Development and Use of the Social Data Analytics Tool (SODATO): Design Propositions, Patterns, and Principles for Big Social Data Analytics*
12. Mark Bruun  
*Essays on Earnings Predictability*
13. Tor Bøe-Lillegraven  
*BUSINESS PARADOXES, BLACK BOXES, AND BIG DATA: BEYOND ORGANIZATIONAL AMBIDEXTERITY*
14. Hadis Khonsary-Atighi  
*ECONOMIC DETERMINANTS OF DOMESTIC INVESTMENT IN AN OIL-BASED ECONOMY: THE CASE OF IRAN (1965-2010)*
15. Maj Lervad Grasten  
*Rule of Law or Rule by Lawyers?  
On the Politics of Translation in Global Governance*
16. Lene Granzau Juel-Jacobsen  
*SUPERMARKEDETS MODUS OPERANDI – en hverdagssociologisk undersøgelse af forholdet mellem rum og handlen og understøtte relationsopbygning?*
17. Christine Thalsgård Henriques  
*In search of entrepreneurial learning – Towards a relational perspective on incubating practices?*
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*Essays in Education, Crime, and Job Displacement*
19. Søren Korsgaard  
*Payments and Central Bank Policy*
20. Marie Kruse Skibsted  
*Empirical Essays in Economics of Education and Labor*
21. Elizabeth Benedict Christensen  
*The Constantly Contingent Sense of Belonging of the 1.5 Generation  
Undocumented Youth  
An Everyday Perspective*



22. Lasse J. Jessen  
*Essays on Discounting Behavior and Gambling Behavior*
23. Kalle Johannes Rose  
*Når stifterviljen dør...  
Et retsøkonomisk bidrag til 200 års  
juridisk konflikt om ejendomsretten*
24. Andreas Søeborg Kirkedal  
*Danish Stød and Automatic Speech  
Recognition*
25. Ida Lunde Jørgensen  
*Institutions and Legitimations in  
Finance for the Arts*
26. Olga Rykov Ibsen  
*An empirical cross-linguistic study of  
directives: A semiotic approach to the  
sentence forms chosen by British,  
Danish and Russian speakers in native  
and ELF contexts*
27. Desi Volker  
*Understanding Interest Rate Volatility*
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*Practice at the Boundaries of Business  
Ethics & Corporate Social Responsibility*
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*Levende læring i kunstneriske  
organisationer  
En undersøgelse af læringsprocesser  
mellem projekt og organisation på  
Aarhus Teater*
30. Leif Christensen  
*Quality of information – The role of  
internal controls and materiality*
31. Olga Zarzecka  
*Tie Content in Professional Networks*
32. Henrik Mahncke  
*De store gaver  
- Filantropiens gensidighedsrelationer i  
teori og praksis*
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*Using the Collective Wisdom of  
Frontline Employees in Strategic Issue  
Management*
34. Yun Liu  
*Essays on Market Design*
35. Denitsa Hazarbassanova Blagoeva  
*The Internationalisation of Service Firms*
36. Manya Jaura Lind  
*Capability development in an off-  
shoring context: How, why and by  
whom*
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*Essays on the Design of Contracts and  
Markets for Power System Flexibility*
38. Andreas Philipp Distel  
*Capabilities for Strategic Adaptation:  
Micro-Foundations, Organizational  
Conditions, and Performance  
Implications*
39. Lavinia Bleoca  
*The Usefulness of Innovation and  
Intellectual Capital in Business  
Performance: The Financial Effects of  
Knowledge Management vs. Disclosure*
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*Economic Organization and Imperfect  
Managerial Knowledge: A Study of the  
Role of Managerial Meta-Knowledge  
in the Management of Distributed  
Knowledge*
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*The Understanding of English Emotion  
Words by Chinese and Japanese  
Speakers of English as a Lingua Franca  
An Empirical Study*
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*The Ministry of Desire - Anxiety and  
entrepreneurship in a bureaucracy*
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*Making Time While Being in Time  
A study of the temporality of  
organizational processes*
44. Valeria Giacomini  
*Contextualizing the cluster Palm oil in  
Southeast Asia in global perspective  
(1880s–1970s)*

- |  |                    |  |
|--|--------------------|--|
| <p>45. Jeanette Willert<br/><i>Managers' use of multiple Management Control Systems: The role and interplay of management control systems and company performance</i></p> <p>46. Mads Vestergaard Jensen<br/><i>Financial Frictions: Implications for Early Option Exercise and Realized Volatility</i></p> <p>47. Mikael Reimer Jensen<br/><i>Interbank Markets and Frictions</i></p> <p>48. Benjamin Faigen<br/><i>Essays on Employee Ownership</i></p> <p>49. Adela Michea<br/><i>Enacting Business Models An Ethnographic Study of an Emerging Business Model Innovation within the Frame of a Manufacturing Company.</i></p> <p>50. Iben Sandal Stjerne<br/><i>Transcending organization in temporary systems Aesthetics' organizing work and employment in Creative Industries</i></p> <p>51. Simon Krogh<br/><i>Anticipating Organizational Change</i></p> <p>52. Sarah Netter<br/><i>Exploring the Sharing Economy</i></p> <p>53. Lene Tolstrup Christensen<br/><i>State-owned enterprises as institutional market actors in the marketization of public service provision: A comparative case study of Danish and Swedish passenger rail 1990–2015</i></p> <p>54. Kyoung(Kay) Sun Park<br/><i>Three Essays on Financial Economics</i></p> | <p><b>2017</b></p> | <p>1. Mari Bjerck<br/><i>Apparel at work. Work uniforms and women in male-dominated manual occupations.</i></p> <p>2. Christoph H. Flöthmann<br/><i>Who Manages Our Supply Chains? Backgrounds, Competencies and Contributions of Human Resources in Supply Chain Management</i></p> <p>3. Aleksandra Anna Rzeźnik<br/><i>Essays in Empirical Asset Pricing</i></p> <p>4. Claes Bäckman<br/><i>Essays on Housing Markets</i></p> <p>5. Kirsti Reitan Andersen<br/><i>Stabilizing Sustainability in the Textile and Fashion Industry</i></p> <p>6. Kira Hoffmann<br/><i>Cost Behavior: An Empirical Analysis of Determinants and Consequences of Asymmetries</i></p> <p>7. Tobin Hanspal<br/><i>Essays in Household Finance</i></p> <p>8. Nina Lange<br/><i>Correlation in Energy Markets</i></p> <p>9. Anjum Fayyaz<br/><i>Donor Interventions and SME Networking in Industrial Clusters in Punjab Province, Pakistan</i></p> <p>10. Magnus Paulsen Hansen<br/><i>Trying the unemployed. Justification and critique, emancipation and coercion towards the 'active society'. A study of contemporary reforms in France and Denmark</i></p> <p>11. Sameer Azizi<br/><i>Corporate Social Responsibility in Afghanistan – a critical case study of the mobile telecommunications industry</i></p> |
|--|--------------------|--|

12. Malene Myhre  
*The internationalization of small and medium-sized enterprises: A qualitative study*
13. Thomas Presskorn-Thygesen  
*The Significance of Normativity – Studies in Post-Kantian Philosophy and Social Theory*
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*Essays on multinational production and international trade*
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Controversies in IFRS in the Aftermath of the 2008 Financial Crisis*
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Regulering i skyggen af Farumsagen*
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*Business Model Innovation  
The Role of Organizational Design*
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*Entrepreneurial Cognition.  
Three essays on entrepreneurial behavior and cognition under risk and uncertainty*
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*Militært chefvirke: Kritik og retfærdiggørelse mellem politik og profession*

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