Essays on Expectations, Power and Social Security

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Proefschrift

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To my parents Anja and Ton, who both showed me that one is never too old to learn

Ἓν οἶδα, ὅτι οὐδὲν οἶδα

I know that I know nothing
Socrates, Athens, 470-399 BC

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[1] In fact, this thesis is some indication of his teaching skills as I simply could not read $\prod_{i=1}^{N} L_i$ when I first arrived in Tilburg.

teresting (yet dark...) world of structural modeling, and I wouldn't have understood any of this without his guidance. Together with Jan Boone, we closely collaborated on Chapter 4 of this thesis, which any reader can tell clearly bears their mark. Frederic also introduced me to the ins and outs of the academic world which will hopefully prove useful in the future. I want to thank Arthur and Frederic for all the knowledge that they shared with me and the time they were willing to bestow on me.

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Last but not least, I would like to thank my boyfriend of ten years, Wouter Roduner, for living with me while I was writing my PhD-thesis. During this time, our home has felt like a place to relax and have fun in. When I was troubled with PhD-affairs Wouter always found a way to make me smile about something else. He also proved an expert in making me forget about PhD-related things by setting up different social activities every weekend. I am convinced that our active private life has kept me fresh and motivated and I think it has been the key to finishing this thesis well in time. Thanks Wout, for everything ;-)!

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Chapter 1

Introduction

Although the introduction is the first - and often the only - section of a thesis that is being read, it is usually the last part written. This is at least the case for my thesis, that apart from this section consists of four separately readable chapters that were independently written[1]. As the research in the chapters developed in different directions, it has become difficult for the reader to discover a common theme. This introduction is meant to point out the similarities between the seemingly unrelated chapters that do exist. The applied research methodology and analyzed agents are for example very comparable. Moreover, apart from links to pensions, ageing and retirement research, it is possible to distinguish three broad research themes in the thesis: expectations, power and social security.

1.1 On the methodology and units of analysis

All four chapters are empirical, meaning that observations from reality are used to derive the main results. I have used several econometric techniques to analyze the available information. Econometrics is defined in this context as using statistical methods in order to characterize relationships between economic variables in reality. In Chapters 2 and 4 novel techniques, or new variations of existing techniques, are developed to answer the relevant research questions, whereas in Chapters 3 and 5

[1]Chapter 4 was written jointly with Jan Boone and Frederic Vermeulen, both at Tilburg University. The other chapters are written without co-authors. I did receive great help from my promotor, Arthur van Soest, my co-promotor, Frederic Vermeulen and many others that will be named in each chapter separately.

well-known methods are applied.

All four papers, which the chapters are based on, analyze information of individuals or households. This implies that the research in this thesis has a micro-economic emphasis. The focus is on individuals in all chapters but Chapter 4, in which we focus on households to investigate differences in preferences between husbands and wives. Micro-data has the advantage that it allows the researcher to analyze fundamental economic decision-making and also, that it enables conclusions on causal relationships. It must be said that only Chapter 5 draws explicit conclusions on causality in this thesis. In Chapters 2 and 4 structural models are estimated. This means that the econometrician has been explicit about which model is assumed for reality and that the structural parameters within this model (such as preference parameters) are estimated using the data at hand. The results in Chapter 3 are based on a reduced-form analysis and should be interpreted as purely associative. Reduced-form regressions require less a-priori assumptions about the data-generating process, but at the same time allow for less bold conclusions.

The fourth chapter is the only using data on individuals from the United States, i.e. individuals participating in the Consumer Expenditures Survey (CEX) conducted by the U.S. Bureau of Labor Statistics. The three other chapters employ data about individuals from The Netherlands, compiled by either Statistics Netherlands (the Socio Economic Panel, SEP, in Chapter 5) or by CentERdata which is affiliated to Tilburg University (the Pensionbarometer in Chapter 2 and 3 and the Dutch Central Bank Household Savings Survey, DHS, in Chapter 3). To collect information CentERdata uses an Internetpanel, which is a group of people that after being selected to participate in the panel regularly answers questions through the Internet. Those who initially do not have access to the Internet are provided with a connection and/or a simple computer in order to keep selection effects as small as possible.

All datasets that I use are survey panel datasets. Surveys have the advantage that although typically smaller than administrative datasets, they contain more qualitative and more subjective information. Specifically in the first part of the thesis on expectations, subjective information is crucial for the analysis. Panel datasets, which are defined as datasets in which the same individuals are interviewed several times in a row, are practical as these provide better opportunities to analyze individual changes over time. This feature, important in Chapter 5, enables the researcher to distinguish between the true influence of a certain characteristic and the effect of being an individual that is likely to have that characteristic. Also, several econometric techniques have been developed specifically for panel datasets, that ensure higher efficiency (i.e. better precision of the estimates). I exploit the panel dimension of the data specifically in Chapters 2 and 5.

1.2 On a common research theme

My PhD research was financed by the Dutch institution that pays out, among other things, old age social security benefits (the *Sociale Verzekeringsbank*). This was arranged by Netspar, the Network for Studies on Pensions, Ageing and Retirement, that has its administrative headquarters at Tilburg University. The research themes of Netspar are reflected in the topics of the chapters in this thesis.

Chapter 2 deals with individual expectations concerning the future of the Dutch old age social security system. In Chapter 3 I investigate whether these subjective expectations help to explain voluntary savings behavior for retirement. The last chapter in this thesis exploits a policy change in employment protection that rendered it less costly to fire workers older than 44 that were in the same job for more than three years. This policy change had an impact on the labor market position of these workers. As an ageing society requires more participation from older individuals in order to keep the social security and health care system financially sustainable, there is a link between Netspar's research focus and this chapter. It is hard to recognize anything pension-related in Chapter 4 as this chapter investigates public good consumption behavior in a non-cooperative household. However, it can be argued that saving for retirement is a public good within the household as longevity risks generate uncertainties about who will consume the savings in old age. The general results discussed in this chapter could thus also apply to a couple that jointly decides on how much to save for old age. Besides pensions, ageing and retirement, three themes are identified that play a role in (a subset of) the chapters in this thesis. These themes, that feature in the thesis' title - Essays on Expectations, Power and Social Security -, are discussed below.

1.2.1 Social security

Social security is defined as general as possible here so that it encompasses concepts used in all four chapters. Social security is the provision of goods by a superior entity that enables all individuals that meet certain criteria to benefit from this provision, and that arranges for the supply of this provision. Typically, we think of social security as being provided by the state, and the goods to be some sort of insurance against undesirable life events. Social security is often installed to mitigate the negative consequences of these events, such as poverty. Think of unemployment insurance, disability benefits and old age social security. Often, the provision of social security generates some undesirable externalities as individuals adapt their behavior to the eligibility requirements or to the mandatory contribution schemes. One example of such behavior is that unemployed individuals who are still eligible for several months of unemployment benefits search less intensively for a job than those who are on the brink of exhausting their benefit (e.g. Lalive, Van Ours and Zweimuller, 2006).

In The Netherlands, social security schemes are divided into employee insurance schemes, that are compulsory for all employees, and national insurance schemes, that everybody living in the country could be eligible for when meeting certain criteria. Chapters 2, 3 and 5 deal with traditional social security. Two broad schemes are dealt with, namely old age social security (good provided: income in old age, provider: Dutch government, beneficiaries: all inhabitants of The Netherlands) and employment protection (good provided: protection against prompt lay-off, provider: Dutch government, beneficiaries: employees). In Chapters 3 and 5 I explicitly address some of the externalities in economic behavior that these schemes could generate. In Chapter 3 I investigate an alternative type of externality; whether uncertainty about the provision of social security changes behavior. I find some evidence that uncertainty about the future of the old age social security system increases savings. In Chapter 5 I find strong evidence that employment protection affects the bargaining position of employers and employees and as such - unintendedly - influences wages.

In Chapter 4, social security in its traditional format does not play a role. However, this chapter discusses voluntary contributions to public goods that are consumed within the household. In this case the good provided is public consumption and the superior agent providing the good is the collective of all members of the household. It is obvious that externalities play a role in any public goods game, as there is an incentive not to contribute when you are expecting your partner to do so. As preferences on which public goods and how much of those goods to purchase differ within the family, the opposite externality is plaguing this game as well.

In order to obtain the allocation of public goods that separate spouses desire both have an incentive to contribute to the goods they prefer, and this could potentially lead to higher total spending on public goods than when the spouses would coordinate their spending.

1.2.2 Expectations

Subjective expectations are dealt with in Chapters 2 and 3. In this thesis an expectation refers to the image an individual has formed over the future realization of a random variable. Economists like to believe that expectations play a role in almost every decision that has an intertemporal dimension. Individual expectations seem crucial when thinking about insurance decisions, where one assesses the trade-off between the cost of insurance, the probability of an adverse event and the cost associated to that event, and investment decisions, where one has to decide on large and often irreversible investments before knowing their return exactly. In real estate markets for example, individuals' perception of the trend in housing prices seems important in determining how much they are willing to pay. Other micro-economic areas in which expectations are assumed to be a factor are buying health insurance (how large is the probability to get cancer?), human capital accumulation (does it pay to go to college?) and long-term decision-making in households (shall we move abroad for his career?).

However, there is still a lot unknown about how individuals come to have certain expectations and also on whether and how they use their expectations in everyday life. It seems likely that not all individuals are always capable of perfectly incorporating all available information in their expectations (if high-educated high-paid bankers in the City couldn't, why would a regular carpenter?). This is probably particularly true when the expectations concern events exogenous to the individual that rarely occur (see for example the studies on natural disasters by Kunreuther and Pauly, 2004, and Viscusi and Zeckhauser, 2008). Chapter 2 in my thesis tries to find out more about how real-life expectations on a rare event, in this case a policy change, are formed. I specifically examine the role of publicity, defined as media attention on the relevant topic, in the expectations formation process. Chapter 3 shows how the generated expectations influence actual savings behavior to see whether expectations are indeed as important as presumed. In both chapters, differences among individuals and demographic groups are explicitly addressed.

A unique longitudinal dataset on social security policy expectations, the Pensionbarometer which is collected by CentERdata in Tilburg and initiated and financed by Netspar, is used in both Chapters 2 and 3. Every three months, a group of respondents are asked for the probabilities they assign to certain policy changes in the Dutch pension system within the next twenty years. Although people do seem to understand the concept of probabilities, there is a remarkable tendency to provide different answers every time the same question is being asked. Filtering out some of the noise, the remaining expectation changes on the aggregate level due to publicity in the media on the relevant topic are minimal. In Chapter 2 I find that higher educated and high income groups hardly adapt their expectations to relevant publicity. Other subgroups do show a publicity reaction. The middle-aged and those who do not often read a newspaper have a relatively high publicity reaction. A potential explanation for this latter finding is that these groups have low quality initial expectations.

Chapter 3 investigates whether and how these individual expectations are related to individual voluntary pension scheme participation. Because the expectations are found to be so noisy in the earlier chapter, this chapter uses average expectations over time to represent people's perception of the future. I find that participation in private pension schemes is higher for those who assign high probabilities to the dismantlement of old age social security. The effects are however not equal for all age and income categories. The young and those with a high income do not seem to prepare much for a policy change, regardless of how likely they believe such a change will be. Although no causal conclusions can be drawn here, these results suggest that subjective expectations are correlated with economic behavior.

1.2.3 Power

Power, or bargaining power, is a crucial element of both Chapters 4 and 5. In organizational science the power of an individual is sometimes defined as the extent to which that individual has control over resources within an organization. Those who can make financial decisions are typically powerful, but also the secretary that knows everyone and everything about the place has considerable power. In economic theory someone's (bargaining) power often depends on the outside options that are available to the individual. A beautiful woman with many suitors can negotiate a better marriage contract than her mediocre fiancé.

Control over resources is important in the non-cooperative model of household consumption and labor supply that is developed in Chapter

4. In this chapter two spouses play a Nash bargaining game, in which they decide how much to contribute to different public goods that are consumed by the household, given the other spouse's contribution. Power matters in this game, as it turns out that the spouse who (potentially) earns much more than the other gets to dictate his/her preferences over the consumption of the public goods. In our model power is a function of relative earning capacity, but also of the estimated preference parameters of the two spouses. If reality was best represented by the so-called unitary model, i.e. if the family's decisions could best be characterized by a single utility function, the relative power of the two spouses should not make a difference. Our model is estimated by focusing on one particular public good - expenses on children's goods. We find that for couples with two or three children we reject the standard unitary model in favor of our non-cooperative approach. Moreover, it turns out that for the majority of these couples, there is a Wife Dictatorship in the sense that the spending pattern is according to her preferences.

Power has a somewhat different connotation in Chapter 5. This paper deals with the wage negotiations between employees on the one hand, and employers on the other hand. The bargaining power that each party brings to the table is partly determined by the relevant outside options. How simple and cheap it is for an employer to dismantle an employment contract, or in other words how easy it is to fire a worker, is important in this respect. An employee's power is furthermore influenced by the extent to which she has control over the effort she exercises in her job (which is an important resource to the employer). In this chapter, the influence of the employer's term of notice is examined. The term of notice is defined here as the amount of time an employer has to notify an employee in advance of her upcoming dismissal. I find that a worker that is protected by a longer term of notice (so that the worker is more costly to fire) earns a significantly higher wage than a comparable individual with a shorter term of notice. Because an exogenous policy change in the term of notice is analyzed in the chapter, it is argued that this relationship should be interpreted causally. It is probable that the higher wages are explained by the enhanced bargaining position of workers with a longer term of notice.

On Expectations

Chapter 2

Have You Heard the News? How Real-life Expectations React to Publicity[1]

2.1 Introduction

Economists like to believe that expectations play a role in almost every decision that economic agents make. In our models, retailers for example have beliefs over competitors' future prices while determining their own price and households are assumed to take their unemployment risk into account before signing a mortgage contract. The academic debate on the actual role of expectations in decision-making is still ongoing though. An expanding literature relates subjective expectations, sometimes measured in probabilities, to individual behavior such as saving, criminal activities and smoking. Interesting papers on the relationship between observed expectations and observed actions include Stephens (2004), Lochner (2007) and Viscusi and Hakes (2008). While I try to contribute to this literature in the Chapter 3 by showing that old age social security expectations are related to pension scheme participation, the current chapter takes one step back.

This study investigates how public information dissemination influences observed expectations for different demographic groups. As more and more evidence is being presented that subjective expectations are

[1]Thanks to Arthur van Soest, Frederic Vermeulen, James Banks, Jan van Ours, Damjan Pfajfar, Stefan Hochguertel, Mauro Mastrogiacomo and several seminar participants for useful comments and suggestions.

important in decision-making, policy-makers would like to know whether expectations can be influenced through information dissemination in the media. This is especially important when individuals' expectations are overly optimistic or overly pessimistic. Two examples in the literature of systematically biased expectations are the yen/dollar exchange rate expectations that were too low for Japanese exporting companies (Ito, 1990) and the probabilities of obtaining certain diseases that were too high for Dutch women (Carman and Kooreman, 2007). In assessing the importance of the publicity reaction, I am answering the call for more knowledge on the underlying mechanism generating expectations by both Manski (2004) in his influential paper on measuring expectations[2] and by Bernanke (2007) in his speech for the NBER monetary economics summer workshop[3].

Unlike the majority of papers in the empirical expectations formation literature (e.g. Gramlich, 1983, Caskey, 1985, Keane and Runkle, 1990, Souleles, 2004 and Lamla and Lein, 2008), this chapter will not base its findings on inflation expectations. Instead I will analyze the influence of publicity on Dutch old age social security expectations, specifically focusing on individuals' expectations regarding the future eligibility age[4]. These expectations are recorded in the Pensionbarometer, a monthly Dutch longitudinal household survey of pension policy expectations, collected by Tilburg University's CentERdata. As U.S. social security expectations are typically best documented, the literature mostly focuses on Americans' perception of their old age benefits (e.g. Dominitz, Manski and Heinz, 2003). Subjective social security expectations have been studied mostly in order to understand the impact of social security policy on retirement savings (e.g. Bernheim and Levin, 1989, Dominitz et al., 2002 and Van der Wiel, 2008).

A major advantage of the Dutch eligibility age expectations over the

[2] "I see a critical need for basic research on expectations formation. Understanding how persons revise their expectations with receipt of new information often is a prerequisite for credible use of econometric decision models to predict behavior." (Manski, 2004, p.1371)

[3] "We must understand better [...] the relationship between policy actions and the formation of inflation expectations." [...] A fuller understanding of the public's learning rules would improve the central bank's capacity [...] to evaluate the implications of its policy decisions and communications strategy."(Bernanke, 2007, p.4)

[4] The Dutch old age social security system, 'AOW', is a universal Pay-As-You-Go pension scheme that is currently rewarded to all Dutch citizens from the age of 65 onwards. The monthly amount an individual receives depends on years of residence in The Netherlands and cohabitation status, and is independent of contributions, income or wealth.

U.S. expectations is that there cannot be valuable private information involved when Dutch individuals consider the future of their old age social security system. Future benefits perceptions in the U.S. are influenced by individual health, job security and income expectations, besides future policy expectations. The same holds for inflation expectations: private information plays an important role as absorbed price information varies across consumers. Unlike the U.S. system, Dutch old age social security is uniform and universal, and expectations regarding its future will therefore only reflect general policy expectations. The information set that individuals could use in forming expectations would for example include the debates and decisions of the Dutch Parliament as well as Statistics Netherlands' estimates of future population growth. This information is all publicly known. A lack of relevant private information is necessary in order to distinguish the publicity reactions of different demographic groups. If unobservable private information would also play a role in the eligibility age expectations, it would be impossible to separate differences in information absorption from differences in the available information set.

Like in most of the expectations formation literature based on survey data (e.g. Carroll, 2003, Mankiw, Reis and Wolfers, 2003, Souleles, 2004, and Branch, 2004), I also observe a high degree of heterogeneity in the level of expectations. The theoretical model developed in this chapter takes this heterogeneity explicitly into account. A method is proposed to assess the influence of publicity on expectations using the second moment of expectation changes. The variance is the relevant unit of analysis here as all publicity reactions are a-priori interesting; i.e. both the positive and negative expectations shifts. The proposed estimator is named the Publicity Reaction Coefficient (PRC) and is estimated for different demographic and media consumption groups.

Public information dissemination, or publicity, is defined here as the information concerning the entity at interest that is disseminated through various public media outlets. Although publicity as such is practically impossible to quantify, newspaper articles are carefully documented in the LexisNexis database. From this I have constructed a dataset containing the weekly frequency of newspaper articles on old age social security, which will be used as a proxy for publicity intensity. Naturally, it is unobserved how many and what newspaper articles individuals read or take notice of when determining expectations. However, those survey periods before which hardly any relevant newspaper articles appeared can be utilized. The observed expectation changes in these surveys are likely to be realizations of the heterogeneous reporting error term which allows me to estimate the value of the reporting error variance per subsample. The Publicity Reaction Coefficient is then estimated by subtracting this error variance from the variance of expectation changes in high-publicity periods.

It turns out that media coverage of the old age social security system typically accounts for between zero and thirty percent of the variance in expectation changes. Rather than publicity, random shifts thus explain the majority of the variance in reported changes in eligibility age expectations, with especially older individuals being more imprecise in their expectation reports. When comparing subgroups, I find that middle-aged, those who infrequently read a newspaper, those who consider themselves below-average informed about Dutch politics and those who read free or populist newspapers have a relatively high PRC. These respondents thus change their expectations more after heavy media attention than other demographic groups. On the contrary older age groups, those with a university degree and those who earn a relatively high income have a low Publicity Reaction Coefficient. These socio-economic groups are thus relatively insensitive to heavy media attention on old age social security. This could be counter-intuitive at first, but several plausible explanations can be given. First, it becomes understandable when the majority of newspaper articles about the old age social security system contains 'old news'. It is after all likely that those with lower quality initial expectations adapt their expectations more whenever already available information is repeated in the media. Second, another explanation could be that those for whom the future of old age social security is more important - those with lower incomes for example - pay more attention to the relevant media exposure.

This chapter will proceed with a short overview of some of the relevant economic literature in Section 2.2. An econometric approach to estimate the Publicity Reaction Coefficient is then proposed in Section 2.3. Furthermore, I describe the Pensionbarometer dataset (Section 2.4) and the information set that was available to respondents during the observation window (Section 2.5). Section 2.6 presents the results of the empirical analysis. Section 2.7 concludes.

2.2 Literature

Before data on subjective expectations were available, economists already formulated models for expectations formation in order to complete their theoretical and structural econometric models. Especially in monetary economics expectations formation is of the utmost importance, and most of the literature thus analyzes inflation expectations. The majority of expectations formation theories have one thing in common; they link the relevant information that is available at time t to what organizations, firms and individuals expect to happen at time $t + 1$. Among the first to formalize a theory on expectations formation were Arrow and Nerlove (1958) who expanded the adaptive expectations theory in which one learns from previous prediction errors. Another well-known example of such a theory is the rational expectations hypothesis proposed by Muth (1961). He introduced the convenient notion that individuals and firms use all the currently available information to correctly compute expectations so that their ex-post prediction errors are orthogonal to this information. Given the evidence of judgement biases in simple evaluation tasks (e.g. Tversky and Kahneman, 1974), the rational expectations assumption is unlikely to hold in reality. Several authors have indeed shown that not all relevant and available information is incorporated in expectations (e.g. Figlewski and Wachtel, 1981, and Carroll, 2003) and that changes in expectations are typically more extreme than ex-post justified (e.g. De Bondt and Thaler (1990)). As most of these studies analyze inflation or exchange rate expectations of professional forecasters, the extent to which expectations of the general public are biased is probably even larger.

Mankiw and Reis (2002) relatively recently introduced a more realistic theory: the sticky expectations model. In this expectations formation model all new information is correctly absorbed in the expectations of a certain fraction of the population, while the other - ignorant - fraction sticks to their previous expectations. Although a vast improvement to the earlier literature on expectations formation, the sticky expectations literature does not discuss which individuals belong to the updating and which belong to the ignorant group. In fact, there is no room for individual heterogeneity as the theory assumes that each period all individuals are equally likely to update their expectations.

Some authors have tested the sticky-expectations model by quantifying the fraction of the population that belongs to the ignorant group, typically analyzing ex-post prediction errors in inflation expectations. Mankiw, Reis and Wolfers (2003) estimate that professional economists in the Livingstone survey update their expectations once every ten months and that households in the Michigan Survey update their inflation expectations once every thirteen months. This comes down to more than ninety percent of individuals ignoring new information revelations each month.

Carroll (2003) finds that the typical household in the Michigan Survey updates inflation expectations roughly once a year, while unemployment expectations appear to be updated more frequently.

Carroll also pays attention to media intensity on the relevant random variable, proxied by the number of newspaper articles published about inflation on the front page of the New York Times and the Washington Post each year. In his paper, Carroll provides a micro-foundation for Mankiw and Reis's sticky expectations, as he proposes that the media are the information channel through which individuals, in different degrees, absorb professional forecasters' inflation expectations. He uses the Michigan Survey to find that the average inflation expectations of households are more accurate, i.e. more like the average expectations of experts, in periods in which there is a lot of news coverage on inflation. Carroll also finds that the updating speed in the total population is faster, or the fraction of the population that has rational expectations, when there has recently been more news coverage. Another example of a study in which publicity plays a direct role is a paper by Lamla and Lein (2008) who examined the effect of both printed media and television coverage on average German consumers' inflation expectations. Their findings support Carolls results that more news indeed leads to household expectations being closer to expert expectations. Lamla and Lein also investigate the effect of the wording of inflation news and find that the expectation bias increased in periods in which the media were blaming the introduction of the euro for significant price increases. Given that both studies analyze average expectations rather than individual expectations, they cannot address if and how demographic groups differ in their uptake of relevant information in the media. As an addition to their work, this study thus focuses on this heterogeneity in the publicity reaction.

While ignoring the publicity element of expectations formation, other authors have examined demographic differences in expectations. Several papers have analyzed the differences in ex-post expectation errors between professional forecasters, or economists, on the one hand and households on the other hand (e.g. Gramlich, 1983). Unexpectedly, this comparison does not always favor the professionals. Gramlich also provides sketchy evidence that low-income and low-education households forecast inflation better. He argues that this counterintuitive finding has its origin in the accelerating inflation during the survey period 1978-1979. Congenital pessimists, which he believes low-income and low-educated individuals are, would have had a forecast advantage in that period. Jonung (1981) was one of the first to decompose household inflation perceptions and expectations demographically and he found that women perceived higher inflation rates than men using Swedish data. Jonung suggests that this might have to do with the different products that men and women buy, as especially food inflation had been relatively high in the survey period. This importance of private information sets in inflation expectations has been a major argument to analyze policy change expectations in this chapter. Not all papers in the expectations formation literature focus on inflation expectations. Souleles (2004) related macro-economic forecast errors to demographic characteristics and he found that the forecast bias decreased in magnitude with age, income and education.

2.3 Expectations formation model

2.3.1 A theoretical model of the information reaction

I consider the following expectations formation model that makes only weak assumptions on the relationship between information, publicity and expectations. Individual i reports expectations at time t over the realization of a random variable y at a specified moment in the future, $t+z$. This expectation will be denoted as $\tilde{y}_{it}^{t+z}$ and is the sum of three components: a function f_i of the absorbed information set that is available at time t, Ω_{it}, a time-invariant component, υ_i, and a reporting error term, ϵ_{it}. The following equation presents this expectations formation model:

$$\tilde{y}_{it}^{t+z} = f_i(\Omega_{it}) + \upsilon_i + \epsilon_{it}.$$

Individual heterogeneity enters the expectations formation formula in many ways. First, individuals do not necessarily absorb all available information. This means that the personal information set Ω_{it} is a subset of the generally available information set Ω_t ($\Omega_{it} \subseteq \Omega_t$). Note that this means that private information is excluded in this expectations formation model; individuals differ in their information absorption but they could in principle all access the same knowledge.

A second source of individual heterogeneity resides in how the absorbed information is translated into an expectation. The information processing function $f_i(\cdot)$ in the expectations formation model therefore has subscript i. Although information absorption and information processing are conceptually different concepts, it is empirically very difficult to distinguish the two effects. Therefore, I will from now on refer to their combined effect on expectations as the information consumption effect.

The expectations formation model includes two additional sources of interpersonal variation. The time-invariant component υ_i represents individuals' inherent inclination towards low or high expectations, i.e. towards pessimism or optimism. Moreover, I make the following assumptions on the distribution of the error term ϵ_{it} that allow for individual heteroscedasticity:

$$\epsilon_{it} \sim n.i.d.(0, \sigma_i^2).$$

The larger the variance of the error term, σ_i^2, the larger is the typical deviation in the expectations report from the information-based expectation. In this model, individuals thus differ in the size of these deviations. A way of thinking about this is that someone with a small error variance is a precise individual and someone with a high error variance is typically sloppy.

One source of heterogeneity in the expectations, namely the individual-specific constant term υ_i, can easily be eliminated by taking first differences. It is important to note that in empirical applications this individual constant term will, besides optimism, also include the influence of initial information on expectations. This is because the expectations formation process will have started (long) before the expectations are first observed by the researcher. The first difference expectations formation model can be found below:

$$\begin{aligned} \tilde{y}_{it}^{t+z} - \tilde{y}_{i,t-1}^{t+z} &= f_i(\Omega_{it}) - f_i(\Omega_{i,t-1}) + \epsilon_{it} - \epsilon_{i,t-1} \\ \Delta_{it}\tilde{y}^{t+z} &= \Delta_{it} f(\Omega) + \Delta_{it}\epsilon. \end{aligned}$$

Note that that the second line in this equation is simply a rewritten version of the first; no additional assumptions have been made. Expectation changes are thus a linear combination of the change in the processed and absorbed information set ($\Delta_{it} f(\Omega)$) and an error component ($\Delta_{it}\epsilon$). As a result of the assumptions on ϵ_{it}, these new errors are independent normal with mean zero and a variance that is twice the individual error variance σ_i^2. The systematic effect of information on expectation changes ($\Delta_{it} f(\Omega)$) is referred to as the information reaction. This chapter will empirically assess the size of this effect and how this size varies over individuals.

2.3.2 Information, publicity and newspaper articles

Public information dissemination, or publicity, is likely to play an important part in how individuals obtain the information that is absorbed in their expectation. The publicity set, Θ_t, is defined here as the set of all information on the realization of the random variable y at time $t+z$ that is disseminated through publicly available media outlets between period $t-1$ and period t. These outlets would include television programmes, Internet blogs and magazine and newspaper articles. The publicity set, Θ_t, can be divided in two separate subsets: the set of truly new information revelations that were unavailable last period, $\Delta_t\Omega$, and the set of repeated information elements that were already known last period, Φ_t. The exact explanation of the publicity set is given in the following equations:

$$\begin{aligned}\Theta_t &= \Delta_t\Omega \cup \Phi_t \\ \Delta_t\Omega &\cap \Phi_t = \emptyset \\ \Delta_t\Omega &\not\subseteq \Omega_{t-1} \\ \Phi_t &\subseteq \Omega_{t-1}.\end{aligned}$$

Note that the publicity set is defined such that each informational element can only enter the set once, so that the number of times that a certain piece of information is mentioned in the media does not play a role in Θ_t.

I assume that the only sources from which individuals derive new information, $\Delta_{it}\Omega$, are the public media outlets. Library visits and say, university lectures, are thus excluded. Note that one does not have to watch a certain television programme oneself to know what it was about. I am allowing for hearing about news from others - colleagues, family, friends - as long as there is only a short period between the broadcasting of the television programme and the informal information dissemination. Lagged publicity effects are however excluded. A person that chooses not to read the newspaper in a certain period cannot decide to read it three months later. This is not such a strong assumption as repetitions of existent information (Φ_t) are an important component of the publicity set (Θ_t). These two assumptions on the newly absorbed information are summarized as follows:

$$\Delta_{it}\Omega \subseteq \Theta_t.$$

Unfortunately, there are two empirical problems concerning the publicity set Θ_t. First, as I am interested in quantifying the information reaction in expectations, a quantitative measure of information dissemination is much more useful than a set of informational elements. This is because each element of the information set will have a distinct influence on expectations and these effects cannot be separated empirically. Second, it is technically impossible to exactly observe all informational elements of the publicity set Θ_t. Even if a researcher would have time to carefully read the thousands of relevant newspaper articles, there is no way to assemble all television programmes that have featured information on the random variable y. Because of these data problems, a new quantitative and observable measure of publicity intensity is introduced, θ_t, which is defined as the number of newspaper articles that mention the relevant random variable y (in this chapter, the old age social security eligibility age) in their content . Newspapers are naturally not the only available media outlet. However, θ_t is restricted to newspaper articles as this is the only information source that can easily be retrieved. The underlying assumption is that the correlation between the publicity intensity across all media outlets is very strong.

A positive relationship between the number of elements in the publicity set Θ_t and the quantitative measure of publicity intensity θ_t is easy to imagine. It is likely that the more newspaper articles appear about a certain subject, the more information about that topic will be disseminated. How strong this relationship is from period to period is hard to say however. One frontpage article by a well-known journalist could have much more influence than a standard press release that has been copied in all large newspapers. I therefore do not further formalize the relationship between the content and quantity of media coverage. The only assumption made is that the publicity set Θ_t is empty when the amount of newspaper articles on the relevant variable y is very low, i.e. when $\theta_t \approx 0$. What is meant with very low will become explicit in Section 2.5.

It turns out that the restrictions on no lagged publicity and on the relationship between Θ_t and θ_t are very useful assumptions. Together they imply that the information-related component of expectation changes - i.e. the information reaction - will equal zero when the media intensity on the random variable is very low. The information reaction in such circumstances is as follows:

$$\Delta_{it} f(\Omega) = 0 \text{ if } \Theta_t = \emptyset \text{ , i.e. when } \theta_t \approx 0.$$

Because this proposition requires important assumptions on how individuals obtain their personal information set from the media, it is appropriate to name the information related component of expectation changes the publicity reaction here rather than the information reaction. Although the two are the same by assumption I want to avoid the possible misinterpretation that the information effect has to do with real information revelations only ($\Delta_t \Omega$) rather than also with repeated elements of the information set (Φ_t).

Note that a lack of publicity is not the only circumstance as a result of which a zero publicity reaction is observed. The publicity reaction could also be zero when 1) the individual does not absorb any new publicity, 2) all elements of the consumed publicity set were already known to the individual or 3) the individual does not change her expectation after processing the newly absorbed information.

2.3.3 Identification of the publicity reaction

The proposition that the informed component of an expectation does not change whenever the publicity on the realization of the relevant variable has been negligible will prove useful in identifying the actual publicity reaction later on. To see this, let me first define two types of time periods: the set L contains periods in which there is hardly any publicity between $t-1$ and t ($t \in L$ if $\theta_t \approx 0$) and the set H contains periods before which some publicity, containing real and/or repeated news, has appeared ($t \in H$ if $\theta_t \gg 0$). For these period subsets, different expectation change expressions can be given.

$$\Delta_{it}\tilde{y}^{t+z} = \begin{cases} \Delta_{it}\epsilon \text{ if } t \in L \\ \Delta_{it}f(\Omega) + \Delta_{it}\epsilon \text{ if } t \in H. \end{cases}$$

It turns out that in low-publicity periods expectation changes are only a realization of the change error, while in high-publicity periods the publicity reaction enters the expression as well. In order for this distinction between L and H to be empirically relevant I need to observe expectation changes in time periods in which the quantity of relevant newspaper articles is low ($\theta_t \approx 0$) and in periods in which it is high ($\theta_t \gg 0$). In Section 2.5 it is explained that I indeed observe such periods for my social security policy expectations.

It is difficult to derive interesting empirical results from the expectations change model in terms of positive and negative changes. This is because both increases and decreases in expectations could be due to publicity reactions but such observations would cancel out in estimations. Instead, the variance of expectation changes which, under assumptions, can provide insights into the degree to which expectations are formed by publicity consumption. In this chapter, the word variance will represent the second moment of expectation changes, i.e. the average squared expectation changes. This is a very intuitive unit of analysis, as it will always increase in the degree of publicity reaction. The traditional variance measure is the second central moment, i.e. the extent to which different observations of a variable typically vary from its mean ($E[(\Delta_{it}\tilde{y}^{t+z} - \overline{\Delta\tilde{y}})^2]$). For the publicity reaction analysis however the second moment ($E[(\Delta_{it}\tilde{y}^{t+z})^2]$) rather than the second central moment is utilized as changes in expectations and not levels are the unit of analysis. Think about an individual that increased her expectation by the same (large) amount for a couple of time periods in a row. The second central moment of this person's expectation changes would equal zero, although it is very well possible that the individual did (strongly) react to publicity. The second moment of her expectation changes will have been relatively large however.

The data at hand does have to satisfy one non-testable assumption in order for the variance to be a suitable unit of analysis. I have to assume that there is no linear relationship between the publicity reaction and the new error term i.e. the covariance between these elements of expectation changes should equal zero:

$$E[\Delta_{it} f(\Omega) \Delta_{it}\epsilon] = Cov(\Delta_{it} f(\Omega), \Delta_{it}\epsilon) = 0.$$

The zero correlation assumption is necessary for the variance of the expectation changes to be the sum of the variance of its components. Note that this relatively weak assumption does not rule out heteroscedasticity. It remains possible for the variance of the change errors ($2\sigma_i^2$) to be a function of individual time-invariant characteristics. If the covariance is indeed zero, the variance of the expectation changes can be decomposed as follows:

$$Var(\Delta_{it}\tilde{y}^{t+z}) = \begin{cases} E[(\Delta_{it}\epsilon)^2] \text{ if } t \in L \\ E[((\Delta_{it} f(\Omega))^2] + E[(\Delta_{it}\epsilon)^2] \text{ if } t \in H. \end{cases}$$

The periods with very little publicity on the variable that individuals have expectations over (L) can now be exploited to identify the proportion of the variance in expectation changes related to publicity. A two-step procedure has to be applied. First, I can estimate the variance of the error component in expectation changes by computing the average squared expectation change in the low-publicity periods. Second, I can decompose the total variance in publicity-rich periods (H) into the variance of the publicity reaction and the variance of the error term which is estimated in the first step. This can be done as the variance of the reporting error term is assumed constant over time, so that the variance of the error term in low-publicity times will be similar to that in high-publicity times. This procedure then generates an estimate of the proportion of the variance in expectation changes that is related to information dissemination.

Using this estimate, I can calculate the Publicity Reaction Coefficient

(PRC). The PRC is defined as the ratio of the estimated systematic variance over the total variance of expectation changes. The higher the PRC, the more important is the role of information dissemination in expectation changes. Equation 1 displays how the Publicity Reaction Coefficient is estimated.

$$PRC = \frac{Var(\Delta_{it} f(\Omega))}{Var(\Delta_{it}\tilde{y}^{t+z})} = \frac{E[(\Delta_{i,t^H}\tilde{y}^{t+z})^2] - E[(\Delta_{i,t^L}\tilde{y}^{t+z})^2]}{E[(\Delta_{i,t^H}\tilde{y}^{t+z})^2]} \tag{1}$$

The Publicity Reaction Coefficient is ideally estimated on the individual level. This can be done when sufficient observations are available per individual in both publicity-rich and publicity-poor periods. Observations in the latter periods are especially important because these are used to estimate the variance of the error term. The dataset that will be used in this chapter does not contain enough of these observations for each individual. Therefore the reporting error variance and consequently the PRC will be estimated per demographic and media consumption subgroup. The subscript i in this section should for my empirical application thus be interpreted as representing different groups rather than persons.

2.3.4 The prediction error variance estimated

Under the assumption of normality of the reporting error term it is furthermore possible to see how the noise variance σ_i^2 is related to demographics. Although this is not central to the publicity reaction analysis, it is nevertheless interesting to see which groups are more prone to errors in their expectation estimates. I choose to model the heteroscedasticity in the variance of $\Delta_{it}\epsilon$ in the multiplicative fashion that is often used. Like in the textbook examples, the assumption is that the individual variance is an exponential function. The squared expectation changes in periods before which no publicity entered the scene ($t \in L$) can then be seen as drawings from the variance distribution. Taking logs on both sides, the model simplifies to a linear equation which can be estimated using OLS. These steps are laid out in the equations below. The γ-coefficients determine which demographic groups make larger errors. The covariates in $x_i^{'}$ include age groups, educational categories, income groups and dummies for gender and marital status.

$$
\begin{aligned}
Var(\Delta_{it}\epsilon) &= E[(\Delta_{i,t^L}\tilde{y}^{t+z})^2] \\
(\Delta_{i,t^L}\tilde{y}^{t+z})^2 &= 2 * \sigma_i^2 = 2 * e^{(\alpha + x_i^{'}\gamma + \eta_{it})} \\
\ln((\Delta_{i,t^L}\tilde{y}^{t+z})^2) &= (\ln(2) + \alpha) + x_i^{'}\gamma + \eta_{it} \\
\eta_{it} &\sim n.i.d.(0, \varsigma^2).
\end{aligned}
$$

2.4 Expectations data

2.4.1 The Pensionbarometer

Although this paper study wishes to say something about the publicity reaction in general, the expectations that will be empirically examined are people's beliefs over the future of the Dutch old age social security eligibility age. This subjective expectation is interesting and useful in the publicity reaction analysis for the following four reasons. First, as the Dutch old age social security system (the 'AOW') is completely universal, i.e. all individuals receive a similar level of benefits from the same age onwards irrespective of contributions paid, all forecasts about potential policy changes refer to the same random variable. The expectations are therefore fully comparable. Second, the majority of individuals expects to grow old and therefore to receive the non-negligible[5] 'AOW' benefits at some point in time. Hence, the future of the old age social security system is of direct consequence to respondents, thereby enhancing the quality of answers provided. Third, as the future of the old age social security system basically depends on a few individuals only (i.e. on 150 members of parliament) and as the debate about its sustainability is followed closely by the media, it seems plausible that all the information that could enter the formation process of eligibility age expectations is publicly available. This implies that differences in expectations only reflect differences in which pieces of information individuals absorb and on how this information is processed, not on which information individuals in principle have access to.

The expectations data analyzed in this chapter are taken from a Dutch survey on pensions' opinions and expectations, the Pensionbarometer,

[5] In 2008 the level of an 'AOW' benefit was equal to gross €997,12 for a single individual and €682,51 for a cohabitating individual. Note that the gross amounts would almost have been equal to net amounts in this case.

which is collected by CentERdata at Tilburg University in The Netherlands. Netspar, a Dutch research network on pensions, ageing and retirement, has initiated and funded the data collection process. The Pensionbarometer is collected mainly to produce longitudinal statistics for confidence-levels in the Dutch pension system. Every month about 500 individuals are asked to answer a few general questions about social security, pensions and retirement via the Internet. The respondents are randomly selected from the CentERdata Internetpanel, which in turn is a random sample of the population drawn from the municipal registry. Whenever a selected individual does not have access to the Internet, CentERdata makes sure that access is arranged. The Pensionbarometer is designed as a panel-type survey, and every respondent receives the questionnaire every three months so that there are in total three recurring groups of respondents. This was done to be able to produce the confidence statistics monthly while not overwhelming the respondents with questions. In this chapter, I use the first thirty-one waves of the Pensionbarometer (16,669 observations). The first wave of the survey was enumerated in May, 2006. The last wave that is included in my analysis was enumerated in November, 2008. A total of 2,573 individuals were interviewed. Individuals participated eight times on average. 296 individuals reply only once (there is replacement sampling) and 707 individuals answer the probability questions ten or more times.

The Pensionbarometer asks a series of four questions about individuals' eligibility age expectations that read as follows:

> "What is the probability (between 0 and 100) that within ten years/ within twenty years the 'AOW'-eligibility age will be ...
> - higher?
> - higher by at least two years?
> - lower?
> - lower by at least two years?"

The questions were always asked in the order above. Four rather then one questions were posed to provide insight in the internal consistency and underlying distribution of the expectations. For longitudinal studies it is however more interesting to examine each question separately as this allows me to focus on changes in expectations. The total sample PRCs will thus be reported for all four probabilities. To be concise, the subgroup Publicity Reaction Coefficients are however reported for the first question on a higher eligibility age only.

Each Pensionbarometer survey, the time horizon for which an individual is asked her expectations is randomly chosen and this time horizon thus varies over the surveys for each individual. The questions either concern a ten or a twenty years time horizon. As there is no clear, cardinal relationship between the ten and twenty year expectations the results for both horizons are reported separately. This leaves me with a first sample of expectation levels with 8,521 ten year observations (2,339 individuals) and 8,148 twenty year observations (2,322 individuals). A second sample of expectation changes is smaller for two reasons. First, the initial Pensionbarometer wave drops out when changes are considered. Second, consecutive expectation answers over different time horizons - when for example first ten year expectations are given and later twenty year expectations - are also ignored. A change in expectations could in such a case namely also reflect a different opinion on the time horizons rather than a publicity reaction. 4,646 first period observations are lost and another 6,716 observations are lost because the time horizon that the individual was asked to provide an expectation over changes from one survey to the next. This ultimately generates a second sample of 2,736 ten year expectation changes (1,296 individuals) and 2,571 twenty year expectation changes (1,283 individuals).

The Pensionbarometer data also contain several important demographic characteristics such as education level, age and income. The Publicity Reaction Coefficient of all relevant subgroups will be estimated in Section 2.6. As the PRC is estimated on the second sample, information on the covariates is reported for this sample only. Three independent variables concern media consumption, which will be interesting when analyzing the publicity reaction. The media consumption data is based on additional questions that were asked to all respondents in June 2008. Unfortunately, this means the media information is available for only 987 respondents. Two questions about direct media consumption read:

> "How often do you read a newspaper (free papers and Internet papers included)?"

and

> "How often do you watch a regular news bulletin on the television?"

The response scales individuals could use were 1) Every day, 2) Multiple times a week, 3) Once a week, 4) At least once a month, 5) Less than once a month. The distribution of answers can be found in Table 2.1.

Table 2.1: Frequencies of informative media consumption - in terms of newspaper and television bulletins - for respondents in second sample.

	Newspaper	Television news
Every day	66%	61%
Multiple times a week	18%	30%
Once a week	7%	5%
At least once a month	2%	2%
Less than once a month	6%	3%
Observations	987	987

The Pensionbarometer sample often reads a newspaper and frequently watches news bulletins on television. More than sixty percent of respondents state that they read the newspaper on a daily basis. It should be noted that newspaper take-up is generally high in The Netherlands. The Dutch Bureau of Statistics reports that in 2006 51% of the population had a paid newspaper subscription. For the empirical analysis of this chapter the consumption data is transformed into two dummies for relatively low media consumption: low newspaper consumption (once a week or less) and low television news consumption (once a week or less). This was done because the latter categories contain relatively little observations. Sensitivity analysis has been performed using a different threshold for the dummy. This had little effects on the results.

I also know which newspaper the respondents most often read, as they were asked to state from which newspaper they typically obtained most information. The frequencies can be found in Table 2.2. The most read category are the free newspapers such as Metro, 'Spits' and 'De Pers' which are distributed in and around public transport and in communal buildings. 12% of respondents indicate that they obtain most information from these sources. Other large categories are two morning newspapers, one more conservative/ populist - 'De Telegraaf' and another more progressive - 'De Volkskrant'[6]. It will be interesting to see whether the slant of a newspaper affects how individuals absorb publicity into their expectations in the empirical section.

Additionally, respondents were asked to rate their own knowledge of current political events. The exact question posed read:

"Compared to the average Dutch person, how well informed

[6] The classification of newspapers according to their tone or background is taken from a newspaper article in the 'NRC Next' on March 4 2009.

Table 2.2: Response frequency for which newspaper the respondents most read in second sample.

	Most read
Free newspapers	12%
‘De Telegraaf’ (populist)	11%
‘De Volkskrant’ (progressive)	10%
‘NRC Handelsblad’ (liberal)	7%
Other	60%
Observations	926

Table 2.3: Response frequency for subjective political informedness for respondents in second sample.

	Being informed
Very bad	3%
Bad	14%
Average	52%
Well	28%
Very well	4%
Observations	987

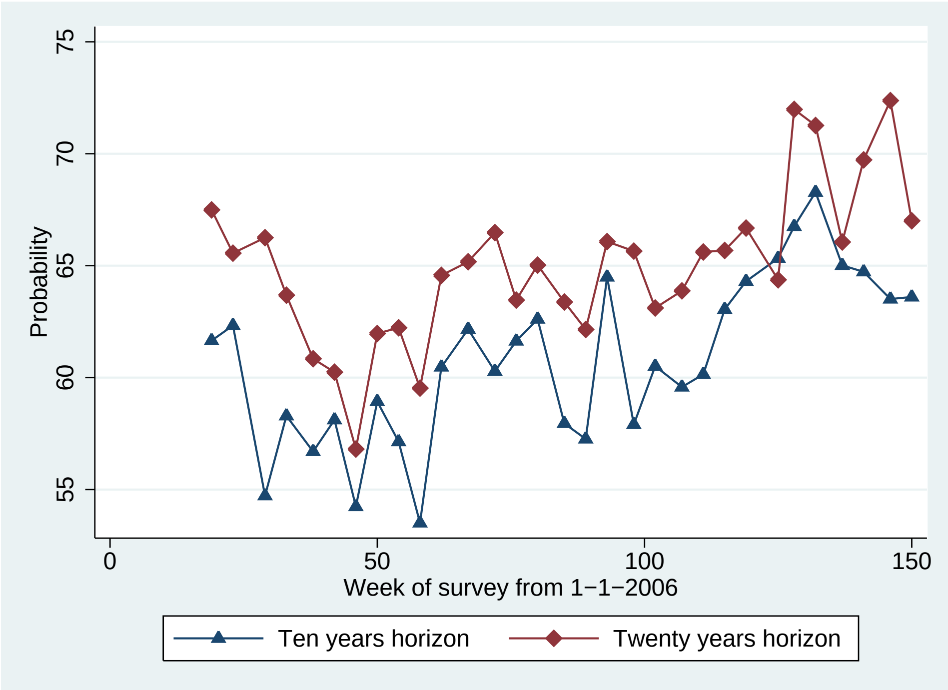
are you about current affairs in Dutch politics?”

Respondents could choose from five answer categories: 1) Very bad, 2) Bad, 3) Average, 4) Well, 5) Very well. Table 2.3 shows that the Pensionbarometer respondents on average consider themselves well-informed, although still sixteen percent believe that they know less than the average Dutchmen. Because of the limited number of observations, in Section 2.6 this data is translated into a dummy for being relatively ill-informed (very bad and bad answer categories). This partition was chosen as it best explains the differences in empirical publicity reactions.

2.4.2 Characteristics of the expectations formation process

Figure 2.1 shows the development of the average subjective expectations in the Pensionbarometer over all waves and both time horizons. Each data-point represents the mean of about 250 observations. The observed average probability of later benefits is well above 50 for all waves and both time horizons. Hence, pessimism prevails. The average does vary over the relevant time horizon. As one would expect, people believe that the prob-

Figure 2.1: Average probability assigned to higher eligibility age for both time horizons (week 1-52: 2006, week 53-104: 2007, week 105-157: 2008).



ability of a policy change within the next twenty years is higher (average is 65) than within the next ten years (average is 61). This difference was consistent in all but one waves. The Pensionbarometer respondents were most optimistic around the end of 2006/ beginning of 2007 (when a new government was being formed) and most pessimistic around the summer of 2008 (when a government appointed committee proposed to raise the eligibility age to 67 years)[7]. The average probability seems considerably volatile; the difference between the smallest and largest average probability is about 15 probability points for both the ten and the twenty year expectations. A histogram in Figure 2.A.1 in the appendix provides more information on the distribution of the twenty year answers. Some clustering of probabilities around 50 is visible. It has been argued that respondents also provide a fifty percent chance response when they have no idea what probability to assign. Manski (2004) however states that

[7]Note that high probabilities are named pessimistic here and low probabilities optimistic out of convenience, not as a normative statement.

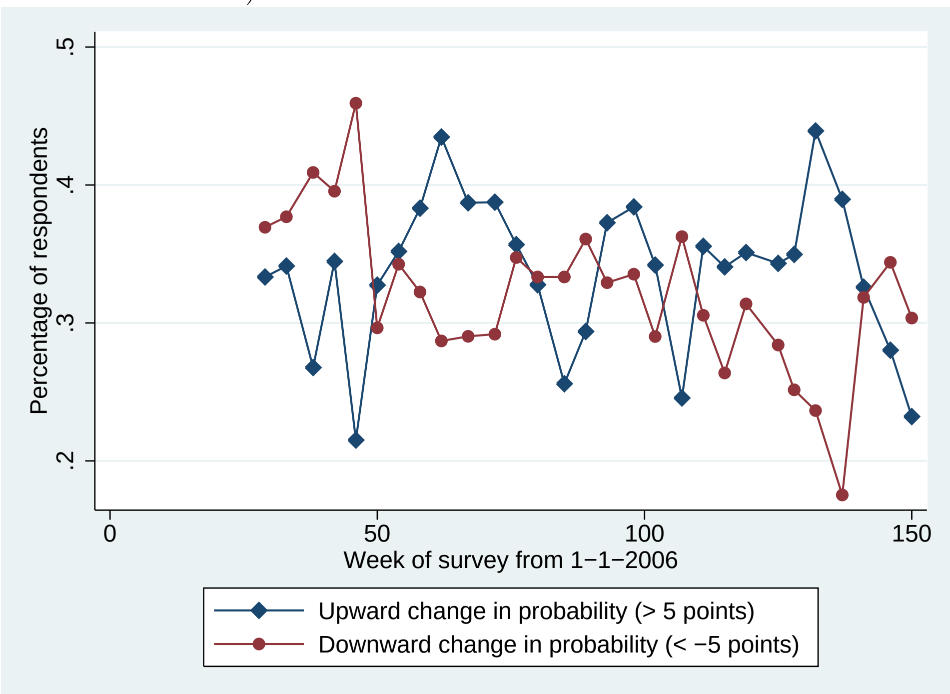
"This concern has largely been laid to rest as empirical evidence has accumulated." (p. 1342). Given that only one percent of respondents always answer 50 to both the ten and twenty year horizon questions I will assume here that all 50 answers represent true opinions.

Although in general respondents are pessimistic about the future of the old age social security eligibility age, considerable heterogeneity in expectations exists. Figure 2.A.2 in the appendix represents the median and the 25th and 75th percentiles for the expectations with a twenty year time horizon. The average interquartile range is substantial and relatively constant over time, at about 40 probability points. This individual heterogeneity in the assessment of future policy changes is to a certain degree persistent. Those individuals who assign a higher (or lower) than average probability in one period are more likely to do so in the next period. The probability of assigning a higher than average probability to a policy change in two consecutive periods is equal to 74% for the ten year expectations and 67% for the twenty year expectations. The probability of assigning a lower than average probability twice in a row is 77% for the ten year expectations and 77% for the twenty year expectations. This means that despite the general tendency to remain less or more optimistic than others, still more than twenty percent of individuals move their expectation to the other side of the distribution from one period to the next.

The vast majority of individuals, also those who remain on one side of the average expectation, alter their stated expectations between consecutive surveys. Figure 2.2 shows the percentages of respondents that change their subjective probabilities by more than five points compared to their previous answer. Absolute changes below six probability points (say from 65 to 60) are ignored as these could just be due to rounding differences. Remember that my second sample only includes observations of individuals who happened to have answered questions for a similar time horizon in two adjacent surveys and that expectation changes are thus measured over a three month period. The average percentage of respondents that display a positive change equals 34% and the average percentage of respondents that display a negative change over all waves equals 32%. About seventy percent of individuals thus significantly update their beliefs within three months. Moreover, only six percent of respondents never change their expectations by more than five probability points.

Table 2.4 finally displays the regression results of two-limit Tobit regressions of the subjective probability of a higher eligibility age on some relevant demographic characteristics for both the ten and the twenty year

Figure 2.2: Percentage of respondents adapting eligibility age expectations since last survey per week (week 1-52: 2006, week 53-104: 2007, week 105-157: 2008).



expectations. A tobit estimation technique was chosen to accommodate censoring of the given probabilities at 0 and 100. In both time horizon regressions, the older one gets, the more optimistic one becomes. Also persistent in both columns; respondents with higher incomes as well as those with a partner are significantly more pessimistic. Gender is also correlated with short-term social security expectations. Women assign higher probabilities to an increase in the eligibility age within ten years. Education moreover has a significant influence on the long-term eligibility age probabilities. Those who completed a higher vocational or an academic education are significantly more pessimistic about the future. Despite the significant covariates, individual heterogeneity and noise seems to be most important in explaining differences in expectation levels as the regressions only explain about 0.4% of the variation in expectations.

2.5 Publicity set

The publicity set Θ_t is defined in Section 2.3 as the set of all informational elements on the realization of the relevant entity (here the old age social security eligibility age) that is disseminated through publicly available media outlets between two survey periods. As it is empirically very difficult to identify the exact elements in and the precise size of the publicity set, Section 3 already introduced the quantity of newspaper articles, θ_t, as a reasonable proxy for the intensity of publicity in a given period. Carroll (2003) and Lamla and Lein (2008) use similar newspaper quantity variables in their papers.

The data on the quantity of newspaper articles is assembled through an online database of written Dutch media content, LexisNexis. Every piece of text that is published in large magazines and newspapers can be retrieved through this database. For each week between the first wave of the Pensionbarometer (week 19 in 2006) and the last reported wave (week 46 in 2008) I have selected all relevant articles that appeared in Dutch national newspapers. Four different series of weekly quantities were generated that could partly overlap: the number of articles containing the acronym 'AOW' in their title, the number of articles containing the acronym 'AOW' in their introduction, the number of articles containing the word 'AOW (eligibility) age' in their total content and the number of articles containing the word 'ageing' in their introduction. Note that 'AOW' is the acronym of the Dutch equivalent of old age social security. Figure 2.3 displays these four series separately. Considerable variation in the number of relevant newspaper articles can be seen, with especially the

Table 2.4: Coefficients and standard errors of two-limit Tobit regressions. Dependent variable: probability of higher eligibility age (between 0 and 100).

	(1) 10 years		(2) 20 years	
	Coef.	S.e.	Coef.	S.e.
Vocational secondary	-4.638*	(1.90)	-1.341	(2.940)
General secondary	-1.317	(2.11)	3.845	(3.124)
Vocational tertiary	0.052	(2.01)	1.430	(3.016)
Higher vocational	1.803	(1.98)	4.921*	(2.950)
Academic education	1.865	(2.21)	7.709***	(3.158)
Age 15-24	-7.106	(35.84)	-0.163	(1.865)
Age 35-44	-6.414***	(1.42)	-6.305***	(1.631)
Age 45-54	-8.283***	(1.36)	-5.281**	(1.616)
Age 55-64	-11.889***	(1.38)	-8.880***	(1.721)
Age >64	-14.394***	(1.40)	-14.086***	(1.761)
Low middle income	1.954	(1.31)	0.385	(1.753)
Middle income	5.828***	(1.42)	4.968***	(1.872)
High middle income	5.350***	(1.48)	4.895**	(1.965)
High income	6.049***	(1.62)	5.824***	(2.094)
Dummy female	2.570**	(0.99)	0.316	(1.307)
Dummy partner	6.318***	(0.99)	6.476***	(1.296)
Observations	8,450		8,087	
Individuals	2,317		2,290	
Censored at 0	288		231	
Censored at 100	1,285		1,491	
Pseudo R^2	0.40%		0.42%	

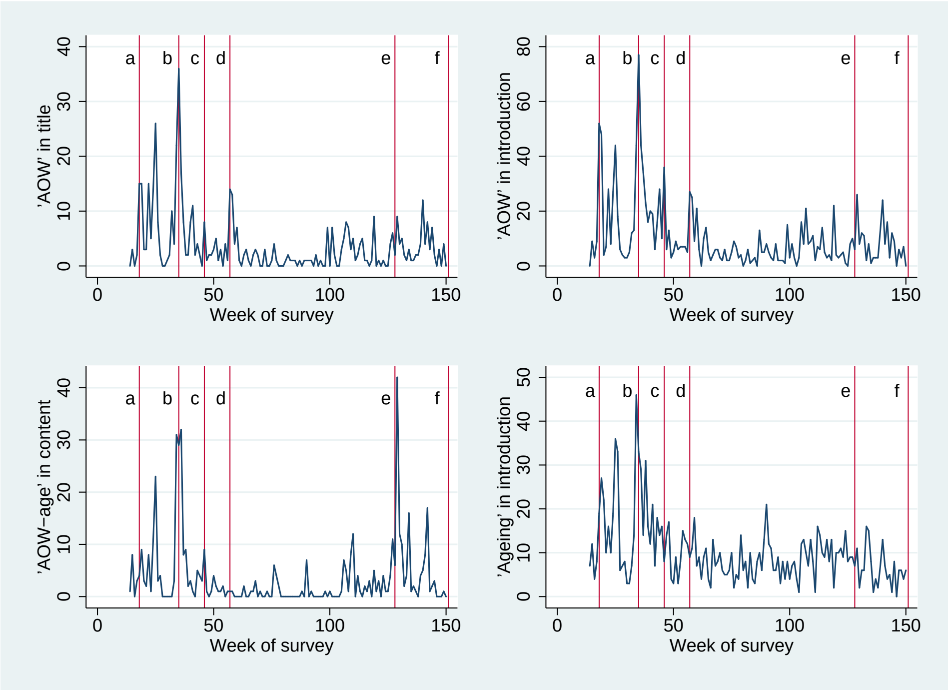
Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Base categories: primary education, age 25-34, low income middle time period low in media attention for all four series.

Some peaks in the publicity quantities have a deducible information

Figure 2.3: Weekly number of newspaper articles in Dutch national media featuring words relevant to the old age social security system and special interest weeks (a-f) (week 1-52: 2006, week 53-104: 2007, week 105-157: 2008).



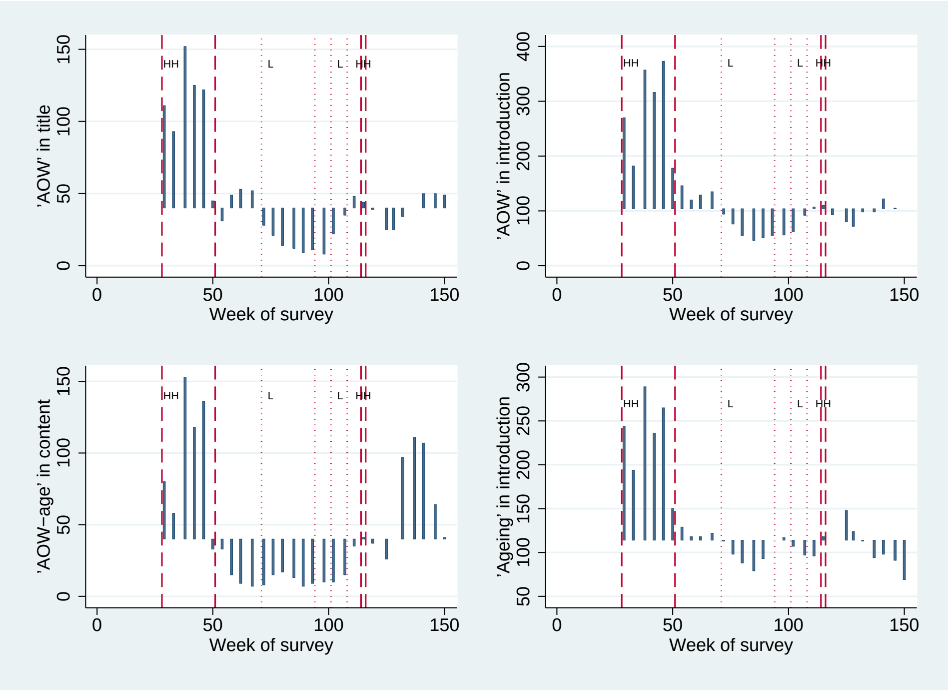
cause and these events are indicated with vertical lines in Figure 2.3. Here is a summary of important events related to old age social security during the period May 2006 - November 2008: a) April 2006: The leader of the social-democrats (Mr. W. Bos) gives a speech in which he promotes higher tax rates for retired individuals. These rates have been lower because those over 64 do not pay old age social security contributions; b) August 2006: The presentation of general election programs by all relevant parties. Most parties, including the liberals and Christian-democrats, promise to keep the old age social security scheme exactly as it is. Some small parties propose to raise the eligibility age. The social-democrats stick to their idea to alter the contributions scheme; c) November 2006: The election results come in. Relatively to polls half-a-year earlier, the Christian-democrats win and the social-democrats loose. It is widely viewed, that this loss is mainly due to the unpopular social security contributions policy proposal; d) February 2007: A coalition government of Christian-democrats and social-democrats is formed and a compromise to increase tax rates for rich retired individuals is revealed; e) June 2008: A special government-appointed committee ('committee Bakker') of prominent social- and Christian-democrats that was asked to think about labor force participation in the long run proposes to e.g. increase the social security eligibility age by two years from 65 to 67 years. The government officially rejects this idea immediately; f) November 2008: Parliament accepts a new taxation law that introduces marginally higher tax-rates for those over 64 and also a tax credit for those between 61 and 65 who are still employed.

Apparently no changes to the old age social security eligibility age were introduced during the discussed period. The contribution scheme was adapted at the very end of the observation window however. The latter could have had spillover effects (positive or negative) to the eligibility expectations. Note that from this short description of events, it becomes clear that for more than a year, between March 2007 and May 2008, the discussion about the future of Dutch old age social security disappeared from the political scene as other matters were more pressing (such as employment protection). This period will be very useful in identifying the variation in expectations that can be attributed to noise.

The quantity of newspaper articles that matters for the publicity reaction in expectations is actually the number of newspaper articles that have appeared since the last time someone participated in the survey, as those will be the potential inputs for changes in expectations. Remember that although the Pensionbarometer is surveyed each month, a single respondent only participates in it every third month. The relevant quantity here is thus the cumulative number of newspaper articles that have been published over these three months since the respondent last answered an expectation question.

To estimate the Publicity Reaction Coefficient my survey span has to include periods in which hardly any newspaper articles appeared ($\theta_t \approx 0$). I will now introduce two different definitions of low-publicity periods, based on the four different series in Figure 2.3. A period is qualified as a standard low-publicity period when the cumulative quantity of articles for all four publicity series was below the median of that variable. The bar plots in Figure 2.4 show these cumulative series per survey week in comparison to their respective medians. The publicity-low periods (L) are those in between the dotted lines. In total eight periods can be classified as information-low using this definition. In the same figure, very-high publicity periods (HH) are identified as those periods before which there

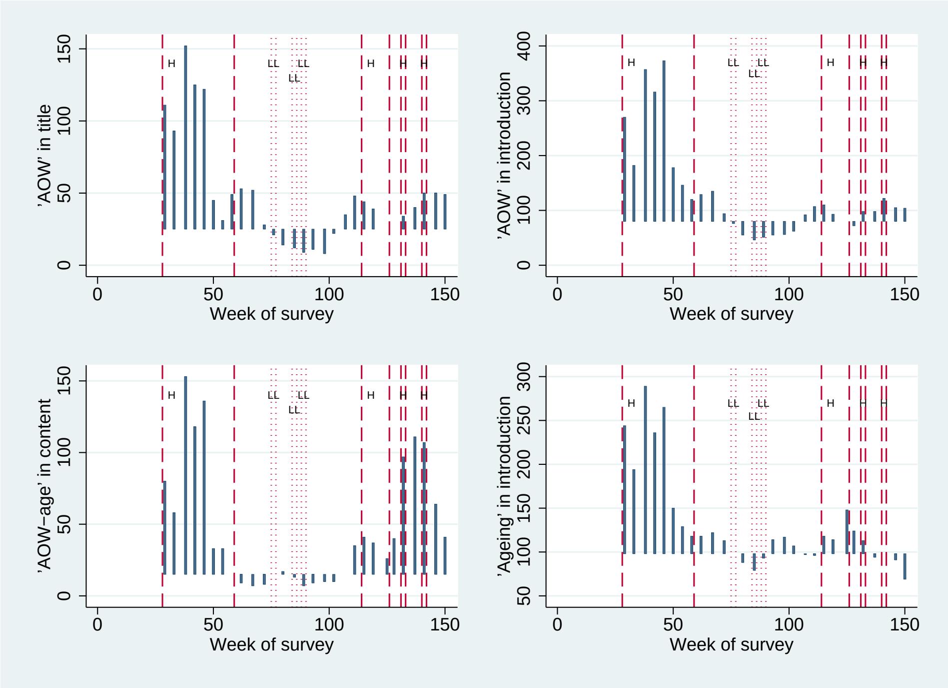
Figure 2.4: Cumulative newspaper articles three months before each survey, relative to *median* per variable. Low publicity periods between dotted lines (L). Very-high publicity periods between dashed lines (HH).



was more than the median level of publicity in all four series. The dataset contains seven of these periods. Table 2.A.1 in the appendix provides a list of all periods and their categories.

For the purpose of sensitivity analysis in Section 2.6 a second definition of publicity-poor periods is introduced, based on the first quartile of the article quantities rather than on the median. Figure 2.5 shows the cumulative number of articles compared to their respective first quartiles. A period is qualified as a very-low publicity period (LL) when the quantity of articles for all four publicity series was below the first quartile of that variable. The very-low publicity periods are those in between the dotted lines in Figure 2.5. Only three periods can be classified as information-low using this definition. In the same figure, high publicity periods (H) are identified as those periods before which - for all four series - more newspaper articles appeared than their respective first quartile. There are thirteen of such high-publicity periods.

Figure 2.5: Cumulative newspaper articles three months before each survey, relative to *first quartile* per variable. Very-low publicity periods between dotted lines (LL). High publicity periods between dashed lines (H).



2.6 Estimation results

2.6.1 Reduced form analysis

This subsection presents reduced-form results of the extent to which the variance of expectation changes and publicity intensity are related. This is done to investigate whether more publicity before a survey indeed leads to larger changes in expectations. Table 2.5 presents the results of OLS regressions of the logarithm of squared changes in the higher eligibility age expectations on publicity intensity and several other covariates that control for individuals' propensity to change their expectations. In Table 2.5 each newspaper articles coefficient was estimated in a separate model, on both the entire sample (column I) or the ten years (column II) and twenty years horizon (column III) subsamples.

The results suggest that higher publicity levels are significantly associated with changes in eligibility age expectations. The effect is non-negligible. One hundred additional newspaper articles that feature the

Table 2.5: Selected coefficients and standard errors of separate OLS models. Dependent variable: logarithm of squared expectation changes in higher eligibility age expectations. Newspaper article series in units of 100.

	(I) All	(II) Ten years	(III) Twenty years
Articles 'AOW'	0.252*	0.255	0.232
in title	(0.114)	(0.155)	(0.171)
Articles 'AOW'	0.115*	0.124	0.099
in introduction	(0.048)	(0.067)	(0.073)
Articles 'AOW-age'	0.006	-0.054	0.064
in total content	(0.093)	(0.129)	(0.141)
Articles 'Ageing'	0.231**	0.227*	0.226
in introduction	(0.076)	(0.105)	(0.113)
Observations	5,275	2,714	2,561

Standard errors in parentheses are clustered at the individual level

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Other covariates: education level, age group, income group, gender and partner

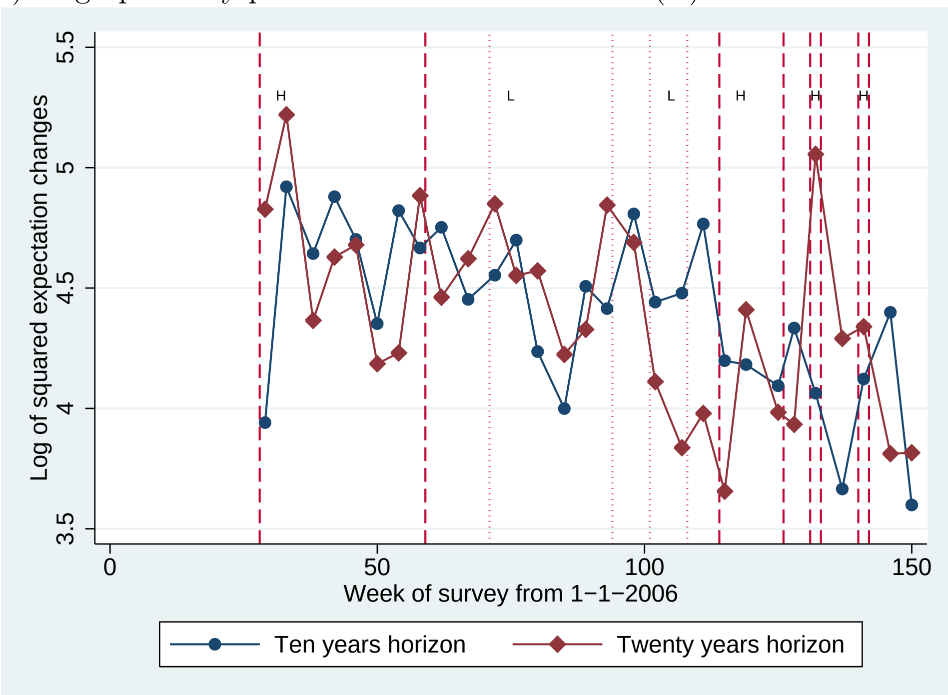
Dutch acronym for old age social security ('AOW') in their title increase changes in expectations by 25 percent. Not only quantity but also the quality of news seems to matter; newspaper articles mentioning 'AOW' in their title have an effect on expectation changes that is more than twice as large as the effect of those mentioning 'AOW' in the introduction. An additional probit-analysis of whether respondents changed their expectations (e.g. with at least 6 probability points) more often in periods of high publicity intensity showed no significant publicity coefficients. This indicates that the results in Table 2.5 are driven by larger expectation changes rather than by more respondents adapting their expectation answers.

2.6.2 Publicity Reaction Coefficient

In this section I estimate the publicity reaction in old age social security expectations using the Dutch Pensionbarometer data. To obtain a first impression, Figure 2.6 depicts average squared expectation changes per survey week for the two different time horizons. As in Figures 2.4 and 2.5, the periods between the dotted lines are classified as low publicity periods (L) and the periods between the dashed lines as high publicity periods (H). Although the extent to which respondents adapted their expectations does seem somewhat lower in the information-poor periods the difference is not overwhelming. This indicates that random shifts in expectation reports account for a substantial proportion of the variance in expectation changes; something that will be supported by the PRC evidence below.

Table 2.6 presents the estimates of the Publicity Reaction Coefficients for the total sample over the two different time horizons and for the four different expectation questions. To remind the reader: the Publicity Reaction Coefficient is defined as the ratio of the estimated systematic variance over the total variance of expectation changes. The estimated systematic variance equals the variance of expectation changes in high publicity periods (H) minus this variance in low publicity periods (L).

Figure 2.6: Average squared change in probability assigned to later eligibility age per survey week. Low publicity periods between dotted lines (L). High publicity periods between dashed lines (H).



$$PRC = \frac{E[(\Delta_{i,t^H}\tilde{y}^{t+z})^2] - E[(\Delta_{i,t^L}\tilde{y}^{t+z})^2]}{E[(\Delta_{i,t^H}\tilde{y}^{t+z})^2]}$$

The PRCs in the first column are estimated exploiting expectation changes in the high publicity and low publicity periods. To check the robustness of these results, the PRC estimates in the second column only use observations in very-high and very-low publicity periods. How high and low and very-high and very-low publicity periods are exactly defined is explained in Section 2.5. In Table 2.A.2 in the appendix one finds the number of observations used per estimate divided into those from publicity-poor and those from publicity-rich periods. Throughout this section, PRC estimates are depicted in normal text when the involved error variance estimate is based on more than 50 observations. Italic text displays Publicity Reaction Coefficients that used error variance estimations based on in between 30 and 50 observations. Finally, a PRC is not displayed whenever less than 30 observations could be used to estimate the associated error variance.

All but three out of sixteen PRC estimates in Table 2.6 are positive, indicating that expectations do change more in periods with a larger publicity set. The publicity reaction is typically larger when the considered time horizon is shorter. Moreover, the probabilities that individuals assign to a lower eligibility age are more sensitive to publicity than the probabilities they assign to a higher eligibility age. For all subsamples however at least seventy percent of the variance in expectation changes was due to random shifts in expectations (as the highest significant PRC-estimate equals 30.5%). This is emphasized by the fact that only three PRCs are significantly different from zero.

Table 2.6: Publicity Reaction Coefficients for the total expectation changes sample

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Prob. for higher age	6.8%	1.8%	20.3%*	-1.7%
	(7.6)	(9.0)	(9.5)	(13.4)
Prob. for two year	3.4%	-2.2%	16.1%*	-4.5%
higher age	(7.7)	(8.2)	(9.5)	(12.9)
Prob. for lower age	13.5%	19.5%	20.5%	9.0%
	(16.4)	(15.6)	(21.1)	(28.2)
Prob. for two year	10.9%	30.5%*	18.8%	26.4%
lower age	(21.6)	(16.0)	(28.2)	(25.6)

Standard errors are bootstrapped and in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

The subsequent tables report Publicity Reaction Coefficients that

Table 2.7: PRCs for the probability of a higher eligibility age for five different age groups

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Age 25-34	42.0%*	-9.0%	*22.4%*	*-73.7%*
	(16.2)	(29.2)	(33.5)	(60.6)
Age 35-44	-24.9%	-11.2%	*-9.6%*	*-92.5%*
	(22.4)	(31.9)	(31.0)	(81.2)
Age 45-54	15.6%	20.1%	33.6%*	39.9%**
	(14.5)	(14.4)	(15.1)	(15.8)
Age 55-64	21.7%	-6.5%	30.9%*	-9.2%
	(13.3)	(17.9)	(15.7)	(26.9)
Age >64	-16.1%	-8.1%	4.3%	2.1%
	(19.0)	(19.4)	(23.5)	(32.6)

Number in italics: error variance based on 30-50 obs
Standard errors are bootstrapped and in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

have been estimated for different partitions of the total sample. To be relatively concise, results are only shown for the higher eligibility age expectations. It should be noted that the subgroup PRC estimates are conditional on the partition in the specified table only and not on other characteristics. This is because there are only a limited number of expectation changes in publicity-poor periods available to estimate the necessary error variance. Estimating the Publicity Reaction Coefficient for smaller subgroups would render these error variance estimates unreliable. The number of observations for each estimate in the subsequent tables can be found in Tables 2.A.3, 2.A.4, 2.A.5, 2.A.6 and 2.A.7 in the appendix.

Table 2.7 reports Publicity Reaction Coefficients for different age groups. Although it is hard to distinguish a clear pattern in the relationship between age and the reaction to publicity, the middle age category (between 45 and 54) displays the largest publicity reaction in three out of four columns. Moreover, the PRC of the middle aged is relatively large and significant when only the very-low and very-high publicity periods are taken into account. Interestingly, the subgroup with the lowest - even negative - publicity reaction are those individuals over the age of 64 that are already receiving old age social security. It can be argued that once

Table 2.8: PRCs for the probability of a higher eligibility age for five different education groups

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Primary	*-49.1%*	*27.9%*	.%	.%
	(60.8)	(26.4)	(.)	(.)
Vocational secondary	8.7%	0.3%	23.5%	6.6%
	(13.0)	(16.3)	(15.3)	(23.7)
General secondary	24.1%	-3.1%	*40.9%**	*-2.3%*
	(19.9)	(24.9)	(19.9)	(42.8)
Vocational tertiary	16.2%	-1.2%	*5.5%*	-2.4%
	(16.9)	(22.9)	(30.2)	(34.5)
Higher vocational	6.3%	2.2%	10.3%	-4.9%
	(15.5)	(19.0)	(21.4)	(30.2)
Academic education	-25.8%	-30.7%	*6.4%*	*-58.3%*
	(31.7)	(35.7)	(32.9)	(68.8)

Number in italics: error variance based on 30-50 obs

No number: error variance based on <30 obs

Standard errors are bootstrapped and in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

someone receives a benefit she is no longer interested in news about the eligibility age as changes in this policy will not apply to her anymore. The same mechanism could explain that those in between 55 and 64 display a significant publicity reaction for a change within ten years (column II) but not for a change within twenty years.

Table 2.8 reports Publicity Reaction Coefficients for education subsamples. There does not appear to be a linear relationship between level of education and the degree to which a respondent reacts to publicity. Hardly any of the PRCs are significantly different from zero. What is striking however is that those respondents who managed to obtain a university degree do not respond to publicity at all; their negative PRCs even indicate that they change their expectations more in publicity-poor than in publicity-rich periods. Perhaps high educated individuals have such high-quality initial expectations that information disseminated through the media does not induce them to change their policy forecasts.

The sample is stratified in five income classes for the estimation of

Table 2.9: PRCs for the probability of a higher eligibility age for five different income groups

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Low income	1.0%	27.6%*	25.8%	*-0.4%*
	(17.5)	(16.6)	(17.7)	(37.1)
Lower middle	19.2%	-11.8%	29.7%*	*-4.5%*
income	(15.6)	(20.1)	(19.4)	(33.0)
Middle income	-9.0%	3.3%	-32.0%	-0.1%
	(18.4)	(19.8)	(30.4)	(29.8)
High middle	33.7%*	-5.7%	37.9%*	-1.3%
income	(11.7)	(23.5)	(17.1)	(33.6)
High income	-2.7%	-17.8%	10.3%	-23.5%
	(19.2)	(22.6)	(27.7)	(35.5)

Number in italics: error variance based on 30-50 obs
Standard errors are bootstrapped and in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

the Publicity Reaction Coefficients in Table 2.9. The ranking of income subgroups in terms of PRC estimates is rather different over the four columns. However, the high middle income category displayed the highest publicity reaction in both ten years-columns, the middle income category displayed the lowest publicity reaction in these columns and the highest income category displayed the lowest publicity reaction in both twenty years-columns. These differences between adjacent groups are surprising and suggest that several opposing factors influence the relationship between income and the publicity reaction. A possible explanation for the low PRCs of the highest income group could be that the future of old age social security does not interest them as they foresee enough alternative income sources.

Table 2.10 reports Publicity Reaction Coefficients for media consumption and self-reported political awareness subgroups. While focusing on newspaper consumption, what catches the attention is that those who read the newspaper once a week or less have the highest publicity reaction. This is the case for all estimates, except for the ten years-horizon in the publicity-poor versus the publicity-rich comparison. A similar picture emerges from the results on individuals' knowledge of current affairs in

Table 2.10: PRCs for different media consumption groups

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Average newspaper consumption	-9.9%	3.1%	14.7%	-6.5%
	(11.3)	(12.0)	(14.2)	(19.8)
Low newspaper consumption	30.9%	-4.3%	*40.8%*	*9.9%*
	(27.0)	(26.4)	(28.0)	(31.5)
Average knowledge of current affairs	-5.1%	-4.0%	22.4%*	-12.3%
	(12.1)	(12.1)	(13.6)	(19.3)
Low knowledge of current affairs	14.3%	31.7%	*-0.6%*	*34.1%*
	(21.4)	(23.8)	(32.5)	(39.0)
Average television news consumption	-2.3%	2.1%	20.7%	-6.6%
	(11.3)	(11.9)	(13.2)	(18.8)
Low television news consumption	*13.7%*	*2.3%*	.%	.%
	(26.2)	(34.9)	(.)	(.)

Number in italics: error variance based on 30-50 obs
No number: error variance based on <30 obs
Standard errors are bootstrapped and in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Dutch politics. Individuals who themselves report that their knowledge is lower than that of the average Dutchman, display higher Publicity Reaction Coefficients. This thus means that ill-informed individuals do pick up on some of the released publicity on the relevant topic and change their expectations accordingly. At first, these results seem counterintuitive. If low media consumption is however correlated with low initial knowledge of old age social security, one can imagine that publicity should induce more changes in the expectations of the low informed group. The television news results are in the same direction, but very imprecise as the number of respondents who report low television news consumption is low.

Finally, Table 2.11 shows PRC estimates for a partition of the dataset into what type of newspaper one most often reads. Four specific categories are selected: those who obtain most information from free newspapers, those who most often read ‘de Telegraaf’ - a populist/ conservative newspaper, those who most often read ‘de Volkskrant’ - a progressive newspaper and those who most often read ‘NRC Handelsblad’ - a liberal

Table 2.11: PRCs for the probability of a higher eligibility age for five different newspaper readership groups

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Free newspapers	19.3%	20.3%	.%	.%
	(25.7)	(26.9)	(.)	(.)
De Telegraaf	*11.8%*	*33.4%*	.%	.%
(Populist)	(32.9)	(21.9)	(.)	(.)
Volkskrant	*0.9%*	-55.5%	.%	.%
(Progressive)	(35.8)	(60.2)	(.)	(.)
NRC Handelsblad	*13.0%*	*-34.5%*	.%	.%
(Liberal)	(31.7)	(62.5)	(.)	(.)
Other	-20.6%	-7.5%	-13.7%	13.0%
	(15.6)	(15.1)	(21.7)	(20.3)

Number in italics: error variance based on 30-50 obs
No number: error variance based on <30 obs
Standard errors are bootstrapped and in parentheses
* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

newspaper. It turns out that those who only read free newspapers and those who often read the populist newspaper have the highest, although insignificant, publicity reactions. This could be because of the specific wording in and content of the articles, but also because the individuals that typically read these newspapers have low-quality initial expectations. Those who read the progressive newspaper do not seem very influenceable by the media in terms of their expectations. Respondents who are in the ‘other newspaper’ category, which includes regional and smaller national newspapers, display the smallest publicity reaction.

All in all, the presented Publicity Reaction Coefficients for the different demographic and media consumption subgroups paint an interesting picture. Three groups are identified that do not adapt their expectations more strongly in high publicity-periods: those over the age of 64, those with a university degree and the highest income group. The middle aged respond relatively heavily to publicity in terms of their expectations, just like those who do not often read a newspaper and those who read free newspapers or ‘de Telegraaf’. Also the respondents that indicate themselves that they know little about Dutch politics display a higher publicity reaction. Since the PRC estimates in this section are only conditional on the characteristic presented in each table, it it unclear what the most important determinants of differential publicity reaction are. More research will be needed to uncover this.

Two explanations of these results come to mind that probably both explain some part of reality. Both involve the publicity set Θ_t that individuals use as an input in their expectations formation process. First, generally less-informed groups are more susceptible to repetitions of old news in the media, Φ_t. The quality of the initial expectations of those with a university degree and the highest income groups is probably higher than the quality of expectations of, say, those who do not often read a newspaper. If the publicity set contains mostly 'old' news those who did not pay enough attention to the news before, should respond to these publicity items by shifting their expectations up or down, depending on which old news elements they absorbed over the last period. Second, as mentioned before, some groups have lower stakes in the future of the old age social security eligibility age so that they will hardly pay attention to the relevant publicity set Θ_t. Those already receiving a benefit and those who are financially well-off probably do not suffer much when the eligibility age goes up and hence will ignore news on the matter.

2.6.3 Variance in expectation change error terms explained

A side product of the estimation of the Publicity Reaction Coefficients is that one can analyze individuals' reporting errors by investigating their expectation changes in the publicity-poor periods. Table 2.12 presents estimation results for regressing the log of squared expectation changes in the low publicity periods as defined in Section 2.5 on several demographic characteristics. Changes in expectations are assumed to be drawings from the reporting error variance here and the presented regressions thus capture the factors that influence heteroscedasticity. The regression coefficients will explain which groups are more likely to provide erroneous expectation answers.

The age group dummies turn out to be among the few significant covariates in explaining how error-prone individuals are. Typically, the younger the respondent the lower the error variance is. This means that especially those over 64 are more inclined to randomly shift their expectations up and down, even if no publicity forces them to do so. In the twenty years-horizon regression, the group of people between 35 and 44

Table 2.12: Coefficients and standard errors of OLS models. Dependent variable: logarithm of squared expectation changes in publicity-poor periods.

	(I) 10 years		(II) 20 years	
	Coef.	S.e.	Coef.	S.e.
Age 25-34	-1.530***	(0.390)	-0.217	(0.420)
Age 35-44	-0.492	(0.364)	-0.701*	(0.409)
Age 45-54	-0.718*	(0.331)	-0.435	(0.368)
Age 55-64	-0.706*	(0.337)	0.300	(0.364)
Vocational secondary	-0.248	(0.503)	-0.362	(0.528)
General secondary	0.210	(0.535)	-0.687	(0.620)
Vocational tertiary	-0.464	(0.532)	-0.627	(0.555)
Higher vocational	-0.208	(0.521)	-0.841	(0.562)
Academic education	-0.184	(0.609)	-0.291	(0.638)
Low middle income	0.001	(0.365)	0.956**	(0.412)
Middle income	0.106	(0.376)	0.929*	(0.431)
High middle income	-0.329	(0.377)	0.517	(0.464)
High income	-0.400	(0.446)	0.549	(0.552)
Dummy female	-0.095	(0.259)	0.228	(0.292)
Dummy partner	0.176	(0.273)	0.069	(0.266)
Observations	740		741	
Individuals	554		559	
R^2	3.4%		3.1%	

Standard errors are corrected for clustering at the individual level

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Base categories: primary education, age >64, low income years old is the only that is significantly less error-prone than the elderly. Income also matters significantly in explaining mistakes in the second column. Those with lower middle incomes and middle incomes have higher estimated error variances. In general, the explanatory power of these regressions is low, indicating that there is large individual heterogeneity in the prediction error variances.

2.7 Conclusion

As empirical evidence is accumulating that subjective expectations influence economic decision-making and that these expectations are sometimes biased, it becomes policy-relevant to know how to influence individuals' expectations. It is likely in this respect that information in the media affects how people feel about the future. This chapter analyzes the role of public information dissemination, or publicity, in a real-life expectations formation process. More specifically: I estimate the degree to which different demographic and media consumption groups adapt their expectations to publicity about the relevant random variable.

To examine demographic differences in the publicity reaction, one needs data of an expectation on which private information has no influence in order to circumvent identification problems. I therefore use the Pensionbarometer, a unique monthly dataset of expectations on the Dutch old age social security eligibility age. As the relevant media attention differed substantially over the survey period, I am able to estimate the proportion of the variance in the eligibility age expectation changes that can be attributed to publicity (the Publicity Reaction Coefficient - PRC) without making strong assumptions on the expectations formation process.

My findings are in line with other empirical papers that suggest media intensity affects the precision and updating of inflation expectations (Carroll, 2003, and Lamla and Lein, 2008). For the total population, the publicity reaction in eligibility age expectations is estimated to be in between zero and thirty percent. Rather than publicity, random shifts thus explain the majority of the variance in changes in eligibility age expectations, with especially older individuals being more imprecise in their expectation reports. The differences in the publicity reaction among subgroups are considerable however. The middle-aged, those who infrequently read newspaper and those who know little about Dutch politics are found to have a relatively high publicity reaction. This can be understood if the majority of newspaper articles about old age social security contains ‘old news’ rather than real information revelations. It is after all likely that those with lower quality initial expectations adapt their expectations more whenever already available information is repeated in the media. Moreover, I find that those over the age of 64 and high income individuals do not adapt their expectations more often in publicity-high periods. It seems as if those with low stakes in the future of the old age social security eligibility age pay little attention to publicity on the matter.

Provided that the explanations that I propose for my PRC results are valid, policy-makers could learn the following about publicity and expectations formation. On the one hand, those who have low quality initial expectations are easily influenced by whatever appears on the subject in newspapers and on television. If policy-makers would like to influence the expectations of the weaker informed group it would be wise to often repeat high-quality information in the media so that eventually even this group hears the news. On the other hand, those who feel they have little to do with the expectation at hand are hard to influence by public information dissemination, as they will choose to ignore the news. This is fine when the disinterest is justified, but rather problematic when it is not.

2.A Appendix

Figure 2.A.1: Histogram of probabilities assigned to higher eligibility age within twenty years.

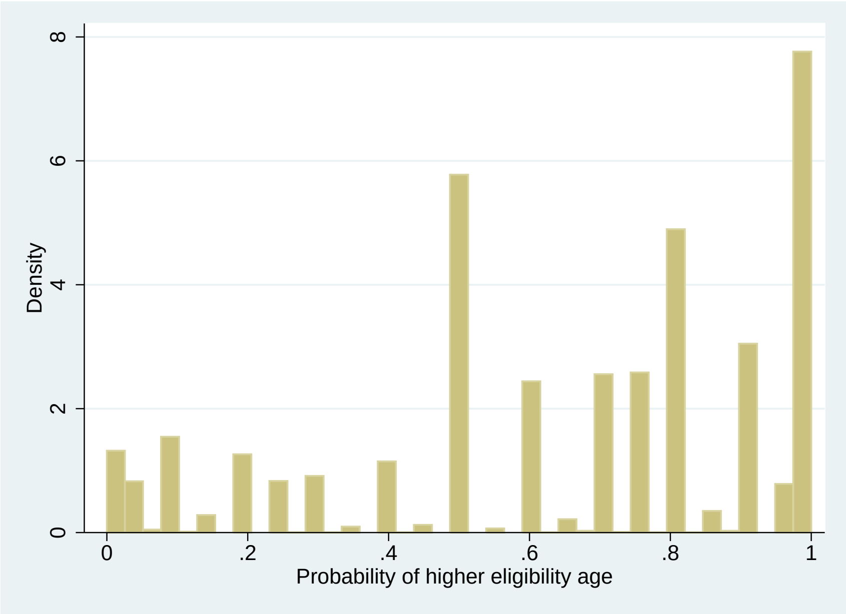


Figure 2.A.2: Interquartile range for probability assigned to higher eligibility age within twenty years (week 1-52: 2006, week 53-75: 2007, week 105-157: 2008).

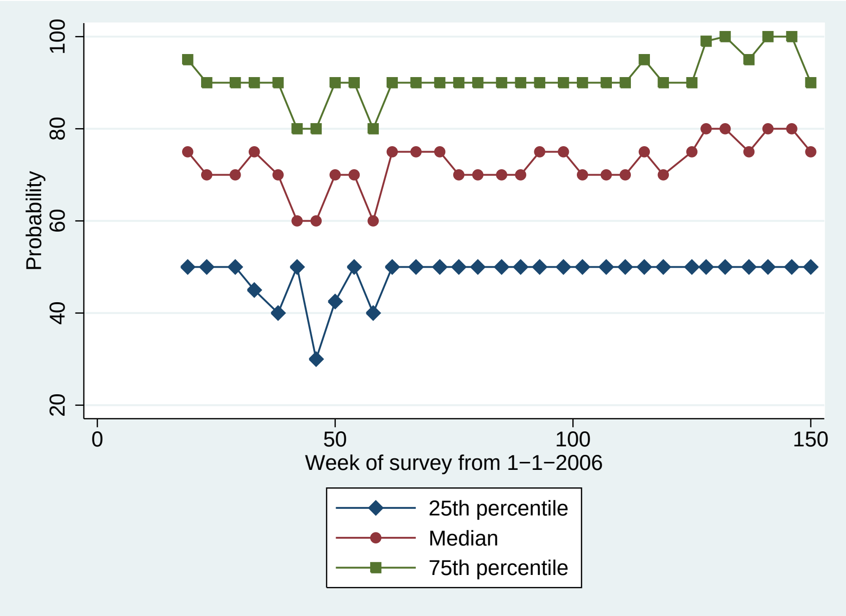


Table 2.A.1: Overview of classification of survey weeks

Week	Year	End survey week	Start media count	Low publicity	Very low publicity	High publicity	Very high publicity
19	2006	13-mei		No	No	No	No
23	2006	10-jun		No	No	No	No
29	2006	22-jul	2-apr	No	No	Yes	Yes
33	2006	19-aug	14-mei	No	No	Yes	Yes
38	2006	23-sep	11-jun	No	No	Yes	Yes
42	2006	21-okt	23-jul	No	No	Yes	Yes
46	2006	18-nov	20-aug	No	No	Yes	Yes
50	2006	16-dec	24-sep	No	No	Yes	Yes
54	2007	13-jan	22-okt	No	No	Yes	No
58	2007	10-feb	19-nov	No	No	Yes	No
62	2007	9-mrt	17-dec	No	No	No	No
67	2007	13-apr	14-jan	No	No	No	No
72	2007	18-mei	11-feb	Yes	No	No	No
76	2007	15-jun	10-mrt	Yes	Yes	No	No
80	2007	13-jul	14-apr	Yes	No	No	No
85	2007	17-aug	19-mei	Yes	Yes	No	No
89	2007	14-sep	16-jun	Yes	Yes	No	No
93	2007	12-okt	14-jul	Yes	No	No	No
98	2007	16-nov	18-aug	No	No	No	No
102	2007	14-dec	15-sep	Yes	No	No	No
107	2008	19-jan	13-okt	Yes	No	No	No
111	2008	16-feb	17-nov	No	No	No	No
115	2008	15-mrt	15-dec	No	No	Yes	Yes
119	2008	12-apr	20-jan	No	No	Yes	No
125	2008	24-mei	17-feb	No	No	Yes	No
128	2008	14-jun	16-mrt	No	No	No	No
132	2008	12-jul	13-apr	No	No	No	No
137	2008	16-aug	25-mei	No	No	No	No
141	2008	13-sep	15-jun	No	No	No	No
146	2008	18-okt	13-jul	No	No	No	No
150	2008	15-nov	17-aug	No	No	No	No

Table 2.A.2: Number of observations in estimation of Publicity Reaction Coefficients for the total expectation changes sample

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Prob. for higher age	1,233	1,155	685	633
	(744)	(742)	(302)	(273)
Prob. for two year	1,232	1,154	684	632
higher age	(744)	(736)	(302)	(272)
Prob. for lower age	1,232	1,152	684	631
	(742)	(734)	(301)	(271)
Prob. for two year	1,231	1,152	683	631
lower age	(741)	(733)	(301)	(270)

Observations in publicity-poor periods in parentheses

Table 2.A.3: Number of observations in estimation of PRCs for the probability of a higher eligibility age for five different age groups

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Age 25-34	166	150	95	83
	(103)	(97)	(48)	(34)
Age 35-44	231	186	133	98
	(116)	(117)	(38)	(38)
Age 45-54	304	268	166	155
	(177)	(198)	(81)	(86)
Age 55-64	279	281	146	150
	(185)	(173)	(71)	(59)
Age >64	253	270	145	147
	(163)	(157)	(64)	(56)

Observations in publicity-poor periods in parentheses

Table 2.A.4: Number of observations in estimation of PRCs for the probability of a higher eligibility age for five different education groups

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Primary	53	68	31	37
	(43)	(47)	(18)	(19)
Vocational secondary	341	312	196	174
	(232)	(207)	(89)	(78)
General secondary	127	151	71	84
	(83)	(85)	(40)	(22)
Vocational tertiary	258	233	136	135
	(146)	(143)	(48)	(51)
Higher vocational	299	264	171	139
	(156)	(180)	(77)	(71)
Academic education	155	127	80	64
	(84)	(80)	(30)	(32)

Observations in publicity-poor periods in parentheses

Table 2.A.5: Number of observations in estimation of PRCs for the probability of a higher eligibility age for five different income groups

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Low income	252	204	149	118
	(150)	(124)	(61)	(43)
Lower middle	223	255	138	131
income	(151)	(142)	(66)	(44)
Middle income	244	207	133	119
	(159)	(178)	(56)	(62)
Higher middle	256	224	134	125
income	(130)	(161)	(62)	(68)
High income	244	256	127	136
	(150)	(136)	(57)	(56)

Observations in publicity-poor periods in parentheses

Table 2.A.6: Number of observations in estimation of PRCs for the probability of a higher eligibility age for newspaper consumption and self-reported political awareness

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Average newspaper consumption	676 (397)	627 (421)	355 (153)	313 (147)
Low newspaper consumption	119 (55)	128 (78)	58 (18)	67 (31)
Average television news consumption	712 (408)	681 (450)	375 (151)	340 (162)
Low television news consumption	83 (44)	74 (49)	38 (20)	40 (16)
Average informedness	662 (384)	640 (425)	340 (145)	328 (151)
Low informedness	133 (68)	115 (74)	73 (26)	52 (27)

Observations in publicity-poor periods in parentheses

Table 2.A.7: Number of observations in estimation of PRCs for the probability of a higher eligibility age for five different newspaper readership groups

	(I) Publicity-poor vs. publicity-rich		(II) Publicity-very-poor vs. publicity-very-rich	
	10 years	20 years	10 years	20 years
Free newspapers	80 (55)	91 (56)	49 (20)	53 (28)
De Telegraaf (Populist)	89 (47)	86 (49)	40 (21)	48 (15)
De Volkskrant (Progressive)	78 (46)	65 (59)	37 (15)	28 (24)
NRC Handelsblad (Liberal)	50 (40)	57 (31)	28 (18)	26 (7)
Other	459 (241)	405 (280)	243 (90)	196 (94)

Observations in publicity-poor periods in parentheses

Chapter 3

Preparing for Policy Changes: Social Security Expectations and Pension Scheme Participation[1]

3.1 Introduction

After several industrialized countries had installed old age social security in the 1940s and 1950s, economists wondered what the effect of these new institutions would be on private savings. A large literature developed that was interested in to what extent public pension schemes crowded out private pension savings. More than fifty years after most Western governments introduced Pay-As-You-Go (PAYG) pension schemes, governments no longer discuss the adornment but rather the demolition of these systems. Particularly the ageing of populations has recently put pressure on policy-makers to cut down on the generosity of their public pension schemes, creating policy uncertainty in first pillar pensions. Therefore, it is now relevant to investigate the response of private savings to a (possible) dismantlement of old age social security.

This chapter tries to answer this question focusing on potential future changes rather than on changes that have already been implemented. Do people sufficiently prepare for the policy changes in PAYG pensions that

[1]Thanks to Arthur van Soest, Frederic Vermeulen, James Banks, Rob Alessie, Norma Coe, Katie Carman, Gema Zamarro and several seminar participants for useful comments and suggestions.

they anticipate? In addition to discovering more about the relationship between public and private pension schemes, this chapter thus also analyzes individuals' capacity to assess how likely policy alterations are and on their capacity to respond accordingly. Knowing whether their citizens are indeed capable of mitigating the negative effects of these policy changes is essential for all governments, and in particular for those that are currently deciding upon policy changes in their old age social security system. The empirical association between subjective expectations regarding two different policy changes and private pension participation is examined for The Netherlands. Though nothing can be concluded on causality, those who are more convinced of future deteriorations are indeed more likely to participate in a private pension fund. These findings suggest that positive probabilities of a deterioration in social security 'crowd in' private pension savings. I define crowding in here as an increase in private savings, in analogy to crowding out which in this context refers to a decrease in private savings due to the installment of old age social security.

In this introduction, some of the relevant literature is discussed. Furthermore, important features of the Dutch pension system are explained. Section 3.2 then sets out a theoretical model of optimal savings under uncertainty over the future of the old age social security system. Subsequently, Section 3.3 explains the estimation strategy and includes a description of the data used in the empirical analysis. Section 3.4 presents the estimation results. Finally, Section 3.5 concludes and gives some policy advice.

3.1.1 The crowding-in of private savings

A relationship between savings and social security was first empirically revealed by Feldstein in 1974. He used aggregate time series to show that social security wealth had halved the rate of private savings in the US. Feldstein argued that consumers had lowered savings because they would be receiving a certain income from PAYG-benefits in retirement. This 'crowding out' effect of old age social security has since then been addressed by many scholars using both macro- and micro-data sources. Results based on individual information provide weak to strong evidence of crowding out effects in various industrialized countries. See Kotlikoff (1979) for the U.S., Jappelli (1995) and Attanasio and Brugiavini (2003) for Italy, Guariglia and Markose (2000) and Attanasio and Rohwedder (2003) for the U.K. and Alessie, Kapteyn and Klijn (1997), Euwals (2000) and Kapteyn, Alessie and Lusardi (2003) for The Netherlands.

In 1974, when Feldstein published his paper, the outlook on the future of social security was very optimistic, as the following quote indicates.

> *"The history of social security shows continually rising benefit levels, a fact that individuals no doubt perceive when they contemplate the order of magnitude of their benefits at retirement age."* (p. 911)

Since the publication of Feldstein's article, increased awareness of ageing altered this outlook on publicly provided old age pensions in most Western countries. An illustration of this can be found in Dominitz, Manski and Heinz (2002 and 2003), who have described expected eligibility for U.S. old age social security using data from the Survey of Economic Expectations (SEE) for 1999 to 2002. They find that a substantial number of respondents believe that the U.S. social security system will collapse entirely within the foreseeable future. In this chapter I show that many Dutch individuals are also pessimistic about the future of Dutch old age social security.

Taking these expectations into account, this chapter investigates the crowding in effect of uncertainty in social security on private savings. In analogy to Martin Feldstein's crowding out effect, a crowding in effect refers to increases in private wealth because of a possible dismantlement of the social security system. To my knowledge, the effects of a potential degeneration of the social security system on wealth accumulation have not been assessed so far.

3.1.2 Expectations measured in subjective probabilities

In this study the response to expectations is analyzed by exploiting variation in the subjective probabilities individuals assign to policy changes. The analysis of explicit subjective expectations is a popular new field of research in economics. Manski (2004) provides an overview of the use of subjective probabilities in modern economics. Individuals seem surprisingly able to translate their knowledge on the idiosyncratic risk they face into a probability that is equal to the ex-post realized probability. For example, Hurd and McGarry (2002) find that elderly who die within two years of participating in the Health and Retirement Study (HRS) had reported a much smaller probability of surviving until the age of 75 than those who actually survived. Subjective probabilities concerning the future occurrence of a change in the social security system will be the independent variables of interest here.

The relationship between subjective expectations and individual savings behavior is analyzed by e.g. Guiso, Jappelli and Terlizzese (1992) and Stephens (2004). Guiso et al. use a subjective measure of income uncertainty in order to quantify the size of precautionary savings. Stephens examines whether job loss expectations affect savings behavior and does not find any effect. He argues this could be because of probability weighting in a loss aversion context (see Bowman et al., 1999) so that individuals with very small job loss probabilities actually over-weight these.

3.1.3 Old age social security in The Netherlands

The Netherlands has a hybrid pension system. All three traditional pillars - the mandatory public pillar, the mandatory occupational pillar and the voluntary private pillar - are of considerable importance in providing an income for the elderly when compared to other countries (see Boersch-Supan, 2004 for more information). The first pillar has a Pay-As-You-Go character, while the second and third pillars are capital funded.

Most employees are obliged to participate in the second pillar, the occupational pension schemes. Annuitization of their assets is compulsory and there is little freedom over the investment choices. The most common formal third pillar pension products are the so-called 'lijfrente-polissen' (translated as life-interest contracts) and 'koopsom-polissen' (translated as lumpsum life-interest contracts). Consumers can buy these annuities from banks and insurers either through regular contributions (life-interest) or through the deposit of a lump-sum (lumpsum life-interest). Contributions to second and third pillar pensions schemes are to a certain degree tax-deductible. Private wealth that is accumulated outside of these designated pension schemes could be called the informal third pillar.

The old age social security arrangement in The Netherlands, the *Algemene Ouderdomswet* (AOW), was established in 1957 to provide a base pension for all inhabitants. Designed on Beveridge's principle, the eligibility for and level of AOW benefits does not depend on contributions paid, income or wealth, only on years of residence in The Netherlands. For each year one has legally resided in The Netherlands between the ages of 15 and 65, one receives two percent of the maximum AOW amount from the age of 65 onwards, irrespective of retirement. The real level of AOW benefits has remained stable since 1979, at approximately seventy percent of the minimum wage. The net level of a couple's combined AOW benefits equals the net minimum wage. In 2007, this came down to a gross monthly allowance of EUR 945 for singles and EUR 648 for pensioners with a partner[2].

Ageing as well as changing economic and political circumstances triggered discussions on the sustainability of PAYG-pensions from the late 1970s onwards. The historical Dutch policy debate is summarized by Van Eekelen (2006). Despite this debate the old age social security arrangement has remained roughly the same as it was when introduced in 1957.

The fact that the PAYG-system is so universal, i.e. each individual in principle receives the same benefits, provides a great research environment. The expectations individuals form over the future of the system thus all relate to the exact same thing and are not confounded by idiosyncratic risk factors. In contrast, the expectations regarding eligibility for U.S. social security reported in Dominitz et al. (2003) combine expectations about macro (such as policy uncertainty) and micro risks (such as unexpected income shocks).

3.2 Theoretical model

The life-cycle model developed here is a three-period model of consumption. All individuals live for three periods. Individuals just face two decisions in their life; what to save in the first period (S_i) and what to consume out of these savings in the second period (z_i). Labor supply and retirement are thus exogenous. This assumption is a simplification of reality, but endogenous retirement would not greatly change my results[3]. In the first period everybody is young and working. Workers receive individual-specific wages (I_i), which are larger than one for everyone. All individuals are retired in periods two and three. I assume a

[2]The system is financed through an earmarked payroll tax, which constitutes 17.9% on the first EUR 30,000 of gross income. Those over 64 do not pay this tax as they are already receiving benefits. In 1997 the AOW tax rate was legally capped so that the remainder of AOW expenditures are now financed through general taxes which are also paid by retirees with a substantial pension income. Whereas in 2001 still all expenditures were financed through the payroll tax, in 2006 this decreased to 70% of AOW expenditures.

[3]If labor supply is an additive function of the utility derived from potential labor income and the disutility experienced from working and if labor demand is a function of workers' productivity, the age at which people retire is independent of the - universal and uniform - PAYG benefit. However, if utility would have a non-additive functional form the retirement age could depend on when and how much benefits an individual receives in the future.

standard additive model as in Browning and Lusardi (1996) that allows for a precautionary savings motive, as this provides clear and plausible results. Lifetime utility therefore additive has a logarithmic functional form.

There exists a government that pays out a fixed old age social security benefit (P) to all citizens. The government raises both a general consumption tax τ and an additional consumption tax θ that equals the contribution to the old age social security system. I choose a consumption tax in order to mirror the tax-deductible nature of pension savings. Individuals that receive a pension benefit do not contribute to the pension system, although they do pay general taxes. Hence, there is a Pay-As-You-Go pension system in place. Note that the government can run a deficit as I assume there exists a financial fund that gathers the excess contributions or pays the excess benefits. For simplicity, it is impossible to leave a bequest in the model and the market interest rate and the discount rate offset each other. Like in the real world, individuals cannot borrow against future old age social security benefits so first period savings should be nonnegative ($S_i \geq 0$).

In the status quo, policy option A, the government pays out the PAYG benefit of level P to all citizens in both periods two and three. The utility functions for all three periods in the status quo are displayed in the equations below.

$$\begin{aligned} u_1^A &= ln((1-(\tau+\theta))(I_i - S_i)) \\ u_2^A &= ln((1-\tau)(z_i + P)) \\ u_3^A &= ln((1-\tau)(S_i - z_i + P)) \end{aligned}$$

Note that the lifetime utility maximization problem can be solved using backward induction; first maximizing over z_i and then over S_i. For expositional reasons I will therefore substitute z_i by the optimal consumption out of savings in the second period in the remainder of this section. One could think of the government's financial fund running out of money because of persistently shrinking generations. At first, the status quo might be maintained as governments can easily borrow on the credit market. After some time though, policy-makers could respond to the ageing of the population by adapting the PAYG pension scheme. The government would then choose one of the following alternative policy options.

- Option B: providing a lower benefit. Individuals will receive a benefit of $(1-\lambda)P$, for $0 < \lambda < 1$.

- Option C: only providing benefits in period three. The government will shift the eligibility age upwards: it will only provide benefits of the original level P in the third period, but not in the second.

Hence, individuals choose an optimal savings level (S_i) in the first period under *uncertainty* over the state of the world in the second and third period. The government decides on a policy option after the first period and this decision is carried out and publicly known in period two and period three. It seems realistic to assume that the government will not change any policy rules for those who are already retired (i.e. between period two and three).

In their first period maximization problem consumers then maximize utility over savings taking into account the probabilities they individually assign to the future occurrence of the different policy options. As each individual is one of many each individual is assumed to be a ‘system-taker’ (think of a price-taker). This means that the subjective probabilities a single individual assigns to different policy options do not influence the actual choice of the government.

$$\begin{aligned} P(\text{Option A: Status quo}) &= \alpha_i \\ P(\text{Option B: Lower benefits}) &= \beta_i \\ P(\text{Option C: Later benefits}) &= \delta_i \end{aligned}$$

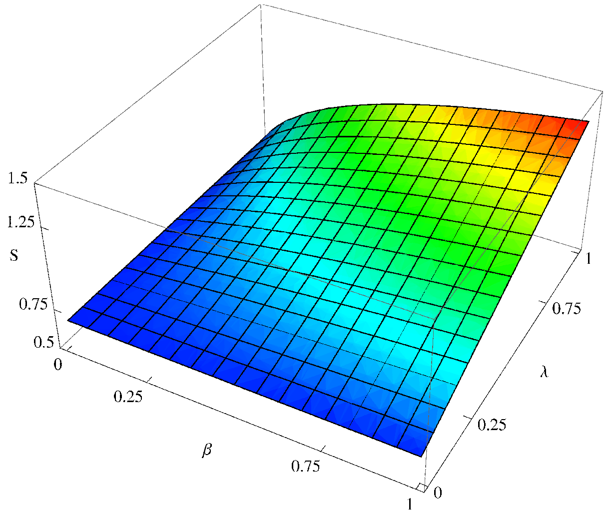
I will analyze uncertainty about two possible policy changes at a time. Individuals thus consider only two possible states of the world. First, the options will be the status quo (policy option A) and a significant policy change, namely lower benefits (B). Note that α_i and β_i add up to one in this case.

$$\max_{S_i} \ln((1-(\tau+\theta))(I_i - S_i)) + 2(1-\beta_i)\ln((1-\tau)(\frac{1}{2}S_i + P)) \\ +2\beta_i \ln((1-\tau)(\frac{1}{2}S_i + (1-\lambda)P))$$

The solution to this maximization problem can be found in appendix 3.A.1 and involves terms which are difficult to interpret at first sight. For clarification, a 3D image of optimal savings is drawn with the probability an individual assigns to lower benefits (β_i) on the x-axis and the cut in the benefit level (λ) on the z-axis in Figure 3.1. The value of the benefit level, P, is normalized to one here and income is set to twice the benefit level.

The figure presents intuitive results, such that if the cut in benefit levels is negligible (i.e. zero) the probability an individual assigns to policy option B does not affect the savings level. Similarly, if someone does

Figure 3.1: Optimal savings when there is uncertainty between the status quo and lower benefits (option B), for different levels of β_i and λ.



not assign any positive value to lower benefit levels coming into effect, the actual size of the cut has no influence on savings. For any positive level of λ however, savings are increasing in β_i, the probability one assigns to lower benefits. Correspondingly, for any positive level of β_i, savings are increasing in λ, the potential cut in benefits.

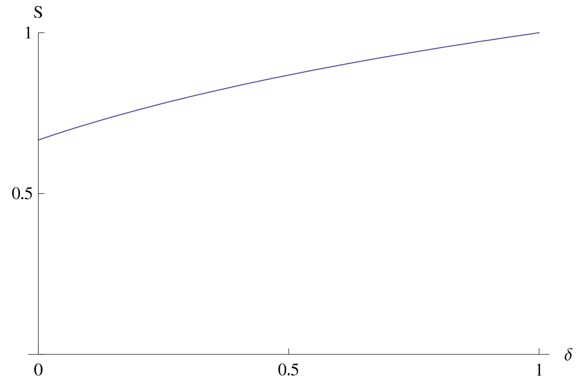
When there is only uncertainty between policy options A, the status quo, and C, an increase in the eligibility age, the maximization problem can be found below. Note that the sum of α_i and δ_i is assumed to be equal to one so that potential changes to the benefits level are ignored.

$$\max_{S_i,z_i} \ln((1-(\tau+\theta))(I_i-S_i)) + 2(1-\delta_i)\ln((1-\tau)(\frac{1}{2}S_i+P)) \\ +2\delta_i \ln((1-\tau)(\frac{1}{2}(S_i+P)))$$

The analytical solution to this problem can be found in Appendix 3.A.1. Figure 3.2 graphically represents optimal saving under uncertainty over the status quo and later benefits, depicting the probability assigned to later benefits, δ_i, on the x-axis. The graph, that has income fixed at twice the benefit level again, is very straightforward. The higher the probability individuals assign to a shift in the social security eligibility age, the higher are optimal savings.

Not surprisingly, assigning a positive probability to any ageing-

Figure 3.2: Optimal savings when there is uncertainty between the status quo and later benefits (option C), for different levels of δ_i.



induced policy change in old age social security should lead to higher savings. This phenomenon could be called crowding in, the opposite of crowding out as introduced by Feldstein (1974).

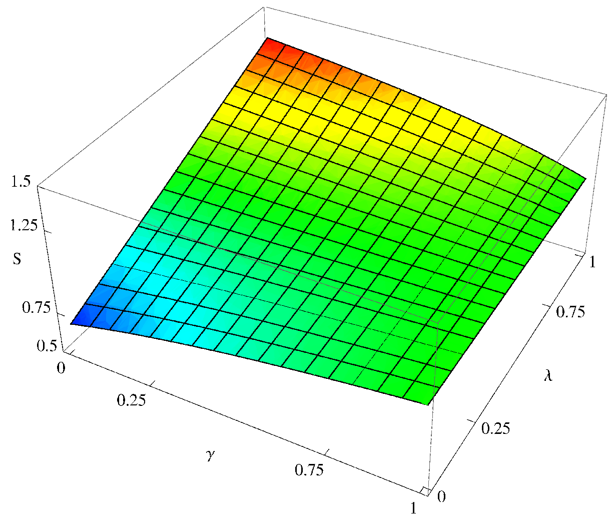
From now on, the situation is considered in which the status quo is no longer sustainable. Saving is analyzed in a world that inhibits uncertainty over which of the two alternative policy options the government will choose, lower benefits (option B) or later benefits (option C). I therefore define γ_i here as $\frac{\delta_i}{\beta_i+\delta_i}$, so as the ratio of the later benefit expectations over the sum of the lower and later benefits expectations. Note that this is necessary in the model as β_i and δ_i can add up to more than one. The pertaining maximization problem is represented by the equation below.

$$\max_{S_i,z_i} \ln((1-(\tau+\theta))(I_i-S_i)) + 2(1-\gamma_i)\ln((1-\tau)(\frac{1}{2}S_i+(1-\lambda)P))$$
$$+2\gamma_i \ln((1-\tau)(\frac{1}{2}(S_i+P)))$$

The analytical solution can be found in Appendix 3.A.1, and optimal savings are graphically presented in Figure 3.3. This graph shows γ_i, the ratio of later benefit expectations over the total expectations of any policy change, on the x-axis and λ, the cut in the benefit-level in case of lower benefits, on the z-axis.

Figure 3.3 uncovers an interesting pattern. The relationship between savings and the parameter γ_i reverses as the benefit cut gets larger. For low cuts ($\lambda < 0.5$), savings increase in γ_i, as the loss associated with only benefits in the third period is larger than the loss associated with lower benefits. For high cuts however ($\lambda > 0.5$), savings actually decrease in

Figure 3.3: Optimal savings when there is uncertainty between lower benefits (option B) and later benefits (option C), for different levels of γ_i and λ_i.



γ_i, as the loss associated with later benefits is now smaller than the loss associated with lower benefits.

The situation in which the government will definitely change the social security system continues to be investigated here. I will argue that the level of private pension savings also depends on the interaction between γ_i - which policy change a person believes is more likely - and her life expectancy. In order to do so the parameter for the cut in benefit levels (λ_i) is personalized to represent differences in the relative individual loss associated with lower (option B) or later (option C) benefits. The idea is that in real-life different survival probabilities render the loss associated with lower and later benefits different across individuals. For example, someone who is certain to die between the age of 65 and 67 is likely to associate a larger loss to a two year increase in the old age benefits eligibility age than to a ten percent cut in the benefits level.

To illustrate this further Figure 3.4 depicts optimal savings for two groups of individuals: the 'long-lived' and the 'short-lived'. The long-lived individuals are certain they will live for three periods, while the short-lived individuals only live for two periods. People know themselves to what group they belong, but the government has no way of differentiating between people in terms of the social security benefit. The long-lived group will therefore receive later benefits with probability γ_i, while the short-lived group will receive no benefit at all with the same probability. Both groups will receive a benefit of level $(1-\lambda_i)P$ with the probability $(1-\gamma_i)$. The analytical expression of optimal savings for the short lived can be found in the appendix. For simplicity, Figure 3.4 includes results for a λ equal to 0.40, which corresponds roughly to a real-life cut in benefit levels of ten percent relative to a two-year increase in the eligibility age.

Figure 3.4: Long-lived and short-lived individuals: Optimal savings when there is uncertainty between lower benefits and later benefits for different levels of γ_i ($\lambda = 0.4$).

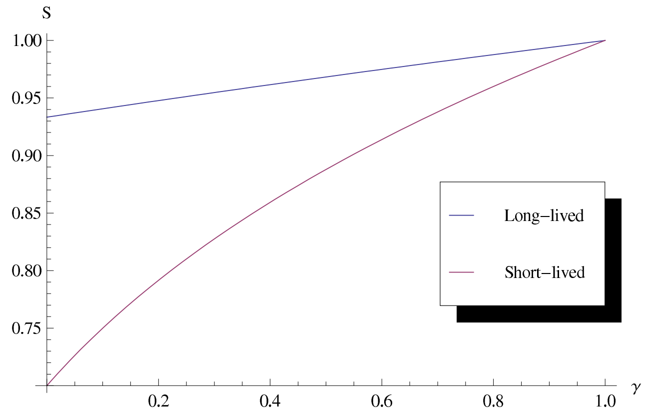


Figure 3.4 shows that savings for the short-lived are much more sensitive to their subjective expectation γ_i, representing the relative likelihood of an increase in the social security eligibility age over a decrease in benefit levels. It can be shown that this difference is larger for higher levels of the benefit cut, λ.

In the empirical analysis there are thus three notions to check: a) Private savings, and hence participation in a private pension scheme, are increasing in the subjective probability individuals assign to lower benefits, β_i. b) Private savings, and participation in a private pension scheme, are also increasing in the subjective probability individuals assign to an increase in the eligibility age, δ_i. c) Private savings, and thus participation in a private pension scheme, are increasing in the relative 'fear' for later benefits (over the sum of lower and later benefits) for short-lived individuals, that assign themselves a relatively low survival probability. To investigate these three notions separately, several models are estimated in Section 3.4.

3.3 Estimation strategy and data

3.3.1 Estimation strategy

Eight single probit regression models are estimated of voluntary participation in a pension scheme. These models all include one or multiple of the expectation parameters: β_i - the probability that benefits will be lowered-, δ_i - the probability that the social security eligibility age will increase - and/or γ_i - the ratio of δ_i over the sum of β_i and δ_i. The dependent variable, participation in a voluntary pension scheme, is binary. The first four estimated models, 1a to 4a, can be found in the equations below. The last four estimated models, 1b to 4b, are identical to their respective a-versions but include interactions of the expectations with a dummy for being subjectively short-lived.

$$\begin{aligned}
y_i^* &= \zeta_{\beta 1} * \beta_i + x_i'\zeta + \epsilon_i \text{ (1a)} \\
y_i^* &= \zeta_{\delta 1} * \delta_i + x_i'\zeta + \epsilon_i \text{ (2a)} \\
y_i^* &= \zeta_{\gamma 1} * \gamma_i + \zeta_{sum} * (\beta_i + \delta_i) + x_i'\zeta + \epsilon_i \text{ (3a)} \\
y_i^* &= \zeta_{\beta 1} * \beta_i + \zeta_{\gamma 1} * \gamma_i + x_i'\zeta + \epsilon_i \text{ (4a)}
\end{aligned}$$

The error terms are all assumed to be i.i.d. drawings from a normal distribution. This implies that it is assumed that the expectations are independent of the error terms. The identification of the expectation coefficients therefore comes from the heterogeneity in expectations among individuals. Although it is possible that some of the heterogeneity in expectations involves general pessimism, this probably does not capture all variation. If pessimism drives everything, the ratio of different policy change expectations should not have a significant coefficient once the sum of the policy change expectations is also included. Pessimism should after all affect the two policy expectations similarly. Nevertheless, I do find significant results for the ζ_γ-coefficient in tables 3.2 and 3.3.

3.3.2 Data

For the empirical analysis data from the Dutch Central Bank Households Savings Survey (DHS), waves 2003 up to 2007[4] is combined with data from the Dutch Pensionbarometer, waves May 2006 up to May 2008. These datasets are both collected by Tilburg University's CentERdata;

[4]Earlier years of the DHS are available but not included as no subjective survival probabilities were administered before 2003.

the first in cooperation with the Dutch Central Bank and the latter in cooperation with Netspar, a Network for Studies on Pensions, Ageing and Retirement. Both datasets are enumerated through the same Internet panel. Much is done to prevent selection effects. The respondents are randomly selected from the general municipalities' administration and when a chosen individual does not have access to the Internet, access is provided. The surveys enumerate partly overlapping subsets of a large pool of respondents so that the datasets could be merged on the identification number of individuals. The information about pension scheme participation and all independent variables except for the expectations are taken from the Dutch Central Bank Households Savings Survey. This survey is designed to investigate savings decisions and enumerates approximately 3000 households each year. The subjective expectations were taken from the Pensionbarometer which is collected mainly to produce longitudinal statistics for confidence levels in the Dutch pension system. Every month on average 500 respondents receive the digital questionnaire although every respondent only receives the questions once every three months. This batch system was introduced to be able to produce the confidence statistics regularly, but not to overwhelm the respondents with questions. In this chapter, the first twenty-four waves of the Pensionbarometer are used of which the last was enumerated in April 2008. A total of 2,554 individuals were interviewed. The sample is limited to those individuals between the ages of 30 and 59 who are participating in a mandatory occupational pension scheme and who have participated in the Pensionbarometer more than once. As a result of the merge and my selection, my sample consists of 3,228 observations and 1,114 individuals.

The dependent variable in the analysis is a dummy on voluntary pension scheme participation. I have taken this variable as the most indicative of one's intention to save voluntarily for old age. As could be already understood from the introduction, most Dutch employees are obliged to participate in occupational pension schemes and total pension wealth is therefore not necessarily a reflection of individual savings decisions. Unfortunately, I do not have detailed information on the level of contributions made into the voluntary (nor the mandatory) pension schemes each year nor of total informal private savings.

The information for this variable is taken from several binary questions in the DHS that ask about voluntary pension scheme contributions each year. An example of such a question is "Did you contribute to a life-interest pension scheme?". Although there is no specific time horizon mentioned in this question, respondents seem to interpret it as relating to

Table 3.1: Percentage of full sample participating in voluntary pension asset types (N=3,228)

Asset type	2003	2004	2005	2006	2007
3rd pillar - life-interest	31%	33%	32%	30%	26%
3rd pillar - lumpsum life-interest	9%	11%	12%	10%	10%
2nd pillar - extra entitlements	1%	2%	1%	2%	1%
2nd pillar - extra payments	2%	2%	1%	1%	1%
One of the above	38%	42%	42%	38%	35 %
N	500	617	697	695	719

a finite time period such as a year. Evidence for this is found in the transition probabilities from one survey to the next: nineteen percent of my sample switch from having contributed to a life-interest to not having contributed whereas this would be impossible if respondents would interpret the question as *ever* having contributed. The dummy in the regressions summarizes whether someone contributed to one designated third pillar pension accounts and/or whether someone voluntarily contributed more than they were obliged to to their occupational pension fund.

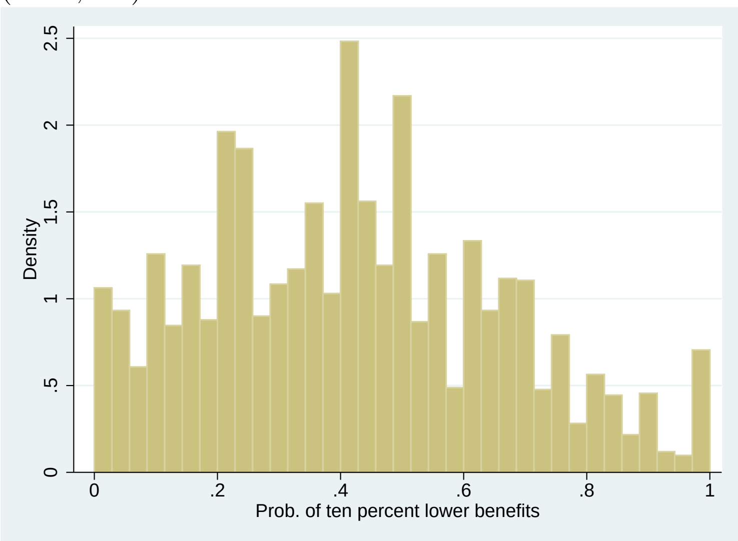
Table 3.1 presents the number of respondents participating in any of these asset schemes per survey year. The variable in the bottom row, whether an individual owns any of the asset types above, is used in the regressions. The most popular voluntary pension savings product is a so-called life-interest contract, a private pension product that requires regular contributions. The second most popular pension product is the lumpsum life-interest contract, a private savings account that is like the life-interest to a certain degree tax-exempt, but that - as the name suggests - requires a lump-sum contribution. Overall, about forty percent of my sample owns at least one type of designated voluntary pension assets. Older individuals (46%) and richer individuals (49%) contribute more often to pension schemes, but young individuals (32%) and low income individuals (27%) still contribute quite often.

The most important independent variables in my analysis are those that represent the policy change expectations, β_i and δ_i. The variables are based on subjective probability answers taken from the Pensionbarometer collected by Tilburg University's CentERdata. Every three months respondents are faced with the following two questions; "What probability (between 0 and 100) do you assign to at least ten percent lower real AOW-benefits within twenty years" and "What probability (between 0 and 100) do you assign to an at least two year higher AOW-eligibility age within twenty years". I use the individual means of the assigned probabilities over the observed 24 monthly waves as my measures for β_i and δ_i. This is done for two reasons. First, as it is unlikely that consumers will change their pension plans every three months according to their updated social security beliefs, I am interested in a relatively stable measure of policy expectations. Second, as Chapter 2 suggests that expectations are measured with a great deal of noise, individual averages will cancel some of that noise out.

Histograms of the β_i (lower benefits probability) and δ_i (later benefits probability) variables are found in Figure 3.5 and Figure 3.6 respectively. It can be seen that the mode of β_i is lower than that of δ_i, suggesting that in general individuals believe that an increase in the eligibility age is the most likely policy change. This is confirmed when checking the average expectations (42 with s.d. 24 and 53 with s.d. 24).

Figure 3.7 displays a histogram of γ_i, the ratio of the estimates of δ_i

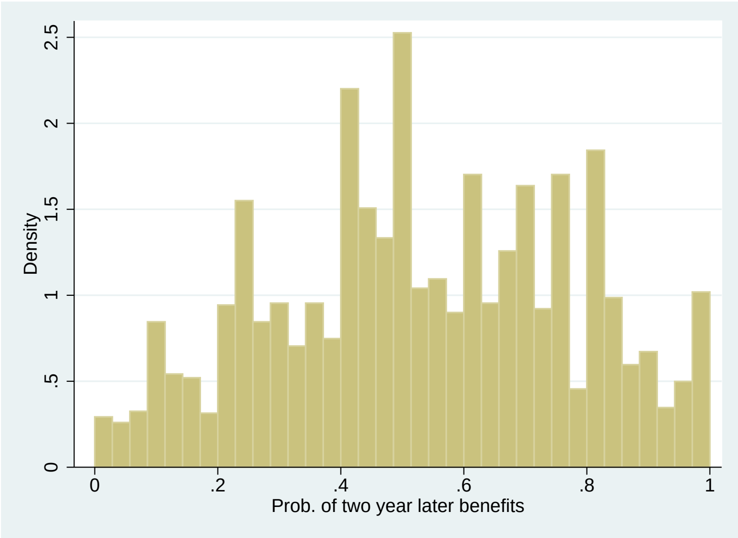
Figure 3.5: Histogram for the probability of lower benefits within twenty years, β_i (N=3,228)



(later benefits probability) over the sum of β_i (lower benefits probability) and δ_i. About 34% of individuals assign an equal probability to both policy changes, but the majority of individuals believes one of the two is more likely to occur with a larger group considering the eligibility age increase more likely. The correlation-coefficient of β_i and δ_i actually is equal to 0.5.

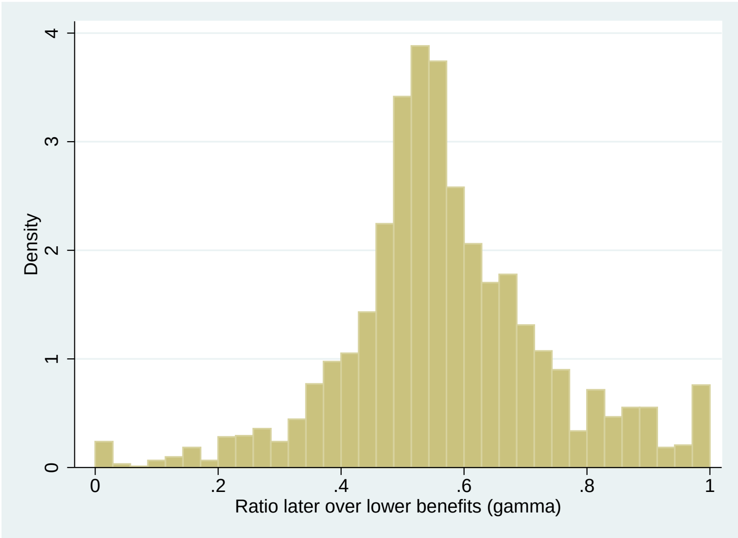
Another important independent variable in the empirical analysis is

Figure 3.6: Histogram for the probability of later benefits within twenty years, δ_i (N=3,228)



the respondent's life expectancy, proxied by a dummy on whether someone is subjectively short-lived. A binary variable was chosen to render the estimation results easy to interpret. The variable is constructed from of a question posed in the Dutch Central Bank Household Savings Survey that asks "What is the probability - between 1 and 10, that you will survive up until the age of 75?". A histogram of answers to this question can be found in the appendix (Figure 3.A.1). The unconditional mean of and median of the survival probability are equal to seven, the mode is equal to eight. In order to construct the short-lived dummy average probabilities are calculated per age and gender group. A graph of these average probabilities can be seen in the appendix (graph 3.A.2). The short-lived dummy is then equal to one if an individual reports a lower probability than the average of his or her appropriate birth year and gender group. This means that 39% of my sample is classified as short-lived. Other controls in the pension fund participation regressions include log gross income, education level, age, a partner dummy, a child dummy and year dummies. Descriptive statistics for six different sub-samples can be found in the appendix (tables 4.1 and 4.2).

Figure 3.7: Histogram of γ_i (the ratio of the estimates of δ_i (lower benefits probability) over the sum of β_i and δ_i (later benefits probability)) in my sample (N=3,228)



3.4 Estimation results

This section discusses the results of the empirical analysis. The appendix contains a table of standard errors (Table 3.A.3). The Akaike's information criteria of these models can be found in Table 3.4.

Table 3.2 displays the relevant coefficients of all eight probit regressions of participation in a voluntary pension scheme. All models are estimated on four different samples; the total sample (column I), the young sample - those between 30 and 39 (column II), a middle aged sample - those between 40 and 49 (column III) and a relatively old sample - those between 50 and 59 (column IV). This was done as there is reason to believe that the degree to which individuals prepare for a policy change in old age social security is different across ages. One could for example argue that the larger the number of years to retirement, the weaker the incentive is to act upon current policy change expectations. Table 3.2 certainly confirms this notion.

The coefficients for the first three models in the first column estimated on the entire sample, in which one type of uncertainty is investigated at a time, illustrate that individuals who assign high policy change probabilities participate more in voluntary pension schemes. This is in accordance with the first two notions developed in Section 3.2. It seems that model 2a, in which there is only policy uncertainty between the status quo and

Table 3.2: Coefficients of probit models. Dependent: participation in voluntary pension scheme. Low age: 30-39 years, Middle age: 40-49 years, High age: 50-59 years. Interactions are with the short-lived dummy.

Model	Variable	(I) All	(II) Low age	(III) Middle age	(IV) High age
1a	Beta	0.241	-0.075	0.301	0.446
2a	Delta	0.507**	0.104	1.139***	0.443
3a	Gamma	0.078	0.263	0.809*	-0.319
	Beta + Delta	0.269**	0.022	0.572***	0.307*
1b	Beta	0.481*	-0.171	0.688*	0.782*
	Interaction	-0.619*	0.248	-0.992*	-0.895*
	Short-lived	0.090	-0.27	0.247	0.203
2b	Delta	0.339	-0.042	1.213***	0.127
	Interaction	0.454	0.402	-0.193	0.875*
	Short-lived	-0.414*	-0.386	-0.043	-0.610*
3b	Gamma	-0.315	0.150	0.572	-0.952*
	Interaction	1.067*	0.343	0.611	1.694**
	Beta + Delta	0.265**	0.021	0.573***	0.285
	Short-lived	-0.778**	-0.35	-0.499	-1.132**
4a	Beta	0.033	-0.162	-0.148	0.306
	Delta	0.464**	0.183	1.200***	0.329
4b	Beta	0.411	-0.196	0.256	0.866**
	Interaction	-1.013***	0.066	-1.040*	-1.600**
	Delta	0.168	0.049	1.107**	-0.204
	Interaction	0.899**	0.381	0.223	1.527**
	Short-lived	-0.221	-0.398	0.179	-0.313
N		3,228	931	1,097	1,198
Ind		1,114	414	422	439

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Standard errors are clustered at the individual level (Table 3.A.3)

All regressions include the short-lived dummy, a constant, log income, education level, age, a child dummy, a partner dummy and year dummies.

an increase in the social security eligibility age, explains most variation in the data. This conclusion can however not be drawn for all age groups. The young react most to their relative 'fear' for a higher eligibility age (γ_i), the middle age category reacts very strongly to their higher eligibility age expectations (δ_i), whereas the oldest age group responds most to their lower benefit level expectations (β_i). The middle age category apparently drives the effects in the overall sample as this group displays the strongest reaction to its policy change expectations. The young on the other hand do not significantly prepare for the anticipated policy changes.

The models 1b, 2b and 3b furthermore include an interaction term between the expectations variables and a dummy for being subjectively short-lived (see Section 3.3 for an explanation of this dummy). In model 1b, where only expectations on the level of benefits is considered, the interaction effect in the total sample is significant and negative, undoing the positive effect of the change probability for the short-lived. Only individuals who believe they have a good chance of being around at age 75 thus prepare for a possible decrease in social security levels. This effect is clear for those older than 39, but is not present for younger individuals. In model 2b, where only uncertainty in the eligibility age is considered, the interaction effect in the entire sample is positive, so that short-lived individuals react more to their eligibility age expectations. The coefficient is however not significant as the interaction has an opposite sign for the middle-aged group. Short-lived individuals between 50 and 59 do significantly let their voluntary pension scheme contribution depend on their δ_i. Focusing on the total sample coefficients in the 3b model, the interaction between being short-lived and the relative fear for an eligibility age increase is large, positive and significant, exactly as theory predicts. The model explains that the expected cost of a higher eligibility age is relatively larger than the cost of lower benefits for individuals that expect to live shorter than others. The empirical effect is clearly driven by the oldest age group, probably because the subjective survival probabilities for this group contain most information.

In model 4a, both policy change expectations are included without the interactions, which takes away some of the effect of lower benefit expectations. However, pension scheme participation is still significantly affected by the probability that the social security eligibility age will go up. Also in this model, the middle age group reacts most distinctly and those under forty hardly react at all. In model 4b, both policy change expectations are included as well as the interactions. The two interactions are both highly significant in the total sample model. Conditioning on the later benefit expectations, short-lived individuals that assign a high probability to lower benefits participate less in pension schemes. Conditioning on the lower benefit expectations, short-lived individuals participate more in pension schemes when they believe a higher eligibility age is likely. When considering the results for the young sample only, it becomes clear that individuals between 30 and 39 do not really prepare for a policy change. Retirement is apparently too far away for this group to take current policy change expectations into account when purchasing pension products. The older two age categories do significantly prepare for policy changes and the interaction effects of being short-lived are most important for the oldest group.

Age is not the only demographic characteristic that is used to divide the sample to estimate the eight models separately. In the last three columns of Table 3.3 the sample is split in three equal income categories to investigate how different income groups react to policy change expectations. There is reason to believe that responses will be different because the Dutch old age social security benefit levels are equal for all (see Section 3.1.3) so that concavity of the utility function could lead to a smaller response to policy change expectations for richer individuals. Column V displays results for low income-individuals - those who earn less than EUR 25,150 a year -, column VI shows coefficients for middle-income individuals - those with a gross yearly income between EUR 25,150 and EUR 38,100 - and the estimates in column VII use information of high income individuals - those with a gross yearly income over EUR 38,100. The appendix contains a table of standard errors (Table 3.A.4) associated to the displayed coefficients. It turns out that responses indeed vary. Broadly speaking, rich individuals indeed do not care so much about upcoming policy changes. The level of social security benefits relative to their own income probably does not justify any additional action if old age benefits might be affected. Middle-income individuals respond much more to their policy change expectations. In particular, the eligibility age expectations explain a substantial degree of pension scheme participation. Moreover, the coefficients of the expectation interaction terms with the short-lived dummy are more significant for those individuals who happen to belong to the middle-income group. Perhaps this group better realizes how the two policy options have different effects on short- and long-lived individuals. Individuals in the low-income group also seem to participate more in voluntary pension schemes when they feel that policy changes are likely. This group is especially worried about the level of pension benefits as the AOW-benefits constitute a very large part of their expected pension income. The interactions with the short-lived dummy are less important.

Table 3.4 provides insights into the fit of the different models for all seven (sub)samples by displaying Akaike's information criteria. The criteria can only be compared within each separate (sub)sample. For the entire sample (column I) model 4b, including both the benefit level (β_i) and eligibility age (δ_i) expectations as well as the interactions with a short-lived dummy, is preferred. This is also the case for the middle age, high age, low income and middle-income sub-samples. For the two subcategories that did not display significant reactions in pension scheme participation according to their policy expectations, the low age and high income groups, this is not the case. Model 2a is preferred when only analyzing young individuals and model 4a is preferred when only analyzing the rich, although the differences in the Akaike's criteria for these two groups are small.

The regression results thus support the three notions introduced in Section 3.2. Participation in voluntary pension schemes is higher for those who assign a high probability to any future change in the PAYG pension scheme. People's expectations regarding a shift in the eligibility age are particularly important. Moreover, conditional on the sum of change expectations, individuals who believe they do not have a long life and who also believe that an increase in the eligibility age is more likely than a decrease in the benefit level, participate more. This can be explained by the fact that the cost of an eligibility age increase will be relatively larger to them. These effects are however not equal for all age and income categories, which can be explained by different time horizons, variation in the relative financial importance of AOW-benefits and differences in the informational value of subjective survival probabilities.

3.5 Conclusion

From the 1960s and 1970s onwards economists have been anxious to know whether private savings in industrialized countries had been crowded out by the installment of old age social security systems in the decades before. The future of Pay-As-You-Go pension schemes seemed prosperous at the time and most studies took "continually rising benefit levels (Feldstein, 1974, p. 911)" into account. During the last two decades however, both economists and policy-makers have realized that Western populations are ageing and that this has consequences for the financial sustainability of old age social security systems. Governments in many OECD-countries are now contemplating how to change their public pension schemes so that

Table 3.3: Coefficients of probit models. Dependent variable: participation in voluntary pension scheme. Low income: lower than EUR 25,150, Middle income: EUR 25,150-EUR 38,100, High income: more than EUR 38,100. Interactions are with the short-lived dummy.

Model	Variable	(I) All	(V) Low inc	(VI) Middle inc	(VII) High inc
1a	Beta	0.241	0.712**	0.333	-0.445
2a	Delta	0.507**	0.640*	0.767**	0.136
3a	Gamma	0.078	-0.201	0.304	0.268
	Beta + Delta	0.269**	0.458**	0.401*	-0.092
1b	Beta	0.481*	1.010**	0.629*	-0.250
	Interaction	-0.619*	-0.750	-0.724	-0.548
	Short-lived	0.090	0.209	0.218	-0.095
2b	Delta	0.339	0.448	0.491	0.109
	Interaction	0.454	0.510	0.710	0.081
	Short-lived	-0.414*	-0.399	-0.475	-0.371
3b	Gamma	-0.315	-0.860*	-0.314	0.123
	Interaction	1.067*	-0.630	1.474*	0.419
	Beta + Delta	0.265**	1.286	0.380*	-0.090
	Short-lived	-0.778**	-0.860*	-0.923*	-0.558
4a	Beta	0.033	0.544	0.010	-0.590*
	Delta	0.464**	0.398	0.762**	0.357
4b	Beta	0.411	1.022**	0.518	-0.347
	Interaction	-1.013***	-1.168*	-1.419**	-0.689
	Delta	0.168	-0.029	0.279	0.237
	Interaction	0.899**	1.048*	1.404**	0.353
	Short-lived	-0.221	-0.195	-0.238	-0.21
N		3,228	1,076	1,076	1,076
Ind		1,114	495	481	421

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Standard errors are clustered at the individual level (Table 3.A.4)

All regressions include the short-lived dummy, a constant, log income, education level, age, a child dummy, a partner dummy and year dummies.

Table 3.4: Akaike Information Criteria per model and sample

Model	Akaike Information Criterion						
	(I)	(II)	(III)	(IV)	(V)	(VI)	(VII)
	All	Low age	Mid age	Hi Age	Low inc	Mid inc	Hi inc
1	4144.8	1127.1	1400.3	1608.2	1212.3	1434.4	1463.7
2	4123.0	1127.0	1362.7	1607.2	1216.5	1417.7	1469.4
3	4131.3	1128.5	1372.5	1603.4	1209.7	1423.8	1469.8
5	4136.9	1128.7	1394.1	1602.0	1209.5	1431.7	1463.4
6	4119.7	1128.0	1364.4	1600.7	1216.4	1415.5	1471.4
7	4119.5	1130.2	1373.1	1589.6	1206.3	1416.8	1471.1
4	4124.9	1128.4	1364.1	1605.9	1210.1	1419.7	1461.9
8	4102.4	1131.4	1360.4	1580.8	1202.5	1405.6	1462.9
N	3,228	931	1,097	1,198	1,076	1,076	1,076

these can last into the (aged) future. Wherever PAYG- pension schemes are still unaltered there thus exists significant uncertainty about future pension policies. Therefore, it is now relevant to investigate how individuals adapt their private savings to a possible dismantlement of public pension provisions.

In this chapter empirical evidence from The Netherlands shows a positive relationship between individuals' expectations about the future of old age social security and their participation in voluntary pension schemes. This increased participation is most pronounced for individuals who assign a higher probability to a potential increase in the eligibility age, but a small positive effect of the subjective probability assigned to a decrease in benefit levels is also found. Which of these policy options an individual believes is more likely to occur also matters. Taking the sum of policy change expectations into account, those who believe that an increase in the eligibility age is more likely than a decrease in the benefit level participate slightly more. This effect is however much stronger for those who assign themselves a low survival probability, probably because an eligibility increase implies relatively larger costs for them.

The empirical effects are however not equal for all age and income categories. The young and those with a relatively high income do not seem to prepare much for a policy change, regardless of how likely they believe such a change will be. Retirement is perhaps too far away for those between 30 and 39 to take current policy expectations into account in their savings decisions. Rich individuals on the other hand probably do not care about the future of public pension benefits as these benefits only amount to a small fraction of their anticipated pension income. The middle-aged and those in the middle-income category seem particularly concerned with an increase in the social security eligibility age, whereas the low-income group responds to a greater extent to their benefit level expectations. The latter makes sense as social security benefits are by far the largest part of anticipated pension income for this group. The different effects for short- and long-lived individuals are most pronounced in the oldest group and in the middle-income group. Apparently, these groups are better aware of the differential impact of later and lower benefits on different survival groups.

Policy-makers and politicians can learn two things from the above results. First, as people's policy expectations matter for their savings behavior, these expectations should be managed with care to the extent that this is possible. The public debate on the future sustainability of old age social security supposedly influences policy expectations and therefore concern should only be raised if this is somehow justified. Second, the results suggest that wealth effects of policy changes in old age social security will partly be offset by individuals' private pension savings. This finding encourages governments to actually carry out necessary changes to the PAYG pension-system if these are generally expected anyhow.

Although my empirical results are suggestive, the case for a crowding in effect would be stronger if the same results would be replicated using level of savings data and longitudinal expectations date, enabling more specific and causal inference respectively. Evidence from other countries about the effect of the (potential) degeneration of the PAYG-pension scheme on private savings is also essential to draw definitive conclusions about the crowding in of savings by policy uncertainty in first pillar pensions. The relatively new SHARE dataset, that collects identical information in several European countries, could be ideal for this purpose.

3.A Appendix

3.A.1 Algebraic results

This appendix displays optimal savings equations resulting from the model described in Section 3.2.

The following savings level is optimal when the government chooses between the status quo (A) with probability $(1-\beta_i)$ and lower benefits (B) with probability β_i (A or B model).

$$S_i^{*,AorB} = \frac{1}{6}(-8P + 2I_i + 6P\lambda - 4P\beta_i\lambda + \sqrt{((-8P + 2I_i + 6P\lambda - 4P\beta_i\lambda)^2 + 12(-4P^2 + 4PI_i + 4P^2\lambda - 4PI_i\lambda + 4PI_i\beta_i\lambda)))}$$

Normalizing the benefit level P to be equal to one and income to be equal to two this becomes:

$$S_i^{*,AorB} = \frac{1}{6}(-4 + 6\lambda - 4\beta_i\lambda + \sqrt{(-4 + 6\lambda - 4\beta_i\lambda)^2 + 12(4 - 4\lambda + 8\beta_i\lambda)})$$

The savings below are optimal when the government chooses between the status quo with probability $(1-\delta_i)$ and an increase in the eligibility age with probability δ_i (A or C model).

$$S_i^{*,AorC} = \frac{1}{6}(-5P - 2\delta_i P + 2I + \sqrt{P^2 + 20\delta_i P^2 + 4\delta_i^2 P^2 + 4PI_i + 16\delta_i PI_i + 4I_i^2})$$

Normalizing the benefit level P to be equal to one and income to be equal to two this becomes:

$$S_i^{*,AorC} = \frac{1}{6}(-1 - 2\delta_i + \sqrt{25 + 52\delta_i + 4\delta_i^2})$$

The following equation displays optimal savings when there will certainly be a policy change but it is uncertain which option the government will choose, either lower benefits with probability $(1-\gamma_i)$ or an increase in the eligibility age with probability $\gamma_i = \frac{\delta_i}{\delta_i + \alpha_i}$ (the B or C model).

$$S_i^{*,BorC} = \frac{1}{6}(-5P + 2I_i - 2P\gamma_i + 2P\lambda + 4P\gamma_i\lambda + \sqrt{((-5P + 2I_i - 2P\gamma_i + 2P\lambda + 4P\gamma_i\lambda)^2 + 12(-2P^2 + 2PI_i + 2PI_i\gamma_i + 2P^2\lambda - 4PI_i\gamma_i\lambda)))}$$

Normalizing the benefit level P to be equal to one and income to be equal to two this becomes:

$$S_i^{*,BorC} = \frac{1}{6}(-1 - 2\gamma_i + 2\lambda + 4\gamma_i\lambda + \sqrt{12(2 + 4\gamma_i + 2\lambda - 8\gamma_i\lambda) + (-1 - 2\gamma_i + 2\lambda + 4\gamma_i\lambda)^2}) \tag{1}$$

The equation for optimal savings of short-lived individuals (those that know they will die after the second period) when there is uncertainty between policy option B (lower benefits) and policy option C (later benefits) is depicted in the following equation.

$$S_i^{*,BorC,short} = \frac{1}{4}(-P + I_i - P\gamma_i + P\lambda + P\gamma_i\lambda + \sqrt{(-P + I - P\gamma_i + P\lambda + P\gamma_i\lambda)^2 + 8(PI_i\gamma - PI_i\gamma_i\lambda)})$$

Normalizing the benefit level P to be equal to one and income to be equal to two this becomes:

$$S_i^{*,BorC,short} = \frac{1}{4}(1 - \gamma_i + \lambda + \gamma_i\lambda + \sqrt{8(2\gamma_i - 2\gamma_i\lambda) + (1 - \gamma_i + \lambda + \gamma_i\lambda)^2}) \tag{2}$$

3.A.2 Figures and tables

Figure 3.A.1: Integer survival probabilities up to age 75 (N=3,228)

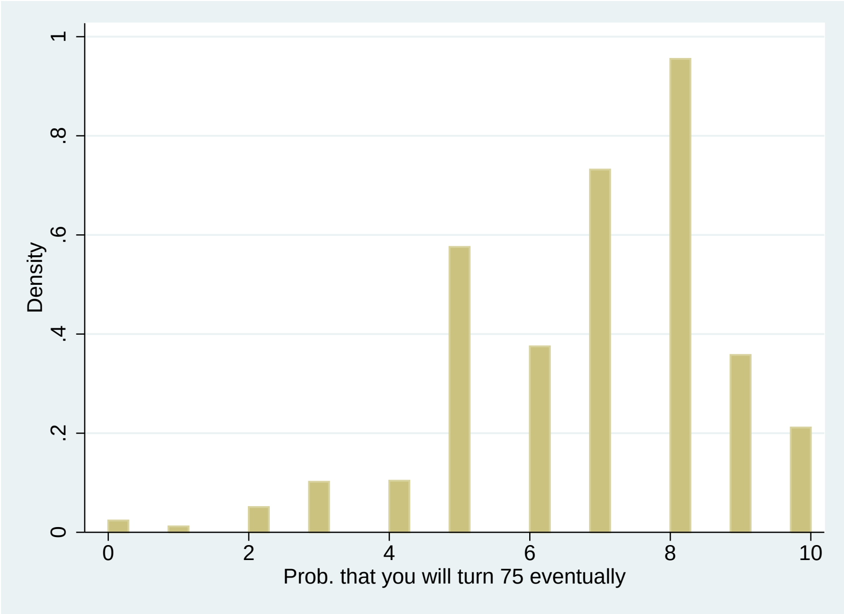


Figure 3.A.2: Lowess smoothing graph of average survival probabilities up to age 75 on a scale from 1 to 10 for different ages and sex (N=3,228)

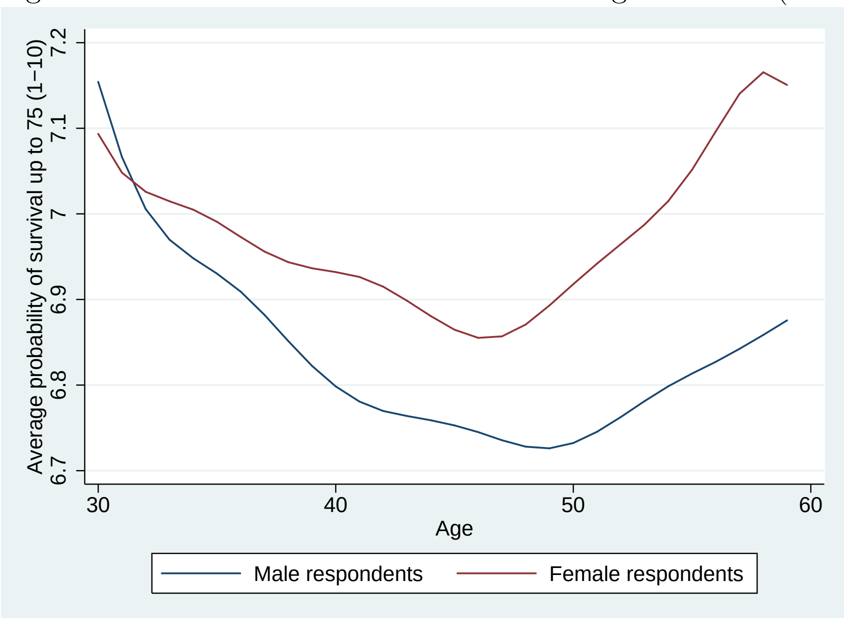


Table 3.A.1: Descriptive statistics of dependent and independent variables in the estimations for different age groups.

Variable	(II) Low age		(III) Middle age		(IV) High age	
	Mean	Sd	Mean	Sd	Mean	Sd
Dummy pension particip.	0.321	0.467	0.365	0.482	0.463	0.499
Prob. of lower benefits (β_i)	0.442	0.230	0.426	0.238	0.395	0.247
Prob. of later benefits (δ_i)	0.539	0.227	0.531	0.230	0.529	0.261
Ratio later over lower (γ_i)	0.557	0.144	0.567	0.163	0.580	0.185
Sum of expectations	0.981	0.395	0.957	0.392	0.924	0.428
Dummy for short-lived	0.408	0.492	0.388	0.488	0.384	0.487
Log of gross income	9.904	1.843	9.742	2.346	9.934	2.149
Age	34.290	2.821	44.813	2.814	54.325	2.827
Dummy for female	0.437	0.496	0.428	0.495	0.356	0.479
Dataset 2003	0.148	0.356	0.160	0.367	0.155	0.362
Dataset 2004	0.189	0.392	0.203	0.403	0.182	0.386
Dataset 2005	0.218	0.413	0.215	0.411	0.215	0.411
Dataset 2006	0.205	0.404	0.213	0.410	0.225	0.418
Dataset 2007	Omitted category					
Primary education	Omitted category					
Secondary education	0.226	0.418	0.342	0.475	0.421	0.494
Vocational education	0.293	0.455	0.248	0.432	0.163	0.370
Tertiary education	0.456	0.498	0.377	0.485	0.383	0.486
Other education	0.021	0.145	0.010	0.100	0.002	0.041
Dummy for child	0.534	0.499	0.603	0.490	0.358	0.480
Dummy for partner	0.716	0.451	0.754	0.431	0.762	0.426

Table 3.A.2: Descriptive statistics of dependent and independent variables in the estimations for different income groups.

Variable	(V) Low inc		(VI) Middle inc		(VII) High inc	
	Mean	Sd	Mean	Sd	Mean	Sd
Dummy pension particip.	0.267	0.442	0.407	0.492	0.493	0.500
Prob. of lower benefits (β_i)	0.421	0.253	0.430	0.241	0.405	0.224
Prob. of later benefits (δ_i)	0.538	0.249	0.549	0.239	0.511	0.233
Ratio later over lower (γ_i)	0.573	0.168	0.569	0.168	0.564	0.165
Sum of expectations	0.959	0.429	0.979	0.407	0.917	0.381
Dummy for short-lived	0.415	0.493	0.389	0.488	0.373	0.484
Log of gross income	8.377	3.198	10.367	0.108	10.836	0.265
Age	44.849	8.650	44.612	8.701	46.482	8.211
Dummy for female	0.724	0.447	0.335	0.472	0.152	0.360
Dataset 2003	0.164	0.370	0.147	0.354	0.154	0.361
Dataset 2004	0.192	0.394	0.185	0.388	0.196	0.397
Dataset 2005	0.213	0.409	0.225	0.418	0.210	0.408
Dataset 2006	0.207	0.406	0.237	0.425	0.202	0.401
Dataset 2007	Omitted category					
Primary education	Omitted category					
Secondary education	0.442	0.497	0.356	0.479	0.215	0.411
Vocational education	0.295	0.456	0.246	0.431	0.148	0.355
Tertiary education	0.209	0.407	0.370	0.483	0.627	0.484
Other education	0.010	0.101	0.019	0.135	0.002	0.043
Dummy for child	0.562	0.496	0.436	0.496	0.478	0.500
Dummy for partner	0.798	0.401	0.675	0.469	0.765	0.424

Table 3.A.3: Coefficients and standard errors of probit models. Dependent variable: participation in voluntary pension scheme.

Model	Variable	All	Low age	Middle age	High age
1a	Beta	0.241	-0.075	0.301	0.446
		(0.159)	(0.268)	(0.269)	(0.260)
2a	Delta	0.507	0.104	1.139	0.443
		(0.163)	(0.281)	(0.286)	(0.246)
3a	Gamma	0.078	0.263	0.809	-0.319
		(0.238)	(0.448)	(0.410)	(0.345)
	Beta + Delta	0.269	0.022	0.572	0.307
		(0.097)	(0.164)	(0.170)	(0.150)
1b	Beta	0.481	-0.171	0.688	0.782
		(0.191)	(0.349)	(0.327)	(0.314)
	Interaction	-0.619	0.248	-0.992	0.895
		(0.268)	(0.462)	(0.450)	(0.448)
	Short-lived	0.090	-0.270	0.228	0.205
		(0.132)	(0.223)	(0.228)	(0.205)
2b	Delta	0.339	-0.042	1.213	0.127
		(0.191)	(0.351)	(0.343)	(0.285)
	Interaction	0.454	0.402	-0.193	0.875
		(0.289)	(0.470)	(0.528)	(0.446)
	Short-lived	-0.414	-0.386	-0.043	-0.610
		(0.172)	(0.277)	(0.312)	(0.260)
3b	Gamma	-0.315	0.150	0.572	-0.952
		(0.267)	(0.543)	(0.470)	(0.390)
	Interaction	1.067	0.343	0.611	1.694
		(0.422)	(0.740)	(0.749)	(0.650)
	Beta + Delta	0.265	0.021	0.573	0.285
		(0.098)	(0.164)	(0.170)	(0.153)
	Short-lived	-0.778	0.350	-0.499	-1.132
		(0.252)	(0.435)	(0.449)	(0.393)
4a	Beta	0.033	-0.162	-0.148	0.306
		(0.176)	(0.301)	(0.295)	(0.286)
	Delta	0.464	0.183	1.200	0.329
		(0.179)	(0.316)	(0.312)	(0.270)
4b	Beta	0.411	-0.196	0.256	0.866
		(0.210)	(0.390)	(0.360)	(0.336)
	Interaction	-1.013	0.066	-1.040	-1.600
		(0.305)	(0.530)	(0.524)	(0.497)
	Delta	0.168	0.049	1.107	-0.204
		(0.209)	(0.395)	(0.373)	(0.304)
	Interaction	0.899	0.381	0.223	1.527
		(0.323)	(0.539)	(0.598)	(0.483)
	Short-lived	-0.221	0.398	0.179	-0.313
		(0.177)	(0.291)	(0.320)	(0.268)

Standard errors are in parentheses

Table 3.A.4: Coefficients and standard errors of probit models. Dependent variable: participation in voluntary pension scheme.

Model	Variable	All	Low inc	Middle inc	High inc
1a	Beta	0.241	0.712	0.333	-0.445
		(0.159)	(0.249)	(0.259)	(0.272)
2a	Delta	0.507	0.640	0.767	0.136
		(0.163)	(0.251)	(0.256)	(0.76)
3a	Gamma	0.078	-0.201	0.304	0.268
		(0.238)	(0.410)	(0.396)	(0.383)
	Beta + Delta	0.269	0.458	0.401	-0.092
		(0.097)	(0.144)	(0.158)	(0.169)
1b	Beta	0.481	1.010	0.629	-0.250
		(0.191)	(0.309)	(0.308)	(0.326)
	Interaction	-0.619	-0.750	-0.724	-0.548
		(0.268)	(0.418)	(0.436)	(0.484)
	Short-lived	0.090	0.209	0.218	-0.095
		(0.132)	(0.216)	(0.224)	(0.220)
2b	Delta	0.339	0.448	0.491	0.109
		(0.191)	(0.419)	(0.461)	(0.522)
	Interaction	0.454	0.510	0.710	0.081
		(0.289)	(0.419)	(0.461)	(0.522)
	Short-lived	-0.414	-0.399	-0.475	-0.371
		(0.172)	(0.257)	(0.283)	(0.296)
3b	Gamma	-0.315	-0.860	-0.314	0.123
		(0.267)	(0.449)	(0.441)	(0.436)
	Interaction	1.067	-0.630	1.474	0.419
		(0.422)	(0.707)	(0.733)	(0.711)
	Beta + Delta	0.265	1.286	0.380	-0.090
		(0.098)	(0.144)	(0.161)	(0.170)
	Short-lived	-0.778	-0.860	-0.923	-0.558
		(0.252)	(0.420)	(0.429)	(0.430)
4a	Beta	0.033	0.544	0.010	-0.590
		(0.176)	(0.290)	(0.290)	(0.287)
	Delta	0.464	0.398	0.762	0.357
		(0.179)	(0.293)	(0.282)	(0.294)
4b	Beta	0.411	1.022	0.518	-0.347
		(0.210)	(0.359)	(0.339)	(0.344)
	Interaction	-1.013	-1.168	-1.419	-0.689
		(0.305)	(0.506)	(0.521)	(0.519)
	Delta	0.168	-0.029	0.279	0.237
		(0.209)	(0.344)	(0.328)	(0.333)
	Interaction	0.899	1.048	1.404	0.353
		(0.323)	(0.513)	(0.530)	(0.558)
	Short-lived	-0.221	-0.195	-0.238	-0.210
		(0.177)	(0.254)	(0.294)	(0.311)

Standard errors are in parentheses

On Power

Chapter 4

Kinky Choices, Dictators and Split Might: A Non-Cooperative Model for Household Consumption and Labor Supply[1]

This chapter is based on Boone, Van der Wiel and Vermeulen (2009).

4.1 Introduction

Common sense suggests that no two persons, and particularly not one man and one woman, are alike. And although it can be argued that two persons that voluntarily form a household will probably do so on the basis of shared interests and preferences, no one will maintain that married individuals have identical preferences on all accounts. This means that (economic) decisions taken jointly in a household will be more complicated than those taken by an individual.

One of the important insights from recent applied microeconometrics is that multi-person households indeed do not behave as single decision makers. This single decision maker assumption, which is fundamental to the standard unitary model of household behavior, is associated with the

[1]Many thanks to my co-authors Jan Boone and Frederic Vermeulen. Moreover, thanks to Arthur van Soest, Arie Kapteyn and Tobias Klein, as well as seminar participants at Tilburg University, for useful comments and suggestions.

theoretical implication that observed demand or labor supply should satisfy the well-known Slutsky conditions. Slutsky symmetry and negativity, though, are usually rejected when confronted with consumption or labor supply data of households (see Fortin and Lacroix, 1997, Browning and Chiappori, 1998, Vermeulen, 2005, and Cherchye and Vermeulen, 2008, for some recent examples). Evidence thus suggests that intra-household bargaining aspects within multi-person households cannot be ignored.

A number of alternatives to the unitary model have been put forward. One strand in the literature assumes that household members only choose Pareto efficient allocations. This is either formalized by means of axiomatic bargaining theory (see, for example, Manser and Brown, 1980, and McElroy and Horney, 1981) or via the so-called collective model (see Chiappori, 1988, 1992, and Apps and Rees, 1988). Another strand of the literature assumes that household members behave non-cooperatively (see Leuthold, 1968, and Ashworth and Ulph, 1981, for seminal contributions and Browning, 2000, and Browning, Chiappori and Lechene, 2009, for more recent work).

One of the main implications of the models that recognize that households do not behave as single decision makers is that intra-household allocations may depend on individual resources of household members rather than on only the household's aggregate means. A leading empirical example in this respect is that children seem to benefit more when the mother (or the caring grandmother) brings relatively more financial resources into the household. This has been consistently shown for both developed and developing nations (e.g. Lundberg, Pollak and Wales, 1997, for the U.K., Thomas, Contreras and Frankenberg, 1997, for Indonesia, and Duflo, 2003, for South Africa). From a policy point of view this dependence on who brings what to the table is very important. It implies that policy makers can alter the intra-household allocation of resources through targeted taxes or transfers.

The targeting issue has been investigated in a collective setting (see Blundell, Chiappori and Meghir, 2005) as well as in a non-cooperative setting. The usual approach in the latter is by considering models that focus on the private provision of public goods (see, among others, Bergstrom, Blume and Varian, 1986, Chen and Woolley, 2001, and Lechene and Preston, 2008). An important restriction in the above non-cooperative models is that the labor incomes of both the husband and the wife, and thus the household's resources, are assumed exogenous. However, it is not difficult to come up with arguments in favor of more realistic models that not only describe the intra-household allocation of resources but also the generation of these resources themselves.

A first aim of the current study is to fill this gap by focusing on the private provision of public goods while also taking into account labor supply decisions and the consequent endogenous individual incomes. We will start out with the most general model and investigate what could happen to public goods demand in this set-up. We define three regimes and show that all Nash equilibria fall in either of these regimes. In the first regime, the husband is dictator (denoted by HD). More specifically, the household's demand for public goods fully reflects the husband's preferences given the household's aggregate resources. The second regime is associated with split might (SM): each spouse's preferences are reflected in the household's spending on public goods. The extent to which this happens depends on the spouses' relative wage rates. Finally, the third regime is associated with the wife being the dictator (WD).

Like in Browning, Chiappori and Lechene (2009), and contrary to Lundberg and Pollak (1993), the regimes are determined 'endogenously' by the spouses' relative wage rates. The main difference between our results and those obtained by Browning et al. (2009) is that we do not have a pure local income pooling phenomenon. In their paper, income pooling refers to the particular situation in which the allocation of a constant amount of resources is independent of which partner brings these resources to the table. As labor supply is endogenous in our model, exogenous shifts in individual income that keep the household's aggregate budget fixed are ruled out, and therefore pure income pooling doesn't exist. Still, there is some sort of local income pooling in both dictatorship regimes: aggregate spending on public goods is as the most powerful spouse would want it to be.

A second contribution of the current study is that the theoretical model for the private provision of public goods with endogenous labor supply will be brought to the data. Contrary to collective models, of which a wide variety of empirical applications are available, empirical evidence for the private provision of public goods remains scarce. This gap will be partly filled in our study by means of an empirical model that focuses on expenditures on children's goods (such as children's clothing, toys or tuition fees). These child related goods are considered to be public goods inside a household. The empirical model does impose more structure than the general model, so that we can derive testable implications against the standard unitary model. The data for the application are drawn from the Consumer Expenditure Survey (CEX) collected by the Bureau of Labor Statistics of the U.S. Department of Labor. Taking the targeting literature into account, we suspect that preferences within the household regarding spending on children's goods will differ. We find that we can reject the standard unitary model in favor of our non-cooperative approach for couples with two or three children under the age of 18. Using the estimated preference parameters, we can then divide households into dictatorship and split might regimes. It turns out that 73% of two-child and 57% of three-child couples spend according to the wife's preferences.

The rest of this chapter unfolds as follows. In Section 2, we propose a general private provision of public goods model with endogenous labor supply and discuss its implications. Section 3 focuses on the empirical specification and the estimation strategy, that will be based on a maximum likelihood approach. The data and estimation results are discussed in Sections 4 and 5. Section 6 concludes.

4.2 The model

We focus on two adult households where f denotes the wife and m the husband.[2] Spouses have to decide on their demands for leisure $l^f, l^m \in [0,1]$ (normalized between zero and one; including the option to be out of the labor market) and on how the household's aggregate resources are allocated to a vector of private (denoted by the vectors $\mathbf{q}^f \in \mathbb{R}_+^{n_f}, \mathbf{q}^m \in \mathbb{R}_+^{n_m}$) and public goods inside the household (denoted by the vector $\mathbf{Q} \in \mathbb{R}_+^{n_p}$). The household's aggregate resources are assumed to be equal to the sum of the individual labor incomes $w^f \left(1 - l^f\right) + w^m \left(1 - l^m\right)$. To keep things simple, we abstract away from non-labor income. Prices of private and public goods are denoted by respectively $\mathbf{p}^f$, $\mathbf{p}^m$ and $\mathbf{P}$.

In what follows, we assume that leisure is a private good in the sense that it does not entail intra-household externalities. Therefore, the wife's and husband's utility functions, representing their own preferences, are given by respectively:

$$u^f \left(l^f, \mathbf{q}^f, \mathbf{Q}\right) \tag{1}$$

and

$$u^m \left(l^m, \mathbf{q}^m, \mathbf{Q}\right). \tag{2}$$

The question now is how the household makes decisions. Following, among others, Browning et al. (2009), we assume a non-cooperative setting by adopting a Nash approach. A Nash equilibrium is defined as follows:

[2] It goes without saying that all results apply to same-sex couples as well. For notational reasons we stick to the traditional husband and wife terminology.

Definition 1 *A* Nash equilibrium *consists of individual leisure, a vector of individual private consumption and individual contributions to public goods* $(l^{i*}, \mathbf{q}^{i*}, \mathbf{Q}^{i*})$ *,* $i = f, m$*, such that for each* i*,* $(l^{i*}, \mathbf{q}^{i*}, \mathbf{Q}^{i*})$ *solves for* $i \neq j$*:*

$$\max_{l^i \in [0,1], \mathbf{q}^i, \mathbf{Q}^i \geq 0} u^i \left(l^i, \mathbf{q}^i, \mathbf{Q}^i + \mathbf{Q}^{j*} \right) \tag{3}$$

$$s.t.\ w^i l^i + \mathbf{p}^{i\prime} \mathbf{q}^i + \mathbf{P}' \mathbf{Q}^i \ = \ w^i.$$

We make the following assumptions on the utility functions. We denote marginal utility with respect to good x by u_x^f or u_x^m and assume that these derivatives are well defined. The first four assumptions are fairly standard. The combination of the first and the third assumption seems to imply that both partners participate in the labor market, as one obtains infinite utility from consuming some private goods and as labor income is the only resource available to pay for these private goods. We come back to this below. The last assumption is most interesting as it implies a conflict within the household.

Assumption *For arbitrary values of* l^i*,* $\mathbf{q}^i$ *(in which* $i = f, m$*) and* $\mathbf{Q}$*, we have that*

1. $\lim_{x \downarrow 0} u_x^i(l^i, \mathbf{q}^i, \mathbf{Q}) = +\infty$ *for leisure* $(x = l^i)$*, each private good* $(x = q_k^i)$ *and each public good* $(x = Q_k)$*,*

2. $\lim_{x \to +\infty} u_x^i(l^i, \mathbf{q}^i, \mathbf{Q}) = 0$ *for each private good* $(x = q_k^i)$ *and each public good* $(x = Q_k)$*,*

3. $u_l^i(1, \mathbf{q}^i, \mathbf{Q}) < +\infty$*,*

4. $u^i(l^i, \mathbf{q}^i, \mathbf{Q})$ *is concave in leisure, each private good and each public good and*

5. *there exist (at least) two public goods* $Q_k, Q_{k'}$ *such that*

$$\frac{u_{Q_k}^f}{u_{Q_{k'}}^f} > \frac{u_{Q_k}^m}{u_{Q_{k'}}^m}.$$

The first assumption says that for each good the marginal utility goes to infinity as the amount of the good goes to zero. The second assumption says that for each private and public good the marginal utility goes to zero as the amount of the good goes to infinity. These assumptions are made for ease of exposition. Since we are interested in corner solutions where one of the partners does not contribute to a public good, we want to avoid corner solutions in private goods and total contributions to public goods.[3] The third assumption implies that u_l^i (for $i = f, m$) is finite if the partner does not work at all ($l^i = 1$). This leaves us with two ways to capture non-participation in the labor market by a partner. First, $w^f = 0$ ($w^m = 0$) implies that the wife (husband) does not participate. Second, in the empirical specification we assume that there are no private goods because of data limitations. Without private goods, non-participation then arises for a range of wages $w^f > 0$. Concavity is sufficient to allow us to use stationary points to characterize a global maximum. Finally, we assume that there is a tension between the partners. They never agree on the overall contributions to all public goods. Under assumptions one to five, we can prove the following. The appendix contains the proof.

Lemma 1 *In Nash equilibrium[4] we have that*

$$\frac{u_l^f}{u_{q_k}^f} = \frac{w^f}{p_k} \quad \textit{and} \quad \frac{u_l^m}{u_{q_k}^m} = \frac{w^m}{p_k}$$

for each private good q_k.

We next define three regimes and show below that all Nash equilibria fall in either of these regimes. In the first regime, the husband is dictator (denoted by HD). More specifically, the household's demand for public goods fully reflects the husband's preferences given the household's aggregate resources. The second regime is associated with split might (SM): each spouse contributes to public goods but not to all of them, and, moreover, has a say on how the household's aggregate resources are allocated. The resulting allocation is however not as any of the spouses would wish it to be. Finally, the third regime is associated with the wife being the dictator (WD). Like in Browning, Chiappori and Lechene (2009), and contrary to Lundberg and Pollak (1993), these regimes are determined

[3] Allowing for such corner solutions adds inequalities to the optimality conditions. This complicates notation without adding insight. The corner solutions for individual contributions to public goods are however interesting as we show below.

[4] It follows from Theorem 1.2 in Fudenberg and Tirole (1991, p. 34) that a pure strategy Nash equilibrium exists in our case as the strategy spaces of both spouses (potential labor supply and consumption) are nonempty, compact, convex subsets of an Euclidean space and as the utility functions of both spouses are continuous in both spouses' strategies and quasi-concave in their own strategy. Multiple equilibria cannot be ruled out.

'endogenously' by the spouses relative wage rates. Note that one special case within the SM regime is the Separate Spheres regime of the latter two papers. Under Separate Spheres, each spouse contributes to strictly different sets of public goods (an example will be given in the empirical application in Section 4.3).

Formally, we have the following (where we use the convention on inequalities with vectors that $\mathbf{x} < \mathbf{y}$ implies that $x_k \leq y_k$ for all k where the inequality is strict for at least one k).

Definition 2 *The three regimes are defined as follows*

HD $u^m_{\mathbf{Q}^m} = \lambda^m \mathbf{P}$ *and* $u^f_{\mathbf{Q}^f} < \lambda^f \mathbf{P}$*,*

SM $u^m_{\mathbf{Q}^m} < \lambda^m \mathbf{P}$ *and* $u^f_{\mathbf{Q}^f} < \lambda^f \mathbf{P}$ *and*

WD $u^m_{\mathbf{Q}^m} < \lambda^m \mathbf{P}$ *and* $u^f_{\mathbf{Q}^f} = \lambda^f \mathbf{P}$*.*

where $\lambda^i = u^i_{q_k}/p_k$ *(from Lemma 1) is the Lagrange multiplier associated with partner* i*'s* $(i = f, m)$ *budget constraint.*

The following proposition demonstrates that the three regimes above are the only ones that can occur in equilibrium. The proof can be found in the appendix.

Proposition 1 *If* $w^m > 0$ *and/or* $w^f > 0$*, then HD, SM and WD are the only possibilities. That is, each Nash equilibrium in (3) is characterized by the equalities in Lemma 1 and the conditions in either HD, SM or WD.*

The example in Browning et al. (2009) demonstrates that the regimes are ordered as HD, (our) SM and WD and that the ordering is a function of the wife's exogenous share of income. We present a similar result, with endogenous incomes, in the next section. Here we consider what we can say about this ordering in our general set-up. The next result (partially) characterizes the ordering of the three regimes in terms of relative wages $\rho = w^f/w^m$. See the appendix for a proof.

Proposition 2 *There exist critical values* $\rho_0 > 0$ *and* $\rho_1 > \rho_0$ *such that the household is in regime HD for each* $w^f/w^m < \rho_0$ *and in regime WD for each* $w^f/w^m > \rho_1$*.*

Our empirical specification in the next section imposes more structure on the utility functions u^m and u^f. This allows us to show that for given

$w^m > 0$ we move through the regimes as w^f increases in the order HD, SM and finally WD. With the general set-up in this section we cannot rule out orderings like HD, SM, HD (again), WD. The robust insight is that when one partner (potentially) has a sufficiently higher wage rate relative to that of the other partner, the household allocation to public goods is determined completely by this partner's preferences - given the household's aggregate resources. That is, one always starts with HD (for low ρ) and ends up with WD (for high ρ).

4.3 Empirical specification and estimation strategy

4.3.1 Empirical specification

We will illustrate the existence of dictatorship and split might regimes by means of a sample of couples with children drawn from the Consumer Expenditure Survey (CEX). This dataset contains not only detailed purchases by households but also information on wages and labor supply of each household member (see the next section for more details). Given the particular data at hand, we will focus on a special case of the general model described above. In this special case the only possible regimes continue to be Husband Dictatorship, Split Might and Wife Dictatorship. Moreover, in the empirical specification the ordering of the regimes is perfectly known. Firstly, it turns out that almost all prime age men in the selected sample work full time. Therefore, we will assume that men's labor supply is exogenously fixed. Wives, on the contrary, are assumed to have a continuous hours choice (including non-participation). Wives' leisure is a private good in the model. Secondly, as is common in budget surveys, expenditures are recorded at the household level. This implies that, for most goods, one cannot observe the spouses' individual consumption of private goods. As a result, we will assume that all consumption is public inside the household. Two public goods will be distinguished: a composite good that relates to expenditures on children's goods (such as clothing, toys and tuition fees) and a composite good relating to other (nondurable) expenditures.

To obtain a tractable empirical specification, we will assume that spouses have preferences that can be represented by Cobb-Douglas utility functions. The utility functions of the husband and the wife are written as follows:

$$u^m = \beta \ln Q_1 + (1-\beta) \ln Q_2 \tag{4}$$
$$u^f = \omega \ln l^f + (1-\omega)(\alpha \ln Q_1 + (1-\alpha) \ln Q_2), \tag{5}$$

where Q_1 (Q_2) is the composite good related to non-children's (children's) goods and l^f is the wife's leisure. Let us denote the respective prices of these public goods by P_1 and P_2. The spouse-specific preference parameters α, β and ω are between zero and one.

Since we want to let the data speak for themselves, we will not make any explicit assumption on the spouses' relative valuation of the public goods. In other words, we will not impose that, say, the wife values the child related public good relatively more than the husband. Therefore, two cases can be distinguished: if $\beta > \alpha$, then the husband values the child related public good less than the wife, while we have the reverse conclusion if $\beta \leq \alpha$.

Let us first focus on the $\beta > \alpha$ scenario. As shown below, this model is associated with three regimes like in the general model in Section 4.2. The first regime is characterized by the husband being the dictator, which implies that the household's aggregate financial resources are entirely allocated according to his preferences. The second regime is associated with split might. Given the assumption that $\beta > \alpha$, the child related public good is entirely financed by the wife, while the other public good is entirely financed by the husband. This situation corresponds to the Separate Spheres case in Browning, Chiappori and Lechene (2009). Finally, there is the regime where the wife will be the dictator. In this regime, the household's aggregate resources are allocated according to her preferences. The specific regime in which a couple will be located will depend on the wife's wage given her husband's wage and the spouses' preference parameters. Let us now explicitly characterize the three regimes when spouses have the above Cobb-Douglas utility functions. It turns out that the three regimes contain four *regions*, as the HD regime contains two distinct situations.

Husband dictator First, for the wife's wage w^f rather small (to be made precise below) the husband dictates the entire allocation over Q_1 and Q_2. This implies that:

$$Q_1 = \frac{\beta y}{P_1} \tag{6}$$
$$Q_2 = \frac{(1-\beta)y}{P_2}, \tag{7}$$

where

$$y = w^m + (1 - l^f)w^f. \tag{8}$$

Now consider the wife's labor supply decision in this situation:

$$\max_{l^f} \omega \ln l^f + (1-\omega)(\alpha \ln \left(\frac{\beta}{P_1}\right) + (1-\alpha) \ln \left(\frac{(1-\beta)}{P_2}\right) + \ln(w^m + (1-l^f)w^f)). \tag{9}$$

The first order condition for l^f can be written as

$$\frac{l^f w^f}{y} = \frac{\omega}{1-\omega}.$$

Solving for l^f, we get

$$l^f = \omega(1 + \frac{w^m}{w^f}). \tag{10}$$

Taking account of the fact that $l^f \leq 1$, the wife will not participate (i.e., $l^f = 1$) if

$$w^f \leq w^m \frac{\omega}{1-\omega}.$$

We will refer to this situation as the first region on the demand curve. In this case demand is characterized by

$$\begin{aligned} Q_1 &= \frac{\beta w^m}{P_1} \\ Q_2 &= \frac{(1-\beta)w^m}{P_2}. \end{aligned}$$

Next consider the case where $w^f \in [w^m \frac{\omega}{1-\omega}, w^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)}]$. The household's aggregate resources are now given by

$$y = (1-\omega)(w^f + w^m). \tag{11}$$

This situation will be referred to as the second region. In this case, the husband stays the dictator and the demand functions for public goods are

$$Q_1 = \frac{\beta(1-\omega)(w^f + w^m)}{P_1} \tag{12}$$

$$Q_2 = \frac{(1-\beta)(1-\omega)(w^f + w^m)}{P_2}. \tag{13}$$

Hence, we find that the husband's contribution to Q_2 is given by

$$\begin{aligned} Q_2^m &= \frac{(1-\beta)(1-\omega)(w^f + w^m) - w^f(1-l^f)}{P_2} \qquad (14) \\ &= \frac{w^m - \beta(1-\omega)(w^m + w^f)}{P_2} \end{aligned}$$

as the wife spends her entire income on Q_2. This phase stops once $Q_2^m = 0$ which happens when $w^f \geq w^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)}$.

Split might Assume that $w^f \in [w^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)}, w^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)}]$. This is the situation in which $Q_2^m = 0$ and $Q_1^f = 0$. Now we have

$$Q_1 = Q_1^m = \frac{w^m}{P_1} \tag{15}$$

$$Q_2 = Q_2^f = \frac{(1-\omega)\, w^f - \omega w^m}{P_2}. \tag{16}$$

This will last until

$$\frac{Q_2}{Q_1} = \frac{1-\alpha}{\alpha}\frac{P_1}{P_2}$$

or, equivalently,

$$w^f = w^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)}. \tag{17}$$

Wife dictator For $w^f \geq w^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)}$, the allocation of the household's aggregate resources is entirely determined by the wife's preferences:

$$Q_1 = \frac{\alpha(1-\omega)(w^f + w^m)}{P_1} \tag{18}$$

$$Q_2 = \frac{(1-\alpha)(1-\omega)(w^f + w^m)}{P_2}. \tag{19}$$

The above characterization of the three regimes applies when $\beta > \alpha$ (i.e., when the wife values the child related public good relatively more than the husband). Figure 4.1 summarizes what happens in this situation for given values of w^m, α, β and ω.[5] Note that the Husband Dictatorship regime can be divided into two regions. The first region, in which the wife decides not to work, is represented by a single allocation between the two public goods, namely the most left dot on the thick line. The second region, in which the wife does work but allocation is according to male preferences, is represented by the remainder of the thick HD-line. A similar characterization can be derived for the situation when $\beta \leq \alpha$ (i.e., when the husband values the child related good relatively more than the wife). The characterizations of the different regimes, in four regions, for

[5] The parameters in the figure are based on the data that we introduce below.

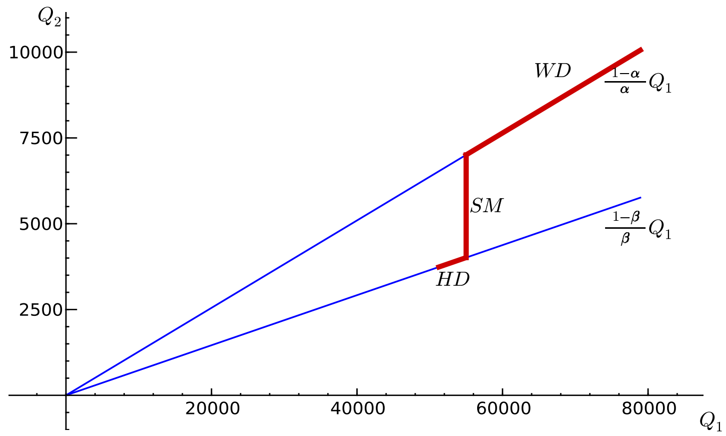


Figure 4.1: The kinky nature of demand for public goods when $\beta > \alpha$. As w^f increases demand moves along the thick line.(Here $w^m = 55.000$, $\alpha = 0.887$, $\beta = 0.932$ and $\omega = 0.191$.)

both situations are summarized as follows:

Situation 1: $\beta > \alpha$

$$l^f = \begin{cases} 1 & \texttt{if} \quad w^f \leq \frac{\omega}{1-\omega} w^m \\ \omega(1+\frac{w^m}{w^f}) & \texttt{if} \quad w^f > \frac{\omega}{1-\omega} w^m \end{cases}$$

$$Q_1 = \begin{cases} \frac{\beta w^m}{P_1} & \texttt{if} & w^f \leq \frac{\omega}{1-\omega} w^m \\ \frac{\beta(1-\omega)(w^f+w^m)}{P_1} & \texttt{if} & w^f \in [w^m \frac{\omega}{1-\omega}, w^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)}] \\ \frac{w^m}{P_1} & \texttt{if} & w^f \in [w^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)}, w^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)}] \\ \frac{\alpha(1-\omega)(w^f+w^m)}{P_1} & \texttt{if} & w^f > w^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)} \end{cases} \tag{20}$$

$$Q_2 = \begin{cases} \frac{(1-\beta) w^m}{P_2} & \texttt{if} & w^f \leq \frac{\omega}{1-\omega} w^m \\ \frac{(1-\beta)(1-\omega)(w^f+w^m)}{P_2} & \texttt{if} & w^f \in [w^m \frac{\omega}{1-\omega}, w^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)}] \\ \frac{(1-\omega)w^f - \omega w^m}{P_2} & \texttt{if} & w^f \in [w^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)}, w^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)}] \\ \frac{(1-\alpha)(1-\omega)(w^f+w^m)}{P_2} & \texttt{if} & w^f > w^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)}. \end{cases}$$

Situation 2: $\beta \leq \alpha$

$$l^f = \begin{cases} 1 & \texttt{if} \quad w^f \leq \frac{\omega}{1-\omega} w^m \\ \omega(1+\frac{w^m}{w^f}) & \texttt{if} \quad w^f > \frac{\omega}{1-\omega} w^m \end{cases}$$

$$Q_1 = \begin{cases} \frac{\beta w^m}{P_1} & \texttt{if} \quad w^f \leq \frac{\omega}{1-\omega} w^m \\ \frac{\beta(1-\omega)(w^f+w^m)}{P_1} & \texttt{if} \quad w^f \in [w^m \frac{\omega}{1-\omega}, w^m \frac{1-(1-\beta)(1-\omega)}{(1-\beta)(1-\omega)}] \\ \frac{(1-\omega)w^f-\omega w^m}{P_1} & \texttt{if} \quad w^f \in [w^m \frac{1-(1-\beta)(1-\omega)}{(1-\beta)(1-\omega)}, w^m \frac{1-(1-\alpha)(1-\omega)}{(1-\alpha)(1-\omega)}] \\ \frac{\alpha(1-\omega)(w^f+w^m)}{P_1} & \texttt{if} \quad w^f > w^m \frac{1-(1-\alpha)(1-\omega)}{(1-\alpha)(1-\omega)} \end{cases} \tag{21}$$

$$Q_2 = \begin{cases} \frac{(1-\beta)w^m}{P_2} & \texttt{if} \quad w^f \leq \frac{\omega}{1-\omega} w^m \\ \frac{(1-\beta)(1-\omega)(w^f+w^m)}{P_2} & \texttt{if} \quad w^f \in [w^m \frac{\omega}{1-\omega}, w^m \frac{1-(1-\beta)(1-\omega)}{(1-\beta)(1-\omega)}] \\ \frac{w^m}{P_2} & \texttt{if} \quad w^f \in [w^m \frac{1-(1-\beta)(1-\omega)}{(1-\beta)(1-\omega)}, w^m \frac{1-(1-\alpha)(1-\omega)}{(1-\alpha)(1-\omega)}] \\ \frac{(1-\alpha)(1-\omega)(w^f+w^m)}{P_2} & \texttt{if} \quad w^f > w^m \frac{1-(1-\alpha)(1-\omega)}{(1-\alpha)(1-\omega)}. \end{cases}$$

4.3.2 Estimation strategy

As is clear from above, the household demand system for female leisure and both public goods has a kinked nature. Figure 4.1 highlights this. Moreover, the kinks are determined endogenously since they depend on the spouses' preference parameters and relative wage rates. In what follows, we will lay down a maximum likelihood estimation strategy to obtain estimates of the unknown α, β and ω parameters. We will estimate these preference parameters separately for various demographic groups, namely for households with different numbers of children.

It is easily seen that in each regime within the situations (20) and (21) there is adding up: adding expenditures on the two composite public goods and female leisure always equals the household's full budget (i.e., $w^f l^f + \mathbf{P}'\mathbf{Q} = w^f + w^m$). This implies that, as in standard demand analysis, one of the goods in the three-good demand system may be deleted. We opt to model the wife's leisure and the child related public good.

The stochastic model specification of our model is as follows. First, we introduce optimization errors in both the observed demand for leisure and in the demand for the child related public good. Let Q_{i2} denote the demand for the child related public good in couple i and let l_i^f be the wife's leisure. These observed demands are assumed to be equal to $f_1^{r,s}\left(w_i^m, w_i^f, P_{i1}, P_{i2}\right) + \epsilon_{i1}$ and $f_2^{r,s}\left(w_i^m, w_i^f, P_{i1}, P_{i2}\right) + \epsilon_{i2}$, where $f_k^{r,1}$ (for goods k =1,2, and for regions r =1,2,3,4) correspond to the left-hand side equations in (20) and $f_k^{r,2}$ to those in (21). Optimization errors are captured by the disturbances ϵ_{ik} (k =1,2). Second, we introduce measurement error in the wife's wage rate. Let us assume that the true wage rate of the wife in couple i is observed with some additive measurement error η_i independent from the true wage. The vector containing both optimization error and measurement error $\begin{pmatrix} \epsilon_{i1} & \epsilon_{i2} & \eta_i \end{pmatrix}'$ is assumed to be drawn from a multivariate normal distribution with mean $\begin{pmatrix} 0 & 0 & 0 \end{pmatrix}'$ and a covariance matrix $\begin{pmatrix} \sigma^2_{\epsilon_1} & 0 & 0 \\ 0 & \sigma^2_{\epsilon_2} & 0 \\ 0 & 0 & \sigma^2_{\eta} \end{pmatrix}$.

Each household's contribution to the likelihood function contains two parts. This first part refers to the probability that a household is observed in one of the four regions in the demand system (i.e. in the two regions in HD, in SM or in WD) and the second part refers to the probability densities of the observed demands for leisure and for the child related good. For expositional reasons we now assume that $\beta > \alpha$. Probabilities associated with the situation where $\beta \leq \alpha$ (denoted by $P_i^{r,2}$ for $r = 1, 2, 3, 4$) can be derived in a similar way.

The probability that couple i is located before the first kink in the demand curves, i.e. in the HD regime without a working wife, can be written as $P\left(w_i^f + \eta_i \leq w_i^m \frac{\omega}{1-\omega}\right)$. As the scale of the measurement error is not identified, we normalize it by assuming that it is equal to the variance of observed wages. Given the distributional assumptions made, this implies that the above probability equals $\Phi\left(w_i^m \frac{\omega}{1-\omega} - w_i^f\right)$. In a similar way, the probabilities associated with the other kinks can be derived. In summary, we have the following probabilities:

$$\begin{aligned}
P_i^{1,1} &= P\left(w_i^f + \eta_i \leq w_i^m \frac{\omega}{1-\omega}\right) = \Phi\left(w_i^m \frac{\omega}{1-\omega} - w_i^f\right) \qquad (22)\\
P_i^{2,1} &= P\left(w_i^m \frac{\omega}{1-\omega} < w_i^f + \eta_i \leq w_i^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)}\right)\\
&= \Phi\left(w_i^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)} - w_i^f\right) - \Phi\left(w_i^m \frac{\omega}{1-\omega} - w_i^f\right)\\
P_i^{3,1} &= P\left(w_i^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)} < w_i^f + \eta_i \leq w_i^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)}\right)\\
&= \Phi\left(w_i^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)} - w_i^f\right) - \Phi\left(w_i^m \frac{1-\beta(1-\omega)}{\beta(1-\omega)} - w_i^f\right)\\
P_i^{4,1} &= 1 - \Phi\left(w_i^m \frac{1-\alpha(1-\omega)}{\alpha(1-\omega)} - w_i^f\right).
\end{aligned}$$

Besides these probabilities, the likelihood function that will be maximized also takes into account the densities of the disturbance terms in the observed leisure and observed expenditures on the child related goods. Let us introduce the binary variable d_0 which equals 1 if the starting values for the preference parameters imply that $\widehat{\beta}_0 > \widehat{\alpha}_0$ and 0 otherwise. Let us further introduce four couple-specific dummy variables d_{i0}^r (r =1,2,3,4) that indicate the region in which a couple locates on the basis of these starting values. Combining all probabilities and densities, and taking account of the above distributional assumptions, one obtains the following loglikelihood function:

$$\begin{aligned}
\log L_0 \;&= d_0 \sum_{i=1}^{n} \left(\sum_{r=1}^{4} \left\{ d_{i0}^{r,1} (\ln P_i^{r,1} + \ln(\frac{1}{\sqrt{2\pi\sigma_{\epsilon_1}^2}} \exp(-\frac{1}{2}\frac{(Q_{i2} - f_1^{r,1})^2}{\sigma_{\epsilon_1}^2}))) \right\} \right. \\
+ \quad & \left. \sum_{r=2}^{4} \left\{ d_{i0}^{r,1} (\ln(\frac{1}{\sqrt{2\pi\sigma_{\epsilon_2}^2}} \exp(-\frac{1}{2}\frac{(l_i^f - f_2^{r,1})^2}{\sigma_{\epsilon_2}^2}))) \right\} \right) \\
+ \quad & (1-d_0) \sum_{i=1}^{n} \left(\sum_{r=1}^{4} \left\{ d_{i0}^{r,2} (\ln P_i^{r,2} + \ln(\frac{1}{\sqrt{2\pi\sigma_{\epsilon_1}^2}} \exp(-\frac{1}{2}\frac{(Q_{i2} - f_1^{r,2})^2}{\sigma_{\epsilon_1}^2}))) \right\} \right. \\
+ \quad & \left. \sum_{r=2}^{4} \left\{ d_{i0}^{r,2} (\ln(\frac{1}{\sqrt{2\pi\sigma_{\epsilon_2}^2}} \exp(-\frac{1}{2}\frac{(l_i^f - f_2^{r,2})^2}{\sigma_{\epsilon_2}^2}))) \right\} \right)
\end{aligned}$$

This loglikelihood function is optimized using an iterative procedure. For given starting values for the unknown preference parameters, denoted by $\widehat{\alpha}_0$, $\widehat{\beta}_0$ and $\widehat{\omega}_0$, the (couple-specific) location of the kinks in the demand system can be calculated. Maximizing the loglikelihood function using these initial locations obtains the first-round estimates of the preference parameters ($\widehat{\alpha}_1$, $\widehat{\beta}_1$ and $\widehat{\omega}_1$). On the basis of these estimates, updated couple-specific locations of the kinks in the demand system can be derived. Furthermore, updated locations of couples on the demand system can be determined in a similar way as above. This also results in updated binary variables d_1 and d_{i1}^r that are defined in a similar way as d_0 and d_{i0}^r, while updated probabilities (22) can be calculated. Finally, all these variables are brought together in the updated loglikelihood function $\log L_1$, which has the same form as $\log L_0$ but this time constructed on the basis of the estimates $\widehat{\alpha}_1$, $\widehat{\beta}_1$ and $\widehat{\omega}_1$. The loglikelihood function $\log L_1$ is again maximized to obtain new estimates $\widehat{\alpha}_2$, $\widehat{\beta}_2$ and $\widehat{\omega}_2$. The iteration procedure is repeated until convergence of the parameter estimates. I.e., until the final maximum likelihood estimates equal the second to last ones which were used to construct the boundaries and probabilities.

Note that the parameters α, β and ω have to be between zero and one to have a coherent system. This requirement turns out to be satisfied in the application below. Note as well that the loglikelihood function is continuous in the parameters whenever for a sequence of iterations

β is always larger than or always lower or equal than α. Given that convergence of the iteration procedure implies the same ordering of β and α when the parameters are about to converge, potential discontinuities do not pose any problem here. As a final remark, we should stress that several local maxima may exist, which is taken into account in the estimation process.

4.4 Data

For the empirical analysis in this chapter we have used the Consumer Expenditure Survey (CEX) collected by the Bureau of Labor Statistics of the U.S. Department of Labor. Like most authors in the consumption literature that use the CEX (e.g., Deaton and Paxson, 1994, Attanasio and Weber, 1995, and Attanasio and Davis, 1996) we have compiled a dataset from the quarterly Interview Survey (IS) that collects data through a recall questionnaire rather than from the biweekly Diary Survey (DS) that collects data through a daily purchase questionnaire.[6] This was done for three reasons. First, the IS contains more observations. Second, the IS was especially designed to collect data on major items of expense. The most substantial elements of child expenses, such as tuition, classify as such. Third, the expenditure component directly related to children is larger - both in terms of absolute and relative value - in the IS than in the DS.

Our CEX dataset comprises of quarterly household observations from the first quarter of 2005 until the first quarter of 2008. The IS has a rotating panel setup in which one household is interviewed a maximum of four times. This means that 89 percent of households in the CEX enters the dataset more than once and 54 percent of households are observed for a full year. We aggregate all family expenses to the quarterly level so that we can also keep the observations of those families that are not observed for a full year. There are 90,955 observations in our waves of the CEX.

We construct a sample that is best suited for our structural approach. It includes observations of all married couples in which the husband works at least 25 hours a week for at least 40 weeks a year and in which neither of the spouses is enrolled in a college or university nor is self-employed. Those families with one or both spouses currently attending college or university are excluded because we would like to be sure that tuition ex-

[6] For a detailed comparison of the Interview Survey and the Diary Survey see Battistin (2004).

Table 4.1: Summary statistics for couples

	One child	Two children	Three children
Number of households	2,596	3,312	1,168
Mean direct child expenditures	\$2,212	\$2,517	\$3,106
Median direct child expenditures	\$800	\$1,270	\$1,588
Mean hourly wage husband	\$25.29	\$28.63	\$30.63
Median hourly wage husband	\$21.06	\$23.42	\$23.67
Mean hourly wage wife	\$17.39	\$18.23	\$17.91
Median hourly wage wife	\$14.92	\$15.44	\$14.58
Wife not in labor force	23.4%	29.8%	34.5%

penses can be classified as direct child expenses. Households that included other adults were dropped. Furthermore, those households in which male wages or female wages fell in the 1st or 99th percentile of the male income distribution were disregarded. Also, those families in which the wife worked more hours than the average male were not included in the sample. Finally, we divide our sample on the basis of the number of children under the age of 18 living in the household. In order for the direct children's expenses to be comparable on an absolute level we differentiate between families with one, two or three children. Because of the many constraints we impose on the data, our three samples are considerably smaller than the total CEX. They consist of 2,596, 3,312 and 1,168 observations for one, two and three child families respectively. Table 4.1 displays summary statistics for the three groups.

For the estimation of the structural parameters in our model we need four variables: l^f, the wife's leisure; Q_2, which contains an estimate of total yearly child expenses per household; w^m, which is the husband's total yearly net labor income; and w^f, which is the wife's full budget being her net wage rate times the maximum number of hours (normalized to one) she could have worked. Not all child expenses are separately observed in the CEX. As we do not know who consumes what it is unclear whether expenses such as sweets or cinema tickets were intended for children or for adults. For some categories it can however be ruled out that the goods were intended for adults. These expense categories include school meals, infant furniture, boys apparel, girls apparel, boys and girls footwear, infants apparel, toys, educational books and supplies, and elementary school, high school and college tuition and fees. Note that we have quarterly expenditure information and that this amount is then multiplied by four to obtain an estimate of yearly expenses in the household.

Price information is also obtained from the Bureau of Labor Statistics. Seasonally adjusted Urban Consumer Price Indices that were reported per calendar month have been used. This means that we do not take regional price variation into account. We compute a monthly CPI for child expenses by averaging the available separate product CPIs and taking the weights that these products have in total child expenses into account. Separate CPIs were available for all components of Q_2, except for baby furniture and school meals. As the quarterly IS interviews take place during all 12 months of the year, our dataset contains 40 monthly values of Q_2's CPI. Because households report expenditures over the previous three months, we have chosen to assign the CPI that pertains to the third month prior to the interview. Cumulative inflation on our basket of child expenses has been 11.9 percent from November 2004 until February 2008. This amounts to a yearly average inflation level of 3.3 percent on Q_2.

The CEX documents total yearly household net income as well as individuals' gross labor incomes. We have employed two alternative ways to estimate the husband's net labor income, w^m. If the household received no non-labor income we derived it as the percentage of household's net income corresponding to his gross labor income share. If the household did receive some non-labor income, we computed his net wage on the basis of his gross wage using an estimated spline relation between gross and net income in the households that did not receive any non-labor income.

Potential female wage income, w^f, is based on the same gross to net conversion as male wage income. The hourly wage rate is moreover computed for all women that are employed but imputed for all women that are currently unemployed or out of the labor force. To do the latter, we estimate a Heckman selection model in which the overidentifying variable is the number of children each women has (the wage imputation is done on the total CEX dataset). The Heckman model corrects for the wife's education level, age group, her state of residence, and for the year of the survey. The potential female wage income, w^f, is then computed by multiplying the (imputed) wage rate times the average weekly number of hours worked by men times 52 (the number of weeks). The female wage rate was imputed in respectively 23.4%, 29.8% and 34.5% of the households in our samples. In other words, these are the percentages of mothers that did not participate in the labor market.

Table 4.2: Summary statistics for singles

	One child	Two children	Three children
Single fathers	212	124	29
Single mothers	965	591	193
Mean child expenditures men	\$1,181	\$1,786	\$2,565
Median child expenditures men	\$528	\$1,264	\$1,696
Mean hourly wage men	\$20.30	\$24.67	\$19.56
Mean child expenditures women	\$1,200	\$1,731	\$1,758
Median child expenditures women	\$600	\$1,080	\$948
Mean hourly wage women	\$15.20	\$15.04	\$14.49

4.5 Estimation results

4.5.1 Stylized facts about child expenses of singles

Using the CEX we also construct similar datasets of single - fulltime employed, non-college going - parents with one, two or three children in order to obtain a first impression of male and female preferences for child expenses. As can be seen in Table 4.2 it turns out that if we do not control for income, single mothers of one child spend on average more on their children than single fathers. Given that the hourly wage rate of the single women is on average lower than that of single men, this is remarkable. In the two and three children samples, the single fathers spend unconditionally more.

If we assume that labor supply of single mothers and single fathers is exogenous[7] (as it is assumed for husbands), then we can estimate the Cobb-Douglas model with a gender taste shifter. Singles are necessarily dictators in this set-up - public good demand is according to their preferences alone. Hence, we estimate a β^f for single women and a β^m for single men using their respective labor incomes. Table 4.3 presents the estimates. It can be seen that in all groups, the female taste for spending on children is higher than the male taste. The difference between the two estimates of β is however only significant at a five percent level in the two children group. The small number of single fathers with three children could account for the lack of significance in their sample.

[7] The selection of parents is such that they should all work 25 hours or more a week, for at least 40 weeks a year.

Table 4.3: Estimation results for singles

	One child		Two children		Three children	
	Est.	St. err.	Est.	St. err.	Est.	St. err.
β^f	0.964	0.003	0.949	0.009	0.939	0.013
β^m	0.974	0.007	0.971	0.004	0.949	0.015

4.5.2 Results for couples with children

Let us now focus on the estimation results for couples with respectively one, two and three child(ren). Preference parameter estimates were obtained by means of the estimation strategy outlined above. The iterated maximum likelihood method worked well: estimates were obtained after only a few iterations. Moreover, the unconstrained preference parameters were between zero and one as theory predicts they should be.[8]

Table 4.4 summarizes the estimation results for the three sets of couples. It is clear from the results that on average wives value the child related public good more than their husbands. This is in line with the results obtained for singles. The table further shows that, ceteris paribus, the marginal utility of the child related public good increases with the number of children. This is the case for both husbands and wives. This makes sense since the data show that the aggregate expenditures on children increase with their number.[9] Also the wife's marginal utility of leisure increases with the number of children. This is moreover reflected by the data given that women on average work less when they have more children.

As is clear from Section 4.3, our non-cooperative model nests the standard unitary model. If α equals β, then the household behaves as if it consists of a single decision maker with preferences that are equal to those of the wife (given that preferences also depend on the wife's leisure). We tested the null hypothesis that α is equal to β for the three sets of couples by means of a Wald test. Interestingly, the null hypothesis is strongly rejected for the couples with two and three children (p-values are close to 0). Their observed behavior can clearly not be captured by a single decision maker's rational preferences. Note though that the null hypothesis could not be rejected for couples with one child (p-value equals 0.47). The ex-

[8]A few local maxima were found. We retained the highest maximum. The results discussed below are qualitatively robust for the different maxima.

[9]This may be remarkable given the public good connotation of the child related expenditures. Still, for a given family constellation, this does not rule out that children's expenses are public in the sense that they appear in both spouses' utility functions.

Table 4.4: Estimation results for couples

	One child		Two children		Three children	
	Est.	St. err.	Est.	St. err.	Est.	St. err.
α	0.922	0.024	0.887	0.009	0.823	0.020
β	0.966	0.040	0.932	0.006	0.894	0.012
ω	0.140	0.006	0.191	0.005	0.222	0.010

Table 4.5: Proportion of couples in the different regimes

	One child	Two children	Three children
Husband dictator	6.82	19.62	32.11
Split might	5.20	7.28	11.30
Wife dictator	87.98	73.10	56.59

penditures on child related goods are relatively low for these households. It could well be the case that the data therefore do not allow for a sharp distinction between the spouses' preference parameters.

On the basis of the above preference parameters, the expected number of couples in each of the three regimes can be calculated. Since $\widehat{\beta} > \widehat{\alpha}$, we derive these numbers on the basis of the right-hand sides of equation (20). The obtained proportions are presented in Table 4.5. Interestingly, all possible regimes contain a significant proportion of the couples in the sample. Moreover, it turns out that most of the households behave as if the wife was the dictator in the household. More specifically, for these families, the allocation of the household's aggregate resources is according to the wife's preferences. Note though that the proportion of households with the wife acting as the dictator decreases with the number of children.

A partial goodness-of-fit test exists, as we know that at least the households in which the wife does not work should be in a Husband Dictatorship according to our model. Table 4.1 presented the relevant percentages for one, two and three-child households. Unfortunately, these percentages are in all three samples larger than the estimated percentages of HD-households. Further research is necessary to improve the goodness-of-fit of our estimates.

4.6 Conclusion

In this chapter, we model the consumption and labor supply behavior of a couple in a non-cooperative setting by adopting a Nash approach. Using minimal assumptions, we prove that demand for public goods is defined by only three regimes. Demand for public goods is either determined by the preferences of one of the partners only (Husband Dictatorship or Wife Dictatorship), or by both spouses having a say on the allocation of income to public goods (Split Might). The particular regime in which a couple locates is shown to depend on the spouses' relative wage rates, which resembles the endogenous regimes (depending on exogenous individual incomes) in Browning, Chiappori and Lechene (2009).

By imposing more structure on the general model, we can derive testable implications on observed demand for public goods and labor supply that allow testing the model against the standard unitary model where a couple behaves as a single decision maker. The model is applied to a sample of couples drawn from the U.S. Consumer Expenditure Survey (CEX) whereby we focus on expenses on children's goods that act as a public good in the spouses' preferences. We find that for couples with two or three children the standard unitary model is strongly rejected in favor of our non-cooperative model. Women apparently like to spend more on goods for their children than men, which would explain findings in the literature that indicate that children benefit more when household resources are owned by women (e.g., Lundberg, Pollak and Wales, 1997, and Duflo, 2003). Using the estimated preference parameters, we can divide households into dictatorship and split might regimes. It turns out that 73% of two-child and 57% of three-child couples spend according to the wife's preferences.

A drawback of the present study is that we assume that the preferences of all husbands and all wives are equal, and that differences in what we observe are only generated by differences in potential wages and by optimization and measurement errors. Given the complexity of our current estimations, this was a necessary first attempt. However, a natural extension of this study will be to allow for unobserved heterogeneity in the preference parameters so that we do not have to assume equal preferences. We are currently working on a random preference version of our model.

4.A Appendix

Proof of Lemma 1
We define for each spouse $i = f, m$

$$\begin{aligned} dl^i &= \varepsilon^i \\ dq_k^i &= -\frac{w^i}{p_k}\varepsilon \\ u^i(\varepsilon^i) &= u^i(l^i + dl^i, \mathbf{q}^i + dq_k^i \iota_k, \mathbf{Q}) \end{aligned}$$

where ι_k denotes a vector which equals 1 (one) at position k and is zero everywhere else. Note that dl^i and dq_k^i are defined in such a way that $\varepsilon^i \neq 0$ is feasible in terms of the budget restriction.

It follows that

$$u^{i\prime}(\varepsilon^i) = u_l^i - u_{q_k}^i \frac{w^i}{p_k}. \tag{23}$$

First, consider the case where $w^i > 0$. Then we prove by contradiction that the equality in the lemma holds. Note that $\frac{u_l^i}{u_{q_k}^i} > \frac{w^i}{p_k}$ implies that utility increases with ε^i contradicting equation (3). We need to be careful though as $\varepsilon^i > 0$ is not possible with $l^i = 1$. However at $l^i = 1$ we have $q_k^i = 0$ (as there is no income to spend on private goods) and thus $u_{q_k}^i = +\infty$ (by assumption 1). Since u_l^i is finite at $l^i = 1$ (by assumption 3), we cannot have $\frac{u_l^i}{u_{q_k}^i} > \frac{w^i}{p_k}$. Further, note that $\frac{u_l^i}{u_{q_k}^i} < \frac{w^i}{p_k}$ implies $\varepsilon^i < 0$ would raise utility again contradicting equation (3). It is not possible to have $\varepsilon^i < 0$ at $l^i = 0$. But due to assumption 1 we cannot have $\frac{u_l^i}{u_{q_k}^i} < \frac{w^i}{p_k}$ at $l^i = 0$.

Second, consider $w^i = 0$. Then we have $l^i = 1, q_k^i = 0$. Due to assumption 1 and 3 we then have

$$\frac{u_l^i}{u_{q_k}^i} = 0 = \frac{w^i}{p_k}$$

and the equality in the lemma holds. *Q.E.D.*

Proof of proposition 1
The proposition implies that we need to rule out two cases. First, we need to rule out that $u_{Q_k}^i > \lambda^i \mathbf{P}_k$ for any partner $i = f, m$ and any public good Q_k. Second, we need to rule out that both $u_{\mathbf{Q}}^f = \lambda^f \mathbf{P}$ and $u_{\mathbf{Q}}^m = \lambda^m \mathbf{P}$.

Suppose (by contradiction) that $u_{Q_k}^i > \lambda^i \mathbf{P}_k$. Since we assumed $w^m > 0$, the husband will contribute to every public good (if the wife does not contribute to any public good). If $w^f > 0$ the wife can contribute to public goods as well. In either case we have $Q_k > 0$ for each public good Q_k and hence $u^i_{Q_k}$ is finite for both partners. Given the Nash assumption that Q_k^{j*} for $j \neq i$ is given, partner i can raise utility by increasing Q_k^i which contradicts equation (3).

Second, assume (by contradiction) that both $u^f_{\mathbf{Q}} = \lambda^f \mathbf{P}$ and $u^m_{\mathbf{Q}} = \lambda^m \mathbf{P}$. This would imply

$$\frac{u^f_{Q_k}}{u^f_{Q_{k'}}} = \frac{P_k}{P_{k'}} = \frac{u^m_{Q_k}}{u^m_{Q_{k'}}}$$

for each pair of public goods $Q_k, Q_{k'}$. However, this contradicts assumption 5. *Q.E.D.*

Proof of proposition 2

Consider $w^m > 0$ and $w^f = 0$. Hence $\mathbf{Q}^f = 0$ as the wife earns no income. Given assumption 1 it is optimal for the husband to contribute to each public good Q_k. Hence $u^f_{Q_k}$ is finite for each public good Q_k. Since $w^f = 0$ implies that $\mathbf{q}^f = 0$, we have $\lambda^f = +\infty$. Hence, indeed we are in regime HD. By continuity this also holds for $w^f > 0$ close enough to zero.

The proof of the existence of ρ_1 is done in the same way but then starting from $w^f > 0, w^m = 0$. This gives us regime WD. By continuity we are also in regime WD for $w^m > 0$ close enough to zero. *Q.E.D.*

Chapter 5

Better Protected, Better Paid: Evidence on how Employment Protection Affects Wages[1]

This chapter is forthcoming as a paper in Labour Economics (Van der Wiel, 2009).

5.1 Introduction

Employment protection legislation (EPL) has been on the (European) political agenda for several decades and continues to be widely debated. On the one hand employers complain that they are incapable of adapting to economic circumstances because of the high cost of firing employees and on the other hand vulnerable groups of employees complain that the firing risk they face is too large. Policy-makers in the meantime have to decide upon an optimal level of protection. In order to make such decisions politicians need to know how different types of employment protection affect labor market outcomes. In order to provide such answers, a large economic literature on employment protection has been developed.

The rationale for installing employment protection legislation is clearly not to manipulate wages. In practice EPL might however have a considerable impact on them. In fact, this impact is a central issue in the academic EPL literature as it is closely related to the effect of employment

[1]Thanks to Arthur van Soest, Jan van Ours, Frederic Vermeulen, Gerard Pfann, Anja Deelen, Bas ter Weel, Davida van Praag, two anonymous referees and several seminar participants for useful comments and suggestions.

protection on employment levels[2]. How these two labor market outcomes are theoretically related depends on the type of employment protection (EP) under investigation, on the relevant labor market institutions and on which further assumptions the respective researcher makes. When one for example considers a competitive economy with individual wage-setting as in Lazear (1990), transferable firing costs such as a severance payment will be shifted to the worker at the onset of an employment contract. In Lazear's model, wages will adjust downwards to take the future firing costs into account and employment will not be affected. Alternatively, one could consider central wage bargaining by a union who cares for incumbent workers only and a general turnover cost associated with dismissals as in Bertola (1990). According to insider/outsider theory, employment protection will then induce unions to bargain for higher than competitive wages. As a result, firms will fire (and hire) fewer workers.

Empirically, only suggestive evidence of the true relationship between employment protection and wages exists. This is partly because most empirical research has analyzed macro-data and composite employment protection indices, which introduces comparability issues and confounding factors into the estimations. This chapter seeks to use the Dutch Socio-Economic Panel (SEP) dataset to empirically establish the causal effect of a specific type of employment protection, namely the employer's term of notice (ToN), on the wage level of employees. The term of notice is defined here as the amount of time an employer has to notify an employee in advance of her upcoming dismissal. Specific groups, notably older workers, are often protected by a longer term of notice. The term of notice is usually ignored in the empirical employment protection literature. This is unfortunate as a long term of notice corresponds to a large number of obligatory wage payments and hence is a substantial firing cost to the employer.

The Dutch labor market has two relevant features that makes it suitable as a research ground for this chapter. First, for each worker, it is possible to calculate the 'dormant' term of notice. I define 'dormant' here as the notice period that would apply in case the employer would soon want to fire the worker. The term of notice can be calculated because the law sets out a formula to do so, which generates variation in the degree of employment protection for workers of different ages and tenure. Second,

[2] I do not discuss the empirical effect of employment protection on employment in this chapter. In order to do a similar fixed effects analysis for employment rather than wages, one would need a dataset containing a substantial number of individuals that are observed over time as being both fired and non-fired. The Dutch socio-economic panel unfortunately features too little of these observations.

in 1999 the legal formula to calculate the term of notice was changed. In short, low-tenured workers of all ages benefited from the introduction of the law on flexibility and security as their employer's term of notice increased while older high-tenured workers experienced a shorter term of notice after the 1st of January 1999. This chapter will exploit this exogenous policy change to answer the causality question and to separate the age, tenure and term of notice effects on wages.

The remainder of this chapter presents evidence of a strong positive causal effect of the term of notice on wages. Each additional month increases wages by three percent. It is also demonstrated that regressions that ignore the multi-collinearity issues involved can overestimate this effect. The theoretical literature provides two plausible reasons for a positive effect of employment protection on wages. First, insider/outsider theory suggests that the bargaining position of insiders is enhanced by better protection when unions only represent incumbent workers. In wage negotiations this could then drive wages up. Second, employment protection creates more incentives for firms and workers to invest in match-specific human capital. These investments will then lead to higher productivity and higher wages. Longitudinal data reflecting individual productivity in all types of jobs would be necessary to ultimately distinguish between the alternative explanations. Unfortunately, this is not available in The Netherlands. Using the alternative data at hand, I do present suggestive evidence that invalidates the investment argument.

This chapter continues as follows. Section 5.2 discusses the most important findings in the employment protection literature on wages. Section 5.3 then lays out the term of notice regulations in The Netherlands and the changes introduced in the law of flexibility and security ('Flexwet'). The empirical strategy is explained in Section 5.4. The utilized data are discussed in Section 5.5 and results will be presented in Section 5.6. Section 5.7 concludes.

5.2 Literature

5.2.1 Theoretical literature

Employment protection, i.e. firing costs, comprises of two elements: *taxes* to be paid outside the job-worker pair and *transfers* from the firm to the worker (Garibaldi and Violante, 2005, p.799). Taxes are broadly defined here: any kind of legal costs associated with layoffs fall under the first type. The term of notice and the severance pay however fall under the second type of employment protection. The length of the term of notice, or the number of additional wage payments, partly determines the size of the transfer firing costs. An extensive theoretical literature looks into the effects of employment protection or firing costs on the labor market. Concerning wages, the predictions of what employment protection actually does vary greatly. Many of these differences result from considering different types of employment protection and different types of labor market institutions. Some of these views are explained in this section, although I do not claim to be exhaustive.

In the employment protection literature (e.g. Oi, 1962, and Bentolila and Bertola, 1990), whether a profit-maximizing firm would like to fire an individual worker primarily depends on the wage of the worker and her match-specific productivity[3]. When the difference between these entities is negative, an employer loses money and she will consider firing the employee. It is costly however to adjust the number of employees downward because of associated firing costs. When wages are fixed and whenever there is a possibility that in the nearby future the worker's productivity could increase, higher firing costs such as a longer term of notice then lower the propensity to fire (and hire) a worker.

Wages are exogenous in the greatest part of this literature. Notable exceptions to this will be discussed here. Bertola (1990) draws from the insider/outsider literature. See Lindbeck and Snower (2001) for an overview. In short, the theory divides the labor market into insiders - incumbent workers who benefit from employment protection - and outsiders - those who do not benefit such as temporary workers and the unemployed. Because it is difficult to get rid of them, insiders have some bargaining power in the wage process and hence demand higher than competitive wages (for a formal model, see Lindbeck and Snower, 1986).

Bertola looks at how employment protection affects wages under different wage setting institutions. First, he investigates labor demand and endogenous wages when wage negotiations take place at the individual level. He finds that under certain assumptions total received wages might not be affected, although outsiders might offer to work for a very low wage in order to become an insider and insiders might afterwards rise wage demands above the competitive level. Second, Bertola assesses wages when there is a wage setting union that cares for everyone in the labor market. He concludes that in this set-up lifetime wages would also remain unaffected. Only in the instance of unions who solely represent working members does employment protection increase total labor income for in-

[3]This chapter considers individual lay-offs that have a financial firing rationale.

siders indefinitely.

Garibaldi and Violante (2005) exploit the idea that a country's wage-setting institutions influence the effect of employment protection on wages in a search and matching framework. The authors built a model with endogenous wage setting behavior by a monopolistic union. Garibaldi and Violante stress that in such a setting the introduction of an exogenous firing cost has two opposing effects on the workers' desired wage level: workers would like to have a higher wage (*the income effect*) but do not enjoy the accompanying higher probability to get fired (*the job security effect*). Whenever the elasticity of the firm's firing probability to wages is low enough, workers will demand higher wages when they are better protected.

Lazear (1990) wrote an influential paper on employment protection, arguing that firing costs do not necessarily affect hirings and firings. He reasons that in a flexible labor market, in the absence of contract and market restrictions, *transfer* employment protection such as the term of notice could be undone by efficient wage setting behavior between workers and firms. He predicts that in a competitive economy with decentralized wage setting, firing costs drive wages down, up to the point where the severance pay and the wages paid during the term of notice can be seen as a delayed payment. Note that Lazear predicts wages to go down at the onset of an employment contract. Pissarides (2001) also suggests a negative effect of employment protection on wages, but argues from the workers point of view (like Bertola and Rogerson, 1997). In his search and matching model the term of notice is endogenous and generates lower wages because risk averse workers accept a lower income during the productive period of a job, in order to receive a higher income during unproductive times.

Because most of the relevant literature thinks about workers as having fixed or at least exogenous productivity, it often ignores another possible route through which employment protection positively affects wages. More employment protection namely enhances the incentives for a firm to invest in a worker and for a worker to invest in firm-specific human capital. Nickell and Layard (1999) briefly describe this mechanism. These human capital investments could pay off in terms of higher productivity and higher wages. Arulampalam, Booth and Bryan (2004) present some indirect empirical evidence that employment protection does increase training of employees. Using a European dataset, they find that those on fixed term contracts take up less training than those on permanent contracts.

While certain theories thus suggest the term of notice could increase wages, others argue it could decrease wages. The most appropriate theories for this empirical study are the ones involving some market imperfections, moderate centralized wage-setting and exogenous employment protection. Lazear's argument is hence not likely to hold in the Dutch economy, in which employers organizations and a small number of labor unions negotiate over wages per industry. For more information on wage setting in The Netherlands see Wallerstein, Golden and Lange (1997).

5.2.2 Empirical literature

The empirical literature has experienced difficulties in establishing a clear relationship between firing costs and wages. Some of the papers discussed in the previous subsection do attempt to present empirical evidence of their models. However, the authors typically only provide suggestive evidence of their theories, mainly because most of them use macro-data and aggregate indices of employment protection that are hard to compare. The estimates are furthermore troubled by confounding factors. Contrary to what his theoretical model predicts, Bertola (1990) for example presents some evidence that the productivity wage gap is actually lower in countries with stricter employment protection. The empirical wage setting literature in its turn often ignores employment protection as it is so hard to quantify. See for example the establishment-level study by Blanchflower, Oswald and Garrett (1990) and the cross-country study of industry wage differentials by Holmlund and Zetterberg (1991). Both papers do suggest substantial insider wage gains.

An interesting firm-level study by Autor, Kerr and Kugler (2007) includes employment protection - i.e. *tax* employment protection - explicitly. The paper exploits U.S. state variation in the adoption of wrongful-discharge protections in order to study firm-level productivity differences. The authors find that the introduction of these laws coincided with a rise in capital investment, non-production worker employment and hence measured a labor productivity increase. Another firm-level analysis was published by Martins (2009), who analyzes a Portuguese policy change in *tax* employment protection that favored firms with twenty or less workers. In 1989 the strict Portuguese rules and regulations considering layoffs were considerably softened, and more so for the smaller firms. Using a large administrative dataset that links employers to employees he finds, among other things, that after the policy change average wages in the smaller firms fell more than in the larger firms. This suggests that in a highly regulated labor market the better protected workers earn more.

The use of micro-data in the empirical employment protection literature is limited. This is unfortunate as micro-data is often a prerequisite for identifying causality and as even in highly regulated economies wages are very heterogeneous across workers of different ages and tenure in the same firms. I am aware of only one micro study on employment protection and wages, which is a paper by Leonardi and Pica (2007). The authors empirically analyze the effect of severance payments on male wages by exploiting an Italian policy change that introduced severance payments for unjust dismissals for firms with less than fifteen employees. This policy change is explained in more detail in Kugler and Pica (2008). Their paper, like mine, thus analyzes *transfer* employment protection rather than legal protection such as in Martins (2009). Leonardi and Pica apply a regression discontinuity design, with the discontinuity being the number of employees, to study entry wages and the tenure wage profile. They use individual wage information from an administrative employers dataset from the Veneto region in Northern Italy. Contrary to their theoretical predictions, the authors find no causal effect of severance payments on entry wages. They do find that the average returns to tenure of previously dismissed workers declined by three percent in the smaller firms, relative to larger firms that did not experience an increase in employment protection. Leonardi and Pica interpret this as partial evidence for Lazear's argument that government-mandated employment protection can be shifted to employees if any employment contract is allowed. Using a Dutch dataset of individuals of all tenures and backgrounds, I find an opposite average effect of employment protection on wages, namely a strongly significant positive one. Additional research should teach us which specific characteristics of the analyzed environments generate these differences.

5.3 Term of notice

There are two paths to dismissal in The Netherlands, and the legal term of notice only applies to the labor office path. This is a relatively slow route that does not require severance payments. The labor office has the discretion to refuse an application, but only does so in a small percent of the cases (i.e. five percent in 2002). Mainly individual lay-offs in small- and medium sized firms and collective lay-offs are dealt with by the labor office, but the path is open to all employers. The other route, through court, is faster and involves substantial severance payments[4]. The cantonal court judge in principle always allows a lay-off but adapts the required severance payment to the specific firing rationale (i.e. a worker receives more if the employer could have done more to prevent the lay-off)[5]. An employer is free to choose a dismissal path and the labor office is thus a credible threat to all workers. Furthermore, nothing changed in the court procedure over the analyzed period. This means that the employer's term of notice and the associated policy change are relevant to all employees on a permanent contract.

The policy change that will be exploited in the fixed effects method below is the introduction of the law on flexibility and security (the 'Flexwet') in The Netherlands on January 1st, 1999. This law intended to diminish differences in the labor market between temporary and permanent workers. On the one hand, employees with temporary contracts received better legal protection than before. On the other hand, employees with permanent contracts lost some rights. The main change for tenured employees was the adaptation and simplification of the legal formula for the term of notice. See Heerma van Voss (1998) and Smitskam and Kronenburg-Willems (2000) for a detailed description of the new law. The calculation of the term of notice before the 1st of January 1999 (old-ToN) was as follows. For every year of tenure an employer had to add a week to the notice period, with a maximum of thirteen weeks. On top of this, workers received an extra week of notice for each year they had worked while being forty-five or older, also with a maximum of thirteen weeks. Note that tenure - which determines the term of notice - is legally defined as the amount of time that has passed since one started working for a certain employer.

The law introduced on the 1st of January 1999 consists of two elements: a new term of notice formula for newly-hired employees and a transitory arrangement for workers that were already employed. The new formula does not depend on age[6] and accommodates fewer term of notice possibilities. For workers employed in between zero and four years

[4]Figure 5.A.1 in the appendix shows the ratio of lay-offs through the labor office over the total number of lay-offs. A constant percentage of about fifty percent of lay-offs are handled by the labor office. A small increase in the ratio can be seen in 1999, the year the legal term of notice was changed. It can be argued that the labor office procedure became more profitable that year because the term of notice went down for the majority of workers.

[5]In a limited amount of cases a cantonal court judge refuses to end the labor contract, for example when a sick worker is dismissed because of her illness.

[6]All age related components of Dutch employment laws are supposed to be phased out in order to satisfy European anti-discrimination treaties.

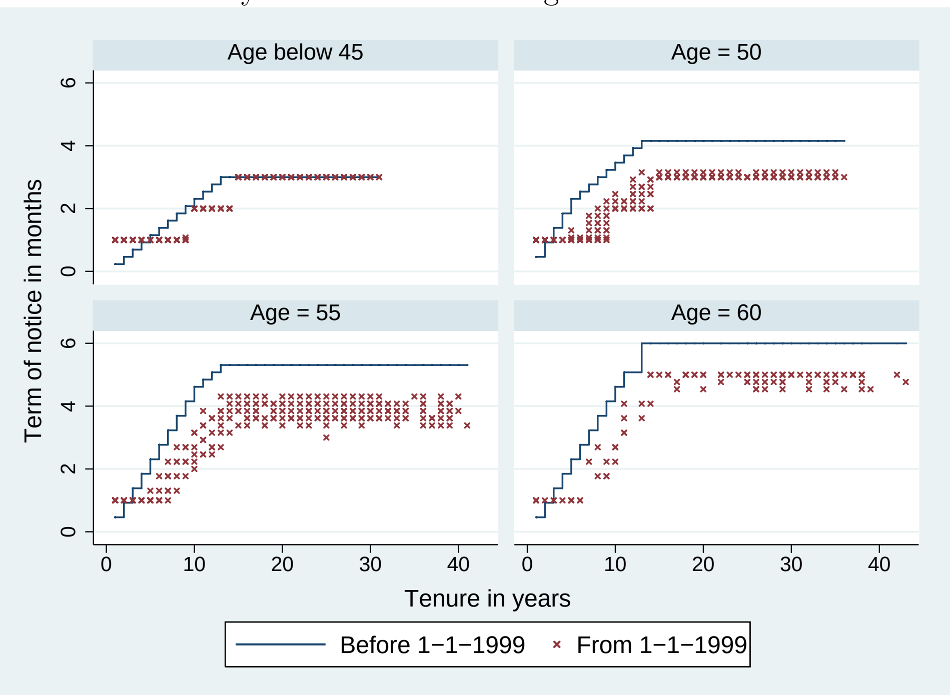
employers face a term of notice of one month. When a worker is employed between five and nine years, her employer will have to notify her two months in advance. If a worker's tenure is between ten and fourteen years, her employer will have to notify her three months in advance. Any tenure longer than fourteen years results in a term of notice of four months.

The 'Flexwet' included an important transitory arrangement for those workers that were already employed by the 1st of January 1999. For these employees, the employer had to calculate both the old term of notice for the employee's tenure and age on the 1st of January 1999 and the new term of notice for the employee's tenure and age at the time of firing. The longest notice period of the two applied. This transitory arrangement was agreed upon so that no workers would suffer a large sudden decline in EP. Another feature of the new scheme was that for both newly hired workers and those in the transitory scheme, the law allows employers to deduct one month of the notice period if they have waited to end the labor agreement until after they have received permission from the labor office. The minimum term of notice should however remain one month, so that the deduction only applies to workers with a tenure above four years. This subtraction is granted in almost all relevant cases (i.e. ninety-six percent in 2002 according to the Ministry of Social Affairs and Employment). In my 'dormant' term of notice calculations I assume that this permission is always granted. Note that I define 'dormant' here as the notice period that would apply in case the employer would soon want to fire the worker. Hence, the introduction of the 'Flexwet' generated a discrete change in EP for many workers at the beginning of 1999, and the transitory scheme ensured a further gradual move towards the new scheme.

The relationship between the old and new term of notice thus depends on age and tenure of the worker and on calendar time. More specifically, besides on tenure and time, it depends on tenure obtained from the age of 45 onwards. Figure 5.1 shows the old- and new-ToN for the respondents in the entire SEP dataset over tenure and selected ages. As the term of notice will be identical for all employees under 45, there is one panel representing them. Additionally, three panels display the relationship between tenure and the term of notice for all 50-, 55-, and 60-year old workers.

As can be seen in the below-45-panel, the old term of notice went up by a week each tenure year and had a maximum of thirteen weeks. How the policy change affected different workers depends on their respective tenure. The new term of notice turns out to be longer than the old term of notice for those employees with a tenure below five years. Young workers

Figure 5.1: Employer's Term of Notice before and after the introduction of the 'Flexwet' by tenure for selected ages of workers



with a tenure from six to fourteen years experience a new term of notice that is shorter than the old-ToN. For young workers with a tenure of both five and a tenure above fourteen years the old and new term of notice are exactly the same.

A different picture emerges for older workers. As can be seen in the panels for the 50-, 55- and 60-year old workers, the old term of notice went up by two weeks each tenure year for at least some of the years and the maximum term of notice was longer than thirteen weeks. The difference between the three panels is determined by how much older one is than 44 as this explains the number of two-week-years in the graph. For employees with a tenure below three years the new term of notice is longer than before. Those who have been working for their employer for exactly three years face the same term of notice before and after the introduction of the 'Flexwet'. The term of notice of workers with a tenure of four years and over is shortened in the new scheme. Wherever there are multiple terms of notice visible for an age-tenure combination this is because of the transition scheme. This arrangement ensures that the term of notice moves closer to the new calculation as the evaluation date is further from the 1st of January 1999. So, the employers of two workers with identical tenures (say fifteen years) and age (say sixty years) but fired at different times after the policy change (say at the 1st of February 1999 and at the 1st of February 2002) will face different terms of notice (in this case five months and four-and-a-half months).

Although the length of the term of notice that applies in the labor office path is set out by law, the 'Flexwet' made it possible to agree upon a different term in a collective wage agreement. For the analysis in this chapter it is important to know to what extent this occurred[7]. Combining information from the Ministry of Social Affairs and Employment and my own calculations I find that after the 1st of January 1999 the legal term of notice applied to eighty percent of workers[8]. If the deviation from the

[7]Smits (2000) and Smits and Samadhan (2002) report that the one month deduction, with a minimum term of notice of one month, prevailed in practically all agreements settled after the 1st of January 1999.

[8]About twenty percent of the workers do not fall under any sort of collective agreement and hence the legal term of notice will always apply to them. The other eighty percent of workers do fall under a collective agreement. I checked 112 (of about 200) collective wage agreements in alphabetical order that were declared to hold for all workers in the relevant sector between 1999 and 2002 and found that roughly seventy percent of these featured the term of notice set out in the 'Flexwet'. In sixteen percent of the cases a term of notice calculation that is related to age and tenure still existed. In fourteen percent of the analyzed collective wage agreements I found a constant term of notice. If I assume that the distribution of workers is equal across the different

labor law, i.e. the measurement error in the real term of notice, is independent of wages, the measurement error leads to an attenuation bias which will drive the coefficient of the term of notice down. Any significant results will hence be underestimating the true causal effect of the term of notice. In the Smits and Samadhan report the distribution of deviations over sectors seems rather equal. Therefore, and because it is hard to know which agreement specifically applies to which worker, I abstract from the deviation in the collective wage agreements and assume that the legal term of notice applies to everyone. In the empirical specification I do control for industry categories.

For older, long tenured employees the introduction of the law on flexibility and security thus resulted in a lower term of notice, but for shorter tenured workers of all ages the term of notice increased. This variation in the direction of the policy change will prove useful in identifying the causal effect of the term of notice on wages in Section 5.6.

5.4 Empirical strategy

The dependent variable in my analysis is the log of the real gross hourly wage rate. Several econometric difficulties trouble the identification of a causal effect on this rate of the term of notice. Everything arises from the fact that for the vast majority of employees in The Netherlands the term of notice is determined by a deterministic function of the total number of years an employee has worked for her employer and the number of years she has done so while being 45 or older. The only other variable influencing the employer's term of notice is calendar time.

$$ToN_{it} = f(\text{Tenure}_{it}, \text{Tenure after } 45_{it}, \text{Time}_t)$$

Note that calendar time both determines under which legal framework a worker is fired and also plays a role in the transition arrangement from the old towards the new scheme. The analyzed wage model can be seen in equation 1, where X_{it} refers to personal characteristics, Z_{it} to employer and job characteristics, υ_i to an individual-specific time-invariant error term and ε_{it} to an i.i.d error term.

$$Wages_{it} = h(ToN_{it}(\text{Tenure}_{it}, \text{Tenure after } 45_{it}, \text{Time}_t), X_{it}, Z_{it}) + \upsilon_i + \varepsilon_{it} \quad (1)$$

agreements it can be concluded that eighty percent of workers on a permanent contract face the legal term of notice after January 1st 1999.

The first econometric problem that arises is that of strong multi-collinearity between the term of notice and the important covariates tenure and age. When one would analyze periods in which the term of notice formula doesn't change, its marginal effect on wages cannot be determined ceteris paribus as changes in the notice period from one year to the next would coincide with changes in tenure and age. To nevertheless answer the research question, one can exploit an exogenous change in the term of notice that does not coincide with a change in tenure and age. To identify a causal relationship between the term of notice and wages this chapter therefore uses the exogenous policy change described in Section 5.3: the introduction of the law on flexibility and security in The Netherlands. It is not necessary to apply any specific type of policy evaluation method however. For reasons explained below, I will estimate a fixed effects regression model which has the additional trait that - through the subtracted average - it automatically incorporates the exogenous variation in the term of notice.

Even when including the exogenous policy change, it is still crucial to correctly control for tenure and tenure beyond the age of 45 so that an estimated term of notice coefficient does not capture any linear or non-linear relationship between these variables and wages. Without imposing any structure on the relationship beforehand I therefore include a full set of all relevant tenure times tenure after 45 dummies. These dummies control for the effect of tenure on wages, for the effect of tenure in older ages on wages and for the interaction effect of the two on wages in the most flexible way. I observe 43 different tenures and 19 different tenures experienced over the age of 45. This leaves me with a total of 513 tenure times tenure beyond 45 dummies[9]. Naturally, other interactions between age and tenure, such as tenure obtained over the age of 25, could potentially also influence wages. However, omitting such interactions will not bias the coefficient of interest as these interactions do not enter the term of notice calculations.

Solving the multi-collinearity and non-linearity problems alone does not result in unbiased estimates of the term of notice coefficient. This is because tenure and subsequently the term of notice suffer from an endogeneity problem. Tenure is an endogenous variable because unobservable characteristics such as work attitude and innate ability influence tenure as well as wages. As a result of the fixed formula for the notice period this endogeneity stains the term of notice variable as well. A permanent

[9] The full set of tenure interactions amounts to less than 45*19=817 dummies as the tenure obtained beyond 45 will always be equal to or smaller than general tenure.

endogeneity problem can and will be addressed by applying a fixed effects type estimator that filters out any time-invariant individual components in the wage regression. Doing so however does not remove a potential non-permanent effect of the unobserved quality of a worker on tenure and wages induced by the 1999 policy change[10].

I will explain the non-permanent endogeneity problem by focusing on two types of workers: low-quality workers that earn a relatively low wage and high-quality workers that earn a relatively high wage. Here I assume that the type is unobserved by the econometrician. Of both types of workers, the relatively high-tenured ones (53 percent of the sample) will have experienced a decrease in their employer's term of notice on the first of January 1999. It is possible that this decrease had a different impact on the two types in terms of layoffs and job switches. Low-quality employees could have been fired sooner than their high-quality colleagues. These fired low-quality workers would then either leave the sample or re-enter the sample in a new job. This would lower the number of low-paid workers in the high term of notice group, resulting in an overestimation of the effect of the term of notice on wages. Note that underestimation is also a possibility, as the 1999 policy change made high-tenured jobs less attractive in terms of employment protection. Hence all workers could have become more likely to voluntarily quit a high-tenured job. High-quality workers potentially quitted their jobs more often than their low-quality colleagues as they receive better alternative job offers. This would increase the number of high-paid workers in the low term of notice group, resulting in an underestimation of the term of notice effect on wages. As it is unclear which job switches and which lay-offs would have taken place without the policy change, the described time-variant type of endogeneity cannot be fully controlled for.

The quantitative importance of the effect can be looked at in more detail. First, it can be checked whether the inclusion of those who get fired and those who change jobs voluntarily drives the term of notice results. To check the robustness of estimates I run a wage model on various restricted samples, such as a sample without anyone that is ever observed as being fired and a sample without anyone that is ever observed to voluntarily switch jobs. The results of these exercises are hopeful and can be found in the sensitivity analysis in Section 5.6. It seems that the term of notice results are robust to the inclusion or exclusion of various groups of mobile workers.

Second, it can be checked whether the number of firings and job

[10] I am grateful to an anonymous referee for pointing this out.

switches before and after 1-1-1999 differed substantially. Table 5.1 shows what has happened to the workers in my sample in the year following their interview. It shows unconditional percentages for those whose term of notice would have or went down and for those whose term of notice would have or went up. The latter category also includes individuals whose term of notice remained unchanged. The first two columns therefore present the percentages for relatively high-tenured workers, with an average tenure of fourteen years, while the last two columns present the percentages for relatively low-tenured workers, with an average tenure of six years. Unsurprisingly, all job turnover percentages are lower for the high-tenured group than for the low-tenured group. Moreover, both groups experienced more dismissals and voluntary job quits after 1-1-1999. Overestimation of the effect of the term of notice on wages requires a substantial number of low-quality, high-tenured workers that are fired because of the policy change. Since only 0.2 percent more high-tenured workers were fired after 1-1-1999, it is unlikely that the non-permanent endogeneity of tenure drives the positive effect on wages found in Section 5.6.

Table 5.1 deserves more explanation. Although theory predicts that

Table 5.1: What happens to employees in the year following inclusion in the sample, before and after 1-1-1999. Sample divided on the basis of tenure and tenure over 45.

	Flexwet decreased ToN (High average tenure)		Flexwet increased ToN or kept constant (Low average tenure)	
	Before	After	Before	After
Same job	93.4%	91.3%	90.0%	82.2%
Fired from job	2.5%	2.7%	5.0%	8.1%
Quitted - Better job	0.6%	0.9%	1.1%	1.6%
Quitted - Personal	3.4%	5.1%	3.9%	8.0%
Total known	3,518	4,601	2,721	4,201
Attrition - Unknown	351	728	324	770

the group who 'suffered' from the policy change by being less protected should display larger increases in job turnover rates this is not what is happening. In fact, the group of individuals with more incentives to stay in their jobs (as their term of notice mostly increased) displayed larger increases in these rates. Apparently, other economic circumstances and policy changes that occurred between 1997 and 2001 affected the job flows of the two groups of workers in the opposite direction. If this is the case, wages might have seen an opposite development as well. Fortunately, it is possible to control for these differential time trends in wages as selection into the higher- or lower term of notice group depends on tenure and tenure beyond the age of 45 only. In the regressions in Section 5.6 differential wage trends are controlled for in a flexible way by adding interaction terms between a dummy for those surveys after 1-1-1999 and the full set of tenure times tenure after 45 dummies. While doing so, the transition arrangement that was part of the 'Flexwet' ensures that there is enough variation in the term of notice after 1999 for identification of its effect on wages. Table 5.2 in Section 5.6 shows that excluding the possibility of differential wage trends from the regressions overestimates the effect of the term of notice of wages.

The 1999 policy change does have one drawback following directly from the advantage laid out above. The separate effect of employment protection on wages of newly hired individuals cannot be analyzed. As Leonardi and Pica (2007) do, it would be interesting to look at this group separately as wages of outsiders (newly hired individuals) could respond differently to a change in employment protection than wages of insiders. As all new employees are under the same term of notice scheme from the 1st of January 1999 onwards, the effect of the change in the term of notice cannot be distinguished from other wage developments in this group. Table 5.2 does show estimates for newly hired individuals only to illustrate that a negative effect of employment protection on wages is unlikely.

Given the discussion above, the term of notice results will be referred to as 'causal' from now on. The causal effect I am interested in can thus be obtained as the marginal effect of the term of notice variable in a regression of wages on the term of notice and the full set of tenure times tenure beyond the age of 45 dummies (and interactions of this set with a dummy for all periods after 1-1-1999). To furthermore control for other determinants of wages I include a full set of age dummies, a full set of time dummies, the job characteristics hours worked and level of occupation and the employer characteristics company size and industry type. Equation 2 is then estimated as a fixed effects linear regression model using the log of the real gross hourly wage rate as the dependent variable. In this equation X_{it} refers to all personal covariates, Z_{it} to employer and job characteristics, $\sum_{t=1998}^{2001} \text{Time}_t$ to a set of time dummies, υ_i to an individual-specific time-invariant error term and ε_{it} to an i.i.d error term.

$$Wages_{it} = \alpha + \beta_{ToN} * ToN_{it} + \gamma' X_{it} + \delta' Z_{it} + \sum_{t=1998}^{2001} \zeta_t \text{Time}_t + \upsilon_i + \varepsilon_{it} \quad (2)$$

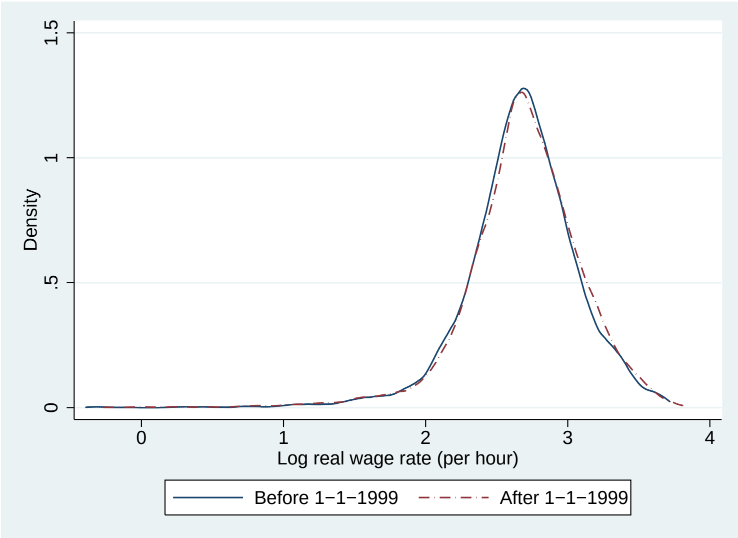
5.5 Data

Five waves of the Dutch Socio-Economic Panel (SEP), a household survey, are used for the empirical analysis of the research question (1997-2001). This longitudinal dataset has been collected annually around April by Statistics Netherlands from 1984 to 2002. I use all available waves before and after the 1999 policy change that contained information on the type of contract a worker was on and on the sector an individual was employed in. Note that a job in this chapter refers to a contract between an employee and an employer and that internal promotions ore demotions thus do not play a role. Only employees with a permanent contract are included in the sample as employers only face a legal term of notice for these workers. This leaves me with a final sample of 17,214 observations. Although the same individuals are observed multiple times in the sample, it is not balanced. 79 percent of the individuals are observed both before and after the policy change, 7 percent only before and 15 percent only after.

The dependent variable in the wage regressions is the logarithm of the real gross hourly wage rate. Net wages are not directly observed. Over the analyzed period the income tax legislation did not change substantially. I use the consumer price index published by Statistics Netherlands to compute real wages (base year is 2005) and the conversion rate between the Dutch guilder and the euro that has been fixed to 2.20371 from the 31st of December 1998 onwards. For simplicity I use this exchange rate for the 1997 and 1998 waves as well. In each survey year, all observations below the 0.1th percentile and above the 99.9th wage percentile are dropped, because rates in these areas seem unlikely. Figure 5.2 shows a kernel density estimation of the cleaned variable both for the period before the introduction of the 'Flexwet' and the period after. The distribution in real wages is very similar across the two periods. The average wage rate before the policy change was (2005 real) €15.59 euro per hour (s.d. 5.83) and after the policy change it was €15.89 (s.d. 6.08). Those with a tenure above two years earn more (€16.63, s.d. 5.72) than those who are new to their employer (€13.41, s.d. 6.07). As expected, those over 45 also earn more on average (€17.76, s.d. 6.56) than their younger counterparts (€14.79, s.d. 5.42).

Figure 5.3 provides information on the average terms of notice in my sample before and after the policy change for different tenures and ages. Before 1999 the aggregate average term of notice was 2.1 months (s.d. 1.5), whereas it was 1.8 months (s.d. 1.1) after the introduction of the

Figure 5.2: Kernel density estimation of the hourly wage rate for the period before and after 1-1-1999 (N=17,214)

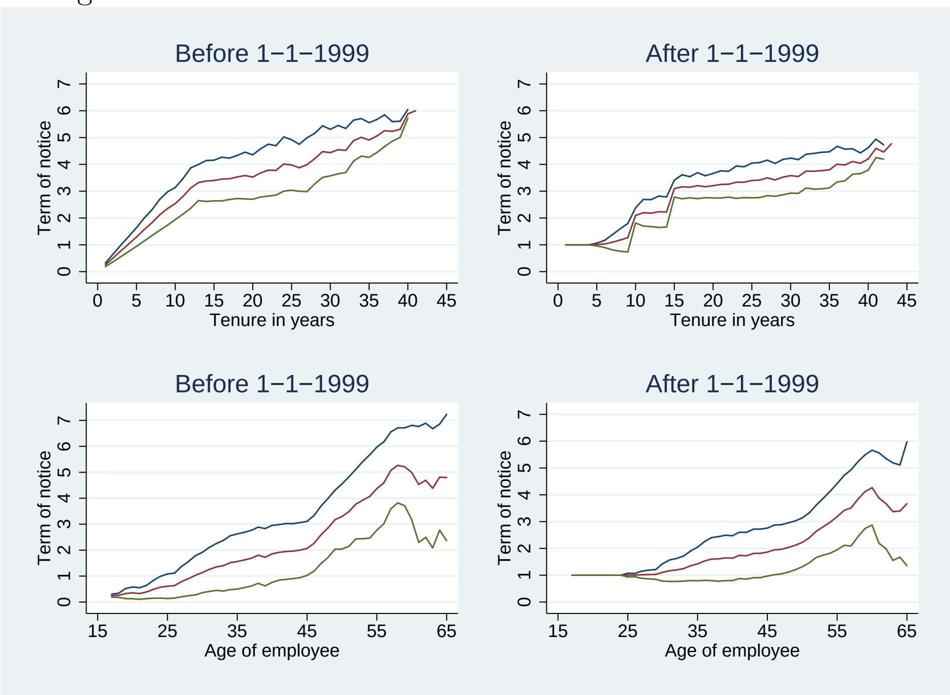


'Flexwet'. Figure 5.3 also presents the ninety percent confidence intervals. It can be seen that the new law increased the minimum and decreased the maximum notice period and that the variation in employment protection across ages and tenures was diminished. In the new scheme, a more distinct stepwise pattern by tenure is also observed.

Figures 5.A.2 and 5.A.3 in the appendix show histograms of the ages and tenures of the workers in my sample both before and after the first of January 1999. Before 1999, the average age is 39.4 years (s.d. 9.4) and the average tenure is 10.2 years (s.d. 8.7). After the first of January 1999, the average age in the sample is 39.9 years (s.d. 9.8) and the average tenure is 9.9 years (s.d. 9.2). A relatively large group of employees are only shortly employed at their employer, and this is more so in the period after 1-1-1999. In the sensitivity analysis in the next section, I check whether only looking at newly hired employees, or only at all individuals other than the newly hired affects the term of notice results. This is not the case.

The regressions presented in table 5.2 include the following control variables of which the descriptive statistics can be found in table 5.A.1 in the appendix: all tenure times tenure over the age of 45 dummies, all age dummies, dummies for the years 1998 to 2001, hours worked, industry and size of the organization in which the worker was employed and level of her occupation. Table 5.A.1 also includes information on a dummy

Figure 5.3: Average employer's Term of Notice and its ninety percent confidence interval in the sample pre 1999 and post 1-1-1999 by tenure and age



representing whether someone took up training paid by the employer and several education dummies. These variables are used in the regressions presented in table 5.3.

5.6 Results

5.6.1 Impact of term of notice on wages

Table 5.2 presents the results of the empirical wage analysis. The table only displays the coefficients and standard errors of the variable of interest - the term of notice. This is done for expositional reasons as the inclusion of all age and tenure times tenure beyond 45 dummies in the regressions makes these difficult to interpret[11]. All specifications include observations of before and after the 1999 policy change. Hence, multicollinearity problems can no longer bother the estimates. Note that in both table 5.2 and 5.3 all standard errors were clustered at the individual level. The regression diagnostics can be found in table 5.A.1 in the appendix.

The coefficient and standard error in the first row refers to a fixed effects linear regression estimate of the log of real gross hourly wages on the term of notice and the described covariates. The model is preferred as it includes interactions between the dummies that determine the term of notice (tenure times tenure after the age of 45 cells) and a dummy for the period after the first of January 1999. As discussed in Section 5.4 this is done to make sure that the term of notice coefficient does not capture any differential wage developments over time other than the introduction of the 'Flexwet'. In the preferred model, the term of notice coefficient equals 0.0324 and is highly significant. This means that for each additional notice month the hourly wage rate of a worker goes up by three percent, ceteris paribus. The estimate in the second row is biased because the fact that wage trends could have been different for low and high-tenured workers is ignored. The result in row two suggests that an additional legal month of notice increases the wage rate, ceteris paribus, by 5.67 percent, which is a substantial overestimation of the impact of the term of notice. In the third row a coefficient and standard error are presented for a pooled OLS regression and the regression estimate therefore suffers from a time-independent endogeneity problem. The effect of the term of notice on wages is only 0.77 percent in this specification and not

[11] All results are however available upon request from the author.

Table 5.2: Term of notice coefficients and standard errors in wage models. Dependent: log of real hourly wage rate.

	Model	Sample	Coefficient	S.e.	N
1	Preferred	Full sample	0.0324***	(0.005)	17,214
2	No interact.	Full sample	0.0567***	(0.007)	17,214
3	OLS	Full sample	0.0077	(0.007)	17,214
4	Preferred	Without 1999	0.0324***	(0.005)	13,807
5	Preferred	Balanced panel	0.0267***	(0.007)	8,650
6	Preferred	Ind before/after	0.0390***	(0.005)	13,582
7	2000 interact.	Full sample	0.0351***	(0.005)	17,214
8	2001 interact.	Full sample	0.0375***	(0.005)	17,214
9	No interact.	Newly hired	0.2084	(0.305)	3,462
10	Preferred	Tenure >1	0.0296***	(0.007)	13,752
11	Preferred	No lay-offs	0.0307***	(0.005)	16,053
12	Preferred	No job switches	0.0356***	(0.006)	14,133
13	Preferred	No lay-offs/switches	0.0337***	(0.006)	13,333

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

All regressions include all tenure times tenure after 45-dummies, all age dummies, year dummies, hours worked, industry type, size of company, level of occupation and a constant.

The preferred model includes interaction terms of the tenure cells and a dummy for after the policy change.

significantly different from zero. Not controlling for the time-invariant endogeneity of tenure can hence underestimate the effect of the term of notice on wages.

To check the robustness of the results in the preferred model it was estimated on several samples and the coefficients obtained in this sensitivity analysis can be found in rows four to thirteen. The analysis clearly confirms that the positive significant effect of EP on wages is robust. The model in the fourth row of table 5.2 does not use observations recorded between April 1998 and March 1999. This was done to ensure that the causal effect of the term of notice on wages is not the result of wage setting behavior anticipating the January 1999 policy change. Although I do not find such anticipation, note that such behavior would not contradict but rather emphasize a causal effect of the term of notice on wages. The fifth line displays the term of notice coefficient for the wage model on a balanced panel sample in which all individuals are observed five years in a row. The estimates are reassuring as the ninety-five percent confidence interval of the significant term of notice coefficient includes 0.0324. The model presented in row six only includes individuals that are observed both before and after the policy change (so that I loose twenty-one percent of the sample). The associated effect of the term of notice is even larger than that found in row one: a four percent wage increase per month. To be more prudent, my preferred model thus includes the individuals that are only observed at either side of the policy change.

Row seven and eight display term of notice coefficients for models in which the possibility for differential time trends with different cut-off years was included. The model in row seven includes interactions with a dummy for all periods after 1-1-2000. The model in row eight includes the same interactions but then with a dummy for all periods after 1-1-2001. It seems that the causal effect found in row one is not driven by my choice of interactions as the coefficients in row seven (0.0351) and eight (0.0375) are even larger. Row nine and ten present the estimates of the wage model on a sample of newly-hired individuals only and on a selected higher-tenured sample respectively. Note that the latter coefficient will be biased because differential wage trends cannot be taken into account and because selection into this group is highly selective. The selection problem also plagues the longer-tenured individuals estimate. However, the positive marginal effects that I find in both models (0.2084, not significant, and 0.0296, significant) are reassuring.

As mentioned in Section 5.4 including individual-specific fixed effects does not solve all endogeneity problems in the data. Although the time-variant type of endogeneity cannot fully be controlled for I can show that the inclusion of marginal groups does not severely affect the term of notice results. In the estimation for row eleven in table 5.2 all those individuals are excluded that are ever observed as being fired from 1996 to 2002. This decreases the sample size to 16,053. The inclusion of this group of people could potentially overestimate the term of notice effect. Indeed the term of notice coefficient in row eleven is lower than in the first row. However, it is precisely estimated at 3.07 percent, which is similar to the estimate in row one. Row twelve presents results for a sample in which those individuals that ever voluntarily switched jobs from 1996 to 2002 are excluded. The inclusion of these workers would perhaps underestimate the term of notice effect, as explained in Section 5.4. The result in row twelve does suggest this is the case. The estimated term of notice coefficient (0.0356) is closer to 0.04 than to 0.03. Row thirteen presents an estimate in which both groups, the ever fired and the ever job switchers, are excluded. This leaves me a sample of 13,333 individuals who either stay in their job over the sample period or who quit their job for personal reasons. Using this subgroup of the population, I estimate a significant effect of the term of notice on wages of around three percent per month (0.0337) as well.

I thus find evidence that the employer's term of notice has a strong positive causal effect on wages. A three percent higher wage rate for each additional month of notice is a relevant and substantial side-effect of this type of employment protection. For a prime-aged worker in my sample, aged 50 with 15 years of tenure in 1998, who experienced a drop in term of notice from 3.15 months to 2.15 months because of the 1999 policy change, this equals a loss in the hourly wage rate of three percent. On a yearly basis this employee, who worked 38 hours a week and earned 19.71 euros, therefore lost €1,262 because of the lower employer's term of notice.

5.6.2 Suggestive evidence of mechanism at play

As discussed in Section 5.2 there are two plausible reasons for a positive effect of employment protection on wages. First, insider/outsider theory suggests that the bargaining position of insiders is enhanced by better protection when unions only represent incumbent workers. In wage negotiations this could then drive wages up. Second, employment protection creates more incentives for firms to invest in workers and for workers to invest in firm-specific human capital. These investments will then lead to higher productivity and higher wages. Using the data at hand I can only provide suggestive evidence on which of the two mechanisms seems more important. For this purpose table 5.3 in included.

In rows one and two I present the term of notice coefficients of mod-

Table 5.3: Term of notice coefficients and standard errors. Dependent row 1-2: dummy for currently following formal training paid by the employer. Dependent row 3-5: log of real hourly wage rate.

	Topic	Model	Coefficient	S.e.	N
1	Training	FE OLS	-0.0075	(0.006)	17,434
2	Training	Probit	-0.0705**	(0.023)	15,475
3	Wages	Low educated	0.0574***	(0.022)	1,864
4	Wages	Middle educated	0.0318***	(0.009)	6,017
5	Wages	High educated	0.0277***	(0.008)	4,017

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

All regressions include the same controls as in table 5.2.

els estimating formal training on the same covariates as in row one of table 5.3. The dependent variable here is a dummy equal to one when the employee is enrolled or has been over the previous year in a course or training program that her employer is paying for. Although investment in a employee-employer match could involve more or other things than formal training, one would expect a positive coefficient of the term of notice on formal training take-up if the investment mechanism would play a significant role. Row one depicts results for a linear probability model and row two for a pooled probit estimation[12]. Both estimates are negative (-0.0075, not significantly, and -0.0705, significantly), suggesting that better protected individuals take up less training offered by their employer. Arulampalam et al. (2004) found a similar effect using the European Community Household Panel.

In rows three to five I present term of notice coefficient estimates for the preferred wage model on three different samples: the sample is divided into those with a low level of education (regular high school or lower), those with a middle level of education (higher level high school or vocational training) and those with a high level of education (higher tertiary education). This was done as I suspected a potential investment mechanism to have differential effects over the unobserved quality of workers. If I assume that there is a strong positive correlation between education

[12]The large number of independent variables rendered estimating a fixed effects binary choice model impossible.

and this quality and if I furthermore assume that the gains from investing in match-specific human capital would be higher for workers with higher abilities, the better-incentives-to-invest effect would ensure a higher term of notice coefficient for the better educated. Note that finding such a higher coefficient could also indicate that better able employees also excel at wage bargaining. However, rows three to five in table 5.3 paint the opposite picture. The term of notice coefficient for the low educated equals 5.74 percent, while for the middle and high educated it is around three percent. Low educated individuals thus benefit more from employment protection in terms of wages. If one indeed believes that higher educated individuals are better able to obtain match-specific human capital, this result also invalidates the investment argument.

These pieces of evidence point in the direction of an improved bargaining position as the mechanism driving the positive effect of employment protection on wages. In the Dutch context, a longer term of notice seems to improve the position of employees in wage negotiations such that they are able to extract a larger part of the rent generated in their job. Note that this conclusion should be treated with caution as the evidence is only suggestive. More research will be necessary to come up with a decisive answer on what exactly explains the positive term of notice coefficients.

5.7 Conclusion

This chapter establishes the causal effect of the employer's term of notice on the wage level of employees. The legal term of notice is defined as the amount of time a firm is required to notify a worker in advance of her upcoming dismissal. As such, the term of notice is an important component of firing costs and thus of employment protection.

In order to find a causal link, I have performed a fixed effects estimation exploiting an exogenous policy change in the term of notice. This procedure corrects for the time-invariant endogeneity of tenure and for the strong multi-collinearity between the term of notice, tenure and age. The latter problem arises because tenure and age, or more specifically tenure and tenure obtained while being 45 or over, are the only inputs in the legal formula that calculates the term of notice. The relevant policy change is the 1999 introduction of the law on flexibility and security ('Flexwet') in The Netherlands. This law altered the calculation of the legal term of notice such that the term of notice of low-tenured individuals of all ages went up and the term of notice of older high-tenured individuals went down. Five waves of the Dutch Socio-Economic Panel (SEP) were used for the empirical analysis (1997-2001) in which a fixed effects linear regression model is estimated using the logarithm of real gross hourly wages. A possible limitation of my approach is that tenure and the term of notice are also influenced by the policy change itself, thereby generating a time-variant endogeneity problem. Robustness analysis however suggests that this type of endogeneity is not driving the results.

The preferred econometric model unveils a very significant, positive causal effect of the 'dormant' term of notice on the real hourly gross wage rate. An increase in the term of notice of one month leads to three percent higher wages. The arguments presented in Section 5.2 describing a negative effect of employment protection on wages thus do not hold in the analyzed context. This was to be expected for Lazear's (1990) efficiency wage argument as the Dutch labor market institutions do not resemble his competitive model with individual wage-setting. The same could be said for Pissarides' (2001) exposition about an endogenous term of notice as the term of notice is decided upon by policy-makers.

The theoretical literature provides two plausible reasons for a positive effect of employment protection on wages. First, insider/outsider theory suggests that the bargaining position of insiders is enhanced by better protection when unions only represent incumbent workers. In wage negotiations this could then drive wages up. Second, employment protection creates more incentives for firms to invest in workers and for workers to invest in firm-specific human capital. These investments will then lead to higher productivity and higher wages. More empirical research needs to be conducted to ultimately decide which of these theories is best describing reality. To do so, detailed information on individual employees such as longitudinal micro-data reflecting individual productivity is needed.

Using the available information instead, I present suggestive evidence that undermines the investment argument. First, better protected employees do not have a higher take-up rate of formal training paid by their employer. If anything, those with a longer term of notice participate less in such courses, which is not what one would expect if the investment mechanism was playing a large role. Second, the effect of the term of notice on wages is much stronger for lower-educated individuals than for middle- and high educated individuals. If one believes that higher educated individuals are better able to obtain match-specific human capital, this also suggests that the investment argument is least important. Hence, the bargaining argument wins the first round in the battle of the mechanisms. In the Dutch context, better employment protection probably improves employees' wage bargaining position such that a larger part of the profits flows to the employee rather than to the employer. Policy makers should be aware of this side-effect of the term of notice on wages. If the wage bargaining argument is indeed stronger, employment protection creates rents for incumbent employees which policy-makers will want to avoid.

5.A Appendix

Figure 5.A.1: The ratio of lay-offs through the labor office over the total number of lay-offs (Source: Ministry of Social Affairs and Employment, 2003)

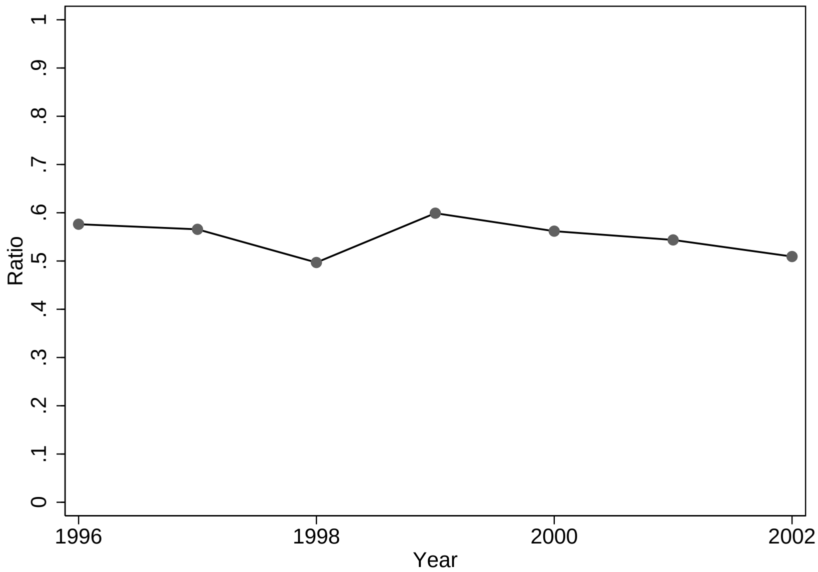


Figure 5.A.2: Histogram of ages in sample

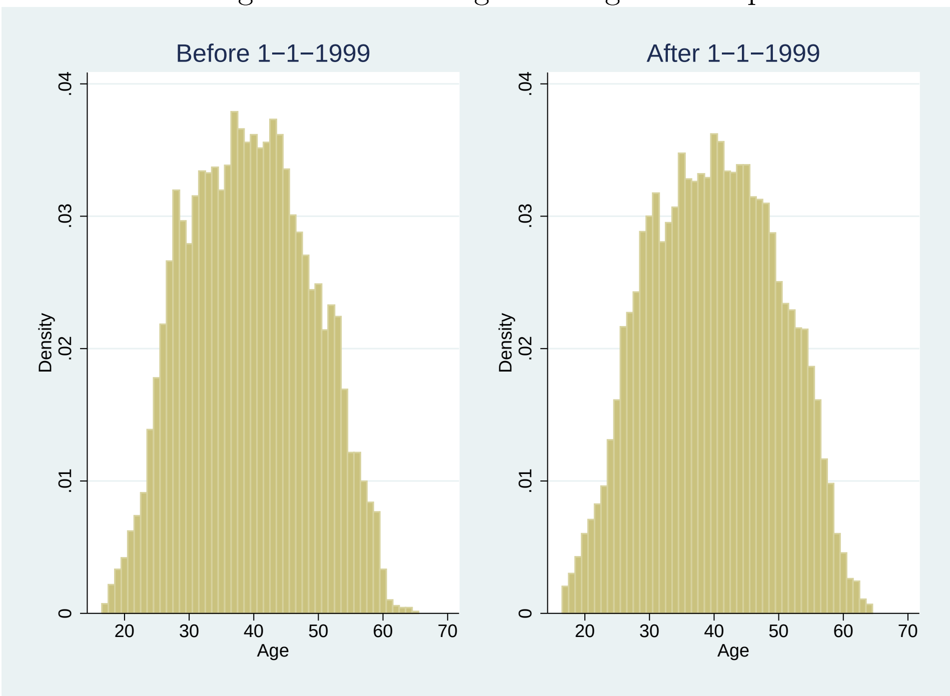


Figure 5.A.3: Histogram of tenure in sample

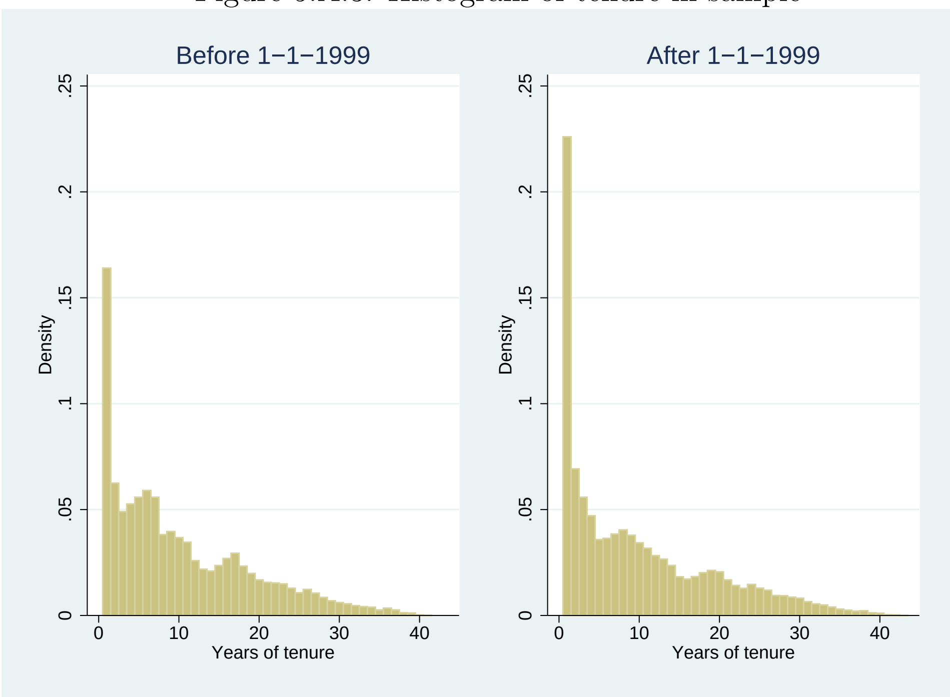


Table 5.A.1: Descriptive statistics of all variables (N=17,214)

	Mean	Sd	Min	Max
Log gross wage rate (per hour)	3.473	0.409	0.387	4.529
Term of notice (months)	1.935	1.273	0.231	6.000
Age (years)	39.694	9.661	17	65
Tenure (years)	10.028	8.996	1	43
Tenure beyond 45	1.629	3.262	0	19
Dataset	1998.995	1.419	1997	2001
Hours worked (per week)	36.231	10.246	12	89
Agriculture	Omitted			
Fisheries	0.000	0.011	0	1
Mining	0.012	0.110	0	1
Industry	0.143	0.350	0	1
Trade	0.006	0.079	0	1
Construction	0.060	0.237	0	1
Retail	0.109	0.312	0	1
Hospitality	0.014	0.119	0	1
Transport	0.058	0.233	0	1
Financial institutions	0.043	0.203	0	1
Real estate	0.098	0.297	0	1
Public services	0.090	0.287	0	1
Education	0.081	0.273	0	1
Health care	0.158	0.364	0	1
Environmental services	0.026	0.158	0	1
Personal services	0.001	0.029	0	1
Other industry	0.099	0.299	0	1
Company 1-19 employees	0.173	0.378	0	1
Company 20-49 employees	0.125	0.330	0	1
Company 50-99 employees	0.100	0.300	0	1
Company >99 employees	Omitted			
Unknown level	0.065	0.246	0	1
Elementary occupations	0.053	0.223	0	1
Lower occupations	0.271	0.445	0	1
Middle occupations	0.342	0.474	0	1
Higher occupations	0.210	0.407	0	1
Academic occupations	Omitted			
Employer-paid training	0.079	0.270	0	1
Lower educated	0.1083	0.311	0	1
Middle educated	0.3495	0.477	0	1
Higher educated	0.2334	0.423	0	1

Table 5.A.2: Regression diagnostics of the models in table 2 and 3.

	Model	Sample	Aic	Ind.	N
1	Preferred	Full sample	-21,228	5,522	17,214
2	No interact.	Full sample	-21,178	5,522	17,214
3	OLS	Full sample	8,445	5,522	17,214
4	Preferred	Without 1999	-17,993	5,396	13,807
5	Preferred	Balanced panel	-7,375	1,730	8,650
6	Preferred	Ind before/after	-14.261	3,189	13,582
7	2000 interact.	Full sample	-21,350	5,522	17,214
8	2001 interact.	Full sample	-21,432	5,522	17,214
9	No interact.	Newly hired	-3,004	2,102	3,462
10	Preferred	Tenure >1	-22,893	4,659	13,752
11	Preferred	No lay-offs	-20,649	5,078	16,053
12	Preferred	No job switches	-18,930	4,565	14,133
13	Preferred	No lay-offs/switches	-18,621	4,248	13,333
1	FE OLS	Training sample	-6,800	5,561	17,434
2	Probit	Training sample	8,964	5,408	15,475
3	Preferred	Low educated	-2,075	612	1,864
4	Preferred	Middle educated	-6,964	1,681	6,017
5	Preferred	High educated	-5,502	1,148	4,017

Akaike's information criterion in fourth column, number of individuals in fifth column.

Chapter 6

Nederlandse Samenvattingen (Abstracts in Dutch)

6.1 Over Verwachtingen

Hoofdstuk 2: Heb je het al Gehoord? Hoe Echte Verwachtingen Reageren op Publiciteit

Steeds meer empirisch bewijs hoopt zich op dat subjectieve verwachtingen van individuen economisch gedrag beinvloeden en dat er soms fouten in deze verwachtingen zitten. Hierdoor wordt het beleidsrelevant om te weten hoe deze verwachtingen gestuurd kunnen worden. Informatie in de publieke media is waarschijnlijk een belangrijk kanaal om het algemene toekomstbeeld van mensen te beinvloeden. Dit hoofdstuk bestudeert de rol van publieke informatie disseminatie, of publiciteit, op een specifiek verwachtingsvormingsproces in de werkelijkheid. Hiervoor analyseer ik een bijzondere dataset met maandelijkse individuele verwachtingen over de toekomstige leeftijd waarop men in Nederland voor het eerst AOW ontvangt. Ik gebruik gegevens die verzameld zijn van mei 2006 tot en met november 2008.

Gemiddeld gezien reageren deze verwachtingen weinig op publiciteit alhoewel er wel grote verschillen tussen subgroepen zichtbaar zijn. Onder andere concludeer ik dat hoger opgeleiden en hogere inkomensgroepen hun verwachtingen bijna niet aanpassen aan nieuwe publiciteit over de toekomst van de AOW. Degenen die niet vaak een krant lezen laten daarentegen wel een relatief hoge publiciteitsreactie zien. Een mogelijke uitleg hiervoor is dat de kwaliteit van de intitiele AOW-verwachtingen van deze groep laag is geweest. Als dat waar is, helpt publiciteit met name slecht geinformeerden de eigen verwachtingen over de toekomst van de AOW te verbeteren.

Hoofdstuk 3: Klaar voor de Verandering: Verwachtingen over Sociale Zekerheid en Spaargedrag

Overheden van rijke landen denken op het moment na hoe zij hun omslagpensioenstelsels zo kunnen aanpassen dat deze ook in de toekomst, met een vergrijsde bevolking, financieel houdbaar blijven. Voor zover beleidsmakers niet al tot versobering van de regelingen besloten hebben, leidt vergrijzing dus tot beleidsonzekerheid in eerste pijler pensioenen. Dit hoofdstuk onderzoekt de invloed van deze beleidsonzekerheid in Nederland op de vrijwillige bijdragen van huishoudens aan pensioenproducten. Hiervoor analyseer ik hoe de verwachtingen van individuen over de toekomst van de AOW samenhangen met of zij vrijwillig bijbetalen aan pensioenproducten.

Het blijkt dat participatie in vrijwillige pensioenproducten hoger is voor diegenen die hogere kansen toewijzen aan een toekomstige versobering van de AOW, en dan met name voor diegenen die geloven dat de AOW-leeftijd binnenkort omhoog zal gaan. Er kan dus geconcludeerd worden dat Nederlandse huishoudens zich voorbereiden op beleidswijzigingen in de eerste pensioen pijler.

6.2 Over Macht

Hoofdstuk 4: Kantige Keuzes, Dictators en Samengaande Macht: Een Non-Cooperatief Model voor Consumptie en Arbeidsaanbod binnen het Huishouden

Het is onwaarschijnlijk dat een koppel het altijd eens is over wat er van het gezamenlijk budget gekocht moet worden. Ook is het onwaarschijnlijk dat een getrouwd stel het altijd eens is over hoeveel uur beiden zouden moeten werken; een beslissing die weer van belang is voor de hoogte van het gezamenlijke budget. In dit hoofdstuk presenteer ik daarom een noncooperatief model van de consumptie en arbeidsaanbod van een koppel dat met elkaar onderhandeld op de 'Nash' manier. Terwijl we weinig aannames doen, kunnen we bewijzen dat de vraag naar publieke goederen binnen het gezin gekarakteriseerd wordt door drie regimes. De vraag wordt ofwel geheel bepaald door de voorkeuren van een van de twee partners (Man Dictatorschap of Vrouw Dictatorschap), of door de voorkeuren van beide partners waarbij de invloed van een partner afhangt van diegenes relatieve loonvoet (Samengaande Macht). De drie regimes zorgen er vervolgens voor dat alle vraagcurves van het koppel naar publieke goederen scherpe knikken vertonen.

Als we een aantal extra aannames doen, kan ons model getest worden op data over de totale vraag naar publieke goederen binnen gezinnen en over het arbeidsaanbod van mannen en vrouwen. We kunnen ons non-cooperatieve model dan testen ten opzichte van een standaard unitair model waarin het koppel zich als een enkele besluitvormer gedraagt. We testen ons model op een steekproef uit de Amerikaanse Consumenten Uitgaven Studie (CEX) waarbij we ons expliciet focussen op de uitgaven aan kinderen. We nemen deze uitgaven aangezien deze publieke goederen zijn, dwz. dat deze gewaardeerd worden door zowel de man als de vrouw terwijl geen van beiden de goederen op gebruiken. Het blijkt dat we het standaard unitaire model voor koppels met twee of drie kinderen sterk kunnen verwerpen in het voordeel van ons non-cooperatieve model. Bovendien blijkt dat in de meerderheid van de stellen, de vrouw de dictator is in de zin dat het uitgavenpatroon aan kindergoederen volgens haar voorkeuren plaatsvindt.

Hoofdstuk 5: Beter Beschermd, Beter Betaald: Empirisch Bewijs dat Ontslagbescherming Lonen Beinvloedt

Dit hoofdstuk stelt de empirische invloed vast van de lengte van de opzegtermijn op het loonniveau van werknemers. Met de opzegtermijn wordt hier de periode bedoeld die een werkgever in acht moet nemen als zij een werknemer via het arbeidsbureau wil ontslaan. In Nederland bepaald een vaste formule in de wet de lengte van deze termijn. De lonen die moeten worden uitbetaald gedurende deze periode zijn een belangrijk element van de ontslagkosten en dus van ontslagbescherming.

Om een causaal verband aan te tonen, maak ik gebruik van een exogene verandering in de berekening van de opzegtermijn die optrad door de invoering van de Nederlandse Flexwet op 1 januari 1999. Ik kan aantonen dat een langere 'latente' opzegtermijn tot hogere lonen leidt. In mijn steekproef, verhoogt een extra maand opzegtermijn lonen met drie procent, ceteris paribus.

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