Risk Aversion & Retirement Decisions:
Using Policy Variation to Identify and Estimate a Structural Model of Retirement*

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Abstract

How do individuals’ retirement decisions respond to changes in retirement benefits? Are the results from structural and reduced-form methods for estimating labor supply responses to changes in retirement benefits consistent with one another? We develop a standard dynamic model of retirement decisions and show that the model implies a relationship between the coefficient of relative risk aversion (\(\gamma\)) and labor supply responses to retirement benefits. This relationship indicates that \(\gamma\) can be estimated via two approaches: (1) a reduced-form approach that relies on identification of labor supply elasticities and (2) a structural approach that explicitly models the micro-foundations of the economic environment. We use administrative data from the Austrian Social Security Database and exploit variation from five pension reforms in Austria between 1984 and 2003 to implement both approaches. Using Cox proportional hazards specifications, we estimate reduced-form income and price elasticities of roughly 0.43 and \(-2.90\) respectively. Next, we implement the widely-used structural approach of estimating \(\gamma\) based on matching retirement patterns across ages. We show that, even though they are unified under the same theoretical model, the two approaches yield different estimates of \(\gamma\) because they are based on different sources of identifying variation. We propose a solution to reconcile these differences while exploiting the policy variation for identification based on estimation via the method of Indirect Inference.

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1 Introduction

With increasing pressure for social security reform in countries around the world, it is important to understand how individuals’ retirement decisions respond to changes in retirement benefits. This task has received significant attention in the economics literature and led to the development of different reduced-form and structural approaches to estimate the responses of retirement decisions to changes in benefits. However, it is unclear whether the different empirical approaches yield consistent or contradictory estimates of labor supply responses to changes in retirement benefits, and hence whether they yield the same implications for policy.

In this paper, we exploit policy variation to (1) identify the causal effects of retirement benefits on retirement decisions and (2) examine the relationship between structural and reduced-form estimates of labor supply responses to changes in retirement benefits. We implement both reduced-form and structural estimation strategies and we highlight important differences in the results. We then develop and implement an estimation strategy based on the method of Indirect Inference (Gourieroux, Monfort and Renault 1993) that reconciles the differences in the empirical approaches by using policy variation to estimate the structural parameters. Specifically, we demonstrate how policy variation can be used to identify and estimate the coefficient of relative risk aversion ($\gamma$). We show that, under widely-used assumptions, a standard dynamic model implies a close link between $\gamma$ and the labor supply responses to changes in retirement benefits.

This relationship between risk aversion and labor supply elasticities indicates two possible approaches to estimating risk aversion. First, one can exploit policy variation to identify the labor supply elasticities and then estimate the coefficient of relative risk aversion by reduced-form methods. Second, one can estimate the coefficient of relative risk aversion by fitting a structural model using moments based on observed retirement patterns, such as retirement hazard rates by age. Note that while other studies have focused on issues of model specification (i.e., what features should be included in the model), the goal of this study is to concentrate on parameter identification within a given class of models. That is, we focus on which moments should be used to estimate the coefficient of relative risk aversion.

An important contribution of this paper is the separate identification of the income and price elasticities from retirement benefits on retirement decisions. Separately identifying these elasticities is important for designing and implementing effective social security
reforms. We identify these elasticities based on variation from multiple pension reforms in Austria which create multiple, independent changes in pension wealth and the price of retirement. We estimate a wealth elasticity of roughly 0.43 and a price elasticity of about −2.90. The relatively larger price elasticity and smaller wealth elasticity contrasts with recent evidence from other social insurance programs. This difference may highlight the anticipated nature of retirement versus the unanticipated nature of events such as unemployment (see Aguiar and Hurst (2005)).

Another key contribution of this paper is to propose an estimation strategy that allows for reconciliation between reduced-form and structural approaches to estimating the risk aversion parameter. We show that the reduced-form and structural estimation strategies yield different estimates of the risk aversion parameter and of labor supply responses to retirement benefits because they rely on different sources of variation for identification. The reduced-form approach relies on within-age across-individual policy variation for identification, while the structural approach relies on across-age observational variation for identification. By contrast, the estimation strategy we propose isolates a common source of identifying variation, the policy variation. This estimation strategy, based on the method of Indirect Inference (II), allows one to estimate the structural model using the reduced-form coefficients as additional moments to be matched, thereby imposing that the structural model be consistent with the micro-level labor supply elasticities.

We illustrate the differences between the reduced-form and structural approaches in two ways. First, we show that the risk aversion parameter implied by the reduced-form elasticities estimated using only the policy variation, \( \hat{\gamma}_{RF} \), differs significantly from the structurally estimated risk aversion parameter based on matching observed retirement hazard rates by age via the Method of Simulated Moments (MSM), \( \hat{\gamma}_{MSM} \). Second, we show that the labor supply elasticities implied by \( \hat{\gamma}_{MSM} \) differ significantly from the reduced-form elasticities estimated using only the policy variation.

The model developed in this paper is a dynamic model of labor force participation at older ages. Under the assumption of additively time-separable utility, the model highlights that the coefficient of relative risk aversion is directly linked to the ratio of wealth and price elasticities in retirement decisions. Chetty (2006) discusses a general relationship between labor supply and risk aversion. The intuition for this relationship can be seen via the following an example. Suppose that an individual receives an exogenous cash grant. If we observe

a relatively large change in his labor supply, then we can infer that his marginal utility of consumption must decline relatively quickly because otherwise the individual would have increased his consumption instead of decreasing his labor supply. Hence a relatively larger wealth elasticity implies a higher degree of curvature, or specifically, a higher $\gamma$. Given the generality of this intuition, we capture this relationship using a parsimonious dynamic model.

Using a parsimonious dynamic model of labor force participation, we demonstrate that this relationship between risk aversion and labor supply elasticities holds regardless of how individuals’ consumption is determined. Because we have data only on earnings and participation histories and specifically lack savings or consumption data, we assume that individuals save at a fixed rate while working to determine consumption while working. Note, however, that the main point of the analysis is not affected by this assumption. While assumptions regarding how consumption is determined will undoubtedly affect the estimated risk aversion and labor supply elasticities, the main point of the analysis is that, without introducing the reduced-form elasticities as additional moments to match as in the II estimation strategy, there is no reason that a model that fits observed retirement patterns across age well should also fit labor supply elasticities identified from policy variation. Indeed, we show that, even with the restriction on saving, the structural model estimated using MSM fits the retirement hazard rates by age quite well but still fails to replicate the reduced-form elasticities.

Although several studies estimate labor supply responses to retirement benefits, the reduced-form and structural studies have tended to pursue their respective empirical approaches largely separately from one another. Thus, for several reasons such as differences in data and institutional settings, it is difficult to use results from existing studies to make comparisons between the empirical approaches. To make this comparison transparent, we pursue both the reduced-form and structural estimation strategies in the same institutional setting using the same data. Specifically, we use social security record data on the census of Austrians from the Austrian Social Security Database. In the reduced-form approach, we

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2The reduced-form approach emphasizes the isolation of specific identifying variation to estimate labor supply elasticities. Friedberg (1999), Song and Manchester (2007) and Haider and Loughran (2008) each use policy variation to provide reduced-form estimates of labor supply responses to changes in retirement incentives. The structural approach emphasizes the development of explicit micro-foundations to model retirement behavior. Stock and Wise (1990), Berkovec and Stern (1991), Lumsdaine, Stock and Wise (1992), Blau (1994), Rust and Phelan (1997), Rust, Buchinsky and Benitez-Silva (2003), and French (2005) each study retirement decisions by explicitly modeling the environment in which the decisions are made.
estimate a Cox Proportional Hazard specification. We exploit variation from five pension reforms in Austria between 1984 and 2003 to identify wealth and price elasticities. Austria therefore presents an appealing setting for this analysis given the available policy variation as well as the relatively simple institutional setting which we describe in more detail below. The reduced-form results indicate a wealth elasticity of roughly 0.4 and a price elasticity of about −2.9. We show that these reduced-form elasticities imply a relatively low coefficient of relative risk aversion, \( \hat{\gamma}_{RF} \approx 0.4 \).

Next, we implement the structural approach by following the widely-used MSM in which we match retirement hazard rates by age. The results indicate a much higher coefficient of relative risk aversion, \( \hat{\gamma}_{MSM} \approx 1.9 \). To emphasize the differences between the two approaches, we estimate the same hazard specification as in the reduced-form using simulated data from the structural model. Comparing the elasticities from the simulated data to those from the real data, we show that the structural model predicts a much larger overall responsiveness to changes in retirement benefits and at the same time overestimates the ratio of wealth and price elasticities. Specifically, the simulated data leads to an estimated wealth elasticity of roughly 3.1 and a price elasticity of about −6.8. These differences highlight that the different estimation strategies, based on different sources of variation (policy variation in the reduced-form approach and differences across age in the structural approach), yield conflicting results.

Given the differences between the two sets of results, we then turn to reconciling the differences using an estimation strategy based on II. Specifically, we augment the moments matched using the MSM estimator (retirement hazard rates by age) with additional moments from the reduced-form model: the wealth and price elasticities estimated by the hazard specification. Thus, we impose the estimated structural parameter \( \gamma \) to be consistent with the reduced-form elasticities which are identified based on the policy variation from the pension reforms.

This paper is organized as follows. In Section 2 we develop a dynamic model of retirement decisions and illustrate the relationship between risk aversion and labor supply. Next, we describe our data and the institutional setting in Austria in Section 3. In Section 4 we present the hazard model and MSM estimates and compare the results. In Section 5, we describe the II estimation strategy and the results. Section 6 concludes.
2 Theoretical Foundations

In this section, we develop a dynamic model of retirement decisions with uncertainty relating to mortality and job separations. Accounting for these sources of uncertainty in a dynamic setting is an important component of capturing uncertainty relating to retirement benefits. The dynamic model that we develop is closely related to previous work in the literature. In particular, Stock and Wise (1990) present a dynamic model of retirement decisions that emphasizes the role of the option value of work in these decisions. In addition to discussing the option value model, Lumsdaine, Stock and Wise (1992) present a dynamic programming model of retirement decisions. Berkovec and Stern (1991) present a dynamic model to examine movement between full-time work, part-time work, and retirement. These models have assumed that consumption is equal to current income.

The intuition behind the model is as follows. In each period, an employed individual must choose whether to retire or whether to continue working. A period in the model corresponds to an individual’s age. When making this decision at the beginning of a period, the individual knows his assets, retirement benefits, and wage. If he chooses to retire, the individual receives his annuitized retirement benefits and faces no remaining uncertainty from the labor market. Based on his assets and retirement benefits, the retired individual chooses his consumption optimally. On the other hand, if an employed individual chooses to continue working, he receives his wage and his consumption is then determined based on his savings and wage income. Importantly, the working individual takes into account the expected continuation value from being able to make a retirement decision at future ages.

We develop the model assuming that the individual follows a fixed consumption rule while working, rather than choosing his consumption optimally. This assumption is driven by the lack of savings or consumption data in the empirical analysis below. While different assumptions regarding consumption decisions will lead to different structural estimates of the risk aversion parameter and labor supply elasticities, the assumptions regarding how consumption is determined do not affect the intuition regarding the relationship between risk aversion and labor supply elasticities.

2.1 The Model

Consider first the optimization problem for an individual who has chosen to retire. Let $R_a(A_a)$ denote the value of retirement at age $a$ for an individual with assets $A_a$ where the subscript $a$ refers to the individual’s age. Once an individual has chosen to retire, the
individual solves the following optimization problem that defines the value of retirement:

\[
R_a(A_t, y^R_a) = \max_{\{A_{t+1}\}_{t=a}^T} u(A_t + y^R_a - k1(a < a_{old})) + \sum_{t=a+1}^T \beta^{t-a} \pi_{t|a} u(A_t + y^R_a - \frac{A_{t+1}}{1+r})
\]

s.t. \( A_{t+1} \geq A \ \forall t \geq a. \)

The variables \( r, \beta \) and \( \pi_{t|a} \) respectively denote the interest rate, the discount factor and the probability of survival to age \( t \) conditional on survival to age \( a \). The function \( u(.) \) captures utility over consumption with \( u' > 0 \) and \( u'' < 0 \). The maximization reflects that the individual chooses his savings in each period \( \{A_{a+1}\}_{t=a}^T \) optimally subject to the borrowing constraint of \( A \leq 0 \). The term \( y^R_a \) denotes the individual’s retirement benefits at age \( a \). Based on the institutional setting in Austria which we discuss in more detail in the next section, these benefits consist of an annual payment from a government-provided pension plus a one-time, employer-provided lump-sum severance payment at the time of retirement. The severance payment is a fraction of the individual’s salary in the year before retirement. This fraction is determined based on the individual’s tenure which is therefore also implicitly included as a state variable. The term \( k \) denotes a claiming cost. If the individual retires at an early age when individuals are only eligible for disability pensions, he must pay the one-time cost of claiming \( k \). This cost can be interpreted as the cost of visiting a physician to be classified as disabled and/or the monetary value of some psychic cost of claiming disability. After age \( a_{old} \), individuals are eligible for old-age pensions and therefore do not face this claiming cost.

Next, consider the problem facing an individual who has chosen to work. As in the case of retirement, the individual must choose his savings optimally. The optimization problem in the case of continuing to work differs from that in the case of retirement in the following respects. First, the working individual must take into account his disutility of work denoted by \( v_a \). Work disutility is increasing with age and each individual is assumed to know the profile of his work disutility across age with certainty. Specifically, prior to facing the first retirement decision, \( v_0 \) is drawn for each individual from a distribution \( \Psi(v) \) defined over \((0, \infty)\). The work disutility profile across ages is then given by \( v_a = v(a, v_0) \ \forall a \). Second, the individual’s income is based on his wage income and any severance pay if he was separated from his previous job and had qualifying tenure. After-tax work income at age \( a \) is denoted by \( y^W_a \). Third, the individual must take into account the continued uncertainty from the labor market. In particular, \( E_a[.] \) captures the individual’s continuation value from being
able to make a retirement decision in the future taking job separation uncertainty into account. Let $W_a(A_a, y_a^W, v_a)$ denote the value of working at age $a$ with assets $A_a$ and work disutility $v_a$. This value function is defined as follows

$$W_a(A_a, y_a^W, v_a) = u(c_a^W) - v_a + \beta \pi_{a+1|a} E_a[D_{a+1}(A_{a+1}, y_{a+1}^R, y_{a+1}^W, v_{a+1})]$$

The individual’s consumption in his flow utility, $c_a^W$, is based on his wage income net of how much of his wage income he saves. Using $\bar{s}$ to denote the individual’s savings rate, consumption is therefore given by $c_a^W = (1 - \bar{s})y_a^W$, and the individual’s savings for next period are given by $A_{a+1} = (1 + r)[A_a + \bar{s}y_a^W]$. The value function $D_a(\ldots)$ captures the value of being in the labor market at age $a$ and having the decision between retiring or continuing to work. When deciding between retirement and work, the individual simply chooses the option that presents the highest value,

$$D_a(A_a, y_a^R, y_a^W, v_a) = \max_{\text{retire, work}} \{R_a(A_a, y_a^R), W_a(A_a, y_a^W, v_a)\}.$$ 

In regard to heterogeneity, work disutility $v$ is allowed to vary across individuals. The interest rate $r$, discount rate $\beta$, survival probabilities $\pi_{a+1|a}$, and consumption-utility function $u(\cdot)$ are restricted to be common across individuals. Additionally, the borrowing constraints and utility functions are assumed to be the same during work and retirement. When considering his retirement benefits at any given age, we assume that the individual forecasts benefits at future potential retirement ages based on the current year’s legislation. Thus, calendar year enters the value functions as an implicit state variable that determines the legislation under which benefits are computed. In this setting, pension reforms then correspond to unanticipated changes in the legislation and hence unanticipated changes to benefits at current and future potential retirement ages.

In this optimal stopping time setting, the individual’s optimal strategy is to set a reservation level for his work preference such that he will retire if his utility from work is smaller than his reservation level. Let $\bar{v}_a$ denote the individual’s reservation level at age $a$.\footnote{Since the value functions $R_a$ and $W_a$ depend on age through the different vectors of survival probabilities at each potential retirement age, the reservation work disutility will also depend on age.} Formally, the reservation level is defined as the preference for work that leaves the individual indifferent between retiring and continuing to work

$$R_a(A_a, y_a^R) = W_a(A_a, y_a^W, \bar{v}_a) \Rightarrow \bar{v}_a = \bar{v}_a(A_a, y_a^R, y_a^W).$$
Given this reservation level, the individual’s retirement decision rule is then

\[
\text{Retire at age } a \text{ if } v \geq \tilde{v}_a.
\]

To compute the value functions used to determine the reservation work utility, we assume that there is a terminal age \( a^* \) at which all working individuals are assumed to retire. With this assumption, the value functions can be computed recursively.

Figure 1 presents an example with computed value functions at age \( a = a^* - 1 \). This figure plots utility on the vertical axis and work disutility on the horizontal axis. Since the value of retirement is independent of work disutility, the value function \( R_a \) is a horizontal line. The value of continued work decreases linearly as work utility increases. This linearity results from two components: first, work disutility enters flow utility at age \( a \) linearly; second, the option value at age \( a = a^* - 1 \) is independent of work disutility since all individuals must retire at age \( a^* \) in the model (i.e. \( D_{a^*} = R_{a^*} \)). The reservation work disutility is then determined by the intersection of the value functions reflecting the indifference between work and retirement that characterizes \( \tilde{v}_a \).

### 2.2 Labor Supply & Risk Aversion

This section develops the intuition relating labor supply to risk aversion in the dynamic model of retirement decisions. Chetty (2006) provides an extensive discussion of this relationship.\(^4\) In particular, Chetty emphasizes that the coefficient of relative risk aversion, \( \gamma \), can be identified based on two components: (1) the ratio of wealth and price elasticities in labor supply and (2) the degree of complementarity between consumption and labor in the utility function. In our setting, we have assumed additively time separable utility. With this assumption, one can fix the degree of complementarity in the per-period utility function to identify \( \gamma \) in terms of labor supply elasticities. We have assumed that work disutility is additively separable from the utility over consumption, so this complementarity is assumed to be 0.\(^5\)

The wealth and price effects from retirement benefits can be seen by examining how

\(^4\)See also Chetty (2008) and Card, Chetty and Weber (2007) for discussions of wealth effects, price effects and risk aversion in the context of unemployment durations.

\(^5\)We assume a zero degree of complementarity based on evidence at job loss (see Browning and Crossley (2001) and Bloemen and Stancanelli (2005)) and retirement (see Hurd and Rohwedder (2006) and Hurst (2008)). This evidence indicates that anticipated and unanticipated changes to labor supply do not lead to significant consumption changes.
the individual’s retirement strategy responds to changes in benefits, wages and wealth at a given age $a$. Differentiating the equation for $\tilde{v}_a$ with respect to $y^R_a$, $y^W_a$ and $A_a$ yields

\[
\begin{align*}
[y^R_a] & : \frac{d\tilde{v}_a}{dy^R_a} = -u'(c^R_a) < 0 \\
[y^W_a] & : \frac{d\tilde{v}_a}{dy^W_a} = u'(c^W_a) > 0 \\
[A_a] & : \frac{d\tilde{v}_a}{dA_a} = u'(c^W_a) - u'(c^R_a) \leq 0
\end{align*}
\]

\[\Rightarrow \frac{d\tilde{v}_a}{dy^R_a} = \frac{\frac{d\tilde{v}_a}{dA_a}}{\frac{d\tilde{v}_a}{dy^W_a}} = \frac{\text{wealth effect}}{\text{price effect}}.\]

The first condition reflects that a small increase in retirement benefits increases the value of retirement through increased current consumption during retirement. This change leads the individual to adjust his reservation level so he is more likely to retire. The second condition illustrates that a small increase in labor market income increases the value of continuing to work though increased consumption while working. This effect leads the individual to adjust his strategy so he is more likely to continue working. The last condition illustrates that a change in wealth affects both the value of retirement and the value of continuing work. The individual adjusts his reservation level to exactly offset the differences in marginal utility so as to remain indifferent between his options. In particular, with perfect credit and insurance markets, individuals would be able to smooth consumption so that $u'(c^R_a) = u'(c^W_a)$ implying $\frac{d\tilde{v}_a}{dA_a} = 0$. However, if an individual is constrained via the borrowing limit, $A_{a+1} \geq A_a$, then it will be the case that $y^R_a < y^W_a$ implies $u'(c^R_a) > u'(c^W_a)$ and hence $\frac{d\tilde{v}_a}{dA_a} < 0$. These conditions can be combined to decompose the labor supply response to a change in retirement benefits into a wealth effect and a price effect. The wealth effect reflects that the change in benefits creates additional wealth for the individual and the price effect reflects that the change in retirement benefits creates a change in the effective wage from working.

Given this decomposition, risk aversion can now be related to labor supply elasticities. In particular, taking a second order Taylor approximation, $u'(c^R_a) = u'(c^W_a) + u''(c^W_a)(c^R_a - c^W_a)$, the coefficient of relative risk aversion can be written as

\[
\gamma = \left( \frac{c^R_a - c^W_a}{c^W_a} \right)^{-1} \left( \frac{y^W_a}{A_a} \right) \left( \frac{\varepsilon^A}{\varepsilon^w} \right)
\]

where $\varepsilon^A = \frac{d\tilde{v}_a}{dA_a}$ denotes the wealth elasticity and $\varepsilon^w = \frac{d\tilde{v}_a}{dy^W_a}$ denotes the price (wage)
elasticity. This equation highlights that a larger wealth elasticity corresponds to a higher degree of curvature. Intuitively, if an individual changes his labor supply significantly upon receipt of an exogenous cash grant, this indicates that the marginal utility of consumption must decline relatively quickly, since otherwise the individual would have increased his consumption instead.

Thus far, we have related risk aversion to labor supply responses to changes in wealth and wages at a given age. Given that this relationship holds at every age, it is possible to generalize the intuition to relate risk aversion to the ratio of wealth and price elasticities based on changes in the profiles of benefits, assets and wages across multiple ages.7 Thus, even though pension reforms typically result in changes in the profiles of benefits across potential retirement ages rather than simpler changes in benefits at a retirement age, such reforms and the resulting wealth and price effects can still be used to identify the risk aversion parameter $\gamma$.

2.3 Discussion

A. Consumption Differences

The relationship between risk aversion and labor supply elasticities draws on concepts that are familiar in the context of studying insurance. In particular, the consumption difference $c^R_a - c^W_a$ reflects differences in consumption across states of employment at a given age. Even though retirement is an anticipated event in the model, the notion of insurance arises from the consideration of unanticipated changes (shocks) to benefits, wages and savings. Intuitively, the degree to which the individual adjusts his labor supply to move between states and offset the effects of these changes on his consumption will depend on how well insured his initial consumption is across the two states. The discrete nature of the employment states and the fact that wages while working are higher than benefits while retired ($y^W_a > y^R_a$) imply that consumption would always be higher if the individual worked

$^6$If $A_a = 0$, $\gamma$ can still be written in terms of the ratio of wealth and price effects,

$$\gamma = \left( \frac{c^R_a - c^W_a}{c^W_a} \right)^{-1} \left( \frac{d c_a / d A_a}{d y_a / d A_a} \right).$$

In this case, the wealth effect cannot be expressed as an elasticity.

$^7$To consider the intuition behind a change in the profile of wealth, it is useful to consider an annuity that pays a constant amount at each age. The wealth effect based on a change in the profile of benefits across multiple ages can then be defined based on a change in the constant amount paid at every age by the annuity. The price effect is based on a change in the individual's wage profile. Chetty (2008) presents a derivation to consider changes across multiple periods using such an annuity.
rather than retired \( (c^W_a > c^R_a) \), though the individual must also take into account the disutility of work. This consumption difference between states does not capture intertemporal consumption differences.\(^8\)

### B. Preference Assumptions

Given the parsimonious nature of the theoretical framework above, it is useful to characterize the conditions in which \( \gamma \) cannot be directly identified using the wealth and price elasticities. We start by noting that a key assumption that allows for the identification of a unique curvature parameter based on the participation outcomes is additive time-separability. Specifically, once utility is assumed to be additively time-separable, it is no longer possible to reconcile any monotonic transformation of the within-period utility function with a given ratio of labor supply elasticities.

While the model above assumes within-period additive separability between consumption utility and the disutility of work, this assumption is not essential for the relationship between \( \gamma \) and the wealth and price elasticities to hold. Specifically, with non-separable within-period utility \( u(c, v) \), one can derive a similar formula for risk aversion as above, but an additional term arises due to the complementarity between between \( c \) and \( v \) (based on the derivative of the marginal utility of consumption with respect to work disutility, \( u_{c,v} \)). In this case, a bound on this degree of complementarity can be assumed so that risk aversion can still be estimated using the wealth and price elasticities (see Chetty (2006)).

Another important assumption regarding preferences relates to state-dependence. To estimate \( \gamma \) based on movements between the states of employment, we must assume that \( \gamma \) is state-independent. In particular, the within-period utility function is assumed to be common across the states of work and retirement (i.e., we assume utility \( u(c) \) for work and retirement rather than allowing for \( u^W(c) \) while working and \( u^R(c) \) while retired). Intuitively, if preferences are state-dependent, the movements between states involve both changes in preferences and changes in the marginal utility of consumption, and the two cannot be disentangled using only participation outcomes.

\(^8\)Several papers have focused on “the consumption drop at retirement.” See Banks, Blundell and Tanner (1998), Bernheim Skinner and Weinberg (2001), Aguiar and Hurst (2005), Hurd and Rohwedder (2008), Blau (2008) and Hurst (2008) for work on this topic. The phrase “consumption drop at retirement” refers to intertemporal consumption changes, i.e., changes in consumption before and after retirement. With concave utility, intertemporal optimization predicts a smooth consumption profile over time. The same intertemporal consumption smoothing behavior holds in the model developed in this paper in the case where consumption while working is chosen optimally rather than via a fixed savings rule. In this case, however, there is still a consumption difference between the states of employment and retirement at a given age.
C. Heterogeneity

Under the preference assumptions discussed above, the model could allow for additional features relating to individual-level heterogeneity in preference parameters or consumption rules. First, in regard to individual-level heterogeneity, consider an example with individual-specific discount factors ($\beta_i$ for the $i$th individual). In this case, the ratio of wealth and price elasticities would vary by individual, but each individual’s ratio could be used to identify the common risk aversion parameter $\gamma$. More generally, one could allow the risk aversion parameter to vary by individual as well. With such heterogeneity, as long as the differences are state-independent, a similar relationship between the risk aversion parameter and the labor supply elasticities holds for each individual. Aggregating across individuals then, one could relate the mean $\gamma$ to the mean elasticity ratio. Such aggregation across individuals will be the focus of the empirical analysis below.

3 Institutional Background & Data

This section discusses the Austrian pension system and the pension reforms that create variation used for identification of elasticities and structural parameters. After discussing the institutional setting, we describe the data and construction of key variables used in the empirical analysis.

3.1 The Austrian Pension System

The Austrian pension system consists of two primary components: government-provided retirement pensions and employer-provided severance payments. A potential third component, private retirement pensions, is virtually non-existent.

We start our description of the pension system with the government-provided pensions as these benefits will be the focus of the empirical analysis. We focus on two forms of retirement pensions: disability pensions and old-age pensions. These pensions are computed based on similar rules. Specifically, an individual’s pension is the product of two elements. The first element is the assessment basis, which corresponds roughly to the average indexed monthly earnings (AIME) used in social security computation in the United States. The assessment basis refers to the last 15 years of earnings. After applying the earnings caps to earnings in each year, the capped earnings in each year are re-valued based on wage adjustment factors. These revaluation factors are intended to adjust for wage inflation so
that existing pensions grow in accordance to wages. After applying the revaluation factors, the capped, revalued earnings are averaged to determine the assessment basis. The second element, the pension coefficient, is then applied to the assessment basis to determine the actual pension level. The pension coefficient corresponds to the percentage of the assessment basis that the individual receives in his pension. This percentage increases to a maximum of 80% based on the number of insurance years and the retirement age. Insurance years correspond to periods of employment as well as periods of unemployment, military service and similar periods of labor market participation. Prior to 2001, disability pensions are computed identically to old-age pensions. In 2001 and after, the pension coefficient used in the disability pension is reduced relative to that of the old-age pension. The reduction in the disability pension coefficient is based on insurance years with lower insurance years receiving larger reductions.

Eligibility for the pensions is as follows. Disability pensions can be claimed at any age, provided that the claimant has been classified as disabled. Generally, an individual is classified as disabled if his working capacity is reduced by more than 50% relative to another individual of similar education. By claiming a retirement pension, the individual essentially exits the labor market. Men are first eligible for old-age pensions at age 60. In addition to being at least age 60, an individual who claims an old-age pension prior to the statutory retirement age, 65, must have 37.5 insurance years or 35 contribution years (years of contributions to the pension system).

Figure 2A presents retirement hazard rates by age. In this figure, retirement hazard are based on claiming either an old-age pension or a disability pension. In this figure, the hazard rates spike at ages 60 and 65 at roughly 80% and 75% respectively; these ages correspond to the minimum early retirement age and the statutory retirement age respectively. To better characterize the population remaining in the labor market, Figure 2B presents the survival function. This figure also highlight the large fraction of individuals leaving the labor market at age 60 and 65 with significant declines at these ages. Importantly, this survival function also highlights a significant amount of retirement prior to age 60. In particular, just under 40% of the sample retires prior to the early retirement age by claiming disability pensions. Figure 2C focuses more directly on disability pensions by presenting the survival function for

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9In our sample, roughly 9% of individuals continue some work within the year after claiming a pension. About 3.5% of old-age-pension claimants continue work, while 12% of disability claimants continue work. After claiming a pension, there is a mandatory 6 month break required to continue work with the same employer, and additionally, there are minimum earnings restrictions. As a result, we focus on the pension claiming decision as an exit from the labor market.

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individuals who claim disability pensions. This figure further emphasizes that individuals enter disability pensions primarily before age 60 and then less so after age 60 since the minimum age for old-age pensions has been passed.

While the government-provided retirement pensions represent the primary component of retirement benefits in Austria, the employer-provided severance payments represent a secondary source of retirement benefits. Individuals qualify for severance pay from their employers at the time of retirement if the individuals have accumulated sufficient years of uninterrupted tenure by retirement. The amounts of the payments are based on tenure as follows: 0 to 9 years of tenure yields no payment, 10 to 14 years of tenure yields a payment of one third of the last year’s salary, 15 to 19 years of tenure yields a payment of one half of the last year’s salary, 20 to 24 years of tenure yields a payment of three quarters of the last year’s salary and 25 or more years of tenure yields a payment of the full last year’s salary. In Section 4, we discuss in more detail how we account for incentives from these severance payments by controlling for tenure.

3.2 Pension Reforms

Between 1984 and 2003, there were five significant pension reforms in Austria in 1985, 1988, 1993, 1996 and 2000. Our detailed knowledge of these reforms and the computation of the pensions is based on Marek (1985, 1987-2003).\textsuperscript{10} Table 2 presents a summary of each reform. Figures 3A & B present benefits-versus-age profiles for different calendar years. These illustrations are meant to demonstrate the variation in pension benefits created by the pension reforms. To make benefits comparable across calendar years, these annual benefits are computed based on an individual who earns the nominal equivalent of 20000 euros in 2003 in each year that he works, and nominal benefits are then adjusted by the CPI to put all benefits in 2003 euros. In these figures, full insurance years (i.e. experience = age - 15, where 15 is the age corresponding to the end of mandatory schooling) is assumed for each age. As a result of this assumption, individuals reach the maximum insurance years at age 60 (45 insurance years) leading to a kink in the benefit schedule at age 60. Additionally, to emphasize some of the early pension reforms, benefits are computed assuming that the individual has worked for the last 10 years at each age, but not prior to that. Finally, to emphasize the changes in incentives for individuals due to the pension

\textsuperscript{10}Ney (2004) and Linerooth-Bayer (2001) provide information on the historical contexts of the reforms. See also Koman, Schuh and Weber (2005) and Hofer and Koman (2006) for studies of the Austrian severance pay and pension systems respectively.
reforms, the benefit-retirement age profiles in each year are computed for a fixed birth cohort (i.e. birth cohort = year - 60). Thus, taking the pension reforms as unanticipated, comparisons across retirement ages within a given calendar year reflect the incentives to retire at different ages, and comparisons across the calendar years reflect changes in these incentives due to the reforms.

The pension reforms generally reduced the generosity of the retirement pension system as government officials felt the pension system was not financially sustainable. This trend is evident by the downward trend in benefits across calendar years illustrated in Figures 3A & B. The pension reforms in the 1980s reduced benefits through changes in the length of the assessment basis. The 1985 reform changed the assessment basis from the last 5 years of an individual’s earnings to the last 10 years. Because wages are generally increasing with age in Austria, this change decreased benefits. The reform was implemented at the start of the 1985 calendar year. The 1988 reform changed the length of the assessment basis from the last 10 years to the last 15 years. This change was phased in between 1988 and 1992 based on birth cohort. Specifically, the legislation determined the length of an individual’s assessment basis based on the year the individual reached age 60. The 1985 and 1988 pension reforms are illustrated in Figure 3A. In particular, benefits decrease between 1984 and 1988 due to the first increase in the length of the assessment basis from 5 to 10 years. Benefits decrease each year from 1988 to 1992 as the second increases in the assessment basis from 10 to 15 years is phased in. As illustrated, these reforms decreased the levels of benefits across potential retirement ages but left the slopes in the profiles unchanged.\(^{11}\)

The reforms in the 1990s continued the reduction in benefits and also specifically aimed to get individuals to retire at later ages. The 1993 reform linked pension coefficients to retirement ages so that the coefficients would rise with both insurance years and retirement ages up to the statutory retirement age, 65.\(^{12}\) The 1993 reform also changed the assessment basis from the last 15 years of earnings to the highest 15 years of earnings. However, this change generally did not affect retirement pension benefits; since wages generally rise with age, the best 15 years of earnings correspond to the last 15 years of earnings for most individuals. This aspect of the reform is likely to have been more relevant for other non-retirement pensions that are also based on an individual’s assessment basis. These changes

\(^{11}\)The increase in the slope of the benefits profile in 1992 at age 65 corresponds to the introduction of a bonus for retirement at the normal retirement age in 1991.

\(^{12}\)Since benefits are computed assuming full insurance years at each age, the illustrated profiles already link benefits to insurance years and retirement age. As a result, the change in benefits due to this aspect of the 1993 pension reform are not evident in the figures.
from the 1993 reform became effective at the start of the 1993 calendar year.

The 1996 and 2000 reforms also focused primarily on changes in pension coefficients. Figure 3B focuses on these later reforms. Specifically, the 1996 reform introduced a bonus/malus system to discourage early retirement (before the statutory age) by penalizing early retirees with reduced pension coefficients. As illustrated in Figure 3B, the comparison between the 1996 (pre-reform) and 1998 (post-reform) benefit profiles highlights the impact of the introduction of the bonus/malus system on retirement incentives. Specifically, this reform decreased the levels of benefits at early retirement ages (the malus) and then increased the slope in the benefit profiles (bonus) to provide increased incentives for later retirement. The 2000 reforms further developed the bonus/malus system by increasing the reductions in pension coefficients for early retirements and also by offering bonus increases in pension coefficients for retirements after the statutory ages. The 2000 reform also affected eligibility by raising the minimum retirement age from 60 to 61.5. The increase was phased-in between October of 2000 and October of 2002. As illustrated, nominal adjustments in later years were lower than inflation so that real benefits declined between 1998 and 2002.

3.3 Data

We use social-security records data from the Austrian Social Security Database, provided by *Synthesis Forschung*. This administrative data covers nearly all individuals employed in Austria between the years 1972 and 2003, with the exceptions relating to tenured public sector employees and self-employed individuals. Observations are in the form of spells that are individual-specific, time-specific and insurer-specific. In the cases of employment, the insurer corresponds to the employer, while in the cases of non-employment such as unemployment or disability, the insurer corresponds to the government agency providing assistance. The time-specific characteristic of an observation means that an observation begins either at the beginning of a new spell (a new individual-insurer match) or on the 1st of January of a year. An observation ends either when that particular spell is terminated during a year, or on the 31st of December of a year.

In addition to being characterized by begin dates and end dates, each spell is also characterized by type. The type of spell refers to a more specific classification within the main categories of employment, unemployment, retirement, and maternity leave. For each spell,

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13 Tenured public sector employees are observed only starting in 1988 or in some cases 1995, and income is not observed for self-employed individuals.
the amount of earned income during the length of the spell is recorded. Specifically, if the spell corresponds to receiving social insurance, no income is recorded for the spell. Income data is top-coded based on the earnings cap for retirement pension computation. Importantly, the social security record data contains all information used in the computation of retirement benefits except insurance years which we are able to impute using the labor market histories.\textsuperscript{14}

The data include some variables specific to individuals and insurers. For each individual, the data include gender, birth date, and nationality. For each of the employers (these may correspond to firms or plants), the data include region and industry classifications. Using the available data on employees and employers, we construct firm-specific tenure.

Our main sample consists of men ages 55 through 65 who are first observed at age 55 in the years 1984 through 2003. Our sample restrictions and the reasons for these restrictions are as follows. We start by focusing on men aged 55 or higher in 2003 (birth cohorts 1948 and earlier). We exclude individuals with less than one year of observed employment time between 1972 and 2003 since these individuals lack sufficient data to compute pension benefits. Next, we exclude foreign nationals as well as those who have spent more than a year as self-employed or as tenured public servants, farmers, or in mining, construction, and railways since these individuals are covered by separate pension systems. Additionally, we exclude individuals who claim non-disability or non-old-age pensions at the time of retirement since these claims may not correspond to retirement decisions.\textsuperscript{15} We exclude men claiming disability pensions before age 55 on the basis that these individuals are likely to be permanently disabled. We also exclude individuals who retire after age 65 since focusing on ages 55 through 65 simplifies recursive computation of the value functions and most retirees (roughly 99%) retire by 65. Next, we exclude remaining individuals with insufficient earnings histories to compute pension benefits and individuals with outlying

\textsuperscript{14}Insurance months are determined using the following imputation for insurance years. Specifically, insurance years are imputed as $\text{InsYrs} = \text{Age} - \text{Edu} - 6 - (\text{time observed not working})$ where $\text{Edu}$ is years of schooling. We observe education for the sample of individuals who experience unemployment and claim benefits during the length of the data. Using this data, we regress education on earnings and quartic polynomials in calendar year and age. We then obtain imputed education using the fitted values from this regression. Using the labor market histories observed in the data, we compute $\text{time observed not working}$. Assuming that education begins at age 6, we combine the predicted education with this information from the observed labor market histories and round up to the nearest year to compute insurance years (years of experience). Insurance months are then given by $\text{InsMths} = 12 \times \text{InsYrs}$.

\textsuperscript{15}The types of pensions claimed are identified in the data. At the time of retirement, other pensions based on, income status, widow status or chronic unemployment may be claimed. We identify men claiming these types of pensions and exclude them from our sample.
observations and missing data. This leaves us with a final sample of 252,907 individuals and 178,997 claimants. Not all individuals are observed to be claimants since some individuals (those at later calendar years) are only observed at younger ages. The sample restrictions are summarized in Table 1.

Using the social security record data, we construct four key variables to capture incentives from the government-provided pensions and the employer-provided severance payments. The first two variables, social security wealth (SSW) and the accrual (ACC) correspond to the pensions. An individual’s social security wealth at a given age is defined as the expected present discounted value of his annual pension benefits if he were to retire at the given age. More precisely, we can write SSW as

$$SSW_{i,a} = \sum_{t=a}^{100} \beta^{t-a} \pi_{t|a} b_i(a)$$

where $b_i(a)$ denotes individual $i$’s annual benefits when retiring at age $a$, $\pi_{t|a}$ denotes the probability of survival until age $t$ conditional on having survived until age $a$ and $\beta = .93$ captures the individual’s discount factor.\(^{16}\) In this definition, we also assume that the maximum age that individuals can live to is age 100. Each individual’s retirement pension is calculated based on the rules of the Austrian pension system and the individual’s observed earnings history. While the social security wealth variable reflects the levels of benefits, the second pension variable, the accrual, reflects the slope of the benefits schedule across potential retirement ages. In particular, an individual’s accrual at a given age $a$ captures the expected change in his social security wealth $SSW_{i,a}$ net of pension contributions from delaying retirement by one additional year. Thus we define the accrual for individual $i$ at age $a$ as

$$ACC_{i,a} = \frac{E_a(SSW_{i,a+1}) - SSW_{i,a}}{SSW_{i,a}}.$$ 

In calculating the individual’s expectation, we assume 1.5% real wage growth to project earnings one year ahead.

Table 3 presents summary statistics by age for key variables from the data used in the empirical analysis. All euro amounts are in 2003 euros; in January 2003, the euro-U.S. dollar exchange rate was 1 euro to roughly 1.06 dollars. The statistics at each age are based on individuals who are not yet retired (i.e. still in the labor market) at the given age, so

\(^{16}\)The survival probabilities are taken from life tables available through Statistics Austria (www.statistik.at). The value of $\beta$ corresponds to a real interest rate of roughly 7.5% which is consistent with the long-term real interest rate in Austria in the mid-1990s.
selection should be taken into account when interpreting profiles across ages within the table. At age 55, the sample consists of 242,402 individuals in the labor market. For this group, the median earnings are roughly 28,000 euros and the median annual benefits are roughly 14,000 euros implying a replacement rate of about 50%. Median earnings increase across the ages indicating that higher income earners tend to retire later. Annual earnings are computed based on the calendar year that an individual reaches the specified age, and this accounts for the earnings dips at ages 60 and 65 because individuals at these ages work only part of a calendar year and then retire once they reach either age 60 or 65. Based on the annual benefits, survival probabilities, and a discount factor of $\beta = .93$, median social security wealth ranges from about 215,000 euros at age 55 to 275,000 euros at age 65 reflecting that higher earners who have yet to retire at the later ages have higher social security wealth. The accrual is close to -10% at each age reflecting the loss in social security wealth from lack of actuarial adjustments. Additionally, the accrual becomes slightly more negative after age 60 reflecting that higher income earners give up more of their social security wealth when they delay claiming their pension.

Asset data is also important for the empirical analysis. Because such wealth data is not available in the social security records data, we use asset data from the Survey of Health, Ageing, and Retirement in Europe (SHARE).\textsuperscript{17} This SHARE dataset has wealth data for individuals in several European countries. We focus on the data collected for Austria in 2005. In particular, we use data on household gross financial assets for 1,391 Austrians ages 50 through 54 in 2005. We present summary statistics characterizing this distribution of assets (in 2003 euros) in the bottom section of Table 3. The data indicates that households have accumulated financial assets roughly equivalent to one-year’s earnings. We discuss how this asset data and the other variables summarized in Table 3 are used in the empirical analysis in more detail below.

4 Reduced-Form vs. Structural Estimates

Before presenting the reduced-form estimates, we present graphical evidence in Section 4.1 illustrating the identifying variation from the pension reforms. In Section 4.2, we specify the reduced-form model, and explain how we use the variation in pension reforms to identify the income and price elasticities using a control function strategy. We then discuss how we use these reduced-form estimates to construct an estimate the coefficient.

\textsuperscript{17}Information on the SHARE dataset can be found at http://www.share-project.org/.
of relative risk aversion using the formula derived in Section 2.2. Finally, in Section 4.3 we present estimates of the structural model using a Method of Simulated Moments estimation strategy based on matching retirement hazard rates by age. We show that the different empirical approaches, based on different sources of identifying variation, yield incompatible estimates of the coefficient of relative risk aversion, even though they are based on the same theoretical model.

### 4.1 Graphical Evidence

Our identification strategy exploits policy variation based on a series of pension reforms in Austria that independently varied the level and slope of pension benefits across ages. Figure 4 presents three time-series for individuals at age 55. The first time-series is the mean accrual at age 55. The second time-series is the median change in social security wealth, where changes are computed relative to the previous year’s legislation. An increase in the first time-series reflects an increase in the price of retirement while a negative value for the second time series reflects a decrease in pension wealth at retirement. The final time-series is the retirement hazard for individuals at age 55. This figure concentrates on individuals at age 55 since the current discussion will be based on two particular pension reforms in 1988 and 1996 that first affect individuals at age 55.

We consider first the identification of income effects from pension benefits on retirement decisions based on the changes in pension wealth. The 1988 pension reform creates variation in pension wealth since the reform phased in a five-year increase in the length of the assessment basis from the last 10 years to the last 15 years of earnings. Since earnings further back in the earnings history are generally lower (i.e., earnings are generally increasing with age), this increase in the length of the assessment basis lowers pension wealth. As illustrated in Figure 4, median pension wealth decreases by roughly 1% with each additional year for the assessment basis. Notice that this reform only affects the level of pension wealth as the accrual is unchanged. Focusing on the retirement responses, the retirement hazard time-series has only a slight decrease at the time of the reform, and this decrease does not persist over the entire phase-in. The lack of distinct changes in the retirement hazard indicate that the wealth effects from pension benefits are likely to be relatively small.

Next, we consider the 1996 pension reform which creates both income and price effects of pension benefits on retirement decisions. This reform increases the penalties for early
retirement (retirement before the statutory age, 65). As a result of these penalties, the mean accrual increases between 1995 and 1997 from roughly -.096 to -.087, reflecting a higher price of retirement. Additionally, the penalties for early retirement reduce pension wealth. Relative to the pre-reform legislation, pension wealth decreases by roughly 0.05 after the reform. While the 1988 pension reform indicates that wealth effects are likely to be relatively small, the 1996 reform indicates that the price effects are likely to be relatively large. Specifically, with this reform that includes price changes in addition to the wealth changes, the hazard falls sharply at the time of the reform from roughly 0.10 to 0.03. These graphical results imply very large elasticities. While it is possible that individuals at age 55 are particularly responsive to financial incentives since age 55 is the first possible age for retirement, it is also possible that there are other confounding changes that make a causal interpretation of the implied elasticities tenuous.

The key to the identification strategy is that the pension reforms create exogenous variation in pension wealth and the accrual that is independent across the reforms. In particular, notice that it is not essential that one pension reform affects only pension wealth while another reform affects both pension wealth and the accrual. This example is simply a special case of independent variation in pension wealth and the accrual across two pension reforms. In the regression analysis below, we pool the exogenous variation in pension wealth and the accrual across the five reforms and across multiple retirement ages to precisely identify the income and price effects from pension benefits on retirement decisions.

While Figure 4 focuses on changes at age 55 to avoid complications from survival bias (recall that age 55 is the first age for retirement), we next turn to illustrating the identifying variation from the pension reforms across all ages and years. To do this, we first regress a retirement indicator, the log of social security wealth and the log of the accrual each on earnings history polynomials and age, year, industry, region, blue and white collar, and change-in-eligibility dummies.\footnote{The regressions are of the form \( Y_{ia} = \delta X_{ia} + \varepsilon_{ia} \) where the subscripts refer to individual \( i \) at age \( a \). The change-in-eligibility dummies capture changes in the mechanical rules governing retirement that are independent from changes in financial incentives. As documented in Table 1, these changes are (1) the introduction of a disability pension at age 57 between 1993 and 2000, the increase in the retirement age from age 60 to 61.5 between 2000 and 2002 and (3) the increased restrictions for claiming disability after 2000.} We then obtain the residual for these three variables and create year-age cell means for each variable. We then plot these cell means in Figure 5. By
controlling flexibly for year, age and income groups, the remaining variation in the residuals comes at the level of year-age-income group interactions. This is the level of variation from the pension reforms which differentially impact different income groups at different ages in different years.

Consistent with Figure 4, the plots in Figure 5 indicate relatively smaller wealth effects and larger price effects. Specifically, the scatter plots show a steeper slope with the accrual residuals than with the social security wealth residuals (−4.784 versus 0.717). Furthermore, in a bivariate regression using the cell means, the estimated coefficients (and standard errors clustered at the year level) on the social security wealth and accrual residuals are respectively 0.849 (0.255) and −4.939 (0.799) respectively. These estimates indicate a wealth -to-price elasticity ratio of roughly 0.17. We now turn to estimating these elasticities more directly in the context of a Cox proportional hazards specification.

4.2 Hazard Model Estimation & Results

To determine the income and price elasticities of retirement benefits on retirement age, we estimate the following Cox Proportional Hazards model on men between the ages of 55 and 65 between 1984 and 2003,

\[ R_i(a) = \tilde{R}(a) \exp\{\beta_1 \ln(SSW_{i,a}) + \beta_2 \ln(ACC_{i,a}) + \delta X_{i,a}\}. \]

In this specification, \( R_i(a) \) denotes the relative hazard for individual \( i \) at age \( a \). The relative hazard is the probability that individual \( i \) retires at age \( a \) conditional on not having retired at an earlier age relative a baseline probability across all individuals at age \( a \). The term \( \tilde{R}(a) \) denotes the baseline hazard rate at age \( a \). This baseline hazard is common across individuals at each age and thus the intuition regarding the baseline hazard closely follows the intuition of age fixed effects in a linear model. As defined in Section 3, \( SSW_{i,a} \) is the expected present value of the individual’s retirement pension if he were to retire at age \( a \), and \( ACC_{i,a} \) is the individual’s expected pension accrual (i.e., the change in \( SSW_{i,a} \) from delaying retirement by an additional year). The term \( X_{i,a} \) refers to covariates for individual \( i \) at age \( a \). We include a base and full set of controls. The base controls include quartic polynomials in calendar year, log annual earnings and log total earnings from the prior 10 years to control for individuals’ earnings histories. The full controls include the base controls as well as dummies for education, industry and region, and quartic polynomials in log annual earnings from each of the prior 10 years. We also
include a quartic tenure polynomial to control for potential heterogeneity in preferences for work that may be correlated with higher levels of job tenure.

This empirical model is based on previous work in the literature. Lumsdaine, Stock and Wise (1992), Coile and Gruber (2000a,b), Gruber and Wise (2004) and others have primarily estimated probit and linear probability models relating pension incentives and retirement decisions. We focus on a hazard model to adopt a more dynamic perspective on each retirement decision as a stopping-time event following a duration of a career. Furthermore, the hazard model presents results precisely in terms of the elasticities we are interested in, whereas the alternative models present coefficients that cannot be easily converted into elasticities. In particular, the coefficients $\beta_1$ and $\beta_2$ relate to the income and price effects that identify the coefficient of relative risk aversion as discussed in Section 2.2. $\beta_1$ captures the elasticity of retirement with respect to pension wealth, and $\beta_2$ captures the elasticity of retirement with respect to the one-year accrual. Notice that in this specification we have defined the price of retirement based only on looking ahead one year. To the extent that individuals are more forward-looking, the price of retirement should take into account the profile of benefits are multiple future ages. The structural model outlined in Section 2.1 captures such forward-looking behavior.

We exploit exogenous variation in retirement benefits created by the five pension reforms in Austria between 1984 and 2003 to identify a causal relationship between retirement benefits and retirement decisions. Specifically, we employ control function methods to use only the variation in pension benefits created by the reforms to identify $\beta_1$ and $\beta_2$ (Heckman and Robb 1985). Without the exogenous variation from the reforms the identification of causal effects is threatened by unobserved heterogeneity in preferences for work. Intuitively, individuals with greater willingness to work may have higher earnings and hence higher pension benefits, thereby creating a correlation between benefits and retirement decisions. We include polynomials in individuals’ earnings histories to control for systematic variation in pension benefits based on earnings histories. Additionally, the baseline hazard controls for changes in the pension benefit schedule that are common across ages. Thus, only the remaining variation in pension benefits, due entirely to the pension reforms, is used to identify the pension wealth and accrual elasticities. In addition, we are able to separately identify both the income and price effects because we observe multiple pension reforms that create independent variation in the level and slope of benefits across retirement ages.

The results from the Cox Proportional Hazards model are presented in Table 4. The first two columns present estimates of the coefficients on log Social Security Wealth (SSW) and
the log accrual rate (ACC) estimated on the entire sample with the base and full controls, respectively. The base results indicate that a 1% increase in pension wealth increases the hazard by 0.44% while a 1% increase in the accrual measure decreases the hazard by roughly 2.5%. After including the full control set, the pension wealth estimate decreases slightly to 0.4% while the estimate for the accrual increases to 3%. Consistent with the graphical evidence presented above, we estimate much higher price effects than wealth effects, on the order of 6-7 times higher.

Recall the hazard rates into retirement were characterized by spikes at ages 60 and 65. In the next two columns, we estimate the model on the sample of individuals 60 and 65 only in order to examine the importance of the proportionality assumption (i.e., that covariate effects are proportionate across ages). Note that the effect of pension wealth is estimated to be slightly smaller at these ages and the effect of the accrual slightly larger, however these differences are not statistically different from the estimates on all ages. Finally, the fifth and sixth columns present estimates of the model allowing for time-varying covariate effects. Specifically, we allow the effects to vary linearly with age. To obtain the estimated effect of a covariate at a given age, multiply the coefficient by age minus 54. For example, the estimated effect of ln(SSW) at age 60 is 0.1095 * (60 - 54) = 0.657. The corresponding estimate of the accrual effect is -2.762. Note that these estimates are similar across all specifications, and in particular the ratio of the wealth to price effect is small. As a result, we will consider the coefficients from the base model estimated on all ages (column 1) to be our baseline estimates.

Recall from Section 2.2 that the coefficient of relative risk aversion can be expressed as the product of three components: (1) the ratio of the income to price elasticities; (2) the wage-to-asset ratio; and (3) the inverse of the consumption difference between work and retirement. Table 5 presents the implied coefficient of relative risk aversion for the range of elasticity ratios reported in Table 4, assuming a wage-to-asset ratio of 1.12 and consumption difference of 0.395. We set the wage-to-asset ratio to 28000/25000, where 28000 is roughly the median wage in our sample and 25000 is the median level of assets in the SHARE data. To benchmark the consumption difference, we use median annual benefits plus 5% of savings while retired (14000+0.05*25000) and median annual benefits net of 10% savings while working (.9*28000). These figures come from the summary statistics shown in Table 3. From Table 5, we see that under our baseline estimates of the income and price elasticities, the implied coefficient of relative risk aversion is relatively small at 0.43. Under a plausible range of elasticity ratios, we see that the formula implies an upper bound of 1
for $\gamma$.

### 4.3 MSM Estimation & Results

In this section, we describe the structural approach to estimating the coefficient of relative risk aversion in the model outlined in Section 2.1. Following French (2005), we fix a number of parameters governing the data generating process of the exogenous state variables ($\chi$), and estimate the set of preference parameters $\theta$ (which includes $\gamma$) conditional on these values. In particular, in the baseline specification we fix the life span $T = 100$ years, the savings rate at $\bar{s} = 10\%$ (based on data from the Austrian Survey of Health, Ageing and Retirement in Europe (SHARE)), the real wage growth rate $g = 1.75\%$ (estimated from the social security records data on individuals ages 50-54), and the interest rate $r = 7.5\%$ (based on inflation minus nominal interest rates in Austria during the sample period). We obtain mortality probabilities $\pi_{a|a-1}$ from life tables for Austria, and we estimate job separation probabilities $\pi_{sep}$ directly from our data. Finally, we also fix the discount factor $\beta = 1/(1 + r) = 0.93$ since it is difficult to distinguish empirically from declining work disutility over time. Thus, $\chi = (T, \bar{s}, g, r, \pi_{a|a-1}, \pi_{sep}, \beta)$. Since we do not observe data on assets, we approximate the initial distribution of assets at age 54 using the Austrian SHARE data; specifically, we sample initial assets for each individual with replacement from the empirical distribution of assets in SHARE.

We parameterize the model presented in Section 2.1 as follows. We assume constant relative risk aversion (CRRA) utility over consumption:

$$ u(c) = \frac{c^{1-\gamma}}{1-\gamma}, \quad \gamma > 0. $$

We assume initial work disutility is drawn from an exponential distribution with mean $\tilde{\eta} = x\eta$, where $\eta > 0$ and $x = u(\bar{c}) - u(\bar{r}\bar{c})$ is a scaling factor for the disutility of work based on income differences between work and retirement (we use $\bar{c} = 30000$ and $\bar{r} = .55$ based on mean wage income and the replacement rate). Work disutility increases linearly with age, with slope $\alpha\tilde{\eta}$ (i.e., $v_a = \alpha\tilde{\eta}(a - 54) + v_{54}$). Thus, the parameters we are interested in estimating are $\theta = (\gamma, \eta, \alpha, \kappa)$, where $\kappa$ equals the monetary cost of claiming a disability pension.

For the estimation, we assume that individuals make decisions with complete knowledge of how pension benefits are calculated in a given calendar year. We assume that their
projections of future benefits are based on that year’s legislation only. Further, we assume that the pension reforms were unanticipated, and that individuals immediately update their calculations based on the new rules. We assume that individuals expect their future earnings to grow at a constant rate per year. In regard to job separations, we assume that the probability of job separation varies only by years of tenure. We assume that separation shocks do not affect wages, so that conditional on separations, wages are still expected to grow at the same constant rate (this is supported by evidence that collective bargaining agreements tend to set wages based on labor market experience rather than tenure). This simplifies the computation of projected pension benefits since we can project pension benefits for individuals at each age based on a single expected earnings path rather than based on multiple paths from different potential histories of job separations.

A common method for estimating θ is the Method of Simulated Moments (MSM). Moment-based estimation strategies match key moments (e.g., retirement hazard rates by age) observed in the sample data with the analogous moments implied by a model parameterized by θ. The goal is to find the value of θ which gives the best “fit” of the model, i.e., by minimizing the (weighted) distance between the observed and predicted moments. Where estimation of the predicted moments is computationally intractable by conventional methods, simulation methods must be employed. MSM approximates such moments using Monte Carlo integration, i.e., by averaging over simulations of the model. Assuming the moment conditions are correctly specified, MSM is consistent for a fixed number of simulations. To estimate the structural model, we employ MSM where we match observed retirement hazard rates by age (weighted by the survival function at each age).

Table 6 displays the structural parameter estimates based on the MSM strategy. Column 1 presents the estimates for the baseline model where χ is fixed as described above. Notice that the structural estimate of the coefficient of relative risk aversion is much higher than 1 at roughly 1.85. Columns 2 and 3 present alternative estimates varying the discount factor β and savings rate s, respectively. Figure 6 presents the actual vs. predicted hazard rates by age for each specification. Note that the three specifications all fit the observed retirement hazards roughly equally well, with the best fit for ages 60 and earlier where the estimation places the most weight.

Column 2 of Table 6 demonstrates that it is difficult to empirically distinguish the discount factor β from the slope of work disutility across ages; in particular, a lower discount factor (less patience) implies that work disutility increases at a slower rate. The cost of
claiming disability also appears to vary with the discount factor; a five percentage point reduction in the discount factor to $\beta = 0.88$ is associated with an increase in the cost of claiming from just under 50% to roughly 55% of the average annual wage at age 55. Further, while it does increase slightly with the decrease in $\beta$, the coefficient of relative risk aversion remains well-above 1.

Since we do not observe data on assets, we cannot match moments based on savings decisions. Thus, identification depends solely on variation in labor supply decisions at the retirement margin, and our results are conditional on the fixed savings rate. Note that the savings rate determines the consumption difference between the states of employment through two mechanisms: first, a lower savings rate mechanically gives higher consumption while working; and second, a lower savings rate implies that less wealth is accumulated at retirement, which decreases consumption at retirement. Thus, a lower savings rate is associated with a larger consumption difference. However, a lower savings rate also implies a higher the wage-to-asset ratio. Since $\gamma$ is inversely related to the consumption difference but is directly related to the wage-to-asset ratio, a lower savings rate has a theoretically ambiguous effect on $\gamma$. Column 3 of Table 6 presents estimates of the structural model assuming a savings rate of 5% in contrast to the baseline savings rate of 10%. Note that $\gamma$ is estimated to be slightly higher than the baseline estimate under the assumption of the lower savings rate. However, the change in the savings rate is not associated with any statistically significant departures from the baseline parameter estimates.

Figure 6 presents plots the predicted and actual hazard rates by age. First, in terms of goodness-of-fit, the figure indicates that the model does fairly well in fitting the across-age moments, especially at early ages. The fit at later ages receives less weight in the estimation since there are relatively few individuals who continue to work after age 60, so this accounts for the larger difference between the actual and predicted moments at later ages. Second, in terms of comparisons across specifications, none of the alternative specifications of the model seem to significantly improve on the baseline model’s fit. This seems to emphasize

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To further explore the relationship between the fixed savings rate assumption and the estimates of $\gamma$, we have estimated two additional specifications of the model: (1) heterogeneous savings rates uniformly distributed between 0.05 and 0.15 and (2) income-correlated heterogeneous savings rates. Specifically, for this second specification, we formed quintiles based on real earnings at age 55. Next, for individuals in the first (lowest) real wage quintile, we drew savings rates from a uniform distribution between 5% and 7%. For the second quintile, savings rates were between 7% and 9%. The savings rates for the remaining quintiles were determined similarly up to a maximum of 15%, so the mean savings rate in the sample remained 10%. The estimates of the coefficient of relative risk aversion did not differ significantly from the baseline estimates in each of these specifications. Because of this lack of a change in the estimates, only the results with the change in the constant savings rate are presented.
that the baseline model is able to fit the across-age retirement pattern quite well.

To see whether the structural MSM estimates of \( \gamma \) are consistent with the reduced-form estimates presented above, we simulate the implied proportional hazard coefficients. Specifically, we simulate the retirement outcomes and earnings histories using \( \hat{\theta}_{MSM} \) and the baseline structural model. Using the simulated data, we construct the social security wealth and accrual measures and estimate the same hazard specifications as those used in the reduced-form analysis. These results are presented in Table 7. It is immediately clear that the estimated structural model overpredicts individuals’ responsiveness to financial incentives: the model implies that the income and price elasticities are roughly 7 and 2 times the actual estimates (see column 1 of Table 4), respectively. Similar to those from the baseline model, the hazard coefficients implied by the other specifications of the models are also significantly different from the reduced-form estimates.

Thus, we see that the structural estimates based on observational data on retirement hazard rates by age are inconsistent with the reduced-form estimates based on quasi-experimental variation from the Austrian pension reforms. In particular, the structural approach attributes too much of the variation in retirement patterns across age to responsiveness to differences in financial incentives. In the next section, we propose a new approach which exploits quasi-experimental variation in financial estimates for retirement to identify and estimate a structural model of retirement.

## 5 Indirect Inference: Reconciling Reduced-Form and Structural Estimates

In this section, we propose an estimation strategy based on the method of Indirect Inference (Gourieroux, Monfort and Renault 1993). Our approach contrasts with the common strategy of matching moments based on observational data on retirement patterns across ages. These studies suffer from the shortcoming that the variation behind such moments may not be exogenous. In particular, variation in pension benefits is driven by differences in individuals’ earnings histories, since pension benefits are a function of past earnings. An implicit assumption is that differences in individuals’ earnings histories are not related to unobservables (e.g., preferences for work). The approach we propose exploits quasi-experimental variation based on pension reforms as a new source of identification. Specifically, the reforms allow us to break apart the relationship between past earnings
and benefits. Thus, we are able to observe individuals with identical earnings histories facing different incentives for retirement. In Section 5.1, we discuss our proposed strategy for identifying and estimating the structural model of retirement outlined in Section 2 based on quasi-experimental variation from the pension reforms. We present our results in Section 5.2, and reconcile the Indirect Inference estimates with the reduced-form and structural estimates from Section 4.

5.1 Estimation Strategy

The method of Indirect Inference (II) can be described as follows. First, specify an auxiliary model. This model is ‘incorrect’ in the sense that optimization of the likelihood does not yield consistent estimates of the parameters of interest, $\theta$. Instead, it provides consistent estimates of some auxiliary parameters $\beta$ such that $\beta_N \overset{p}{\rightarrow} \beta(\theta_0)$ when the data are generated under the true value of $\theta = \theta_0$. This binding function $\beta = \beta(\theta)$ links the auxiliary (reduced-form) and structural parameters. II consists of finding the value of $\theta$ that minimizes the distance between $\hat{\beta}_N$ (the estimates on the observed data) and an estimate of $\beta(\theta)$. This estimate is obtained by simulating $S$ paths of the model conditional on $\theta$ and estimating the auxiliary model on the simulated retirement decisions of the pooled sample (of size $SN$). Thus, the II estimator is the following minimum distance estimator:

$$\hat{\theta}_{II} = \arg \min_{\theta} \left( \hat{\beta}_N - \tilde{\beta}_{SN}(\theta) \right)^T W \left( \hat{\beta}_N - \tilde{\beta}_{SN}(\theta) \right),$$

for some weighting matrix $W$. Gourieroux et al. (1993) show that the II is a consistent estimator of $\theta$ for $N \to \infty$. Identification requires $\dim(\beta) \geq \dim(\theta)$, and the binding function should be a one-to-one function mapping $\theta$ to $\beta$.

We apply a variation on the II estimator by combining it with the more traditional MSM estimator in order to identify and estimate all the parameters in $\theta$. Specifically, our II estimator matches the following moments: retirement hazard rates by age, and the coefficients $\beta_1$ and $\beta_2$ from the proportional hazard model above (where the coefficients are weighted by the inverse of their standard errors). We retain the retirement hazard rates as moments in order to identify the remaining parameters governing work disutility (level and slope) and the psychic cost of claiming a disability pension. Intuitively, while $\gamma$ (identified from the quasi-experimental variation based on the pension reforms - within age, across individuals) captures the financial incentives relating to retirement decisions, the remaining parameters (identified from observed retirement patterns across ages) capture non-financial
incentives relating to the evolution of disutility of work (including the discontinuity in psychic cost of claiming benefits) across ages. Since $\alpha$, $\eta$, and $\kappa$ soak up all of the variation in retirement hazards that is not related to the quasi-experimental variation based on the pension reforms (similar to a control function strategy), this ensures that the quasi-experimental variation identifying the income and price elasticities will be used to estimate $\gamma$.

5.2 Indirect Inference Estimates

Table 8, Panel A, displays the structural parameter estimates based on the II strategy. Column 1 presents the estimates for the baseline model where $\chi$ is fixed as described in Section 4.3. In particular, for this specification we assume a discount factor of 0.93 and savings rate of 10%. The estimated coefficient of relative risk aversion in the baseline model using the II strategy is roughly 0.55, which is very close to the reduced form estimate of 0.43. As was the case for the MSM estimates, $\hat{\gamma}$ increases with a lower discount factor and savings rate, but it is always less than 1. Since $\gamma$ is estimated to be lower to match the proportional hazard coefficients, the other parameters must adjust to match the observed retirement patterns across age. In particular, note that the II estimate of mean work disutility is much higher than the MSM estimate (0.79 vs. 0.14) while the slope is flatter (0.38 vs. 1.52). Finally, since differences between the levels of wages and benefits are smaller with a lower degree of risk aversion (i.e. $(15000)^\gamma - (10000)^\gamma$ is increasing in $\gamma$), a larger fixed cost $\kappa$ is required to fit the spike in the retirement hazard at age 60; indeed, $\kappa$ is estimated to be nearly three times as large as the corresponding MSM estimate (40077 vs. 13816).

Columns 2 and 3 present alternative estimates varying the discount factor and savings rate, respectively. As with the MSM estimates presented in Table 6, the estimates in column 2 of Table 8 illustrated the difficulty of separately identifying the discount rate and the slope in the disutility of work across ages. With a lower $\beta$, the estimated slope coefficient $\alpha$ decreases by roughly 50% from the baseline estimate and the fixed costs increases slightly to roughly 42,700. Intuitively, if the value of work at future ages is reduced because of a reduction in $\beta$, the model compensates by decreasing the rate at which work becomes less pleasurable. Column 3 of Table 8 presents the parameter estimates when the savings rate is reduced from the baseline value of 0.10 to 0.05. In this case the estimated $\gamma$ is slightly less than the corresponding baseline estimate while the estimated $\eta$ is slightly
higher than the corresponding baseline estimate. The estimates for $\alpha$ and $\kappa$ are virtually unchanged from the respective estimates. Intuitively, a slightly lower savings rate seems to make retirement relatively less attractive since the marginal gains from continuing to work are relatively larger and individuals consume less leisure since their wealth is lower from reduced wealth accumulation. As a result, the model compensates by increasing the mean disutility of work. Additionally, the slightly lower savings rate implies a slightly larger consumption difference since consumption while working has increased and consumption at retirement has decreased due to the lower wealth accumulation. Since these changes can dominate any changes in the wage-to-asset ratio, this leads to a slightly lower estimate of the coefficient of relative risk aversion.

Figure 7 presents actual vs. predicted retirement hazard rates by age for each specification. Since the II strategy fits both retirement hazard rates and proportional hazard (PH) coefficients, by construction the predicted hazard rates using the II estimates must fit the model worse than using the MSM estimates which ignore the PH coefficients. Indeed, the model overpredicts retirement at earlier ages, particularly at age 55, and underpredicts retirement just before age 60. Like the MSM-estimated model, the II-estimated model underpredicts retirement after age 60; in both estimations, the moments at later ages receive significantly less weight since there are relatively few individuals working past age 60.

Finally, Panel B of Table 8 displays the implied PH coefficients. The II estimation strategy is based on matching these coefficients to their reduced-form counterparts presented in column 1 of Table 4. The II estimates of the levels of the PH coefficients for both specifications are much closer to the actual coefficients than the MSM estimates. The II estimates of both the income and price elasticities are roughly half the size of their MSM counterparts, and fit the actual estimated elasticities much better (1.40 vs. 0.44, and -2.92 vs. -2.90) than the MSM estimated elasticities presented in Table 7.

5.3 Reconciling the Estimates

Comparing the estimated parameters and elasticities across the reduced-form, MSM and II approaches highlights differences in the sources of identifying variation and also the roles of various modeling assumptions. As evident in the graphical analysis and institutional background figures, the reduced-form approach and the MSM approach exploit two different sources of identifying variation. In the proportional hazard setting, the specification includes sets of income and year polynomials as well as a baseline hazard that
is common across all individuals at each age. These variables control for the non-reform variation thereby isolating the remaining pension reform-related variation to identify the income and price elasticities, $\beta_{SSW}$ and $\beta_{ACC}$ respectively. This identifying variation is illustrated in the residual plots presented in Figures 5A & B. In contrast, by fitting only the retirement patterns across age, the MSM approach uses only across age variation to identify the structural parameters. This identifying variation is captured in Figure 2A (the hazard rates) and the across-age summary statistics presented in Table 3. In particular, this across-age variation is unrelated to the variation from the pension reforms, and thus it is not surprising that the elasticities predicted from the MSM estimates are not similar to the elasticities estimated in the reduced-form setting. The II approach reconciles these differences by matching the reduced-form elasticities as additional moments. This allows the reform-related variation that identified the reduced-form elasticities to drive the estimation of the structural model.

In examining the reduced-form and II estimates, it is clear that some differences in the elasticities persist. This may be accounted for by some modeling assumptions behind the structural model. Specifically, while it is able to match the price elasticity $\beta_{ACC}$, the II estimation overestimates the wealth elasticity $\beta_{SSW}$. This difference can be explained by either a correlation between the disutility of work and wages or heterogeneity in $\beta_{SSW}$. Because the model assumes that the disutility of work is independently distributed across all individuals, the model predicts that only the higher wage earners continue working beyond age 60. These individuals are likely to have small changes in their social security wealth because they have reached their maximum pension benefits. As a result, they may seem responsive to relatively smaller changes in pension wealth thereby leading to a larger estimated wealth elasticity. To account for this, the model could include a positive correlation between wages and the disutility of work so that some higher wage individuals are predicted to retire earlier and some lower wage individuals (with relatively larger changes in pension wealth) are predicted to retire later. This could then lead to a lower predicted $\beta_{SSW}$ in the II estimation. The higher $\beta_{SSW}$ in the II estimation could also reflect underlying heterogeneity in $\beta_{SSW}$; i.e. higher wage individuals have relatively higher responsiveness to change in their pension wealth. As it stands now, the model does not permit this heterogeneity since $\gamma$ is assumed to be constant across individuals. However, the estimation could be generalized to allow for heterogeneity in $\gamma$ and the ratio of individual level wealth and price elasticities would be related to each individual’s $\gamma$.

Lastly, it is useful to focus on similarities and differences relating to the elasticity ratio
First, comparing the MSM and II estimates, these two empirical approaches yield a similar estimate of the elasticity ratio even though the levels of the elasticities differ. The similarity in the elasticity ratio combined with the differences in the estimates of the coefficient of relative risk aversion highlights the different predictions regarding consumption from the two approaches. In particular, the MSM estimation predicts a smaller consumption difference than the II estimation. Intuitively, the MSM estimation yields a relatively larger $\gamma$ based on sensitivity to changes in benefits and retirement around age 60 while the II estimation yields a lower $\gamma$ based on matching the elasticities. As a result of the larger $\gamma$, the MSM estimation predicts more consumption smoothing throughout retirement and hence lower consumption at retirement than the II estimation. Because consumption while working is fixed based on the fixed savings rate assumption, the lower consumption at retirement in the MSM estimation implies a smaller consumption difference. Second, comparing the reduced-form and II estimates, the II estimation yields a higher elasticity ratio. This difference in the elasticity ratio combined with the similarity in the estimates of the coefficient of relative risk aversion highlight that the II estimates imply a smaller consumption difference than that used in the reduced-form setting. If the II estimates of the consumption difference were used in the reduced-form setting, the implied estimates of the coefficient of relative risk aversion would be smaller than the estimates presented in Table 5. Thus, some differences between the reduced-form and II estimates will persist given the differences in the elasticity ratios.

6 Conclusion

Using policy variation from multiple pension reforms in Austria, we estimate the income and price effects of retirement benefits on retirement decisions and show that the coefficient of relative risk aversion is likely to be much lower than previous structural estimates based on observational variation and the Method of Simulated Moments (MSM). Reduced-form estimates based on a Cox proportional hazard specification indicate income and price elasticities of roughly 0.43 and -2.90 respectively. The relatively large role of the price elasticity contrasts with evidence from other social insurance programs such as unemployment (Card et al. 2007, Chetty 2008) and disability insurance (Autor and Duggan 2007). Moreover, these elasticities indicate a relatively low degree of curvature, i.e. $\gamma < 1$. We then show that, in contrast, the commonly used structural estimation strategy based on matching retirement patterns across ages yields a much higher estimate of the
degree of risk aversion, i.e. $\gamma \approx 2$. This implies that, relative to reduced-form predictions, structural models estimated using this MSM strategy are likely to overstate individuals’ responsiveness to financial incentives from retirement benefits. To reconcile these differences, we propose an alternative approach based on the method of Indirect Inference developed by Gourieroux, Monfort and Renault (1993). This estimation strategy matches labor supply elasticities which are precisely identified using the policy variation and provides estimates of the coefficient of relative risk aversion that are directly in line with the reduced-form approach. This methodology can be applied in several other settings with policy changes to estimate structural parameters while also reconciling any differences between structural and reduced-form approaches.
References


Notes: This figures is based on a computed example at age 65 using the following parameter values: \( \gamma = 0.75, A = 15000, y^W = 40000, y^R = 30000 \) and \( v \) in \([0,10]\).
Notes: These figures are based on the sample of all individuals claiming retirement pensions after 1984 (394934 individuals and 275,379 claimants).
Notes: Benefits are computed under the following assumptions: full insurance years at each age, fixed birth cohort across retirement ages, earnings history with positive earnings in last 10 years, nominal earnings in each year equal to 20000 euros in 2003. All nominal benefits in each calendar year are adjusted to 2003 euros. Please see the text for more details.
Notes: This figure is based on data at age 55 only. Please see Table 2 for the sample restrictions. The change in log social security wealth is computed relative to the previous year’s legislation. Please see the text for more details.
Notes: Standard errors for the slope coefficients are clustered at the year level. Points are labeled by year.age.
Notes: The baseline model uses a discount factor of $\beta=0.93$ and a fixed savings rate of 0.10. The lower discount factor is $\beta=0.88$ and the lower savings rate is 0.05. Please see the text for more details.
Notes: The baseline model uses a discount factor of $\beta=0.93$ and a fixed savings rate of 0.10. The lower discount factor is $\beta=0.88$ and the lower savings rate is 0.05. Please see the text for more details.
## Summary of Austrian Pension Reforms - 1984 - 2003

<table>
<thead>
<tr>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td>change in assessment basis from last 5 years to last 10 years of earnings</td>
<td>change in assessment basis from last 10 years to last 15 years of earnings, phased in 1988-1992</td>
<td>change in assessment basis from last 15 to best 15 years of earnings</td>
<td>introduction of bonus / malus system (lower pension coefficient to penalize early retirement)</td>
<td>development of bonus / malus system (increased penalties for early retirement)</td>
</tr>
<tr>
<td>change in revaluation factors used in assessment basis</td>
<td>linking pension coefficient to retirement age</td>
<td>introduction of early retirement due to reduced working capacity at age 57</td>
<td>increase in minimum retirement age from 60 to 61.5, phased in 2000 - 2002</td>
<td>increased restrictions for claiming disability pension</td>
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<tr>
<td>elimination of early retirement due to reduced working capacity at age 57</td>
<td></td>
<td></td>
<td></td>
<td>elimination of early retirement due to reduced working capacity at age 57</td>
</tr>
</tbody>
</table>

Notes: Please see text for more details regarding the pension reforms.
Table 2
Sample Restrictions, Initial Sample (Males, Birth Cohorts ≥ 1948): 2403454

<table>
<thead>
<tr>
<th>Sample Restriction</th>
<th>Sample After Restriction</th>
<th># of Individuals Excluded</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. Less than 1 year of employment in 1972-2003</td>
<td>1512323</td>
<td>891131</td>
</tr>
<tr>
<td>2. Non-Austrian nationality</td>
<td>1417209</td>
<td>95114</td>
</tr>
<tr>
<td>3. Public servants, mining, rail, farmers, construction for 1 or more years</td>
<td>1075285</td>
<td>341924</td>
</tr>
<tr>
<td>4. Self-employed for 1 or more years</td>
<td>744597</td>
<td>330688</td>
</tr>
<tr>
<td>5. Claiming non-old-age or non-disability pensions</td>
<td>720308</td>
<td>24289</td>
</tr>
<tr>
<td>6. Claiming before age 55</td>
<td>648305</td>
<td>72003</td>
</tr>
<tr>
<td>7. Claiming or last observed before 1984</td>
<td>394934</td>
<td>253371</td>
</tr>
<tr>
<td>8. Age &lt; 55, or Age &gt; 65 in 1984 - 2003, &amp; Age &gt; Claim Age (if Claiming)</td>
<td>355805</td>
<td>39129</td>
</tr>
<tr>
<td>9. Missing Pension Variables &amp; First Observed at Age &gt; 55</td>
<td>254130</td>
<td>101675</td>
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<tr>
<td>10. Outliers &amp; Missing Earnings and Industry Data</td>
<td>252907</td>
<td>1223</td>
</tr>
</tbody>
</table>

Final Sample, Ages 55-65 in Years 1984-2003, First Observed at Age 55

| # of Individuals                  | 252907 |
| # of Claimants                   | 178997 |
| # of Observations                | 1101443|

Notes: The number of claimants in the final sample is less than the number of individuals in the sample since younger individuals in the later years of the sample have yet to claim pensions. Further details regarding the samples and restrictions are contained in the text.
Table 3
Summary Statistics by Age

<table>
<thead>
<tr>
<th>Age</th>
<th>Annual Earnings Mean</th>
<th>Annual Earnings Median</th>
<th>Annual Benefits Mean</th>
<th>Annual Benefits Median</th>
<th>SSW Mean</th>
<th>SSW Median</th>
<th>ACC Mean</th>
<th>ACC Median</th>
</tr>
</thead>
<tbody>
<tr>
<td>55, N=242,402</td>
<td>33747.97</td>
<td>27878.50</td>
<td>23303.60</td>
<td>34174.19</td>
<td>28283.65</td>
<td>24977.03</td>
<td>34449.05</td>
<td>28534.26</td>
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<tr>
<td>56, N=197,959</td>
<td>34174.19</td>
<td>28283.65</td>
<td>24977.03</td>
<td>34449.05</td>
<td>28534.26</td>
<td>26154.48</td>
<td>35177.93</td>
<td>29086.64</td>
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<tr>
<td>57, N=172,739</td>
<td>34449.05</td>
<td>28534.26</td>
<td>26154.48</td>
<td>35177.93</td>
<td>29086.64</td>
<td>27776.99</td>
<td>35177.93</td>
<td>29086.64</td>
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<tr>
<td>58, N=144,321</td>
<td>35177.93</td>
<td>29086.64</td>
<td>27776.99</td>
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<td>29086.64</td>
<td>27776.99</td>
<td>35177.93</td>
<td>29086.64</td>
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<tr>
<td>59, N=123,954</td>
<td>35177.93</td>
<td>29086.64</td>
<td>27776.99</td>
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<td>29086.64</td>
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<td>60, N=108183</td>
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<td>29086.64</td>
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<td>27776.99</td>
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<td>61, N=16,268</td>
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<td>62, N=7,657</td>
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<td>29086.64</td>
<td>27776.99</td>
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<tr>
<td>63, N=4,565</td>
<td>35177.93</td>
<td>29086.64</td>
<td>27776.99</td>
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<td>27776.99</td>
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<td>64, N=2,793</td>
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<tr>
<td>65, N=1,787</td>
<td>35177.93</td>
<td>29086.64</td>
<td>27776.99</td>
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<td>29086.64</td>
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Percentile

<table>
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<tr>
<th>Percentile</th>
<th>10</th>
<th>25</th>
<th>50</th>
<th>75</th>
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<tbody>
<tr>
<td>Assets</td>
<td>0.00</td>
<td>4930.6</td>
<td>24884.33</td>
<td>76296.25</td>
<td>160007.4</td>
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Notes: The statistics shown for earnings, annual benefits, SSW and assets are in 2003 euros. Annual earnings are computed based on the calendar year that an individual reaches the specified age. SSW is computed assuming $\beta=.93$. The asset statistics are based on household gross financial assets from SHARE-Austria data. We use information from 1,465 individuals ages 50 through 54 from the SHARE-Austria data.
### Table 4
Hazard Model Estimates

<table>
<thead>
<tr>
<th></th>
<th>All Ages Base Controls</th>
<th>All Ages Full Controls</th>
<th>Ages 60 &amp; 65 Base Controls</th>
<th>Ages 60 &amp; 65 Full Controls</th>
<th>Time-Varying Covariates Base Controls</th>
<th>Time-Varying Covariates Full Controls</th>
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</thead>
<tbody>
<tr>
<td>ln(SSW)</td>
<td>0.4389</td>
<td>0.4013</td>
<td>0.3253</td>
<td>0.2626</td>
<td>0.1097</td>
<td>0.1018</td>
</tr>
<tr>
<td></td>
<td>(0.0775)</td>
<td>(0.0962)</td>
<td>(0.0402)</td>
<td>(0.0466)</td>
<td>(0.0125)</td>
<td>(0.0138)</td>
</tr>
<tr>
<td>ln(ACC)</td>
<td>-2.8972</td>
<td>-3.3815</td>
<td>-2.8575</td>
<td>-3.7060</td>
<td>-0.4804</td>
<td>-0.4334</td>
</tr>
<tr>
<td></td>
<td>(0.8502)</td>
<td>(1.6025)</td>
<td>(1.1183)</td>
<td>(1.5927)</td>
<td>(0.1857)</td>
<td>(0.2683)</td>
</tr>
<tr>
<td>ln(SSW)\text{ln ACC}</td>
<td>0.151</td>
<td>0.119</td>
<td>0.114</td>
<td>0.0708</td>
<td>0.228</td>
<td>0.235</td>
</tr>
<tr>
<td></td>
<td>(0.0477)</td>
<td>(0.0557)</td>
<td>(0.0389)</td>
<td>(0.0312)</td>
<td>(0.0857)</td>
<td>(0.142)</td>
</tr>
<tr>
<td>lnSSW</td>
<td>_{60} \text{ln ACC</td>
<td>}_{60}</td>
<td>0.439</td>
<td>0.401</td>
<td>0.325</td>
<td>0.263</td>
</tr>
<tr>
<td>lnACC</td>
<td>_{60} \text{ln ACC</td>
<td>}_{60}</td>
<td>-2.897</td>
<td>-3.381</td>
<td>-2.857</td>
<td>-3.706</td>
</tr>
</tbody>
</table>

**Observations**: 1101444 1101444 121182 121182 1101444 1101444 1101444 1101444
**Individuals**: 252907 252907 119118 119118 252907 252907
**Retirements**: 178997 178997 96982 96982 178997 178997

**Notes**: Standard errors clustered by year are shown in parentheses. All coefficient estimates should be interpreted as changes in the baseline retirement hazard. All specifications include the following base controls: education dummies, a quadratic polynomial in tenure, and quartic polynomials in calendar year, log annual earnings, and log total earnings in the prior 10 years. All specifications also include a censored dummy (current tenure begun in 1972 or earlier) and the interactions between this dummy and each of the severance pay and tenure variables. The full controls specifications include the base controls, industry and region dummies, and quartic polynomials in log earnings from each of the prior 10 years. Please see text for more details.
Table 5
Risk Aversion Estimate from Reduced-form

<table>
<thead>
<tr>
<th>γ</th>
<th>Elasticity Ratio (β_{SSW}/β_{ACC})</th>
</tr>
</thead>
<tbody>
<tr>
<td>-0.25</td>
<td>0.71</td>
</tr>
<tr>
<td>-0.20</td>
<td>0.57</td>
</tr>
<tr>
<td>-0.15</td>
<td>0.43</td>
</tr>
<tr>
<td>-0.10</td>
<td>0.28</td>
</tr>
<tr>
<td>-0.05</td>
<td>0.14</td>
</tr>
</tbody>
</table>

Notes: These calculations are based on a consumption difference between retirement and employment of 0.395 (=(14000+0.05*25000-.9*28000)/(.9*28000)) and a wage-to-asset ratio of roughly 1.12 (=28000/25000). The numbers used in these calculations come from the medians listed in the summary statistics (Table 3). Please see text for more details.
Table 6
Structural Parameter Estimates
Estimation based on Matching Retirement Hazard Rates

<table>
<thead>
<tr>
<th></th>
<th>Baseline</th>
<th>Lower Discount Factor ((\beta = 0.88))</th>
<th>Lower Savings Rate ((s = 0.05))</th>
</tr>
</thead>
<tbody>
<tr>
<td>Curvature of Consumption Utility: (\gamma)</td>
<td>1.8493</td>
<td>2.1226</td>
<td>2.0190</td>
</tr>
<tr>
<td></td>
<td>(1.7170, 2.1050)</td>
<td>(1.8710, 2.5760)</td>
<td>(1.7970, 2.3210)</td>
</tr>
<tr>
<td>Distribution of Work Disutility: (\eta)</td>
<td>0.1363</td>
<td>0.0890</td>
<td>0.1375</td>
</tr>
<tr>
<td></td>
<td>(0.1220, 0.1440)</td>
<td>(0.0870, 0.1150)</td>
<td>(0.1240, 0.1450)</td>
</tr>
<tr>
<td>Slope of Work Disutility: (\alpha)</td>
<td>1.5163</td>
<td>1.3272</td>
<td>1.5138</td>
</tr>
<tr>
<td></td>
<td>(1.3100, 1.6150)</td>
<td>(1.0300, 1.4070)</td>
<td>(1.2700, 1.5850)</td>
</tr>
<tr>
<td>Disability Pension Fixed Cost: (\kappa)</td>
<td>13815.97</td>
<td>15195.91</td>
<td>12548.74</td>
</tr>
<tr>
<td></td>
<td>(12268.01, 14034.74)</td>
<td>(14424.33, 18624.77)</td>
<td>(11358.07, 13321.85)</td>
</tr>
</tbody>
</table>

Notes: 95% confidence intervals are shown in parentheses below the parameter estimates; confidence intervals are based on the bootstrapped distributions of parameter estimates that were calculated using 100 replications in which individuals were drawn with replacement. Estimates are based on the same sample used to estimate the proportional hazard specifications in Table 4. The baseline specification is based on a discount factor of \(\beta = 0.93\), a real interest rate of \(r = 0.075\), a fixed savings rate of \(s=0.10\) and a fixed wage growth rate of 0.0175. Please see the text for more details.
<table>
<thead>
<tr>
<th></th>
<th>Baseline</th>
<th>Lower Discount Factor ( $\beta = 0.88$)</th>
<th>Lower Savings Rate ( $s = 0.05$)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\beta_{SSW}$</td>
<td>3.0478</td>
<td>2.9793</td>
<td>3.1278</td>
</tr>
<tr>
<td>$\beta_{ACC}$</td>
<td>-6.8033</td>
<td>-7.1911</td>
<td>-8.0154</td>
</tr>
</tbody>
</table>

Table 7: Hazard Coefficient Estimates from Structural Model

Notes: Estimates are based on a the same sample used to estimate the proportional hazard specifications in Table 4. The baseline specification is based on a discount factor of $\beta = 0.93$, a real interest rate of $r = 0.075$, a fixed savings rate of $s=0.10$ and a fixed wage growth rate of 0.0175. Please see the text for more details.
Table 8
Indirect Inference Estimates
Estimation based on Matching Retirement Hazard Rates and Proportional Hazard Coefficients (All Ages)

<table>
<thead>
<tr>
<th></th>
<th>Baseline</th>
<th>Lower Discount Factor (β = 0.88)</th>
<th>Lower Savings Rate (s = 0.05)</th>
</tr>
</thead>
<tbody>
<tr>
<td>(A) Parameter Estimates:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Curvature of Consumption Utility: γ</td>
<td>0.5536</td>
<td>0.5865</td>
<td>0.4601</td>
</tr>
<tr>
<td></td>
<td>(0.4080, 0.5910)</td>
<td>(0.5140, 1.0830)</td>
<td>(0.3860, 0.7800)</td>
</tr>
<tr>
<td>Distribution of Work Disutility: η</td>
<td>0.7915</td>
<td>0.7301</td>
<td>0.8664</td>
</tr>
<tr>
<td></td>
<td>(0.7390, 0.9510)</td>
<td>(0.6930, 1.0530)</td>
<td>(0.6950, 1.0240)</td>
</tr>
<tr>
<td>Slope of Work Disutility: α</td>
<td>0.3847</td>
<td>0.2149</td>
<td>0.3911</td>
</tr>
<tr>
<td></td>
<td>(0.3320, 0.4410)</td>
<td>(0.1370, 0.2410)</td>
<td>(0.2940, 0.4880)</td>
</tr>
<tr>
<td>Disability Pension Fixed Cost: κ</td>
<td>40077.79</td>
<td>42720.21</td>
<td>40419.24</td>
</tr>
<tr>
<td></td>
<td>(34701.39, 42989.22)</td>
<td>(39509.99, 52331.75)</td>
<td>(34657.94, 46010.27)</td>
</tr>
<tr>
<td>(B) Proportional Hazard Coefficients:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Coefficients with All Ages</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>β&lt;sub&gt;SSW&lt;/sub&gt;</td>
<td>1.3977</td>
<td>1.4596</td>
<td>1.2579</td>
</tr>
<tr>
<td>β&lt;sub&gt;ACC&lt;/sub&gt;</td>
<td>-2.9209</td>
<td>-2.9853</td>
<td>-3.2309</td>
</tr>
</tbody>
</table>

Notes: 95% confidence intervals are shown in parentheses below the parameter estimates; confidence intervals are based on the bootstrapped distributions of parameter estimates that were calculated using 100 replications in which individuals were drawn with replacement. Estimates are based on a the same sample used to estimate the proportional hazard specifications in Table 4. The baseline specification is based on a discount factor of β = 0.93, a real interest rate of r = 0.075, a fixed savings rate of s=0.10 and a fixed wage growth rate of 0.0175. Please see the text for more details.