How much and how fast do investors respond to equity premium changes? Evidence from wealth taxation

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Abstract

Using administrative panel data on Norwegian investors' portfolios, we document strong but slow portfolio allocation responses to a persistent wealth-tax-induced shock to the equity premium. Short-run responses resemble the weak sensitivity documented using surveys. The longer-run responses are much larger and can be rationalized by moderate risk aversion. We document that equity premium shocks affect stock market entry but not exits, suggesting that entry costs dominate participation costs. Our finding of slow responses supports the asset-pricing literature that uses adjustment frictions to explain important asset-pricing puzzles, and have implications for optimal capital taxation when tax rates differ across assets.

JEL: G11, G12, G51, H20, H31

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1 Introduction

How investor portfolio allocation responds to changes in return differentials is of critical importance to many debates in financial economics and public finance. Frictionless portfolio models, such as Merton (1969), on which modern portfolio theory and optimal capital taxation are built,[1] predict an immediate and sizable response when investors have a realistic degree of relative risk aversion. Yet, such responses are not found empirically. This constitutes a puzzle with at least two potential explanations: either risk aversion is much higher than what is normally thought or frictions introduce costs that limit or delay portfolio adjustments. Distinguishing between these potential explanations is crucial because they have very different implications for modeling investor behavior. More generally, understanding why investors appear insensitive to return differentials is important for resolving asset pricing puzzles and matters for optimal capital taxation when tax rates differ across assets (Saez and Stantcheva, 2018).

A recent literature (see, e.g., Gabaix and Koijen 2021; Bacchetta, Davenport, and van Wincoop 2021; Bacchetta and Van Wincoop 2010; Duffie 2010) singles out portfolio adjustment frictions as an important reason why investors' portfolio allocation is insensitive to variation in return differentials such as the equity risk premium. This literature shows theoretically that adjustment frictions are critical for explaining a rich set of asset pricing puzzles, such as stock price volatility, momentum, post-earnings announcement drift, and the forward discount puzzle. However, providing empirical evidence to support the building blocks of these models is challenging due to limited sources of identifying variation in the equity premium. Time series variation is particularly hard to come by, and it takes a very long time series to merely estimate the level of the equity premium (Merton, 1980). Identifying changes is even harder. One solution has been to use cross-sectional variation in beliefs about the equity premium elicited from surveys. This approach, however, suffers from measurement error problems and, as we argue, from ignoring the dynamics of how investors respond to return shocks.

Our solution is to exploit a policy-induced change in the equity risk premium in Norway. Prior to 1998, Norwegian households faced a higher marginal wealth tax rate on safe assets than on risky assets, which created a "wealth-tax risk premium" of about 30 basis points. In 1998, a tax reform equalized the marginal rates, and thereby removed the wealth-tax risk premium. We exploit the fact that only households with wealth above a certain threshold (i.e., wealth-tax payers) were directly affected in a differences-in-differences framework.

This quasi-experimental setting is unique in that it provides a clear time series shock to the equity premium that does not need to be estimated. In addition, we have a well-defined treatment and control group, allowing us to difference out the effects of other macroeconomic shocks. Importantly, the Norwegian administrative data contain detailed information on the components of a household's wealth, and thus portfolio allocation, over time. This allows us to fully exploit the time series variation in the equity premium by studying the *dynamics* of how investors respond. We find this to be decisive in reconciling seemingly contradictory survey-based

[1] On optimal capital income taxation, see, e.g., Domar and Musgrave (1944), Mossin (1968), Stiglitz (1975), and Sandmo (1977), as well as the more recent contributions by Boadway and Spiritus (2021), Gerritsen et al. (2020), Guvenen et al. (2019), and Boar and Knowles (2022) when returns on risky assets are heterogeneous across investors.

evidence with work-horse portfolio models. Our setting has other attractive features as well. First, since wealth taxes are assessed on the stock of wealth, they only affect expected returns and not volatility. The counter-example is capital income taxes, whose effect on volatility offsets the effect on expected returns (Mossin 1968, Gordon 1985). The second feature is that wealth tax reforms are rare, easy to understand, and subject to widespread media coverage. Households should thus have had several opportunities to learn about the effect of the reform on after-tax returns and, because reforms are rare, expect it to be long-lasting. This creates the kind of setting that favors the frictionless models in which we would expect immediate and strong responses. Finding a slow response in this setting would constitute strong evidence of adjustment frictions.

We find that the 1998 wealth-tax-induced 30 basis point reduction in the risk premium substantially lowers the share of financial wealth allocated to stocks. Crucially, the adjustment process is gradual: within two years, the risky share drops by a modest 0.5 percentage points. If this were the total response, it would require a very large coefficient of relative risk aversion of 15 to be consistent with a frictionless Merton (1969) model. However, the response continues to grow. After 5–6 years, the share has dropped by about 2 percentage points, which can be rationalized by a contained relative risk aversion of the investors' value function of 1.81. In other words, investors react slowly to long-lasting perturbations in the equity premium, and these responses converge to a reasonable frictionless benchmark over time. These findings are consistent with models with portfolio adjustment frictions such as Gârleanu and Pedersen (2013) who study a partial equilibrium portfolio model with quadratic adjustment costs. They show that the stock market share adjusts slowly to changes in the equity premium, gradually reaching the optimal frictionless share. They are also consistent with models of heterogeneous costly attention in which less attentive investors respond slower to the equity premium shock (Duffie, 2010).[2]

Guided by theory, we expand the evidence by documenting heterogeneous responses. First, standard portfolio models imply that investors with higher risk aversion not only have lower risky shares but also respond less to changes in the equity premium. We test this ubiquitous tenet of portfolio models by using pre-reform portfolio shares as a proxy for risk tolerance. We find that—in all years after the reform—more risk-tolerant investors respond more. While they respond more strongly, as predicted by theory, they do not respond faster.

Second, to better understand what drives the sluggishness of the responses, we rely on work by Calvet, Campbell, and Sodini (2009) who study how portfolio rebalancing speed depends on household observables. Using their estimated coefficients, we predict heterogeneous rebalancing speeds in our sample, and use these predicted values as a proxy index for investor-specific portfolio adjustment costs. This reveals a remarkable difference: households in the bottom quartile of adjustment cost respond more than twice as fast as the remaining sample. Interestingly, they do not have stronger cumulative responses—only faster. Since Gabaix and Koijen (2021) demonstrate that it is the weighted average speed that matters in determining aggregate stock price dynamics,

[2]Sluggishness in portfolio adjustments can be due to "sticky actions", in which agents pay a cost to change their portfolio allocation, "sticky information", in which they pay a cost to change their information (e.g. an attention cost to observe their portfolio holdings as in Alvarez et al. (2012) and Duffie (2010) or to obtain the information on the after-tax equity premium). The sticky information model has been used by Bacchetta and Van Wincoop (2010) and Duffie (2010) while Gârleanu and Pedersen (2013) rely on the sticky action model. They have observationally similar implications, and we make no attempt to distinguish between them.

we examine the relationship between these weights and adjustment speed. Our exercise shows that those who hold more of the stock market respond much faster to equity premium shocks and face lower adjustment costs. However, a considerable share of stock market wealth is held by high-adjustment-cost households. Hence, sluggish portfolio adjustment is likely important in determining aggregate stock price dynamics.

We document that households treated by the equity premium reduction were on an upward-sloping trajectory in terms of their stock market exposure before the 1998 reform. In many settings, one would attribute these pre-trends to unobservable factors and adjust for them by detrending the main estimate (see, e.g., Jakobsen, Jakobsen, Kleven, and Zucman 2020). However, we believe that this is inappropriate in our setting since the wealth-tax risk premium had only been in place since 1992. We argue that the observed pre-trends are reflecting delayed responses to the 1992-reform's effect on the equity premium. Hence, pre-reform trends do not provide a valid counterfactual for the post-1998 period, and adjusting for them would exaggerate the treatment effects. Our first argument rests on the finding that the post-1998 responses are sluggish. This means that we should expect to also find sluggish responses in the pre-period materializing as an upward-sloping trend. Our second argument is that detrending our estimates would imply an unreasonably low coefficient of relative risk aversion of 0.5. Finally, the positive trend we see during 1992–1997 is not part of a longer-run trend. While our main data start in 1993, we obtain supplementary tax data that allow us to observe whether a household purchased stocks in any given year. These data go back further in time and show no pre-trends prior to the 1992 introduction of the wealth-tax risk premium. However, consistent with our other findings, purchasing behavior increases sharply in 1992 and trends upward until 1998 when the wealth-tax risk premium was removed. Hence, the observed "pre-trends" appear to be causally driven by the 1992 introduction.

Finally, we study how stock market participation reacts to persistent changes in the equity premium. Interestingly, we find a slow but strong increase in participation in response to the initial 1992 shock. This increase in participation stops following the 1998 reform. Importantly, there is no symmetry: households who were previously nudged into participation barely exit. The 1998 removal of the wealth-tax risk premium has a modest and delayed effect on stock market exits. Theoretically, these dynamics are consistent with the presence of both one-time entry costs and per-period participation costs that cause asymmetric effects of increases and decreases in the equity premium on participation. The fact that the 1998 removal of the wealth-tax risk premium has a modest effect on exits implies that per-period costs are low relative to entry costs. We quantify this by modeling the entry and exit decision in a Merton (1969) framework. This allows us to provide novel estimates of both one-time entry and per-period stock market participation costs.[3] The estimated entry cost of \$800 and recurring participation cost of \$89 can be easily reconciled with existing estimates of participation costs and moderate risk aversion.

Related literature. A growing survey-based literature relies on elicited beliefs to study how sensitive portfolio allocation is to the equity risk premium (Vissing-Jorgensen 2003; Ameriks, Kézdi, Lee, and Shapiro 2020; Amromin and Sharpe 2014; Kézdi and Willis 2009; Dominitz and

[3] See concurrent work by Choukhmane and de Silva (2021) for very similar findings by exploiting default options in retirement-saving plans to distinguish between frictions and risk aversion.

Manski 2007; Giglio, Maggiori, Stroebel, and Utkus 2021; Beutel and Weber 2022).[4] Common to these papers is that they find perplexingly small sensitivities of the risky share of financial wealth to beliefs about the equity risk premium. In a recent study, Giglio et al. (2021) use U.S. data on Vanguard retail investors to regress the risky portfolio share on the current belief about the equity premium. They estimate a sensitivity of around 0.7, which is very close to the immediate response that we find, but an order of magnitude below our long-run estimate. Their low estimate could be evidence of frictions, and indeed Giglio et al. (2021) document that investors that are less likely to be subject to portfolio frictions display stronger sensitivities. Notably, however, these point estimates are imprecise and still require a large risk aversion parameter to be consistent with the frictionless benchmark. Furthermore, the larger sensitivity of some investors may reflect a higher risk tolerance rather than a lower exposure to portfolio adjustment frictions.

In general, besides frictions, a low estimated risky-share sensitivity may be due to the presence of measurement error in elicited beliefs. If for any reason, the *reported* expected risk premium differs from the expectation according to which the agent is optimizing, attenuation bias will drive the estimated sensitivity toward zero. Because the exact source of this measurement error problem is hard to pin down, it is difficult to find the appropriate econometric technique to address it. The approach in Giglio et al. (2021) is to rely on two different elicitations of beliefs, and then remove measurement error not common to the two elicitations. However, this has a rather small effect on the estimated sensitivity, either because measurement errors tend to be correlated across elicitations or because other frictions play an important role.[5] Beutel and Weber (2022) improve upon the IV strategy by using a randomized information experiment, which also addresses potential reverse causality (Chaudhry, 2022). They find that the risky-share sensitivity remains low. They attribute this insensitivity to the fact that the risky share is typically constrained to be between 0 and 1, which attenuates OLS estimates toward zero in the presence of extreme beliefs. We view this emphasis on constrained responses as complementary to our emphasis on sluggish responses.

We make three important empirical contributions. First, our setting has the unique ingredient of a clearly defined shock to the equity premium that affects some but not all households. This means that we do not need to rely on belief elicitation to obtain variation in the equity premium, and thus our results are unaffected by typical survey-based econometric issues. Second, we also improve upon identification. In our setting, we can take out household fixed effects and thereby control for time-invariant confounders. Hence, whether reported beliefs covary with constant traits that independently affect portfolio allocation (e.g., risk aversion) does not matter.

Beyond improvements related to measurement and identification, a key strength of our paper is the ability to observe dynamic responses to risk premium changes. We show that this is critical for understanding the long-run sensitivity of portfolios to the risk premium. Gârleanu and

[4]See also the literature that studies portfolio choices more broadly, e.g., Arrondel, Calvo Pardo, and Tas 2014; Hanspal, Weber, and Wohlfart 2020; Choi and Robertson 2020; Merkle and Weber 2014.

[5]The presence of non-classical measurement error is less acknowledged. For example, respondents may confound expected returns and risk by reporting more pessimistic expectations when they perceive the variance to be higher. This effectively produces an omitted variables problem that leads to a downward bias in estimated sensitivities. This issue may be amplified by controlling for the expected variance (Pischke, 2007), and also biases the estimated variance sensitivity (Griliches, 1986). The results in Beutel and Weber (2022) suggest this is a possibility.

Pedersen (2013) and Bacchetta, Davenport, and van Wincoop (2021) show that long-run responses correspond to the frictionless Merton (1969) model predictions. Hence, by contrasting immediate and long-run responses, we document the extent of the adjustment frictions that Gârleanu and Pedersen (2013) and Bacchetta et al. (2021) emphasize.

Our paper is directly related to the recent wave of papers that rely on portfolio adjustment frictions to explain aggregate fluctuations in asset prices—what Gabaix and Koijen (2021) label the "inelastic market hypothesis" (see also Duffie 2010, Bacchetta and Van Wincoop 2010 and Bacchetta et al. 2021). In a frictionless world, a small deviation of the equity premium from its average would trigger a swift reallocation of investor portfolios that would quickly reabsorb the initial price differentials. With adjustment frictions, however, portfolio reallocations are slow, fueling large changes in equilibrium asset prices (as in Duffie 2010). We provide direct evidence of sluggish adjustment, lending empirical support to the key building block of this literature.

We also contribute to the literature on taxation and household portfolio choice initially reviewed by Poterba (2002). More recent contributions include Poterba and Samwick (2003), Alan, Atalay, Crossley, and Jeon (2010), and Desai and Dharmapala (2011) who study the effects of capital income taxation on investment in tax-preferred assets. The central contribution of our paper is to obtain marginal tax rate variation that is directly related to—but does not affect—the riskiness of the asset. The income-bracket-related variation in effective marginal tax rates on different asset classes in the U.S. and Canada favors assets whose return realization may be delayed or accrued within retirement accounts rather than simply riskier assets.

This paper also contributes to a growing empirical literature on household responses to capital taxation,[6] and wealth taxation in particular (Seim 2017, Londoño-Vélez and Ávila-Mahecha 2018, Zoutman 2018, Jakobsen et al. 2020, Durán-Cabré María et al. 2019, Brülhart et al. 2019, Ring 2020, Ring and Thoresen 2021, Berg and Hebous 2021). While some studies consider effects on portfolio allocation (Durán-Cabré María et al. 2019, Ring 2020), they do not exploit variation in the equity premium as we do.

Our evidence of sluggish responses to changes in the equity premium is relevant to the recent theoretical literature on optimal capital taxation (e.g., Boadway and Spiritus, 2021; Gerritsen et al., 2020; Saez and Stantcheva 2018). These models assume no portfolio adjustment frictions and no participation costs. This is potentially problematic, as one can infer from optimal labor income taxation, where participation is important (Lehmann et al. 2011; Jacquet et al. 2013).

The paper proceeds as follows. Section 2 presents a simple conceptual framework, with no frictions other than participation costs, which we use to clarify the link between the wealth tax reform and the equity premium, and to get an idea of what the risky share sensitivity would be in the absence of portfolio adjustment frictions. Section 3 describes the data and the tax reform. Section 4 presents our empirical strategy and discusses identification. Section 5 shows the results for the portfolio share and stock market participation. Section 6 relates our findings to portfolio models with frictions. Section 7 puts the results in perspective and concludes.

[6] On the empirical front, see, e.g., recent contributions by Boissel and Matray (2021); Nekoei and Seim (2018); Arefeva et al. (2021); Glogowsky (2021); Lavecchia and Tazhitdinova (2021); Martínez-Toledano (2020); Agrawal et al. (2020); Tsoutsoura (2015); Dray et al. (2022); on the theoretical front, see, e.g., Boadway and Spiritus (2021); Gerritsen et al. (2020); Gaillard and Wangner (2021); Guvenen et al. (2019); and Boar and Knowles (2022) when returns on risky assets are heterogenous across investors.

2 Conceptual Framework

In this section, we illustrate how differential wealth taxation affects investors' portfolio share in stocks (the risky share) and their incentives to participate in the stock market. We do this in the simple setting of a two-asset Merton (1969) model that we extend to include participation costs but no portfolio share adjustment cost, as in Gârleanu and Pedersen (2013).

2.1 Effect of Taxation on the Portfolio Stock Market Share

In the case of two assets—a risky asset (stocks) and a safe asset (riskless bonds)—wealth taxation can affect portfolio allocation to the extent that it affects relative returns. To illustrate this, assume that the capital income from both stocks and bonds is taxed at the same rate τ and that, in addition to a tax on income from capital, also the stock of wealth in the two assets is taxed at the rate τ_w. The net of tax expected return on stocks and bonds, r_s^n and r_f^n, can then be written as

$$r_s^n = r_s(1-\tau) - \tau_w d \quad \text{and} \quad r_f^n = r_f(1-\tau) - \tau_w,$$

where r_s and r_f are the (pre-tax) expected returns on stocks and risk-free assets, and $0 \leqq 1-d \leqq 1$ is a valuation discount that reflects the structure of wealth taxation in Norway. When $d = 1$, wealth in stocks and in bonds is valued at market prices for the sake of the wealth tax; when $d < 1$ stocks enjoy a valuation discount that increases the net of tax return on stocks compared to bonds. We define this wealth-tax risk premium as $\Delta = \tau_w(1-d)$. The corresponding net-of-tax equity premium is

$$r_e^n = (r_s - r_f)(1-\tau) + \Delta, \tag{1}$$

which may also be written as $r_e^n = r_e(1-\tau) + \Delta$.

A wealth-tax-induced equity premium Δ emerges when $0 \leqq d < 1$; it disappears when $d = 1$ and thus all assets are equally valued for the purpose of wealth taxation. In Norway, over our sample period, capital income is taxed at the same constant rate (28%) and there is no exemptions on this source of income. Thus, without loss of generality, assume $\tau = 0$, $\tau_w > 0$ and $0 < d \leqq 1$.

For an investor that participates in the stock market, the relationship between the wealth-tax-induced shift in the risk premium and the risky share of financial wealth can be illustrated in the Merton (1969) model as:

$$\alpha = \frac{1}{\gamma}\frac{r_e^n}{\sigma_s^2} = \frac{1}{\gamma}\frac{r_e + \Delta}{\sigma_s^2}, \tag{2}$$

where α is the risky share, γ is the coefficient of relative risk aversion (assumed constant), and σ_s^2 is the variance of stock returns.[7] We see that wealth taxation only affects the equity premium, and thus the risky share, through the wealth-tax risk premium, Δ. In our empirical setting, $0 < d < 1$ is in the years before the 1998 reform, which increases the equity premium by $\Delta = \tau_w(1-d)$. The

[7] This formula is simplified by the immaterial assumption that wealth taxes are assessed before returns are realized. In practice, wealth taxes are assessed on end-of-year wealth holdings, which would include any non-consumed gains. This implies that $r_e^n = r_e + \Delta - \tau_w r_e$. This only differs by the expression for r_e^n in equation 2 by the second-order term $\tau_w r_e$. Further, the variance of after-tax stock returns is $(1-\tau_w)^2\sigma_s^2$, which, since τ_w is about 1 percentage point, is very close to σ_s^2.

1998 reform sets $d = 1$, imparting a shock to the equity premium of $-\tau(1 - d_{pre1998}) < 0$. Hence, as stocks become less rewarding compared to safe assets, the Merton (1969) model predicts that investors should decrease the share of their wealth allocated to risky assets.

To get a sense of the magnitude of the sensitivity of the portfolio share to the equity premium, consider a marginal change in Δ caused by a decrease in the risky-asset valuation discount, d. The effect on the optimal portfolio share is:

$$\frac{\partial \alpha}{\partial \Delta} = \frac{\partial \alpha}{\partial r_e^n}\frac{\partial r_e^n}{\partial \Delta} = \frac{1}{\gamma \sigma_s^2}. \tag{3}$$

Because $\frac{\partial r_e^n}{\partial \Delta} = 1$, it is directly informative of the sensitivity of the portfolio share to the equity premium. Notice that differently from a capital income tax, the wealth tax affects returns additively. Hence a change in the valuation discount affects the equity premium while leaving the variance of stock returns virtually unchanged.[8] Assuming a variance of stock returns of 0.04, which is in the ballpark of the variance of historical stock market returns in Norway (see Fagereng, Gottlieb, and Guiso 2017), the portfolio share allocated to stocks would decrease by between 6 and 13 times the fall in Δ for realistic values of relative risk aversion between 2 and 4. Thus, according to a reasonably calibrated frictionless Merton model, an increase in the wealth tax rate on risky assets of around 30 basis points—the increase brought by the 1998 tax reform that we study—could lower the risky portfolio share by about 1.8 to 5 percentage points. These calculations based on the Merton model provide a useful frictionless benchmark against which we can compare our estimates.

2.2 Effect on Stock Market Participation

If participation in the stock market entails some fixed cost, a shock to the after-tax equity premium affects whether households participate in the stock market. However, the exact effect will crucially depend on the nature of participation costs. In the simple two-asset portfolio model, assume that a fixed cost of ϕ is required for a household to participate in the stock market. Assume ϕ is the sum of a one-time fixed entry cost ϕ^E that an individual pays once in his life at first entry into the market and a per-period cost ϕ^P that is incurred every period the investor is in the market. Then an investor with initial wealth of w, that did not participate in the past, enters the market in the current period if

$$Eu([\alpha^* r_e^n + r_f - \tau_w]w]) - u'((r_f - \tau_w)w)\phi > u((r_f - \tau_w)w). \tag{4}$$

An investor who is already in the market decides to stay in the current period if

$$Eu([\alpha^* r_e^n + r_f - \tau_w]w]) - u'((r_f - \tau_w)w)\phi^P > u((r_f - \tau_w)w). \tag{5}$$

The left hand sides of equations (4) and (5) measure the investor's expected utility if she par-

[8] This is subject to a slight caveat in the Norwegian setting. If we account for the fact that wealth taxes are assessed on the stock of wealth plus its one-year return, the volatility of after-wealth tax returns is $(1 - \tau_w)^2 \sigma_s^2$. Thus if $\sigma_s^2 = 0.04$, then a 30 basis points change in the wealth tax on stocks would only change the variance by about 0.00016. Thus any changes to the (after-tax) volatility will be negligible.

ticipates in the stock market, allocates the optimal share $\alpha^* = \frac{r_e^n}{\gamma\sigma_s^2}$ to stocks, and pays the participation cost ϕ evaluated at the marginal utility of risk-free wealth if she enters for the first time (equation 4); or pays $\phi^P < \phi$ if she already entered in the past (equation 5). The right-hand side is the utility the investor achieves if she allocates her entire wealth to the risk-free asset. Taking a first-order approximation of the expected utility on the left hand side around $\alpha^* = 0$, the participation condition requires that $\alpha^* r_e^n w > \phi$, or equivalently, $w > \frac{\phi}{\alpha^* r_e^n}$ for a first-time entrant, and $w > \frac{\phi^P}{\alpha^* r_e^n}$ for an investor that is re-entering or staying in the market. Substituting in for the optimal share using equation (2) defines the wealth threshold, $\bar{w} = \frac{\phi\gamma\sigma^2}{(r^{en})^2}$ for entering the market for the first time; and $\bar{w}^P = \frac{\phi^P\gamma\sigma^2}{(r_e^n)^2}$ defines the threshold for staying or re-entering. Both thresholds are decreasing in the net of tax equity premium and increasing in the cost of participation, the risk aversion of the investor, and the variance of stock returns.

The above discussion shows that an increase in the net of tax equity premium encourages entry into the market, the more so the higher is the portfolio share invested in stocks upon entry. Once entry has occurred, however, the wealth threshold for staying in the stock market is lower because $\phi^P < \phi$ implies $\bar{w}^P < \bar{w}$. Hence, all else equal, a drop in the equity premium, such as the one produced by the wealth tax reversal that occurred in 1998, will discourage first-time entry but have a contained effect on the propensity to exit the stock market. Indeed, if $\phi^P = 0$, then nobody among those that already participated in the market would exit after the 1998 tax reversal.

Assume that wealth in the relevant population is distributed with cumulative density function $G(w)$. Then, all else equal, the stock market participation rate is

$$\pi(\bar{w}) = 1 - G(\bar{w}) = 1 - G\left(\frac{\phi\gamma\sigma^2}{(r_e^n)^2}\right). \tag{6}$$

Differentiating with respect to Δ, we find that a marginal *increase* in the tax-induced equity premium from a more generous valuation discount on stocks has the following effect on participation:

$$\frac{\partial\pi(\bar{w})}{\partial\Delta} = \frac{\partial\pi(\bar{w})}{\partial r_e^n}\frac{\partial r_e^n}{\partial\Delta} = -G'(\bar{w})\frac{d\bar{w}}{dr_e^n} = 2G'(\bar{w})\frac{\phi\gamma\sigma^2}{(r_e^n)^3}. \tag{7}$$

Since $\frac{\partial r_e^n}{\partial\Delta} = 1$, the participation response to the valuation discount on equity identifies the response of participation to an increase in the equity premium. Similarly, a marginal *decrease* in the tax-induced equity premium, stemming from a less generous valuation discount on stocks, causes a decrease in participation of $\frac{\partial\pi(\bar{w})}{\partial\Delta} = \frac{\partial\pi(\bar{w})}{\partial r_e^n} = 2G'(\bar{w})\frac{\phi^P\gamma\sigma^2}{(r_e^n)^3}$.

Finally, the average stock market share across investors who participate in the stock market and those who do not—the unconditional share—is $\alpha_u = \pi(\bar{w})\times\alpha$. The effect on the unconditional share of a change in the equity premium is

$$\frac{\partial\alpha_u}{\partial r_e^n} = \pi(\bar{w})\frac{\partial\alpha}{\partial r_e^n} + \alpha\frac{\partial\pi(\bar{w})}{\partial r_e^n} \tag{8}$$

We will estimate unconditional shares and stock market participation and obtain estimates of the effects of the equity premium on α_u and $\pi(\bar{w})$. In section 5.6, we use the estimates of $\frac{\partial\alpha_u}{\partial r_e^n}$

and $\frac{\partial \pi(\bar{w})}{\partial r_e^n}$ to obtain estimates of the implied risk aversion and the participation costs.

3 The Data and the Tax Reform

3.1 Data Sources

Our data consist of several administrative registers from Norway. They cover the entire population and provide detailed information on households' wealth and portfolio allocation since 1993. Because Norway levies an annual net wealth tax, the tax authorities collect data on an exhaustive range of wealth components. For most households, the main components of taxable wealth are housing wealth, financial wealth, and (the negative of) any debts. Importantly, while the progressive nature of the wealth tax exempts a majority of households, the tax authorities collect data on the asset holdings of *all* households. While we do not observe wealth in defined-contribution (DC) plans, this is not a concern in our setting due to the late adoption of DC plans in Norway, after the end of our sample period.[9]

For our case, it is particularly useful that financial wealth components are reported directly by financial institutions, at their prevailing market values, to the tax authorities. This is typically done at the asset or asset-class level, allowing us to distinguish between investments in "safe" assets (such as bank deposits or government bonds) and "risky" assets (such as public equity held directly or through mutual funds). This provides us with an accurate measure of households' financial portfolio allocation, where there is little scope for error caused by tax evasion or recall bias.[10] Particularly for households near the wealth-tax thresholds, from whom we obtain identifying variation, there is little evidence to suggest substantial evasion (Ring, 2020).

Finally, because the data sources cover the entire population, sample selection issues are unlikely to play a meaningful role. Sample attrition is minimal, and limited to rare events such as migration or death. We also obtain data on individual characteristics, such as incomes (reported by employers and listed on tax returns), education (through the national education database), and age (from the central population register) that we use in our analyses. The general features of the data sources are discussed in Fagereng et al. (2020) and Ring (2020).

[9] In Norway, employment-related defined contribution plans only allow for employer (not employee) contributions. These plans were only available as of 2001. The standard pension saving scheme consisted of a government-provided defined-benefit pension. In 2002, only 3% of workers had a DC and by 2003, only 8% (FAFO, 2014). Hence, even by the end of our sample period, at most 8% of taxpayers (and likely much fewer since many were already retired) had at most two years of accumulated contributions in their DC plans. Hence, the importance of DC plans is small relative to that documented by, e.g., Bach et al. (2020). This means that there is no grounds for concerns that the responses we observe in investors' stock portfolio is offset by changes within their DC stock portfolios.

[10] Financial assets not held domestically must be self-reported as it is not typical for foreign financial institutions to report asset holdings directly to the Norwegian tax authorities. However, foreign assets typically account for a small share of household portfolios. Based on reported values, foreign assets account for only 3% of the financial wealth of households in the 50th to 90th percentile of the wealth distribution (Fagereng et al., 2020). For households in the top 0.01% of the wealth distribution, this figure only grows to about 8%. However, since our identification comes from exposure to the wealth tax, which is levied at a relatively low threshold in Norway, the statistics for households in the 50th to 90th percentile are more relevant. Additional research by Alstadsæter et al. (2019) shows evidence consistent with very limited tax evasion for households below the very top.

3.2 Variable Definitions

We group financial wealth into risky and safe assets. Risky assets consist of a household's stock market wealth (SMW). This is the sum of directly-held listed stocks and stocks that are held indirectly through mutual funds. Because, as argued, defined contribution plans are irrelevant in Norway over the sample period, our data capture all households' stockholdings to which the wealth tax reform applies.

Safe assets consist of bank deposits (e.g., checking accounts, savings accounts, and other low-risk interest-bearing products offered by banks) as well as direct and indirect bond holdings.

Taxable net wealth (TNW) is computed by the tax authorities, and is the sum of financial wealth, housing wealth, other real assets, such as vehicles and art, and any outstanding claims (e.g., owed wages), minus debt. TNW serves as the tax base for the wealth tax. Importantly, some assets enter at a discounted value. This includes housing wealth, which enters at about 30% of estimated market value during our time period,[11] as well as stock market wealth. The presence of this stock market wealth discount and its removal in 1998 is discussed in more detail in the next section.

3.3 Wealth Taxation and the 1998 Reform

Norway imposes an annual, progressive wealth tax on its residents. During our sample period of 1993–2003, any taxable net wealth that exceeds a certain threshold is subject to a tax rate that starts at about 1%. While these thresholds have grown over time, they were effectively fixed during the period that we study.

Wealth tax formula. During 1993–1997, wealth taxes are accrued by household i in year t according to the following formula:

$$\begin{aligned} wtax_{i,t} &= \tau_{w1}\,\mathbf{1}[TNW_{it} > T_t^1](TNW_{it} - T_t^1) \\ &+ (\tau_{w2} - \tau_{w1})\mathbf{1}[TNW_{it} > T_t^2](TNW_{it} - T_t^2) \\ &+ (\tau_{w3} - \tau_{w2})\mathbf{1}[TNW_{it} > T_t^3](TNW_{it} - T_t^3), \end{aligned} \tag{9}$$

where $\tau_{wj,}, j = 1, 2, 3$, are the nominal wealth tax rates applied to any wealth in excess of the corresponding wealth-tax thresholds $T_t^j, j = 1, 2, 3$ and $\mathbf{1}[.]$ denotes the indicator function.[12] These nominal rates were 1.1%, 1.3% and 1.5%, and the thresholds were set at NOK $120,000$, $235,000$, and $530,000$ (about \$17,000; \$35,500 and \$75,000).[13] This means that households with TNW

[11]As discussed in Ring (2020), this approximation was not very good, and lead to a new model that was introduced in 2010, after the end of our sample period. While housing assets are discounted, debts are not.

[12]There are a few minor complications. Some households qualified for slightly higher thresholds due to, e.g., the number of dependent children. Generally, these thresholds were about NOK 25,000 to NOK 30,000 lower. We account for this in our empirical setting by applying these lower thresholds to qualifying households ("tax class 2", or, in Norwegian, "skatteklasse 2"). In addition, some low-income wealth-tax payers faced lower marginal rates due to a tax ceiling. However, this only affected about 0.03% of our sample. Finally, during 1992 and 1993, there was no third bracket, which implied that the top rate was 1.3% and the the max wealth-tax-induced equity risk premium was 32.5 basis points.

[13]This uses 1997 prices and the exchange rate of 7 NOK per USD. If we inflate the thresholds to 2021 prices, they are about \$28,000, \$59,000, and \$267,000 at 2021 prices. Further note that the third tax bracket was not in effect

below NOK 120,000 paid no wealth taxes. Households with TNW above NOK 530,000 would face a marginal wealth tax of 1.5%, but the first 120,000 would go untaxed, the next 115,000 would be taxed at a rate of 1.1%, and the following 295,000 would be taxed at a rate of 1.3%.

Since some assets enter at a discount, the effective marginal wealth tax rates are sometimes below the nominal rates, τ_{wj}. In particular, while safe assets, such as deposits or bonds enter without a discount, risky assets (stocks and mutual funds)[14], were subject to a one quarter value reduction ($d \approx 25\%$) before contributing to TNW during 1992–1997. This deflated the marginal tax rates on these assets to 0.825%, 0.975%, and 1.125%, respectively.

Over this period, the presence of these valuation discounts on risky but not on safe assets created an (additional) marginal equity premium (Δ) of 27.5, 32.5, and 37.5 basis points for taxpayers above the corresponding three wealth tax thresholds.

The 1998 reform. The reform scrapped the second threshold and thus effectively set the marginal tax rates at 0.9%, 0.9% and 1.1%. Importantly, it removed the valuation discount on risky assets. This *equalized* the marginal tax rates. This implies that the households in the bottom, top, and intermediate wealth tax bins (as of 1997) saw a reduction in the wealth-tax-induced risk premium (Δ, using the notation of section 2) of 27.5, 32.5, and 37.5 basis points, respectively.

We illustrate the effect of the 1998 reform on the wealth tax rates in Figure 1. The top, red and solid line corresponds to the effective marginal tax rate on safe assets for households in the top wealth tax bin. The corresponding dashed, red line provides the lower, marginal rate on risky assets. The difference between these two lines in 1997 provides the induced risk premium of 0.375%. In 1998, this difference was eliminated, and thus the two lines join. Similarly, the sets of green and blue lines describe the change in the marginal tax rates for households in the intermediate and lower wealth tax bins.

This reform offers an attractive natural experiment to study the effect of changing the equity risk premium. First, wealth taxes derive most of the revenue from the stock, rather than a given year's return. Hence, as argued in section 2, differently from shifts in the equity premium arising from taxes on *returns* on wealth, wealth tax-induced variation barely affects the volatility of returns (Gordon, 1985). Since volatility is held virtually constant, this allows us to isolate the pure effect of the equity premium on the portfolio allocation. Second, during the 1992-1997 period, the tax rules that gave rise to a wealth-tax equity premium stayed constant for most taxpayers, and so did the wealth tax thresholds and marginal rates.[15] Similarly, after the 1998 change, the new rules stayed fixed until 2004.

The reform is well-suited for testing frictionless models of portfolio allocation. First of all, the wealth tax rules were effectively constant for many years before and after the reform. This should both give investors ample time to understand and respond to change in the equity premium. Secondly, the equity premium change should be salient, as the adverse effect on after-tax stock

during 1993; hence $T^3_{1993} = \infty$.

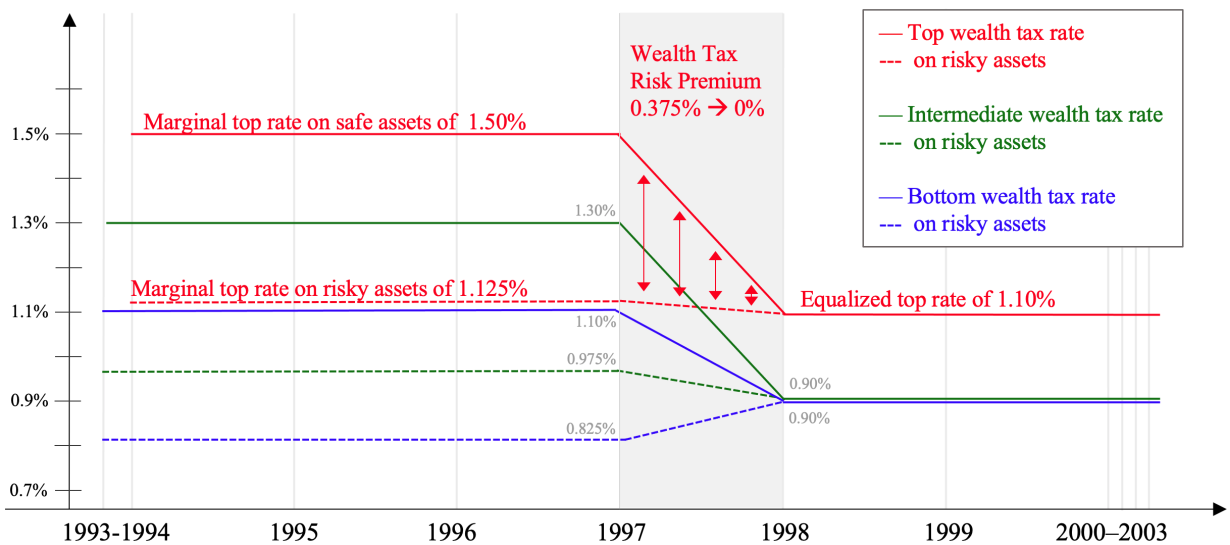
[14]For mutual funds that consist of both stocks and bonds, the tax authorities only apply the valuation discount to the stock component. We only include this stock component in our measure of stock market wealth, and the bond component in our measure of safe assets.

[15]The only change was the introduction of the third threshold (as indicated by the last term in equation 9) in 1994.

returns saw widespread media coverage (see Appendix A.2). Hence, inattention and information gathering costs, such as those modeled by Abel et al. (2007) and Alvarez et al. (2012), are unlikely to be amplified in our setting. Understanding the effects of the reform should in fact be quite straightforward: the Norwegian tax authorities provide detailed pre-filled tax returns each year with the relevant information. Prior to the 1998 reform these documents would show that stocks enter at a discount into the wealth-tax base, but that starting in 1998 stocks enter at their market value and are thus fully subjected to the wealth tax. In summary, the reform induces persistent and salient variation in the equity premium and thus provides a reasonable setting for testing frictionless models of portfolio allocation.

Figure 1: The Empirical Setting

This figure describes the quasi-experimental variation in the marginal tax rates on safe and risky assets. Prior to 1998, there were three wealth tax thresholds, above which households faced marginal wealth tax rates of 1.1%, 1.3%, and 1.5%. Assets held in stocks or mutual funds were subject to a 25% valuation discount, which lowered the marginal tax rates on these assets to 0.825%, 0.975%, and 1.125%, respectively. As of 1998, the marginal tax rates were 0.9%, 0.9%, and 1.1%. The nominal threshold cut offs were effectively unchanged other than that the second threshold ceased to matter. As of 1998, there was no valuation discount on stocks or mutual funds, which equalized marginal tax rates on safe and risky assets. Households remaining in the top, intermediate, and bottom tax bins, would see a reduction in the wealth-tax risk premium of 0.375 pp., 0.325 pp., and 0.275 pp., respectively. The wealth tax thresholds were NOK 120,000, 235,000, and 530,000. As of 1998, the second threshold is (effectively) removed.



Importantly, the 1998 reform was effectively a reversal of the 1992 reform that introduced the wealth tax risk premium. A lack of asset-class level data for the years prior to 1993 prohibits us from fully exploiting this initial reform using our difference-in-difference strategy. However, our empirical results are consistent with a positive effect of the introduction of the tax-induced equity premium in 1992 on the stock portfolio share in the subsequent years. In particular, the fact that the risk premium was introduced in 1992, and that our pre-trends are likely driven by delayed responses to this introduction is a key reason for why our baseline estimates do not adjust for pre-trends. In section 5.5, we show robust evidence supporting this strategy. While there have been other years with a wealth-tax-induced risk premium, the 1998 reform offers the cleanest experiment. For example, while the equity discount was reintroduced in 2005, this change was reversed by 2008 and occurred simultaneously with increased wealth tax thresholds. The short-livedness of the changes makes it difficult to pin down the effects of changes in the equity premium because, as we document, investors adjust their portfolio slowly.

Finally, we note that while the presence of a wealth tax threshold facilitates identification by comparing households initially above and below the threshold, it may also cause the "average" and "marginal" after-tax equity premia to differ. However, this wedge does not qualitatively affect our findings, as we discuss in Appendix A.3.

3.4 Sample Characteristics

We focus on data for the years 1993 to 2003. This time interval includes five years before and five years after the 1998 wealth tax reform.[16] For our analysis, we exclude households who owned private equity (non-listed stocks) in the four years leading up to the 1998 reform. We do so to stay consistent with the approach in the literature that studies the portfolio effects of the equity premium based on surveys or brokerage accounts data, where private equity wealth is typically excluded. It also allows us to avoid any potential confounding due to a shifting of assets out of limited liability companies (LLCs) and onto personal balance sheets.

Table 1: Summary Statistics

This table provides summary statistics for the main variables used in the analyses as of 1997. Taxable net wealth (TNW), Financial wealth (FW), and labor income are provided in NOK. USD/NOK exchange rate averaged about 7 during 1997. SMP denotes stock market participation and conditional SMS is the SMS for stock market participants. College is a dummy variable, and Biz/Law/Econ is a dummy variable for whether the highest-level of education is in any of these fields. Labor income includes labor-income related transfers such as pension income.

	Full sample		Wealth-tax payers		Non-payers	
	Mean	SD	Mean	SD	Mean	SD
TNW	160,948	234,755	365,153	151,239	-32,690	92,822
FW	91,562	111,612	147,542	130,170	38,437	48,934
Stock Market Participation	0.288		0.371		0.210	
Conditional Stock Market Share	0.242	0.239	0.213	0.222	0.290	0.257
Age	51.40	11.80	57.43	10.56	45.69	9.93
College	0.221		0.194		0.247	
Biz/Law/Econ	0.025		0.0183		0.030	
Labor Income	289,484	156,084	279,457	158,249	299,190	153,302
Observations	171,262		83,358		87,904	

We further only keep households with a positive and stably assessed housing value for at least three years leading up to the reform and who held at least NOK 5,000 (around $700) of financial wealth. These filters exclude zero-asset and zero-liability households with zero taxable net wealth who would be poor controls for wealth-tax payers.[17] We also restrict our sample to

[16] We omit 2004 since, during this year, the government announced that the wealth-tax risk premium would be reintroduced in 2005. We are unable to go further back than 1993 since 1993 is the first year with detailed enough data to distinguish between safe and risky asset holdings.

[17] During this time period, the first wealth tax threshold was fairly low at only NOK 120,000. Hence, to obtain a reasonably large control group, we also include households with negative taxable net wealth. This does not imply that they had negative market-value net wealth, however, due to the fact that the tax value of housing typically only corresponded to about 30% of the market value. Hence, for households with around zero taxable net wealth, there are two groups of households: (i) those with no house and zero market-value net wealth and (ii) those with a house but positive market-value net wealth. This first group is arguably a poor control group

households who either consistently paid or did not pay a wealth tax for four years leading up to the reform and to ensure that their pre-reform marginal tax rates were stable. This selection criteria is rooted in Jakobsen et al. (2020) who exploit a similar empirical setting. We further focus on individuals aged 30 years or above to avoid the imminent expectation of large changes to a household's financial situation caused by labor-market entry. Finally, we restrict our sample to households with at least minus NOK$250,000$ of taxable net wealth in 1997. This produces a sample that is close to evenly split between households who did and did not pay a wealth tax.

Table 1 shows summary statistics of the wealth and portfolio variables for our sample. Compared to non-payers, wealth-tax payers are older, more likely to participate in the stock market (37% compared to 21%), but conditional on participating, they invest a smaller share of financial wealth in stocks (21% compared to 29%), but have comparable labor income and education levels. Non-payers have negative taxable net wealth (TNW) on average, reflecting the fact that housing wealth is discounted but mortgage debts enter in their entirety (see Ring and Thoresen 2021 and Bjørneby et al. 2020 for more details). In Appendix Table A.2, we further split wealth-tax payers into their corresponding wealth-tax bins (i.e., above the first, second, or third threshold in 1997), and provide the same summary statistics. This shows that wealth-tax payers are older, but age is not increasing above the first threshold whereas stock market participation (SMP), the stock market share (SMS), and labor income are steadily increasing with the wealth-tax bin.

4 Empirical Approach and Identification

Our approach to identify the effect of a change in the equity premium relies on a difference-in-difference (DiD) strategy. The DiD strategy is the standard approach in the wealth tax literature for exploiting changes in marginal tax rates. A regression discontinuity design does not work since we do not observe marginal portfolio allocation (only the total allocation), hence we would not expect a discontinuous treatment effect.[18]

Our DiD strategy is described by the following reduced-form regression equation:

$$y_{i,t} = f_i + \eta_{c,t} + \sum_{s \neq 1997} \mathbf{1}[s = t] \times \beta_s \times Treat_{i,1997} + \upsilon_{i,t}, \tag{10}$$

where the outcome variable $y_{i,t}$ is either the year t household i portfolio share in stocks or an indicator for stock market participation; f_i is an individual-level fixed effect; and $\eta_{c,t}$ are year fixed effects estimated separately for different birth-cohorts, where c denotes the cohort.[19] These cohort-specific time effects capture jointly life-cycle portfolio effects as well as calendar time effects specific to the cohort. $Treat_{i,1997}$ takes the value 1 if i was above the first tax threshold in 1997. To ensure minimal changes in wealth-tax risk premium in the pre-reform period, we restrict the

for wealth tax payers who necessarily have positive market-value net wealth. Hence to obtain more homogeneity across treatment and control groups, we drop households who do not own a house.

[18]There are additional issues related to potential unobserved evasion being more prominent near the threshold, a regression discontinuity design would overweight these households. In addition, households very close to thresholds, whether above or below, may have shared similar beliefs about whether they would pay a wealth tax in the long run: since the fixed cost of stock market entry is relatively important, this may cause the treatment effect to be somewhat smooth near the boundary and thus hard to identify in a RDD framework.

[19]We group households into 10-year cohort bins based on the average age of household adults in 1997.

sample to the households for which $Treat_{i,t}$ is constant for four years prior to the 1998 reform (i.e $t = 1994, ..., 1997$).

Because the specification allows us to estimate year-specific responses to the reform (β_t), we can examine whether there is gradual portfolio adjustment to the 1998 change in the equity premium. In a frictionless world, we would have that $\beta_{1998} = \beta_{1999} = ... = \beta_{2003}$. In a world with adjustment frictions in the portfolio share (e.g., Gârleanu and Pedersen 2013 or Bacchetta and Van Wincoop 2010), we would expect to find that β_t is (in absolute value) increasing in t for $t > 1998$. This would be consistent with gradual portfolio adjustment, where the adjustment trajectory is steeper when adjustment costs are higher.

To directly estimate the cumulative response to a unitary change in the equity premium, Γ_t, we multiply $Treat_{i,1997}$ with the reform-induced reduction in the after-tax risk premium. This is simply the negative of the pre-reform wealth-tax risk premium, $-WTRP_{i,1997}$, which varies at the wealth-tax bin level, $b \in$ {"below first wealth-tax threshold", "above first threshold", "above second threshold", "above third threshold"}.

$$y_{i,t} = f_i + \eta_{a,t} \quad + \sum_{s \neq 1997} \mathbf{1}[s = t] \times \Gamma_t \times (-WTRP_{i,1997}) \times Treat_{i,1997} + \upsilon_{i,t}. \tag{11}$$

This specification makes it clear that the empirical variation in the (expected) equity premium is about 37 basis points (the highest wealth-tax risk premium in 1997). Since almost no households have a 100% ex-ante stock market share, we thus do not face the attenuation issues documented by Beutel and Weber (2022).[20] In the specifications that investigate heterogeneous effects according to some discrete, time-invariant household characteristic, h, we estimate the cohort fixed effects ($\eta_{a,t}$) and the sensitivities (ε_t) separately for each h. To ease the readability of regression tables and reduce noise, the underlying regression equations are adjusted to only let post-reform responses vary at the biennial level, i.e., 1998–1999, 2000–2001, and 2002–2003. The year-by-year coefficients are reported in the figures. In the specifications that adjust for pre-trends, we include the term $\zeta \cdot Treat_{i,1997} \cdot t$, where the additional parameter, ζ, is estimated by omitting also the years 1994–1996 from the set of years for which we estimate year-specific responses as opposed to just 1997 in our baseline specification. The specifications that adjust for pre-trends estimate ζ separately for each b when using equation 11 and separately for each h when estimating heterogeneous responses.

We also note that a recent literature shows that standard two-way fixed effects approaches may in some cases identify non-convex averages of the underlying treatment effects in the presence of heterogeneity (Goodman-Bacon 2021; Baker et al. 2022; Sun and Shapiro 2022). An important source of this problem is heterogeneity over time in treatment effects (e.g., sluggish responses). Time heterogeneity is not an issue in our setting since we explicitly estimate dynamic effects. However, as we show in section 5.3, there is heterogeneity in cumulative effects across groups. Fortunately, we find that the pooled estimates are indeed convex averages of group-specific effects. Nevertheless, we show that our main estimates are virtually unchanged when re-estimated with the

[20]Beutel and Weber (2022) show that belief outliers combined with portfolio constraints attenuate estimates of Γ. For example, the inability of households with very high equity premium beliefs to hold a stock market share of more than 100% attenuates OLS estimates toward zero. This implies that a simple calibration exercise may greatly overstate the implied risk aversion. While the 0% lower bound may in principle be an issue, our finding of a fairly contained coefficient of relative risk aversion suggests that, if present, it is inconsequential in our setting.

heterogeneity-robust estimation package *did_multiplegt* by De Chaisemartin and d'Haultfoeuille (2020).

4.1 Discussion of Identification Strategy

Our identification strategy relies on comparing households who were subject to the wealth tax prior to the 1998 reform to those who were not. In addition, when we directly estimate the implied effect of changing the risk premium using equation 11, we also exploit the fact that households in higher tax brackets were treated more. The mechanical implication of our approach is that treatment and control groups differ in terms of their 1997 TNW, and thus also in terms of its components such as financial wealth. Our differences-in-differences strategy follows Jakobsen et al. (2020) and thus relies on the assumption that households whose 1997 TNW was above the wealth tax thresholds did not *differentially* change their portfolio allocation for reasons unrelated to the wealth tax reform.[21] In section 5, together with the main results, we discuss objections to this assumption. Here we address three potential issues.

Mean reversion. Comparing households above and below tax thresholds when marginal tax rates change is ubiquitous in the public finance literature (see, e.g., Gruber and Saez 2002). One key concern in this literature is mean reversion. This is because treatment is assigned based on past values of the outcome variable (Weber, 2014; Jakobsen and Søgaard, 2020), which may have a transitory (mean-reverting) component. Our strategy does not suffer from this problem directly, as there could be mean reversion in taxable net wealth even if the stock market share remains constant. However, to the extent that mean reversion is driven by changes in stock market wealth, this issue would apply to our setting. The issue is ameliorated by conditioning on households' being subject to the wealth tax for several years prior to the reform. We test for whether there is a residual issue by assigning treatment in 1995, rather than 1997, and plot the dynamics of the stock market share prior to and after the reform in 1997 in Appendix Figure A.1. This allows us to test whether there is mean reversion after treatment assignment. If there is, we should observe a reduction in the stock market share during 1996–1997. We find no evidence of this. In fact, the dynamic pattern of adjustments is nearly identical to our main results, which shows that mean reversion is unlikely to play a role in our empirical framework.

Pre-trends. We find that our treatment group (wealth-tax payers) were increasing their stock market share more than the control group leading up to the 1998 reform. However, as soon as the reform occurred, this relationship reversed. This is reassuring in the sense that treated households are unlikely to have been on a pre-existing trend that would explain our findings.[22]

[21] Jakobsen et al. (2020) study the effect of abolishing the Danish wealth tax on reported wealth holdings. Their treatment group consists of ex-ante wealth-tax payers who experience a drop in marginal tax rates. The control group consists of households initially below the wealth-tax threshold who do not experience such a change. While the Jakobsen et al. (2020) setting offers a uniform reduction in marginal tax rates across asset classes, rather than a differential change as in our case, the otherwise-similar setting motivates us to follow their difference-in-differences identification strategy closely.

[22] More rapid yearly reductions in the risky share as one grows older would otherwise be a concern as wealth-tax payers, on average, are older. At any rate, our empirical specification allows for cohort-specific age effects (equation 11)

The presence of an upward-sloping pre-trend is not surprising given the fact that the wealth-tax risk premium was introduced in 1992. Hence, we are at least partially observing gradual responses to the 1992 equity premium introduction. Importantly, this does not violate the implicit parallel trends assumption of our identification strategy. In fact, adjusting for the pre-trend by detrending our estimates would inflate the causal effect of the 1998 reform, essentially by double counting the treatment effect. As we discuss in section 5.5, it is indeed likely that the increasing stock market share for the treated households compared to the control group during the pre-1998 reform period reflects only delayed responses to the 1992 introduction of the wealth-tax risk premium. [23]

Passive movements. In our setting, treated households have higher initial stock market shares than other households. Hence, their portfolio share tends to be more sensitive to aggregate stock price movements. To verify that this is not driving our results, we show that our findings are robust to controlling for the initial (1993) stock market share interacted with year dummies to capture aggregate stock price variation (see Appendix Figure A.4). We also employ a more direct approach, following Calvet et al. (2009), which removes variation in the stock market share caused by aggregate market movements, i.e., "passive" movements in stock market portfolio shares, leaving movements that reflect "active" portfolio allocation decisions. More specifically, we normalize the stock market share to zero in 1997 and, for any year t, we deduct the change in the stock market share that would arise if the investor was simply holding the market portfolio. We do this both for the case when dividends are not reinvested (as in Calvet et al. 2009) and for the case when dividends are invested (as is the case for many mutual funds). As we discuss further in Section 5.2, both these approaches produce estimates of the dynamic responses that are virtually identical to our baseline estimates.

5 Results

Before discussing our main results based on our differences-in-differences specification, we start showing some preliminary descriptive visual evidence.

5.1 Descriptive Evidence

Figure 2 plots how household stock market shares vary with taxable net wealth before and after the 1998 reform.[24] To capture the overall effect of changes in the equity premium, we plot the values of the unconditional stock portfolio share. The figure reveals that beyond a level

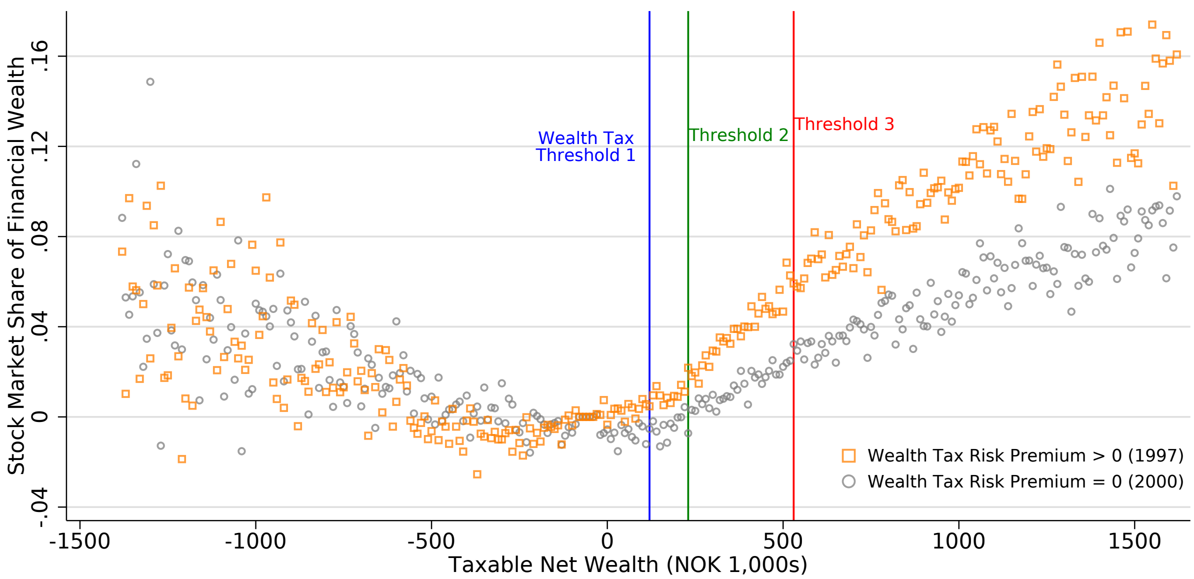
[23]Our qualitative findings of strong responses to equity premium changes do not hinge on the pre-trend being entirely caused by the 1992 introduction of the wealth tax risk premium. If we assume that there is some long-run underlying trend (that we should adjust for), then, if we detrend the post-period estimates, we are doing two things: (i) we remove the trend from the raw post-1997 estimates, and (ii) we add the first-period treatment effect to the second-period effects. Hence, if we multiply the detrended second-period estimates by 0.5, we obtain the average sensitivity to the wealth tax risk premium, where the average is across an introduction and a removal. This approach is agnostic to whether or not there exists an underlying long-run trend. By visual inspection of Figure 3, which shows the dynamic effects on the stock market share, we see that 0.5 times the detrended estimates are indeed slightly larger in magnitude than the raw estimates, hence our qualitative findings would remain unchanged.

[24]This graphical analysis is inspired by a similar, more formal approach in Jakobsen and Søgaard (2020)

difference (essentially, a year fixed effect), households below the first wealth tax threshold had similar portfolio allocations before and after the reform. However, once we consider households who paid a wealth tax, we see that their stock market shares were much higher prior to the reform when the wealth tax subsidized the equity risk premium. We also see that the difference between 1993 and 1998 portfolio shares diverges as taxable net wealth increases. This is consistent with the fact that the wealth-tax risk premium was increasing in taxable wealth.

Figure 2: Wealth-tax Payers Had Higher Stock Market Shares when the Wealth-tax Risk Premium Was in Place

This figure plots the average stock market share ($SMS_{i,t}$) of financial wealth within NOK 10,000 bins of taxable net wealth ($TNW_{i,t}$). Grey (circular) scatter points provide the within-bin mean SMS during 2000, which is 3 years after the wealth-tax risk premium was removed. Orange (square) scatter points provide SMS for 1997, when the wealth-tax risk premium was in place. Within each year, SMS is normalized to an average of zero for households with TNW $\in [-300\ 000, -250\ 000]$. This normalization removes the gap of 3.9 pp. higher SMS during 2000 (i.e., takes out a year fixed effect). The blue, green, and red vertical lines indicate the wealth tax thresholds. During 1997, being located to the right of any of these lines is associated with an increasingly higher wealth-tax risk premium.



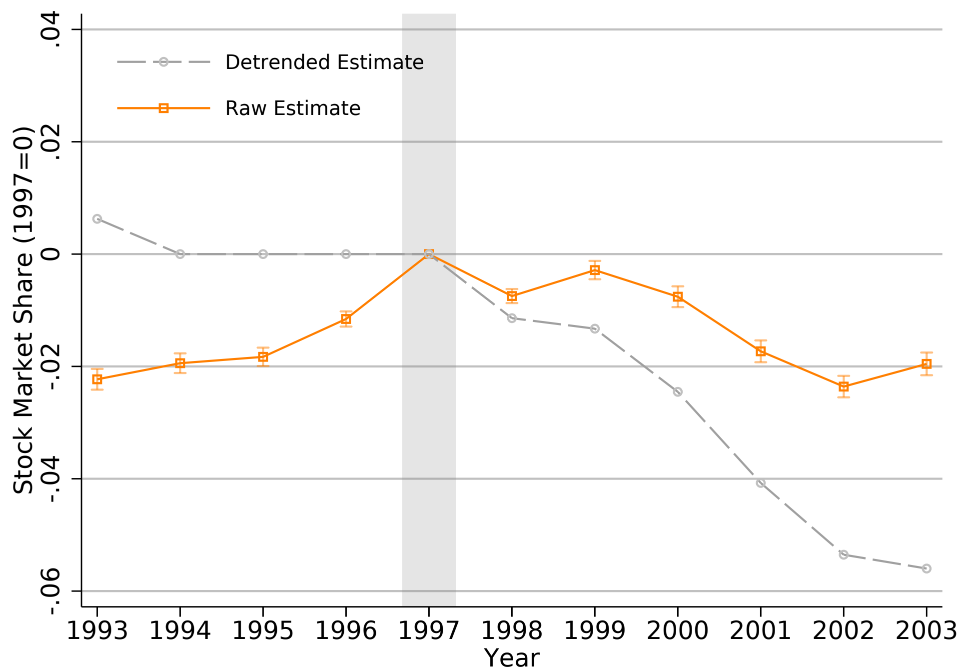
In other words, the figure shows that wealth-tax paying households allocate more wealth to the stock market when the wealth-tax risk premium was in place (in 1997) relative to when it was not (in 2000). This is highly suggestive evidence that individual investors adjust their portfolios in response to changes in the equity risk premium. Our subsequent analyses rely on a difference-in-differences framework that allows us to take out household fixed effects and study dynamic responses.

5.2 Effects on Portfolio Allocation

Figure 3 visualizes the effect of the reduction in the wealth-tax-induced risk premium on households that are initially above the wealth tax threshold ($Treated_{i,1997} = 1$) on their unconditional stock market share. The orange line with hollow squares provides the differences in stock market shares between treated households and controls, estimated with equation 10. This specification takes out household fixed effects and normalizes the difference between treated and controls to be zero in 1997. We see that treated households were on an upward-sloping trend prior to 1997, but that this relationship reversed immediately following the reform. While treated households increase their stock portfolio share more than non-treated during 1993–1997, the 1998 reform-induced reduction in the risk premium triggered a noticeable effect on their portfolio allocation by inverting the previous pattern. Visual inspection of Figure 3 shows that households need about 5 years to adjust to equity premium perturbations. While there are no directly comparable findings in the literature, Choukhmane and de Silva (2021) find similar dynamics when studying how individuals substitute away from default options in employer-sponsored retirement schemes.

Figure 3: Dynamic Effect of a wealth-tax-induced Reduction in the Equity Premium on the Stock Market Share of Financial Wealth

This figure shows the reduced-form effect of the reduction in the wealth-tax-induced risk premium on the stock market share of financial wealth. We employ a differences-in-differences methodology, in which treated households are those initially above the wealth tax threshold ($Treat_{i,1997} = 1$). The orange line with hollow squares provides non-detrended estimates using equation (10). The gray line with hollow circles provides detrended estimates, where the years 1994, 1995, and 1996 are omitted to estimate the time trend for the treated. Capped horizontal lines provide 95% confidence intervals for the non-detrended estimates, where the standard errors are clustered at the household level.



The gray dashed line provides detrended estimates for the post-period. The underlying specification assumes that the counterfactual during 1998–2003 is that the pre-period trend would continue. Accordingly, the detrended estimates of the effect of the 1998 reform are much larger in magnitude. This detrending adjustment is appropriate to the extent that we expect that, absent the treatment, treated households would have continued to increase their stock market shares differentially more than the control group. However, it is not obvious that we should make this adjustment: the fact that the wealth-tax-induced risk premium arose in 1992 suggests that the observed positive pre-trend may simply be a delayed response to the increased after-tax risk premium. If delayed responses to the 1992 reform can fully account for the observed pre-1998 dynamics of the risky share, then we should not adjust for pre-trends. This is because the delayed response would seize once the wealth-tax risk premium is removed. An intermediate possibility is that the pre-1998 pattern reflects both the response to the 1992 increase in the equity premium as well as some pre-existing trend in which case the true response to the change in the equity premium in the years following the 1998 reform would lie between the two lines.

It is our interpretation that pre-trends are primarily driven by delayed responses to the 1992 introduction of the wealth tax risk premium. The fact that the pre-trends and post-1998-reform responses are virtually symmetrical is consistent with this interpretation. As the 1992 reform was effectively reversed by the 1997 reform, households bring their stock portfolio shares back to the pre-1992 levels. In section 5.5, we use supplementary data available over a longer horizon that supports this interpretation.

Table 2 provides our main estimates of the sensitivity of the stock market share to changes in the equity premium. These are estimated using equation 11, which enriches equation 10 (which is used for the figures) by also exploiting higher treatment intensity for households in higher wealth-tax brackets. The specification also allows us to directly estimate the sensitivity of the portfolio share to a unitary change in the risk premium. Anticipating the evidence in section 5.5, we show regression estimates that do not adjust for pre-trends. To improve precision and increase the readability of our tables, we report biennial estimates by allowing post-reform responses to vary at a two-year frequency, i.e., 1998–1999, 2000–2001, and 2002–2003. The first column shows the baseline estimate for the whole sample. The other columns show heterogeneity effects for subgroups of investors, which are discussed in the next subsection.

The coefficients in Table 2 should be read as showing cumulative responses to a unitary increase in the equity risk premium. We find that the response to the change in the equity premium induced by the new tax regime is gradual, consistent with portfolio adjustment frictions. The response in the first two years is only 24% of the cumulative response as of 2002–2003. We use this ratio as a measure of response speed and report it at the bottom of the table. As Figure 3 shows, the cumulative response continues to strengthen to about 2001 and then stabilizes, suggesting that the cumulative sensitivity estimate for the last sample years, 2002–2003, provides a reasonably good approximation to the complete long-run response. This cumulative estimate is 6.97, which is a material response. In section 5.6, we use the estimated cumulative response of the unconditional stock portfolio share to infer the implied coefficient of relative risk aversion. We establish that it falls in the range 1.8–2.8, which is consistent with the effects implied by a standard frictionless Merton (1969) portfolio model.

In sum, our estimates suggest that investors do respond strongly to changes in the equity premium—at least when these changes are relatively persistent, and only when enough time has been allowed to observe the full portfolio adjustment as in our setting. The immediate response is instead weak and far below the values implied by Mertonian portfolio models with no adjustment costs.

The sluggish portfolio response evident from column (1) of Table 2 is consistent with the partial-equilibrium portfolio model of Gârleanu and Pedersen (2013) where investors face quadratic portfolio adjustment costs, and stock returns show predictable variation. In this model, investors respond slowly to changes in expected returns. The current portfolio share is a linear combination

Table 2: The Sensitivity of the Stock Market Share to Wealth-tax-induced Variation in the Equity Premium

Notes: This table shows the estimated cumulative effect of a unit increase in the equity premium on the stock market share (SMS) of financial wealth. Column (1) provides the results from estimating equation (11) when estimating treatment effects at the biennial level, and without adjusting for pre-trends. Columns (2) and (3) provide results on heterogeneous responses. Estimates in a and b columns are for values below and above a percentile cut-off value for the splitting variable, and are estimated in the same regression. Columns (2a) and (2b) consider the heterogeneity with respect to the SMS in 1993; which prior to the split at the 75th percentile has been residualized with respect to age, income, and financial wealth. Columns (3a) and (3b) split households by portfolio adjustment cost (PAC) (see equation 12), interpretable as an index of adjustment costs. We define "response speed" as the first two years' response divided by 2002–2003 cumulative response. Standard errors are clustered at the household level. Stars indicate significance at the 10%, 5%, and 1% levels.

		Heterogeneous effects by			
		$\widetilde{SMS}_{1993}$		Portfolio Adjustment Cost	
	Full sample	< p75	> p75	< p25	> p25
	(1)	(2a)	(2b)	(3a)	(3b)
1998-1999	1.63***	1.88***	3.72***	2.95***	1.34***
	(0.20)	(0.22)	(0.47)	(0.39)	(0.23)
2000-2001	3.92***	4.50***	7.07***	4.46***	3.79***
	(0.28)	(0.30)	(0.51)	(0.52)	(0.32)
2002-2003	6.91***	6.27***	16.36***	5.52***	6.97***
	(0.29)	(0.31)	(0.68)	(0.55)	(0.33)
Response speed	0.24	0.30	0.24	0.53	0.19
Mean FW, 1998–2003	171,937	193,930	108,932	263,482	138,512
Mean SMS, 1998–2003	0.088	0.082	0.103	0.124	0.075
N	1,881,975	1,879,216		1,878,318	

of the past share and a target share that varies with the expected returns, where the weight on the past share is increasing in the cost of adjustment. Other models with frictions, for example due to costly information acquisition or portfolio adjustment, have similar implications (see Gabaix 2019 for a review). In the absence of adjustment frictions, portfolio allocation responds instantaneously to movements in the equity premium. The presence of frictions dampens the immediate response while its cumulative magnitude increases as investors adjust, resembling the pattern in Figure 3.

The finding that households are slow to adjust to changes in the equity risk premium may help explain the low sensitivity found in the survey-based literature. Giglio et al. (2021) show that there is substantial time-series variation in the expected risk premium: the average expectation was 5% in 2017, dipped to around 2% following the onset of the COVID pandemic, and had risen to above 6% by February 2020. Thus, to the extent that households' portfolio adjustment is sluggish, typical cross-sectional regressions of concurrent stock market shares on expected returns may provide low estimated sensitivities. In addition, if one exploits time variation in elicited risk premia that is largely transitory and portfolio adjustments are costly, then the investor's optimal response may be to stick to the current allocation. Only long lasting changes in the equity premium would be informative of investors' frictionless sensitivity. Indeed, Gârleanu and Pedersen (2013) show that portfolio sensitivities are stronger if the change is long lasting.

Our estimates rely on the identifying assumption that households affected by the reform did not *differentially* change their portfolio allocation compared to unaffected households (ex-ante non-payers of the wealth tax) for reasons unrelated to the wealth tax reform. One objection to this assumption is that because households affected by the reform are also wealthier, differential portfolio behavior of the treated compared to the non-treated may reflect different sensitivities of wealthier households to stock market movements after the 1998 reform. This would imply that we capture "passive" movements in the portfolio share rather than slow "active" portfolio adjustments. We find this to be unlikely. Firstly, we show in Appendix Figure A.4 that our results are virtually unaffected by including 1993 (the earliest we observe) stock market shares as a control variable. This initial share is interacted with year dummies. Hence, this control term addresses the possibility that wealth-tax payers, with higher initial stock market exposure, may have been on differential (possibly nonlinear) trends for reasons unrelated to changes in the wealth-tax risk premium. Secondly, we find that our results are robust to using a measure of the stock market share that is purged of passive market movements following the methodology discussed in Section 4.1. Appendix Figure A.2 shows that removing passive variation in stock market shares leaves our results unchanged, implying that the entire difference between treated and untreated households arises from differential active portfolio adjustments.

As shown by Grossman and Laroque (1990), in the presence of costly to trade housing wealth, the portfolio share in stocks may be sensitive to transaction costs when trading occurs. In particular, the investor behaves in a significantly more risk averse manner around the time housing transactions.[25] To make sure that the differential portfolio share dynamics that we observe by comparing treated and non-treated households is not a reflection of differences in the frequency of housing transactions, we perform a robustness check where we include controls for whether a household sold of bought housing in the surround years. As shown in Figure A.3, this has no impact on our results.

5.3 Heterogeneity in Responses

We perform additional analyses that contrast the response of specific groups for whom theory predicts different responses to a change in the equity risk premium.

Risk tolerance. The Merton (1969) model implies that more risk-tolerant investors should respond more strongly to a variation in the equity premium. Columns (2a) and (2b) of Table 2 split investors on the basis of the stock portfolio share in 1993, well before the reform, which we use as a proxy for the investor's risk tolerance. To get closer to a proxy for risk tolerance, we use the residualized portfolio shares, $\widetilde{SMS}_{1993}$, which are created by removing variation that can be explained by age, financial wealth, and income levels that can affect the portfolio share due to life-cycle effects. Consistent with the model predictions, we find that households with a (residual) portfolio share above the 75th percentile exhibit a cumulative response of 16.36, which is more than 2.5 times the cumulative response exhibited by the remaining sample.

Notice that while the difference between high and low-risk-tolerance investors holds in all years in the sample, the speed of the response—measured by the ratio of the 1998-1999 response

[25] When the desired wealth composition is achieved following a housing transaction, the investor wants to avoid further trading (and trading costs). To avoid further trading, the investor adopts a more cautious investment strategy, with a lower share in stocks.

to the cumulative response in 2002-2003—is similar for the two groups: 30% for the below and 23% for the above-75th-percentile group. This suggests that our risk tolerance proxy allows us to split households based on risk aversion while keeping traits that are correlated with adjustment costs constant.

Portfolio adjustment costs. As a second exercise, we split the sample on the basis of household portfolio adjustment cost which we infer from predicted portfolio rebalancing speed. Calvet, Campbell, and Sodini (2009) define rebalancing speed as the fraction of passive changes in the stock market share due to variation in stock prices that is undone by active trading. This index is informative of the size of adjustment costs. We construct this index according to Calvet, Campbell, and Sodini (2009). They estimate portfolio rebalancing equations allowing rebalancing strength to depend on household observables and report the estimated coefficients. We use these coefficient estimates to impute the index to the households in our sample. While we do not have access to the exact same set of variables used by Calvet, Campbell, and Sodini (2009),[26] our predicted measure has similar variation: the 5th to 95th percentile range is 0.25, which is close to the range of their measure of 0.33.

Let PRS_i denote the household predicted rebalancing speed. We define an index of household i portfolio adjustment cost as

$$PAC_i = [1 - PRS_i/max_i(PRS_i)] \tag{12}$$

conveniently belonging to the [0,1] interval. We use this to study heterogeneity in portfolio sensitivities to changes to the equity premium in columns (3a) and (3b) of Table 2. Our results reveal that low-adjustment-cost households (bottom quartile of PAC) do in fact respond faster to changes in the equity premium than households facing higher portfolio adjustment costs. The portfolio share coefficients of the low-cost group reveal a significantly stronger response in the initial years following the reform (more than twice as large), but a slightly smaller cumulative response. The ratio of the initial to final cumulative response is more than 50% for the low-adjustment-cost group and 20% for the higher-cost group in column (3b). These findings are consistent with low-adjustment-cost households needing less time to get closer to their long-run response to the change in the equity premium.

5.4 Effects on Stock Market Participation

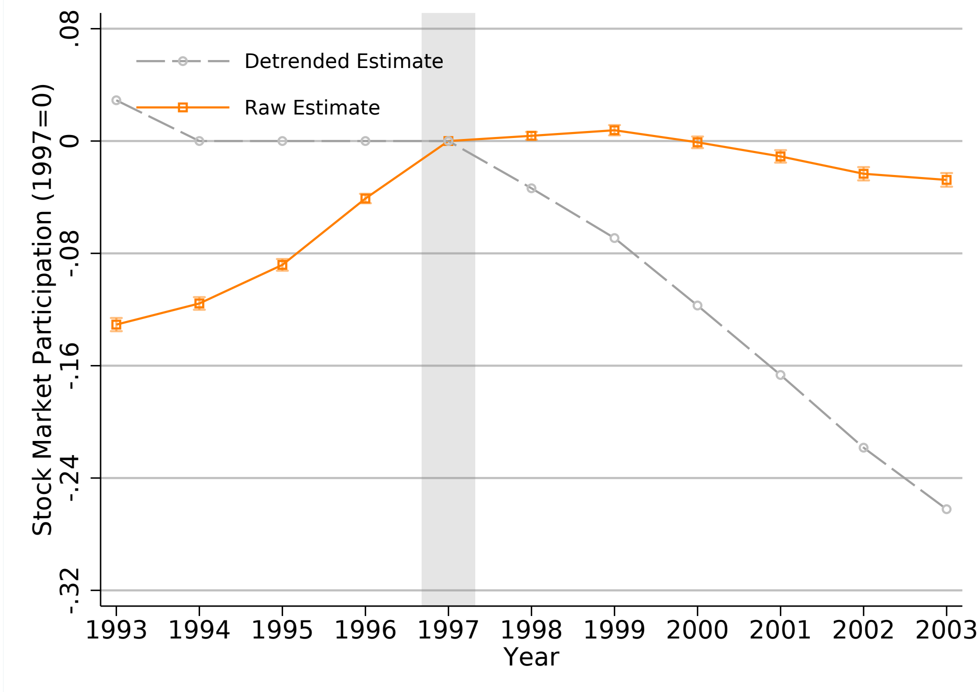
We present our main findings on stock market participation in Figure 4. This figure provides the cumulative effect on participation by comparing ex-ante wealth-tax payers with non-payers. The difference in participation rates is normalized to be zero in 1997, the year prior to the reform.

[26]See their Table IV. We do not have access to the same portfolio characteristics, so these are ignored. For the financial characteristics, we use gross labor income assuming a 35% tax rate to get disposable income; we ignore private pension income; we also ignore changes in log financial wealth; unemployment dummy is set to 1 if anyone in the household receives unemployment income; retired dummy is set to 1 if the mean household age is above 67; we omit the student dummy (age restrictions mean that we are unlikely to have many of these in our sample); and we use the number of household adults rather than family size. We use the coefficients from the first column in their Table IV to predict the adjustment speed measure.

We show both pre-trend adjusted (gray dashed line) and unadjusted (solid orange line) estimates. Focusing on our preferred, unadjusted estimates, we see that stock market participation gradually rose by about 13 percentage points in the years prior to the 1998 reform. This mirrors our findings on the stock market share, which we interpreted as a gradual adjustment to the 1992 introduction of the tax equity premium. Interestingly, we find that the response to the reversal of the equity premium in 1998 is essentially flat for the first three years after the reform and only declines in the last three sample years.

Figure 4: The Effect on Stock Market Participation

This figure shows the effect of the reduction in the wealth-tax-induced risk premium on stock market participation ($1[SMS_t > 0]$), by comparing households initially above the wealth tax threshold ($Treat_{i,1997} = 1$) with those below. The orange line with hollow squares provides differences between the treatment and control group, estimated using equation (10). The gray line with hollow circles provides the detrended differences. Table A.1 provides the estimated sensitivities from equation 11, i.e., the implied responses of a unitary change in the equity risk premium. Capped horizontal lines provide 95% confidence intervals for the detrended estimates, where the standard errors are clustered at the household level.



The presence of a wealth-tax equity premium before 1998 induced considerable growth in the participation rate. When this incentive was removed in 1998, the participation growth immediately halts and, after some time, turns slightly negative. However, there is no symmetric reversal, which suggests that those who were nudged into participation do not reverse their decision to participate. This asymmetry is consistent with the presence of a one-time fixed participation cost that non-participants incur only upon their first entry into the market (as in Alan 2006). The fact that participation does drop slightly after a few years following the 1998 reform, suggests that investors also face some per-period fixed participation cost. As the optimal portfolio share adjusts slowly downward, some investors find it desirable to exit and save on the per-period cost.

In Appendix Figure A.5, we decompose the participation effect into entries and exits from the stock market. This shows that our findings on participation are fully driven by increased entry rates prior to 1998 and decreased entry rates after the wealth-tax risk premium was removed. Exit rates are hardly affected by the wealth-tax risk premium. These findings highlight the importance of asymmetric participation costs: once households have incurred the larger one-time entry cost, perturbations in the equity premium are unlikely to induce an exit, as there is less to save, in terms of the per-period cost, by exiting.

Importantly, the fact that participation responds slowly to the equity premium increase is consistent with inattentive investors who accrue gradually to the market as in Duffie (2010), but also with the fact that investors' wealth is a slow-evolving variable, and thus it takes time for it to cross the threshold that triggers participation.

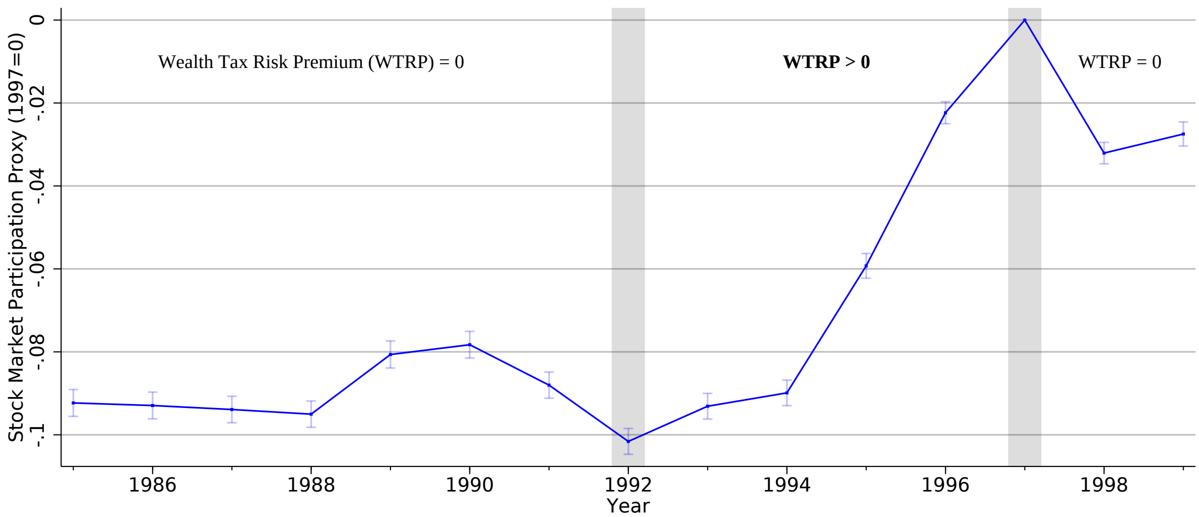
In Appendix Table A.1 we show estimates of equation (11) and of the cumulative marginal effects on participation of the 1992 increase in the equity premium, which is 44.2 (SE=0.75, see first row). Similarly, we estimate at 9.1 (SE=0.76) the cumulative marginal effect on participation of the decrease in the equity premium following the 1998 reform (last row). The underlying assumption is that the cumulative increase in participation before 1998 (after 1997) is the causal effect of the 1992 introduction (1998 reversal) of the wealth-tax risk premium. In section 5.6 below, we discuss how we can use these estimates to recover the implied one-time and per-period participation costs.

5.5 Attributing pre-trends to the 1992 introduction of the wealth tax risk premium

Our empirical findings clearly show that households who were affected by the drop in the wealth-tax risk premium in 1998 were on upward-sloping trajectories in terms of both participation and their stock market share in the years before. While we have presented both pre-trend adjusted and unadjusted results, our preferred approach does not adjust for pre-trends. This approach is justified under the assumption that pre-trends are causally driven by the earlier 1992 introduction of the wealth tax risk premium. Because our main data begin in 1993 we cannot use our main outcome variables (e.g., stock market share of financial wealth) to test the hypothesis that the 1993–1997 effects are part of a longer-run pre-trend starting well before 1992.

Figure 5: Long-run Pre-trends

The figure uses stock-saving tax deductions (independent of the wealth-tax scheme), which extend back to 1985, as a proxy for stock market saving to show pre-trends for a longer horizon. We regress an indicator for claiming stock-saving deductions at the individual level (*AMS fradrag*) on year and household fixed effects as well as year fixed effects interacted with the treatment dummy, $Treat_{i,1997}$, and report the estimated coefficients on $Treat_{i,1997}$ in the figure. Capped lines provide 95% confidence intervals.



However, some data on investment behavior is available for the years prior to 1993. Specifically, during 1982–1999, all income-tax paying households were eligible for a so-called stock saving deduction ("AMS fradrag" in Norwegian). Importantly, this stock-saving incentive was independent from the wealth tax scheme. Individual taxpayers could deduct annual purchases of stocks and mutual fund shares up to at most NOK 5,000, and reduce their income tax bill by 15% of the amount (Statistics Norway, 1997).[27] The potential income tax reduction was modest at NOK 750 (USD 100). However, we observe all these deductions in our data going back before 1993, and can use them to trace "active" stock market investing due either to entry into the market by non-participants or to increased exposure by the existing participants.[28]

We produce Figure 5 by re-estimating our baseline specification (equation 10), using an indicator for claiming the stock-saving deduction as the left-hand-side variable. We use the same sample of households and define treatment as before. Hence, we estimate the effect of paying a wealth tax in 1997 on whether households claimed the stock-saving deduction during the (extended) 1985 to 1999 period. This allows us to understand whether the pre-trends in portfolio choices observed in Figures 3 and 4 are driven by long-run trends from before the 1992 tax reform, or if they are causally driven by the 1992 introduction of the wealth tax risk premium.

The findings, shown in Figure 5, are striking. From 1992 to 1997, wealth-tax payers were increasingly more likely to actively purchase stocks (i.e., obtain the stock-saving deduction). The process reverts when the wealth-tax risk premium vanishes in 1998. The 1992–1997 trend closely

[27] The deduction is claimed the same year as the stocks are purchased. If the stocks are sold within five years, the investor must repay the tax reduction through a negative deduction in the year of the subsequent sale.

[28] In, e.g., 1997, when our variables overlap, we find that 55% of entrants (stock market share goes from zero to positive) are captured by a dummy for obtaining a stock-saving deduction. This is found by regressing a stock-saving deduction indicator variable on the stock market participation indicator among the previous years nonparticipants, where the coefficient on participation is 0.55 (t-statistic = 465.63)

resembles that found when considering stock market shares and participation. Crucially, this long-run analysis reveals that the stark 1992–1997 trend is not part of a longer-run trend. During the 1985–1992 period, the difference in active stock purchases between wealth-tax payers and non-payers is fairly constant and the cumulative change is weakly negative.

In summary, our findings suggest that the 1992–1997 pre-trends in Figures 3 and 4 are driven by the 1992 introduction of the wealth-tax risk premium. This means that these trends should not be used to infer counterfactual trends during 1993–2003. Accordingly, our main specification should not adjust for pre-trends.

5.6 Backing out investor risk aversion and participation costs

We use our empirical findings to get a sense of what the underlying participation costs are and to infer the value of the coefficient of relative risk aversion. We do these calculations using both the cumulative response to the 1992 and to the 1998 changes in the equity premium. This allows us to infer both one-time and per-period participation costs as well as the implied coefficient of relative risk aversion. We discuss the details of this exercise in Appendix A.4, and we show the resulting estimates of participation costs in Table A.3

Risk aversion. To back out risk aversion, we need an estimate of the implied effect on the conditional stock market share. We obtain this from our estimates on participation and the unconditional share via the following identity: $\alpha_u = \pi \times \alpha$, which implies that

$$\frac{\partial \alpha_u}{\partial r_e^n} = \pi \frac{\partial \alpha}{\partial r_e^n} + \alpha \frac{\partial \pi}{\partial r_e^n}. \tag{13}$$

If we focus on the pre-1997 period, the cumulative difference gives us an estimate of the effect of the 1992 introduction of the wealth tax risk premium of 7.35:

$$\pi_{1992} \frac{\partial \alpha}{\partial r_e^n} + \alpha_{1992} \frac{\partial \pi}{\partial r_e^n} = 7.35,$$

where π_{1992} and α_{1992} is the participation rate and the stock market share among the participants in 1992, respectively. As discussed in section 5.4, our estimate of $\frac{\partial \pi}{\partial r_e}$ from the initial 1992 increase in the equity premium is 44.20 (Appendix Table A.1, first line). Using the estimated $\frac{\partial \pi}{\partial r_e^n}$, we compute the implied cumulative marginal response of the conditional share as

$$\frac{\partial \alpha}{\partial r_e^n} = (7.35 - \alpha_{1992} \times 44.2)/\pi_{1992}. \tag{14}$$

To retrieve the investor risk aversion we invoke Merton formula, assuming $\alpha = \frac{r_e^n}{\gamma \sigma^2}$ and thus $\frac{\partial \alpha}{\partial r_e^n} = \frac{1}{\gamma \sigma^2}$. In our context, α is the share of stocks in financial wealth. The Merton formula applies to the share of stocks on a broad notion of wealth, which also includes real assets, most significantly housing. When housing is costly to trade, as in Grossman and Laroque (1990) and Alvarez et al. (2012), the Merton formula for the share of stocks in financial wealth still holds, but the relative risk aversion in the Merton formula must to be interpreted as the relative risk aversion of the value function (see Alvarez et al., 2012, Online Appendix, p. 13). Additionally, if housing is risky and its returns are correlated with stock returns, the formula would also contain a hedging term which should be taken into account when retrieving the implied risk aversion from our empirical estimates. Empirically, Jordà et al. (2019) show that the correlation between housing and stock returns in Norway is essentially zero[29]. Hence we ignore the hedging term and infer that $\frac{\partial \alpha}{\partial r_e^n} = \frac{1}{\gamma(z)\sigma^2}$, with the understanding that $\gamma(z)$ captures the risk aversion of the value function which varies with the financial wealth/ housing wealth ratio, z (Grossman and Laroque (1990) and Alvarez et al. (2012)). The sensitivity of γ to z is particularly high around the time when consumers buy or sell houses. As discussed in section 5.2, we account for this dependence with appropriate indicators of housing purchases/sales, thus purging our estimates from the most relevant effect of variation in the financial/housing wealth ratio. We can back out γ by imposing $\sigma^2 = 0.04$, which is the variance of the Norwegian stock market index (Fagereng et al. 2017):

$$\gamma_1 = \frac{1}{0.04} \frac{\pi_{1992}}{7.35 - \alpha_{1992} \times 44.2}. \tag{15}$$

For the post-1998 period, the estimated cumulative marginal effect of the unconditional stock market share is 6.9 and the estimated $\frac{\partial \pi}{\partial r_e^n}$ from a *reduction* in the equity premium is 8.9 (Appendix Table A.1, last line). Following a similar procedure as above we obtain a second estimate of the risk aversion parameter as:

$$\gamma_2 = \frac{1}{0.04} \frac{\pi_{1997}}{6.9 - \alpha_{1997} \times 8.9} \tag{16}$$

Table 3 shows the two point estimates of the risk aversion parameter and their bootstrapped standard errors. The two values are 2.81 (SE = 1.13) when using the estimated reponses to the 1992 reform and 1.81 (SE = 0.09) when using the responses to the 1998 reform. Both estimates are statistically significant, but the latter is much more precise. Statistically, we cannot reject the null that the two values are equal (last column). In sum, our estimates of the cumulative response of the share either to an increase or a decrease of the equity premium are consistent with moderate levels of risk aversion.

Table 3: Implied Coefficient of Relative Risk Aversion

This table provides the coefficient of relative risk aversion implied by our empirical findings in the Merton (1969) portfolio model. See text for description. Standard errors are obtained from a 200-repetition bootstrap procedure.

	(1) Effect of 1992 Reform	(2) Effect of 1998 Reform	(3) Difference
Impled coefficient of relative risk aversion, γ	2.8123	1.8090	-1.0033
(bootstrapped SEs)	(1.1386)	(0.0914)	(1.1071)
Empirical estimates used	1993–1997, Cumulative	1997–2003, Cumulative	

Participation costs. Assuming a coefficient of relative risk aversion of 2.3 (the average of the two estimates in Table 3 above), we use the estimated sensitivity of participation to the

[29]The correlation coefficient is 0.025 over the period 1870-2019 and negative at -0.075 after WWII. In neither case they are statistically significant (p-values 0.52 and 0.77 respectively).

equity premium of 8.9 to infer the size of participation costs. In Appendix A.4, we illustrate the details of the exercise and show the sensitivity of the estimates to changing the risk aversion parameter (Table A.3). Our main estimates entail a per-period cost, ϕ^P, of NOK 625 or about \$89. This magnitude of the per-period cost is fairly modest and comparable to other estimates in the literature (see, e.g., Fagereng et al. 2017 who estimate per-period participation costs for Norway between \$65 and \$109 using a structural approach, Vissing-Jorgensen 2002 who finds a participation cost for the US of \$350, Paiella 2001 who documents participation costs in the \$70 to \$140 range).

Similarly, using the estimated $\frac{\partial \pi}{\partial r_e^n}$ of 44.20 (the cumulative response to a unitary increase in the equity premium following the 1992 reform, see Appendix Table A.1, first line) we find $\phi = 5,598$ NOK (\$800). This estimate includes both the one-time and the per-period cost. Deducting the previous estimate of per-period participation cost, we find an implied one-time entry cost of NOK 4,974 (\$710). This is about 1.7% of the average labor income in our sample, which is in line with the 2% entry cost estimated by Alan (2006) in a calibrated life-cycle model.

The calculations above rely on the interpretation that pre-trends are causally driven by the 1992 introduction of the wealth-tax risk premium. If this were not the case and we should adjust for this pre-trend, then the response to the 1998 drop in the risk premium is given by the gray dotted line in Figure 4. This shows a steep and large decline, with a drop in the stock market participation rate of 26 percentage points. This implies a sensitivity of participation to the risk premium of about 90, which is considerable. To rationalize a value for $\frac{\partial \pi}{\partial r_e^n}$ as large as this, one would need a per-period participation cost of about NOK 6,729 or \$897. This is too large to be credible (see, e.g., Choukhmane and de Silva 2021 who find little support for material *per-period* costs), which strengthens our interpretation of the trend in the share and participation before 1998 being the causal response to the 1992 tax reform. A further issue with using the detrended estimates is that they implicitly assume that households either did not respond at all to the initial increase or responded immediately in 1992 (prior to when our panel begins). We argue that this is inconsistent with the evidence presented in section 5.5.

6 Relating our estimates to models with adjustment frictions

Our empirical design speaks directly to recent models of equilibrium asset prices that highlight the importance of frictions for understanding stock price and exchange rate movements. For instance, the inelastic market hypothesis of Gabaix and Koijen (2021) is directly related to the sensitivity of stock portfolio shares to the equity premium and thus to our estimated responses. Gabaix and Koijen (2021) posit that the portfolio share in stocks α_i of some investor i is determined by a simple rule $\alpha_i = \theta_i e^{k_i \hat{\pi}}$ where θ_i is a baseline stock portfolio share, $\hat{\pi}$ is the deviation of the equity premium from its average, and k_i is the responsiveness of investor i to movements in the equity premium. The price elasticity of stock demand for investor i (using Gabaix and Koijen 2021 notation) is then $\zeta_i = 1 - \theta_i + k_i \delta$, which is directly increasing in the responsiveness of the investor's portfolio share to equity premium changes.

In Gabaix and Koijen (2021), investors are institutional investors, and low values of k_i are meant to reflect ubiquitous, binding mandates to stick to pre-determined portfolio allocation rules. Our evidence suggests that mandates may mirror the adjustment pattern of individual investors served by the funds rather than being an independent source of sluggishness in portfolio adjustments.

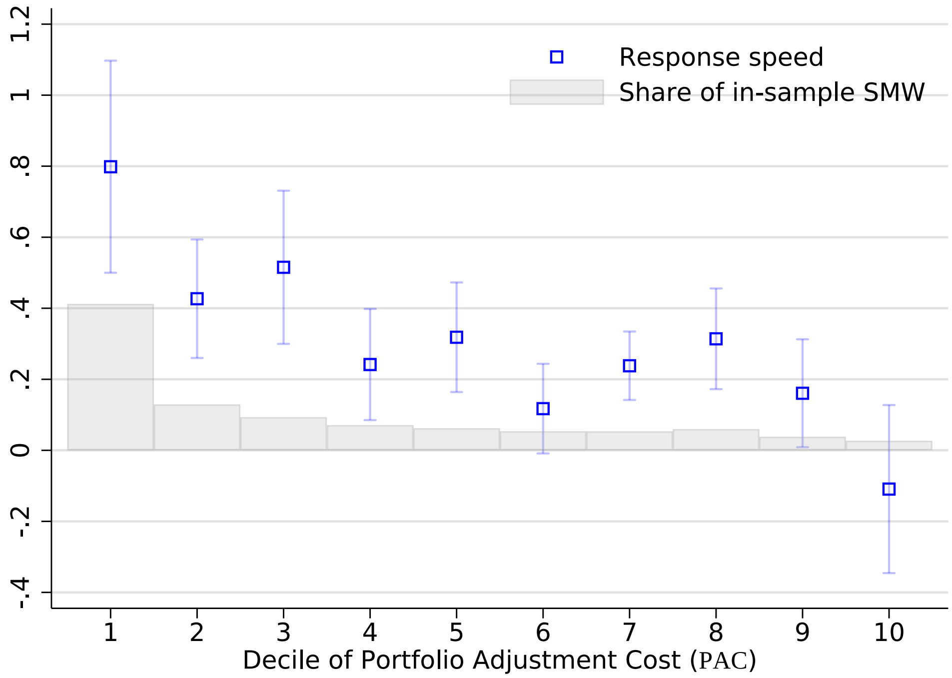
Gabaix and Koijen (2021) also show that broad market movements in stock prices depend on the weighted-average responsiveness, where the weights are the shares of the stock market owned by each investor. Hence, to understand market movements, one needs to study how adjustment speed covaries with stock market wealth. Accordingly, we use the Calvet et al. (2009) index of adjustment costs to split households into 10 deciles. For each decile, we calculate the share of in-sample stock market wealth owned by households in that decile. We then estimate our main specification in Table 2, column (1), allowing dynamic responses to vary at the predicted adjustment-cost decile. We use this to compute the speed at which households respond to the equity premium shock. We present our results in Figure 6, which has two key findings. First, the estimated response speed (blue squares) shows considerable heterogeneity driven by portfolio adjustment costs. Crucially, the response speed is significantly higher for households with lower portfolio adjustment costs (higher rebalancing strength according to the Calvet et al. 2009 index). In particular, for households in the lowest PAC decile, 80% of the total response to the equity premium change takes place within the first two years—compared to less than 20% among households in the highest five or so deciles of adjustment costs.

Second, it is visually clear that the speed of response to the equity premium shock is strongly correlated with stock market wealth. Households with lower adjustment costs hold a larger share of the total stock market (grey bars). Those in the bottom decile of portfolio adjustment costs hold about 40% of the stock market wealth and, as noticed, their response speed is quite high.

Our findings imply that the equal-weighted and share-weighted average response speeds will be quite different. Indeed, we calculate that the share-weighted observed response speed is 0.51, which is more than twice as much as the equal-weighted speed of 0.24 (which can be inferred from column 1 of Table 2). Importantly, however, this does not mean that frictions do not matter. For households in the highest 9 deciles of PAC, who own a combined 60% of the market, the response is quite slow. Most of these households' two-year responses are well below 40% of their long-run responses.

Figure 6: Heterogeneity in the Portfolio Response Speed to Equity Premium Shocks

This figure shows how stock market share (SMS) response speed (measured as the 1998–1999 effect divided by the 2002–2003 cumulative effect) varies across deciles of portfolio adjustment cost (PAC), defined in equation (12). The gray bars provide the share of total (in-sample) stock market wealth (SMW) held by each of the deciles. The average response speed weighted by the share of in-sample SMW is 0.51. The underlying specification does not adjust for pre-trends.



Overall, the results shown so far support the idea that households face significant portfolio adjustment frictions that considerably slow down their response to changes in the equity premium. Once households are allowed enough time to adjust their portfolio, the size of the response is as predicted by Mertonian portfolio models and in line with a relatively contained coefficient of relative risk aversion. But the convergence to the Mertonian benchmark takes years, even when the change in the equity premium is salient and long lasting, as is the case in our setting.

7 Discussion and conclusions

In this paper, we have presented novel evidence on how households respond to changes in the equity risk premium. We use long-lived and salient variation in the equity premium arising from wealth taxation, which side-steps important, attenuating measurement-error issues in the literature that relies on survey-based measures of stock returns beliefs. Our data allow us to show the dynamics of the responses over several years—a feature that is essential to identify the portfolio adjustment process in the presence of adjustment frictions. We document that investors do respond to long-lasting changes in the equity premium but the speed is slow. However, once investors are given enough time, their final response size is of the same order of magnitude as that of Merton-type frictionless models calibrated with reasonable levels of risk aversion. Our evidence lends strong support to the burgeoning asset pricing literature that invokes portfolio adjustment frictions as the cause for estimated inelastic stock market demand.

We can use our estimates to set a bound on the size of the monetary equivalent of the adjustment cost. A back of the envelope calculation suggests that for the average household, the monetary benefit from adjusting to the change in the equity premium in the same year of the 1998 reform, rather than spreading the response over the following 5 years, is \$3.2.[30] Hence, even very small adjustment costs in excess of this contained amount can cause responses to be very sluggish. Yet, the benefits from faster adjustment are not equal across investors. Importantly, we show evidence that the speed of adjustment is positively correlated with investors' ownership share of the stock market. As in Gabaix and Koijen (2021) this implies that stock market price elasticity depends heavily on the distribution of stocks in the population: if stock market ownership is highly concentrated, the market becomes more elastic and thus less volatile.

Our findings have implications for optimal capital taxation as well. It may be optimal to tax different types of assets at different rates (see, e.g., Scheuer 2013). For example, policymakers may wish to incentivize households to tilt their savings towards risky assets, perhaps to foster more entrepreneurship and growth. If so, the extent to which differential taxes have behavioral responses (i.e., there is portfolio reallocation) matters. If there are very small responses, which would be the case if households are highly risk averse, then lowering the tax rate on risky assets would only serve as a tax break for stock market participants and would not tilt savings toward riskier assets. Our findings show that the behavioral responses are sizable and in line with those implied by a reasonably parameterized model. In addition, our findings imply that one should treat with caution estimates of portfolio reallocation that rely on a short sample period, as they may understate true long-run behavioral responses. This is far from obvious in the context of public finance. While one may expect total savings to respond sluggishly, this sluggishness occurs entirely without frictions. Capital taxation affects the level and growth of consumption, which produces a sluggish effect on the stock of savings (Saez and Stantcheva, 2018). Portfolio composition, on the other hand, could (in principle) be immediately adjusted. Hence, it is not obvious that differential capital taxation will have sluggish effects.

While we provide evidence of the existence of adjustment frictions, we make no attempt to identify the nature of the friction. For example, these adjustment frictions could be driven by information gathering costs, by investor inattention, and behavioral frictions (e.g., procrastination or willingness to only use new savings to change the asset allocation), or be due to portfolio re-adjustment requiring costly effort, as in Gârleanu and Pedersen (2013). Identifying the nature of the friction is important to design policies aimed at reducing market volatility. We leave this task for future research.

[30] This is computed as $\sum_{t=1998}^{T=2003}(\alpha_T - \alpha_t) \times \Delta \times W_{1997}$ where α_T is the stock portfolio share prevailing in the last year of sample, Δ is the (average) wealth-tax risk premium, and W_{1997} is the average stock of financial wealth prior to the reform. Hence, it measures the total extra return the investor would earn by adjusting the share to its "long run" value α_T already in the first year of the tax reform. In a quadratic adjustment cost model with adjustment cost parameter θ the above measure is equal to $\sum_{t=1998}^{T=2003}\theta^{t-1997}|\alpha_T - \alpha_{1997}|$, increasing in the adjustment cost parameter θ.

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A Appendix

A.1 Additional Figures and Tables

Figure A.1: Quasi-placebo Effects when Assigning Treatment in 1995

This figure shows the effect of the reduction in the wealth-tax-induced risk premium on households initially above the wealth tax threshold in 1995 ($Treat_{i,1995} = 1$), rather than 1997 (as in the main specification), on their stock market share. The implicit placebo test is whether we observe a sharp reduction in the SMS during 1996 and 1997 preceding the actual risk-premium reduction that occurs in 1998. The orange line provides detrended estimates, in which the years 1994—1997 are omitted to estimate the time trend for the treated. The purpose of the exercise is to identify the (intent-to-treat) effect when using earlier (1995) treatment status as the main treatment variable while acknowledging that treatment does not go into effect until 1998. The gray line provides the raw differences between the treatment and control group. Differences are normalized to be zero in 1997 (while treatment is assigned as of 1995). Capped horizontal lines provide 95% confidence intervals, where the standard errors are clustered at the household level.

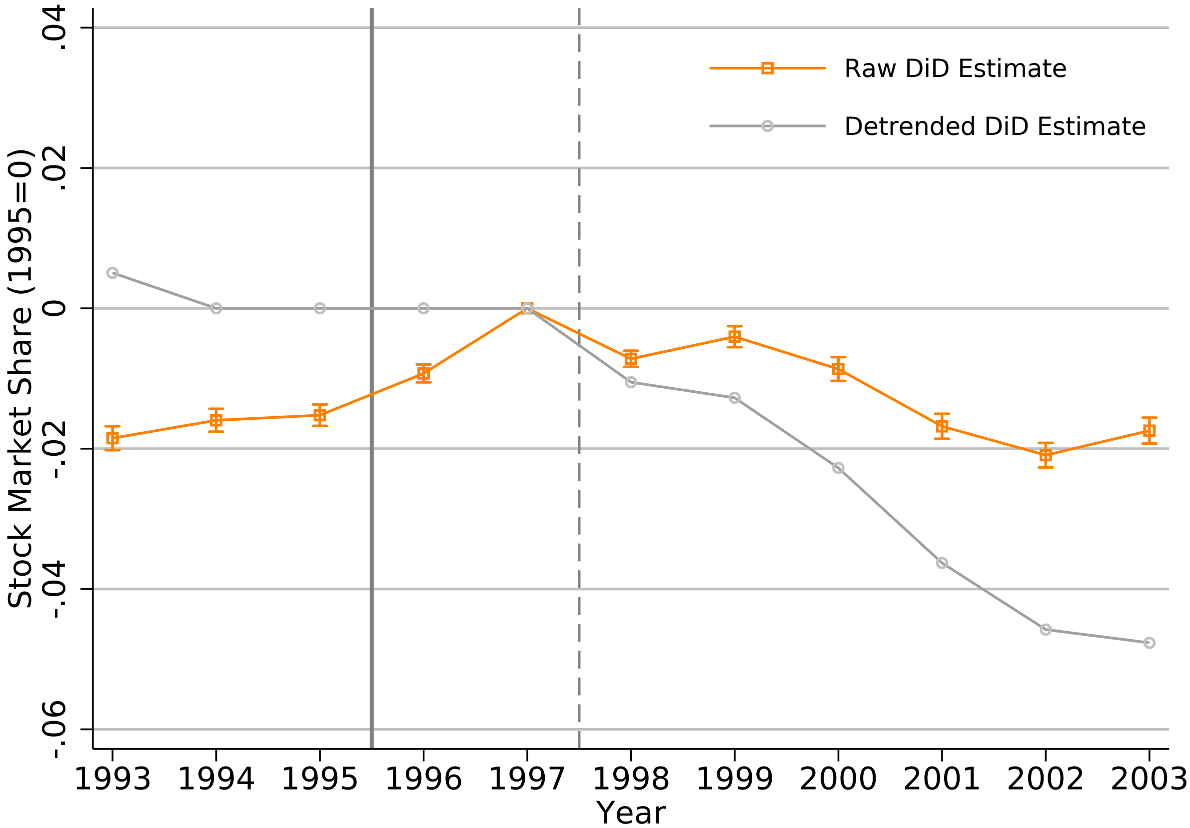


Figure A.2: Robustness to Isolating Active Portfolio Adjustment

This figure shows the effect of the reduction in the wealth-tax-induced risk premium on households initially above the wealth tax threshold in 1997. The solid orange line provides our baseline estimate. The short-dashed (triangles) and long-dashed (circles) lines remove changes driven by cumulative market returns since 1997. That is, $SMS^{active}_{i,t} = SMS_{\{i,t\}} - SMS_{i,1997} \times R^s_{1997,t}/[SMS_{i,1997} \times R^s_{1997,t} + (1 - SMS_{i,1997}) \times R^f_{1997,t}]$. $R^s_{1997,t}$is the cumulative return of holding the Norwegian stock market from 1997 to t. When $t < 1997$, we divide by the cumulative returns from t to 1997. We use two different measures of stock market returns. The short-dashed lines (triangles) assumes that dividends are reinvested (Source: OECD, FRED SPASTT01NOM661N). The long-dashed lines (circles) does not assume that dividends are reinvested (see, e.g., Calvet et al. 2009, source: S&P Global Equity Indices). $R_{1997,t}$ is a measure of the cumulative risk free return since 1997, proxied for by the average bank deposits rate per Statistics Norway's statistics.

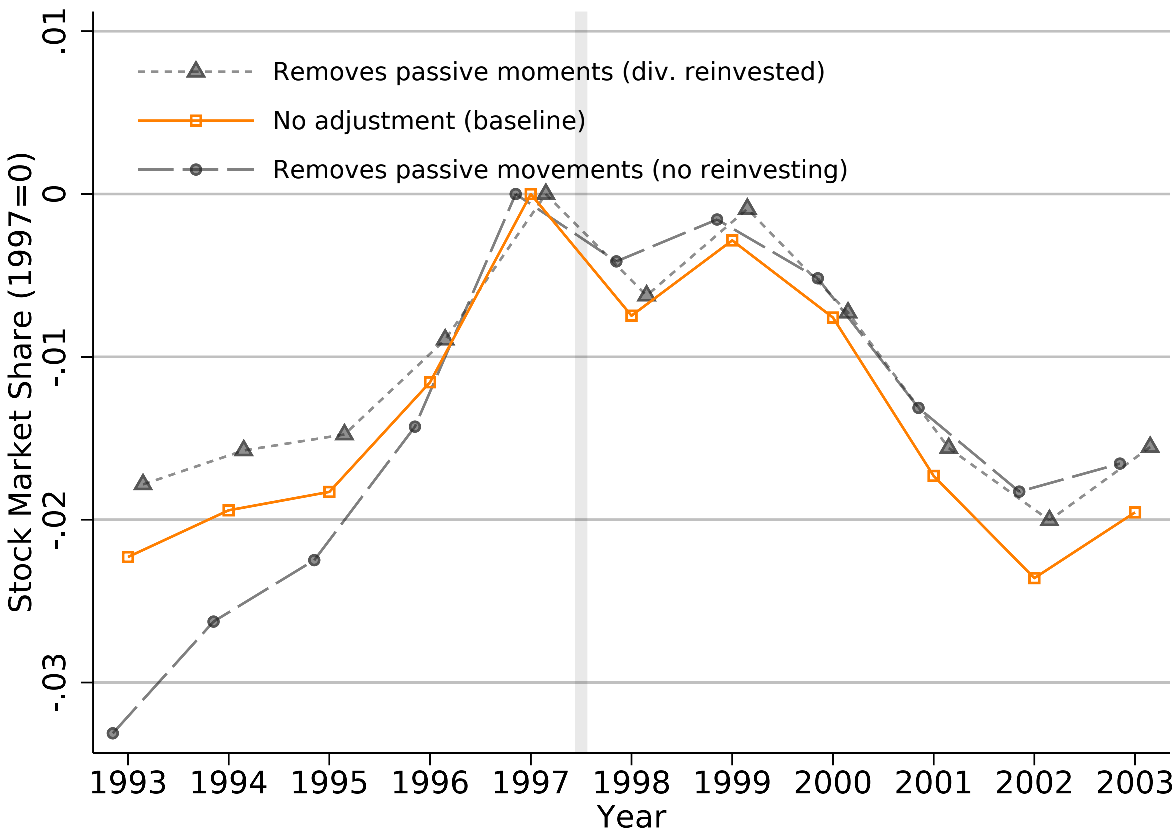


Figure A.3: Robustness to Controlling for Timing of Housing Transactions

This figure shows the effect of the reduction in the wealth-tax-induced risk premium on households initially above the wealth tax threshold in 1997. The solid orange line provides our baseline estimate. The short-dashed (triangles) line shows the estimated coefficient once we flexibly control for changes to homeownership and the relative timing of transactions real estate markets. The homeownership control is simply a dummy for whether the household at time t owns a house according to their tax returns. For transactions, we construct the variable NextTransactionYear which takes a value of zero if there is no transaction between $t$and 2004. If there is a transaction, it takes the value of the year closest to t. Hence, if a household transacts in 2002, this NextTransactionYear takes a value of 2002 for all the preceding years in our sample. We then take out fixed effects at the $t \times NextTransYear$ level. This implies that in the year 2000, for example, we estimate a fixed effects for each of the transaction years 2000, 2001, 2002, 2003, and 2004, and also a fixed effect for households that do not transact during 2000–2004. We also include a variable PreviousTransactionYear, that is similarly constructrued but only accounts for transactions occurring strictly before t.

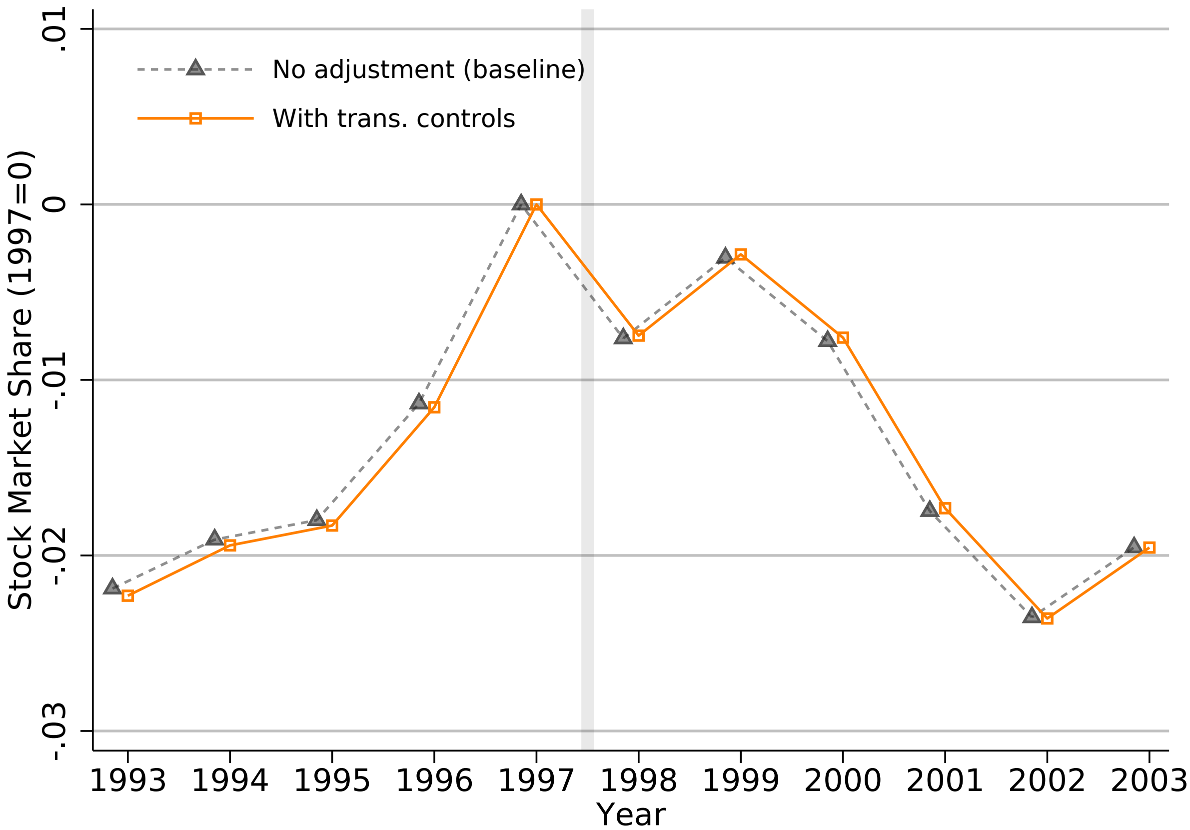


Figure A.4: Robustness to Controlling for Ex-Ante Stock Market Exposure

This figure shows the reduced-form effect of the reduction in the wealth-tax-induced risk premium on households initially above the wealth tax threshold ($Treat_{i,1997} = 1$) on their stock market share. The underlying specification drops 1993 observations but includes the household stock market share as of 1993 as a control variable: the orange line with hollow squares provides non-detrended estimates using equation 10, but with the additional term $\zeta_t SMS_{i,1993}$, and is only estimated for $t > 1993$. This specification is meant to address differential, possibly non-linear trends based on ex-ante stock market exposure. Capped horizontal lines provide 95% confidence intervals for the non-detrended estimates, where the standard errors are clustered at the household level.

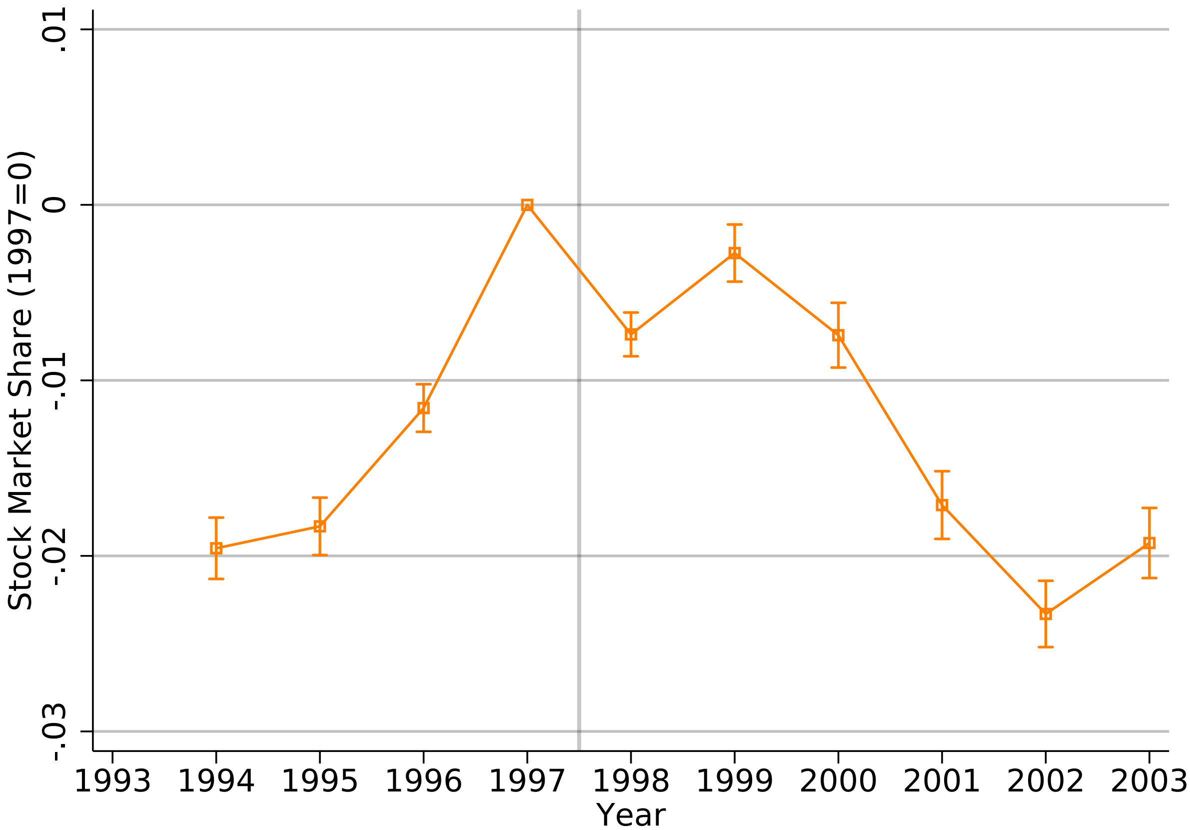


Table A.1: The Effect on Stock Market Participation

Notes: This table shows the estimated effect of a change in the risk premium on stock market participation, estimated using equation 11, without adjusting for pre-trends. Standard errors are clustered at the household level. Stars indicate significance at the 10%, 5%, and 1% levels.

(A) Cumulative Effects

	1992-reform	1998-reform
	(1)	(2)
Cumulative Effect	44.20***	9.16***
	(0.75)	(0.76)
Period	1993–1997	1998–2003
Implied by estimate	$\hat{\varepsilon}_{1993}$	$\hat{\varepsilon}_{2003}$

(B) Underlying Regression Estimates

	(1)
$\hat{\varepsilon}_{1993}$	44.20***
	(0.75)
$\hat{\varepsilon}_{1994}$	39.14***
	(0.71)
$\hat{\varepsilon}_{1995}$	29.74***
	(0.53)
$\hat{\varepsilon}_{1996}$	13.83***
	0.53
$\hat{\varepsilon}_{1997}$	–
	(.)
$\hat{\varepsilon}_{1998}$	-1.19**
	(0.44)
$\hat{\varepsilon}_{1999}$	-2.49***
	(0.57)
$\hat{\varepsilon}_{2000}$	0.20
	(0.64)
$\hat{\varepsilon}_{2001}$	3.50***
	(0.69)
$\hat{\varepsilon}_{2002}$	7.61***
	(0.74)
$\hat{\varepsilon}_{2003}$	9.16***
	(0.76)
N	1,883,765

Figure A.5: Effect on Stock Market Entry and Exit

This figure provides the estimated yearly reduced-form effects on stock market entry and exit. Entry is defined as $max(SMP_t - SMP_{t-1}, 0)$ and the estimates are provided by the hollow squares (solid line). Exit is defined as $max(-[SMP_t - SMP_{t-1}], 0)$ and the estimates are provided by the hollow circles (dashed line). The two separate regressions are performed by estimating equation 10.

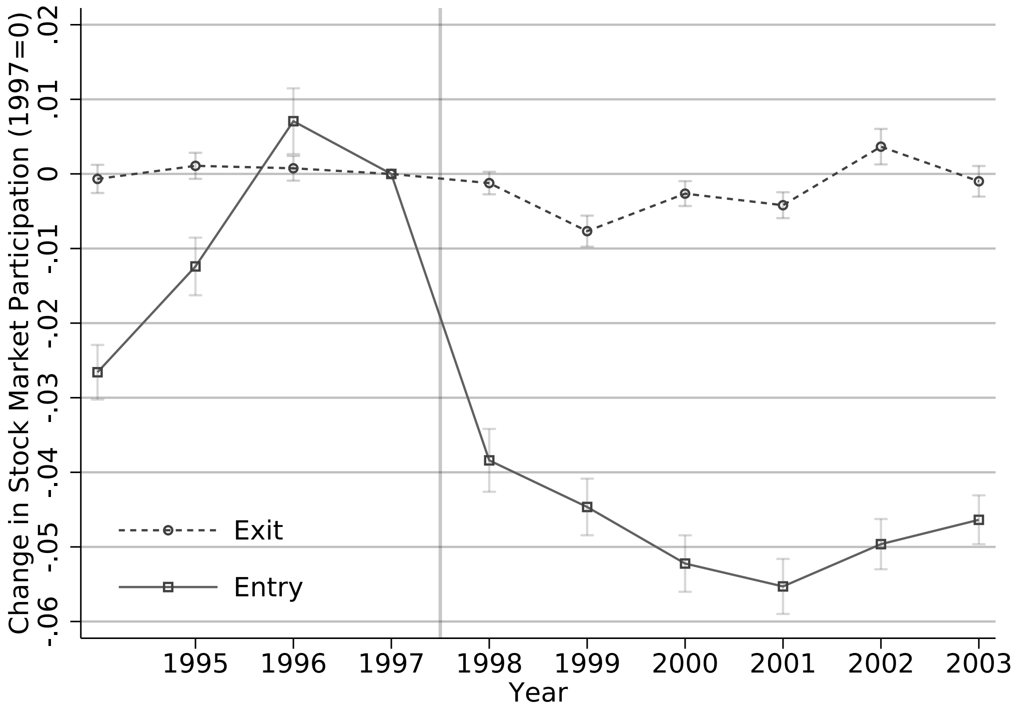


Figure A.6: Heterogeneity-robust estimation

This figure provides the estimated yearly reduced-form effects the stock market share using the estimator by De Chaisemartin and d'Haultfoeuille (2020).

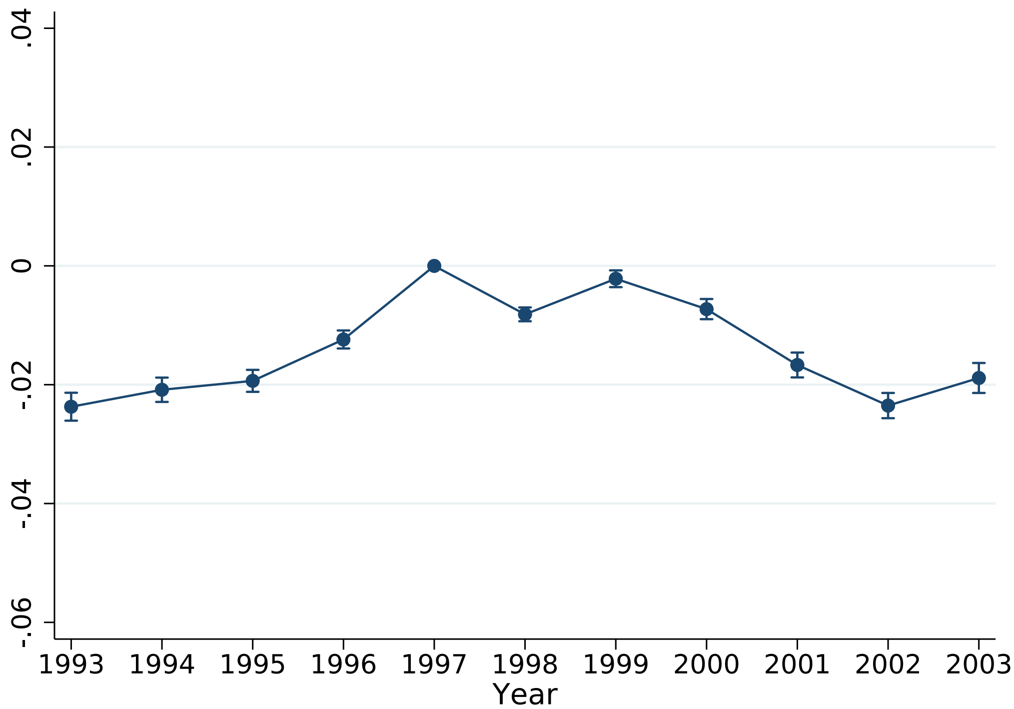


Table A.2: Summary Statistics by Wealth Tax Bins

This table provides summary statistics for the main variables used in the analyses. Taxable Net Wealth (TNW), Financial Wealth (FW), and Labor Income are provided in NOK. USD/NOK exchange rate was approximately 7.5 in Dec 1997.

	Non-payers	Above First Threshold	Above Second	Above Third
	Mean	Mean	Mean	Mean
TNW	-32,691	187,252	363,122	618,852
FW	38,437	59,259	138,950	300,541
SMS	0.061	0.059	0.083	0.091
Mutual funds/GFW	0.041	0.043	0.064	0.068
DH Stocks/GFW	0.020	0.016	0.019	0.023
SMP	0.21	0.23	0.38	0.51
Age	45.69	57.71	57.25	57.55
College	0.25	0.14	0.20	0.26
Biz/Law/Econ	0.03	0.01	0.02	0.03
Labor Income	299,190	229,492	284,661	326,644
Observations	87,866	19,229	50,473	13,890

A.2 Media Coverage of Wealth Tax Reform

Searches in Norwegian newspaper archives provide an abundance of articles about the implications of the wealth tax reform. Below we translate excerpts from some of these articles, choosing an array of newspapers to underline the breadth of coverage (different local as well as national outlets.[31])

- Finnmark Dagblad (northern-most county's regional newspaper), 1997/07/16, writes that "Minister of Finance Jens Stoltenberg will change the tax system. In addition to lower wealth taxes and higher thresholds, he warns about a reduction in the stock discount. This means that he wishes to make it less profitable to invest in stocks rather than to save in, for example, a bank. Today, the stock discount works such that if you place your savings in public equity, you get a discount of 25 percent. This means that you are only taxed on 75% of the real value [...]"

- Aftenposten (large, old national newspaper), 1997/10/14, writes that the new national budget is both "sweet and sour for the households." "Wealth: The government reduces the wealth tax for persons as part of a larger tax package where the counter-balance is a reduction in the so-called stock discount (see page 6). But for ordinary folk without stocks and with wealth, the tax cut implies a net gain."

- Lofotposten (Northern-central regional newspaper), 1997/10/14, writes that "tobacco more expensive, but less taxes," and follow by explaining that "stocks in listed companies shall no longer be valued at 75 percent of the market value." They explain that "With these changes, the owning of stocks and saving in banks become equals."

- Bergens Tidende (Regional newspaper based in 2nd largest city, Bergen), 1997/10/14, describe the new national budget and how it will affect the large stock owners: "According to the new proposal [for the national budget], the stock discount shall e removed, which means that the stocks' full value shall be used at taxation. As a counter-act, the highest marginal wealth tax rate is reduced from 1.5 to 1.1 percent [...]"

- Dagens Næringsliv (large national business-focused newspaper), 1997/11/18, write in their headline that it will become "more expensive for employers [...]. They elaborate and say that "the so-called stock discount for wealth taxation of listed stocks is being removed [...]

- Dagsavisen (national labor-movement newspaper), 1997/12/10, write in their outline of the new tax policies that "In the [coalition government] budget, the tax discount on listed stocks was [...] removed. It was earlier at 25 percent. [...] At the same time, the tax rate was lowered by 1.5 to 1.1 percent.

- Nordlys (northern-Norway regional newspaper), 1998/08/01, write that the political parties are preparing for a tax-treasure hunt. "The stock discount will become this fall's political trend word, so you might as well learn its contents [...] Last year, the labor party government

[31] Original articles, written in Norwegian, are available through either the national public library, wow.nb.no, or the newspaper archive subscription service *atekst*.

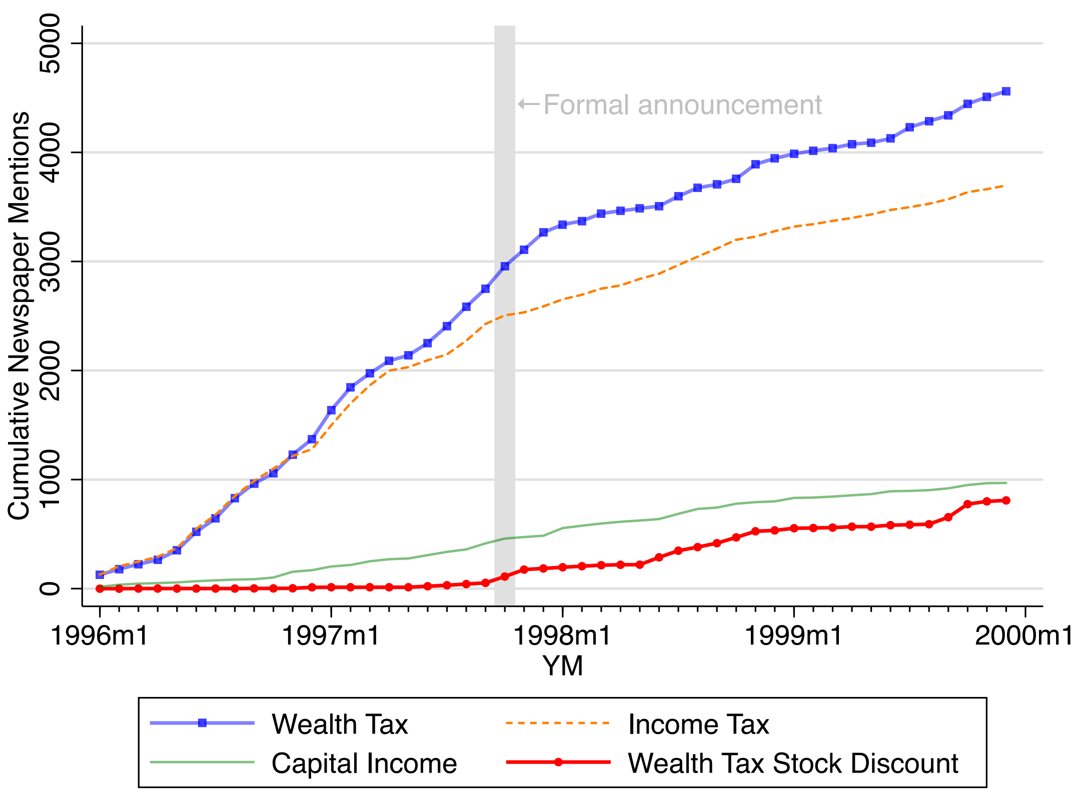
had had enough. the rich had gotten rich enough, they better pay a wealth tax on their stocks [..]"

We also examine the media coverage more formally. Figure A.7 provides the cumulative number of times different tax-related keywords appear in newspapers. we rely on the digitized archives provided by the Norwegian National Library that covers virtually all Norwegian newspapers during 1996-1999. We see that for most of 1996, "income tax" and "wealth tax" were mentioned almost an identical number of times. Starting in early 1997, and accelerating around the time the new tax policy was formally announced in October 1997, the cumulative mentions diverge—with "wealth tax" taking the lead. We also see that there are few changes during 1997 in how often capital income is mentioned, but that the terminology "stock discount", which refers to the fact that stocks are tax-favored relative to safe assets, starts appearing more regularly as of October 1997.

This exercise underlines the salience of the tax reform. Even prior to the reform, wealth taxation is mentioned as often as income taxation, and following the reform wealth taxation is mentioned considerably more.

Figure A.7: Media Coverage of Tax Reform

This figure shows the cumulative mentioning (starting Jan 1st 1996) of different tax-related keywords. The blue line with square symbols provides the cumulative number of times the word "formuesskatt" (translated: wealth tax) has appeared. The counting procedure relies on the National Library's methodology: if the keyword appears multiple times in a given newspaper issue (e.g., the Monday issue), then it is only counted once. The orange dashed line provides the cumulative number of times "inntektsskatt*" (translated: income tax) is mentioned. The asterix allows for definitive and plural variations. The solid green line considers "kapitalinntekt" (translated:capital income). The circle-connected red line considers the mentioning of "aksjerabatt*" (translated: [wealth tax] stock discount). The underlying data is publicly available at www.nb.no, and was accessed on 12/4/2021.



A.3 Changes in the average versus the marginal equity premium

The Merton model provides predictions regarding the marginal effect on portfolio allocation of a marginal change in the equity risk premium. The ideal experiment is perhaps that the risk-free rate decreases exogenously for a random subset of investors who then face a higher equity risk premium. In such an experiment, we would expect the full substitution effect from risk-free to risky assets as predicted by the Merton model.

In our setting however, this effect could be muted. This is because the wealth tax scheme during 1992–1997 had differential effects on the marginal and average after-tax equity premium faced by investors. Consider first the 1992 introduction of the wealth tax risk premium. This government introduced a 25% tax assessment discount on stocks. This assesment discount implied that, for wealth tax payes, 1 NOK of stocks had a 25% lower wealth tax effect than 1 NOK of safe assets, which is what caused the wealth tax risk premium. However, as investors respond to this wealth-tax induced equity premium by substituting away from deposits and bonds towards stocks, their taxable net wealth decreases. At some point, the investor may have reduced their stock holdings enough to bring their taxable wealth below the wealth tax threshold. Once taxable net wealth is below the threshold, further substitution away from safe assets toward risky assets would not lower the wealth tax bill, and, hence, there is no longer an incentive to substitute. This means that in 1997, investors might not have as high of a risky share as what the Merton model would predict given the after-tax equity premium. Accordingly, once the wealth-tax risk premium is removed in 1998, households might also not respond fully as subscribed by the Merton Model: while their final risky share, observed in 2003, would be the one implied by the Merton model, the baseline equity share in 1997 may have been "too low," especially if these households had very low risk aversion. Importantly, however, the households who would have had a muted optimal share as of 1997 should have taxable net wealth equal to the 1997 wealth tax threshold. They would not be located well above the threshold, because then they would still, on the margin, have faced a positive wealth tax risk premium. While it is of course impossible to locate exactly at this threshold through portfolio allocation responses (e.g., uncertainty about income flows and realized returns), we would expect the muted investors to be located near or around the tax threshold.

To conclude, the fact that the wealth tax differentially affected the marginal and average (after-tax) equity premium will potentially cause a downward bias in the *magnitude* of the long-run portfolio response. It should not affect the dynamics. Hence, our qualitative finding that households have sizable portfolio responses when given enough time is, if anything, strengthened by the fact that some households may have had muted portfolio shares as of 1997.

A.4 Estimating implied entry and exits stock market participation costs

Recall from section 2.2 that $\frac{d\pi(\bar{w})}{dr_e^n} = \frac{\phi 2G'(\bar{w})\gamma\sigma^2}{(r_e^n)^3}$ for a first-time entrant and $\frac{d\pi(\bar{w})}{dr_e^n} = \frac{\phi^P 2G'(\bar{w})\gamma\sigma^2}{(r_e^n)^3}$ for an investor who exits or one that re-enters or chooses to continue to stay in the market. Both expressions are measured when the change in the equity premium occurs. We do three exercises. First, we use our estimates of the effect of a unitary increase and a unitary decrease in the equity premium to obtain estimates of the investors risk relative risk aversion. We have shown in section 5.6 how this can be done obtaining two point estimates of relative risk aversion:1.81 when the variation in the equity premium due to the 1998 reform and 2.81 when using the estimated responses to the 1992 introduction of a tax induced equity premium. To recover the risk aversion parameter, we use the stock portfolio share conditional on participation and the participation rate for the treated before the reform,in 1997 and in 1992. We observe the 1997 values but not the 1992 ones as data start in 1993. Hence we use the values for the conditional share and the participation rate in 1993 (0.142 and 0.145 respectively), but estimate the 1992 participation rate to be 0.12 by subtracting from 0.145 the average additional increase in participation among the treated in the first two years following the reform to account for some effect in between 1992 and 1993.

Second, we use the above expressions for $\frac{d\pi(\bar{w})}{dr_e^n}$ to get a sense of the size of the effect of a unit change in the equity premium on participation, as implied by the model. Third, we use the estimates of $\frac{d\pi(\bar{w})}{dr_e^n}$ together with their analytical expressions to obtain an estimate of the implied total entry cost, ϕ, (the sum of the one-time and the per-period costs) and the per-period cost, ϕ^p. Namely, we exploit the estimated cumulative response to the drop in the equity premium starting in 1998 to pin down the per-period cost as $\phi^P = \frac{(r^e)^3}{2G'_{1997}(\bar{w})\gamma\sigma^2}\frac{d\pi(\bar{w})}{dr_e^n}$, where $\frac{d\pi(\bar{w})}{dr_e^n} = 9.1$ is our estimated cumulative participation response to the 1998 reversal of the tax equity premium. Similarly, we use the estimated cumulative response to the increase in the equity premium following the 1992 reform to estimate $\phi = \frac{(r_e)^3}{2G'_{1992}(\bar{w})\gamma\sigma^2}\frac{d\pi(\bar{w})}{dr_e^n}$, where $\frac{d\pi(\bar{w})}{dr_e^n} = 44.20$, and $G'(\bar{w})$ is computed at the time of the reform.

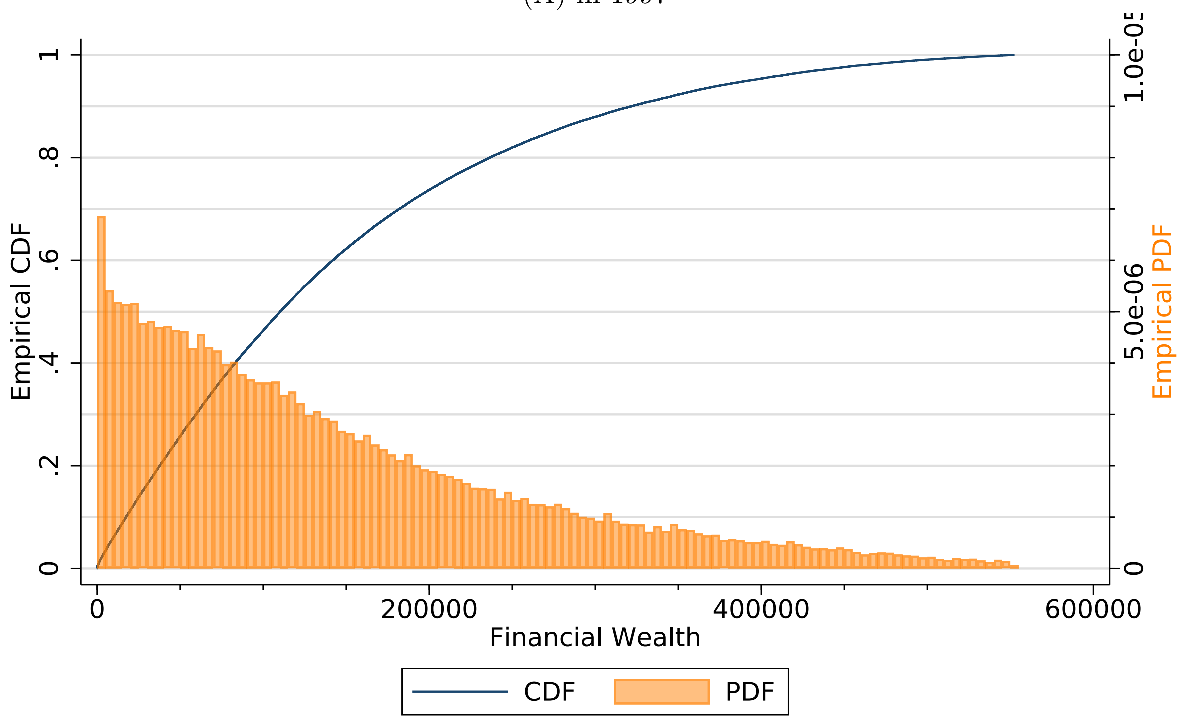
For the last two exercises, we need to make assumptions about $G(w)$. Rather than assuming a specific distribution, we approximate $G(w)$ with the empirical distribution of financial wealth. Figure A.8 plots both the CDF and the probability density (PDF) of financial wealth for households affected by the tax reforms in 1992 and 1997.

In the data, the average stock market participation rate in 1997 is 37% among the treated households. Using the simple participation model, the wealth cut-off, $\bar{w}$, should be such that $0.37 = 1 - G_{1997}(\bar{w})$; thus, $\bar{w}$ corresponds to the value of wealth where the empirical $CDF = 0.63$ in 1997. At this value of wealth, we find that the empirical PDF is about 2.6×10^{-6}. This is our approximation for $G'_{1997}(\bar{w})$ when calibrating the marginal effect of a drop in the equity premium or estimating per-period participation costs after the drop in the premium in 1998. We use the same procedure to compute the empirical PDF in 1993— the first year for which we have the data after the 1992 reform—when the participation rate was 14.5% for wealth-tax payers. In this case our approximation for $G'_{1992}(\bar{w})$ is 1.4×10^{-6}, which we use when estimating entry costs in response to the 1992 reform or calibrating the marginal effect of an increase in the equity premium.

Figure A.8: Financial Wealth Distribution

The bars in this figure show the empirical probability distribution function (PDF) by NOK 10,000 bins of Financial Wealth (FW) as for the treatment group. The corresponding y-axis is on the right hand side. The concave line (left-hand-side axis) shows the empirical cumulative distribution function (CDF). For the purpose of this figure, FW is trimmed at its 99th percentile. Panel (A) shows the distribution of financial wealth in 1997, and Panel (B) shows the distribution as of 1993. The stock market participation among the treated was 37% in 1997 and 15% in 1993. For 1997, the $\bar{w}$for which $1 - G(\bar{w}) = 37\%$ gives $G'(\bar{w}) = 2.672 \times 10^{-6}$, the latter is obtained from the empirical PDF. For 1993, associated with $1 - G(\bar{w}) = 15\%$, is a $\bar{w}$ of about 235,000, with a corresponding $G'(\bar{w}) = 1.443 \times 10^{-6}$.

(A) in 1997



(B) in 1993

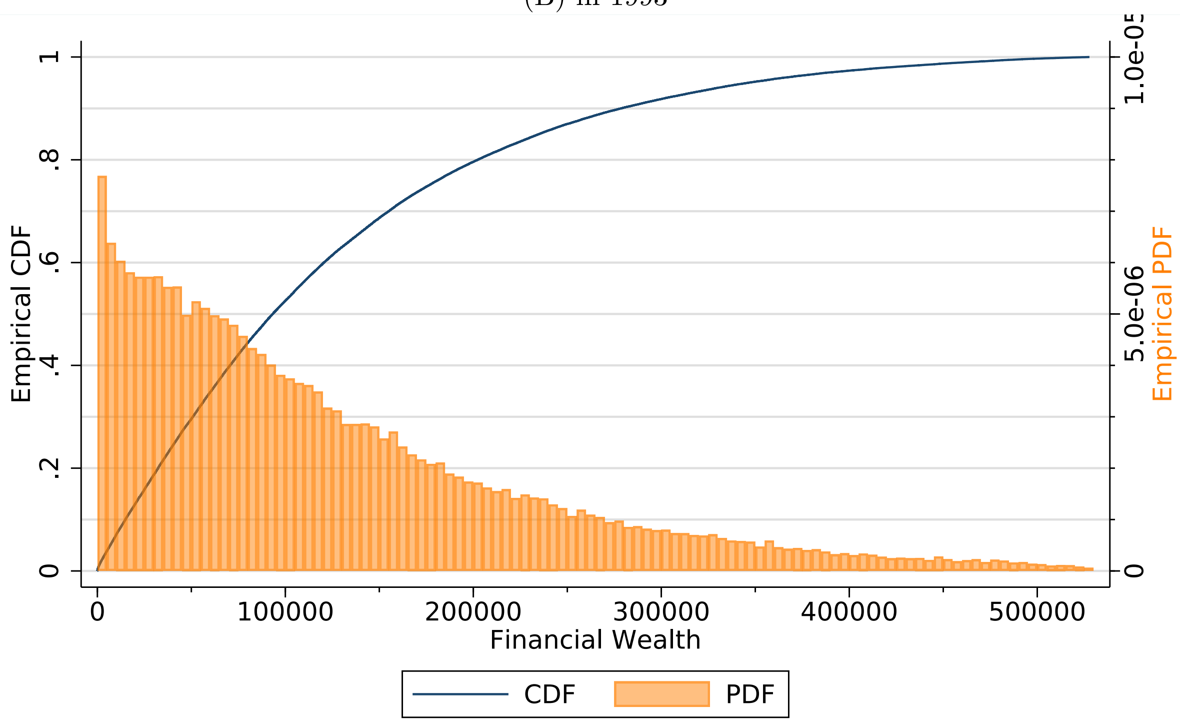


Table A.3: Calibrated Marginal Effects of Equity Premium Changes and Estimated Participation Cost

Notes: See text for description of the methodology. This table uses a USD/NOK conversion rate of 7.

Risk aversion	Calibrated marginal effects: $\frac{d\pi(\bar{w})}{dr^e}$		Implied participation cost (USD)	
	1992 reform (1)	1998 reform (2)	one-time entry, ϕ (3)	per-period, ϕ^P (4)
3	57.4	14.3	615.8	68.7
2.31	44.2	10.99	799.7	89.2
2.81	53.8	13.4	657.4	73.6
1.81	34.6	8.6	1020.7	113.9

The first two columns of Table A.3 show the calibrated marginal effects of the 1992 increase in the equity premium and the 1998 reversal on participation for various levels of risk aversion, including the two risk aversion parameters implied by the empirical effect on the portfolio share, as well as the average of the two parameters of 2.31. For this calculation, we need an assumption on the size of the participation costs ϕ and ϕ^P. Assume that the one-time fixed participation is around 2% of permanent income (as estimated by Alan 2006 in a calibrated life cycle model). In our context, this amounts to about 5,600 NOK ($800), using the average labor income of wealth-tax payers (NOK 280,000, Table 1 in the text) as a proxy for permanent income. As for the per-period cost, an approximate value consistent with existing estimates is around $100 (NOK 700), which is in the ballpark of the values of per-period costs estimated by Fagereng et al. (2017) for Norway using a structural estimation approach or Vissing-Jorgensen (2002) on US data who finds a participation cost of $350 and Paiella (2001) documenting participation costs in the $70 to $140 range. With these values, and further assuming a baseline equity premium of 3.2% and a variance of stock returns of 0.04 (as in Fagereng et al., 2017), the marginal effect on participation of a unit increase in the equity premium implied by the model is 44.2 for a risk aversion of 2.3; and 53.8 for a risk aversion of 2.81; and 34.6 for risk aversion of 1.81. The calibrated marginal effect of a unit decrease in the equity premium around is 8.6 when risk aversion is 1.81 and 13.4 when risk aversion is 2.81.

Columns (3) and (4) show the estimated total one-time entry and per-period costs, ϕ and ϕ^P, implied by our estimates of the marginal effects, again for different levels of risk aversion. For risk aversion of 2.31, we estimate the entry cost at NOK 5,598 ($799.7) and the per-period participation cost at NOK 624.7 ($89.2). For risk aversion of 2.81 the two figures are respectively $657.4 and $73.4; and for risk aversion of 1.81 they are $1020.7 and $113.9, respectively. These values are all in the ballpark of existing estimates.