

WEALTH TAXATION AND CHARITABLE GIVING*

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Abstract

We study how tax incentives affect giving using two quasi-experiments from Norway. First, using a shock to wealth tax exposure, we estimate the semi-elasticity of giving with respect to the after-tax rate of return. Inconsistent with the notion that households accelerate giving to reduce future taxes, we find that a 1% wealth tax reduces giving by 26%. We also show that wealth taxation reduces the likelihood of households starting to give. There is no effect on exits from charitable giving, which implies that entry costs (but not participation costs) are important for understanding giving behavior. Second, using bunching at an income-tax deduction threshold, we estimate a modest compensated own-price elasticity of giving of -0.44. This elasticity exhibits only minor heterogeneity with respect to income and wealth, but is considerably larger for religious than nonreligious giving. We develop a simple life-cycle model with endogenous charitable giving to interpret our combined findings. The calibrated model exhibits intertemporally inelastic consumption preferences with an EIS of only 0.08. The calibrated preference parameters imply that crowd-out effects of capital taxation on charitable giving may be sizable in settings where the level of giving is high.

Keywords: Charitable giving, wealth taxation, capital taxation, intertemporal substitution

JEL codes: H24, H31, H41, D64

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1 Introduction

The use of tax incentives to promote charitable giving is ubiquitous. These incentives typically take the form of an income tax deduction that lowers the after-tax marginal price of giving (Saez 2004; Diamond 2006; List 2011). Such policies are in place in nearly all OECD countries (OECD, 2020) and have produced fertile grounds for a large empirical literature. The main focus of this literature has been to estimate the after-tax own-price elasticity of giving that is needed for determining the optimal tax incentives such as the degree of tax deductibility (Saez, 2004). The indirect effects of other types of taxation, such as those on household savings, have been neglected. This is despite a surging interest in reducing wealth inequality through policies such as more comprehensive capital taxation (Bastani and Waldenström 2020b; Scheuer and Slemrod 2021; Saez and Zucman 2019a). Importantly, these policies may also curb households' willingness to voluntarily redistribute their wealth through charitable giving. Yet, these cross effects have seen little empirical attention nor played a role in the growing optimal capital taxation literature (see, e.g., Saez and Stantcheva 2018; Straub and Werning 2020; Rotberg and Steinberg 2021; Broer et al. 2021; Guvenen et al. 2021; Gaillard and Wangner 2021; Boar and Knowles 2022).

This paucity of empirical evidence is problematic as there are clearly defined but theoretically ambiguous links between capital taxation and household giving behavior. Capital taxation in the form of a wealth, capital income, or capital gains tax reduces the after-tax return on savings that households can achieve. This renders current consumption as well as charitable giving relatively more attractive than saving for the future, which causes households to give more through an intertemporal substitution effect. This substitution effect behaves as a pseudo avoidance strategy in which households give more today in order to reduce life-time taxes. Working in the opposite direction is the income effect. A tax on savings reduces after-tax wealth and disposable income, causing households to choose to give less. In sum, whether the giving channel accelerates the redistributive potential of capital taxation or represses it is theoretically ambiguous. Whether these linkages between giving and capital taxation should be a first-order concern in optimal taxation is an open, unexplored question.

Empirically studying the effect of capital taxation on giving is challenging due to (i) a scarcity of identifying variation in the after-tax return on savings that is both exogenous and plausibly uncorrelated with other determinants of charitable behavior; and (ii) limited data on household giving. In many settings, charitable giving is self-reported, which leaves it unclear whether one is observing changes to reporting or actual giving behavior (Tazhitdinova, 2018), which is a distinction that generally matters for welfare analyses (Chetty, 2009). We overcome these challenges by exploiting quasi-experimental variation in the annual taxation of net wealth combined with comprehensive third-party-reported data on charitable giving.

We present new empirical evidence on capital taxation and charitable giving on two fronts

and tie the main results together in a simple life-cycle model that incorporates charitable giving. *First*, we provide novel empirical evidence on how capital taxation affects charitable giving. We obtain identifying variation in both the average and marginal after-tax rates of return on savings from substantial changes to the wealth tax treatment of housing wealth. Starting in 2013, the tax authorities began removing the preferential treatment of secondary housing wealth, while leaving it in place for primary homes. This meant that secondary home owners saw a large accounting increase in their taxable wealth, which increased their probability of paying a wealth tax and more than doubled their average wealth tax bill. This extensive-margin shock lowers marginal after-tax rates of returns on wealth and the intensive-margin shock lowers the average after-tax return on wealth. The nature of the reform allows us to control for overall estimated housing wealth while obtaining identification from pre-reform portfolio allocation into secondary versus primary housing wealth. We use this variation in a difference-in-differences (DiD) instrumental variables (IV) framework.

For many years prior to the reform, treated and control households were on identical trajectories in terms of their giving behavior. However, as soon as the reform occurs in 2013, we find a sharp reduction in the giving of secondary home owners. Using our IV specification, we find that a one percentage point increase in the tax rate on wealth decreases giving by 26%.

Theoretically, our negative giving estimate is the combination of two forces: A negative income effect of paying more in wealth taxes on the intensive margin and a positive intertemporal substitution effect from extensive-margin wealth taxation increasing the relative price of future giving. We empirically disentangle these two forces by exploiting the first-stage heterogeneity caused by progressive taxation (as in [Gruber and Saez 2002](#)). We find that the negative effect of wealth taxation on giving is, in accordance with theory, driven by households paying more in wealth taxes on the intensive margin. Interestingly, we find no evidence of offsetting positive effects of paying a wealth tax on the extensive margin. Inconsistent with intertemporal substitution effects being important, our point estimate is economically and statistically close to zero. In other words, the pseudo avoidance strategy of giving more now to reduce life-time wealth taxes is not present in our data.

We further examine whether capital taxation affects *whether* households give. We show that the average treated households becomes about 0.8 percentage points (2.4%) less likely to give by. This finding indicates the presence of fixed costs of giving. By reducing the optimal giving amount, fewer households find it worthwhile to give. Interestingly, we find that the participation effect is entirely driven by reduced entry into giving. This indicates that fixed costs of giving consist entirely of a one-time entry cost as opposed to per-period participation costs.

Second, we use a bunching framework to estimate the elasticity of charitable giving with respect to its after-tax price. This elasticity, although nominally unrelated to wealth taxation, provides a crucial empirical moment for calibrating a structural model of giving. In Norway, charitable giving is deductible in the income tax base, up to a fixed deduction limit, and charitable

organizations report giving amounts directly to the tax authorities. Importantly, the presence of a deduction limit creates a large discontinuity in the marginal after-tax price of giving, which allows us to make novel use of a bunching framework to infer the compensated after-tax own-price elasticity of giving. While there is clear evidence that households bunch at the deduction limits, the implied elasticity is economically modest at about -0.44.

Furthermore, we investigate whether it is reasonable to assume that the compensated own-price elasticity is a constant parameter. By using regression-based techniques to uncover bunching heterogeneity (Bastani and Waldenström, 2020a), we find that the magnitude of the elasticity is decreasing in income and wealth, but that this heterogeneity is economically small. We further find no evidence that wealth taxation causally affects the compensated own-price elasticity. Together, these findings support the common assumption of a constant compensated elasticity.

While our main analyses consider charitable giving as a uniform good, we exploit the granularity of our data to study differential effects of tax incentives based on the type of giving. In terms of the responsiveness to wealth taxation, we find both religious and nonreligious giving to be highly responsive. Among nonreligious giving, the effects are largely explained by a reduction in giving to organizations focused on international issues such as humanitarian aid, which highlights potential unintended far-reaching spillover effects of capital taxation. In terms of the responsiveness of giving to its after-tax own price, we find that religious giving is ten times as elastic as nonreligious giving.

Finally, we combine our empirical findings to inform a life-cycle model that endogenizes charitable giving. We assume that households' per-period utility is given by

$$\frac{c_t^{1-1/\sigma}}{1-1/\sigma} + \kappa \frac{g_t^{1-1/\varepsilon}}{1-1/\varepsilon}, \quad (1)$$

where c_t is consumption, σ is the elasticity of intertemporal substitution, g_t is charitable giving, and ε is a key preference parameter governing the general elasticity of giving. We first show theoretically that giving responses to wealth taxation are jointly determined by σ and ε , while the compensated own-price elasticity is driven only by ε . Accordingly, we use the bunching evidence to determine that $\varepsilon = 0.44$ in our calibration. We then find that the EIS (σ) must equal 0.0815 to replicate our findings on how wealth taxes affect giving. To identify the EIS, we seek to match the empirically estimated effect of changing a linear wealth tax rate and remain agnostic as to the differential effects of changing marginal and average tax rates. Hence, the low calibrated EIS provides partially independent evidence that intertemporal substitution effects are weak. The calibration exercise highlights how our findings from two different research designs inform the core parameters needed to model how charitable giving responds to a wide range of tax incentives. We illustrate the applicability of our calibrated model by simulating the partial-equilibrium response of a removal of the income-tax deductibility of giving. The effect of this

removal corresponds to an uncompensated own-price elasticity of -0.49, which is quite close to our estimated compensated elasticity of -0.44.

An important question is whether the effect of wealth taxation on giving is large enough to warrant attention in optimal tax models and from policymakers. In terms of a propensity to give out of wealth tax payments, our estimate is in fact economically modest at -0.012. However, this low level effect is driven by modest charitable giving in a country with ambitious social support systems, such as Norway. Hence, we further use our model to calculate dollar for dollar crowd-out effects in a high-giving environment such as the U.S., where giving as a fraction of GDP is more than ten times higher than in most European countries (OECD, 2020). By recalibrating κ such that giving equals 5% of gross income (List, 2011), we find a considerably higher crowd-out effect: each additional \$1 of annual wealth tax revenue would reduce the revenues of charitable organizations by \$0.28. We further find that the low EIS drives this large crowd-out. Setting the EIS equal to unity (i.e., log utility), each \$1 of annual wealth tax *increases* giving by \$0.033. The takeaway from this exercise is that spillover effects on charitable giving may indeed play a first-order role in determining the total redistributive effect of wealth taxation when the EIS is low, which is what this paper and other quasi-experimental studies find (e.g., Best et al. 2020), and when the overall level of giving is high, as in the U.S. and the U.K.

While our baseline calibration ignores fixed costs of giving, we find that our key finding of a low EIS is highly robust to modeling entry costs. Calibrating the EIS to those who already give, we estimate a virtually identical EIS. In a further exercise, we allow for heterogeneity in the strength of the giving motive (κ) and explicitly model the decision to enter into charitable giving. This allows us to use our empirical estimates on participation to calibrate entry costs of about NOK 35,000 (USD 5,500). This cost is roughly equal to one third of the present value of charitable giving over the life-cycle of marginal givers. We find that the average nonparticipant would give NOK 575 if entry costs were removed, which would increase total giving by about 36%.

Literature. Our paper contributes to a growing empirical literature on the effects of capital taxation (see, e.g., Lavecchia and Tazhitdinova 2021; Agersnap and Zidar 2021; Glogowsky 2021; Nekoei and Seim 2021; Martínez-Toledano 2020; Boissel and Matray 2021; Arefeva, Davis, Ghent, and Park 2021; Agrawal, Foremny, and Martinez-Toledano 2020; Wong 2020; Dray, Landais, and Stantcheva 2022; Londoño-Vélez and Tortarolo 2022) and particularly the literature on household responses to wealth taxation (Seim 2017; Zoutman 2018; Durán-Cabré, Esteller-Moré, and Mas-Montserrat 2019; Londoño-Vélez and Ávila-Mahecha 2020a; Londoño-Vélez and Ávila-Mahecha 2020b; Jakobsen et al. 2020; Brülhart, Gruber, Krapf, and Schmidheiny 2021; Ring 2020; Berg and Hebous 2021). Our main contribution to the wealth tax literature is to consider the effect on charitable giving. This contribution is important for three reasons. (i) Any effect on giving may amplify or muddle the intended redistributive effects of wealth taxation. (ii) There is no direct evidence on how wealth taxation affects consumption, charitable giving included. (iii) In our

setting, charitable giving is third-party reported, which allows us to isolate real responses. The existing wealth tax literature primarily either focuses on evasion or considers combined evasion, avoidance, and real responses to wealth taxes. A central finding in the wealth tax literature is that reported wealth is very sensitive to taxation. Extrapolating from this, one might expect to find that a way in which households reduce their wealth tax burden is to increase their giving. Our findings do not support this notion, which is consistent with households primarily adjusting through evasion or avoidance (Londoño-Vélez and Ávila-Mahecha, 2020a).

We further add to the body of research on the role of tax incentives in charitable giving. This literature is particularly concerned with the own-price elasticity of giving (see, e.g., Feldstein 1975; Randolph 1995; Auten, Sieg, and Clotfelter 2002; Meer 2014; Meer and Priday 2020; Bakija and Heim 2011; Fack and Landais 2010; Duquette 2016; Almunia, Guceri, Lockwood, and Scharf 2020; Hungerman and Ottoni-Wilhelm 2021; Cage and Guillot 2021). Our most direct contribution is to estimate the after-tax own-price elasticity using a methodology that is new to this literature, in combination with third-party reported data on giving. Few papers in this literature exploit non-linear price schedules, as we do, likely because exemption caps are typically not fixed, but depend on taxable income, as in the U.S. federal tax code.¹ Since we employ third-party reported data on giving, we are able to focus on actual giving rather than itemization responses to tax incentives (Meer and Priday, 2020). Consequently, our price elasticity of -0.44 is considerably smaller in magnitude than the elasticity of around -1 found in several of the analyses based on U.S. data.² We further provide important evidence on elasticity heterogeneity. Our finding that richer, higher-income households are, if anything, *less* price elastic suggests that the opposite finding in other settings could be driven by itemization rather than real giving responses.

Our main addition to the charitable giving literature is to consider the effects of capital taxation.³ By documenting how capital taxation affects giving, we shed new light on the intertemporal aspects of charitable giving (see, e.g., Breman 2011; Andreoni and Serra-Garcia 2021; and Meier 2007). In particular, our finding of a weak intertemporal substitution effect implies that households care not only about how much they give but also *when* they give it. This has implications for how one should model preferences for giving. Our paper also provides novel evidence on income effects. By studying the intensive-margin response to wealth taxation, we provide information on the marginal propensity to give out of unearned income. To our knowledge, few such quasi-experimental estimates exist (see Drouvelis, Isen, and Marx 2019 for an overview).

¹A notable exception is Hungerman and Ottoni-Wilhelm (2021) who exploit a state tax exemption threshold to estimate the elasticity of giving to a specific university.

²We note that the U.S. evidence does not unambiguously point to large estimates. While Bakija and Heim (2011) conclude that the price elasticity is close to -1, Randolph (1995) reports estimates ranging from -0.3 to -0.5. Our estimate is larger than the intensive margin elasticity of -0.2 found by Almunia, Guceri, Lockwood, and Scharf (2020) in the U.K. but close to that found by Fack and Landais (2010) for France.

³Cage and Guillot (2021) exploit a wealth tax reform in order to obtain identifying variation in the relative price for political and charitable giving, as opposed to the after-tax return on savings (as we do).

We also study whether capital taxation affects the own-price elasticity of giving. This is related to the notion that behavioral elasticities are not immutable parameters, but can rather be influenced by various policy instruments at the tax authorities’ disposal (Slemrod and Kopczuk, 2002). While, e.g., Fack and Landais (2016) document the effect of tax enforcement on the price elasticity of giving, there is no evidence on whether nominally unrelated tax parameters, such as the tax rate on wealth, may alter this elasticity.

Relatedly, we also contribute to the literature that studies crowd-out effects in charitable giving (see, e.g., Deryugina and Marx 2021; Gruber and Hungerman 2007; Andreoni and Payne 2003; Okten and Weisbrod 2000; Meer 2017; Payne 1998; Nyborg and Rege 2003; Hungerman 2005; and Boberg-Fazlić and Sharp 2017). This literature is particularly concerned with how government spending crowds out private giving. However, little attention is given to how the *financing* of government spending through taxing household savings may play an additional role.

The paper proceeds as follows. Section 2 introduces a simple life-cycle model with charitable giving that highlights the relationship between our reduced-form findings and structural primitives. Section 3 introduces the data and institutional setting. Section 4 considers the effect of wealth taxation on giving. Section 5 uses a bunching approach to estimate the own-price elasticity. In section 6, we calibrate the life-cycle model and discuss applications of it. Section 7 concludes.

2 Conceptual Framework

In this section, we introduce a simple partial-equilibrium model of charitable giving. The comparative statics that follow from the model demonstrate the connection between giving, wealth taxation, and the own-price elasticity of giving. We subsequently calibrate the model in section 6 to discuss the implications of our quasi-experimental findings.

Model. Suppose an agent optimally chooses consumption, c_t , annual giving, g_t , and savings, s_{t+1} . The agent derives per-period utility from consumption and giving. As in Hungerman and Ottoni-Wilhelm (2021) and Almunia et al. (2020), we use additively separable preferences and parameterize “warm-glow” utility from giving as $\kappa \frac{g_t^{1-1/\varepsilon}}{1-1/\varepsilon}$, where κ is a utility weight. This parametrization is convenient as it implies a constant compensated own-price elasticity equal to ε . Our empirical setting allows us to estimate ε and to test whether the constant elasticity assumption is reasonable. Unlike other work, we do not assume quasi-linear preferences. We allow for a finite elasticity of intertemporal substitution (σ) by parameterizing $u(c_t) = \frac{c_t^{1-1/\sigma}}{1-1/\sigma}$.

We define the following household optimization program:

$$\max_{\{c_t, s_{t+1}, g_t\}_{t=0}^T} \sum_{t=0}^T \beta^t \left[\frac{c_t^{1-1/\sigma}}{1-1/\sigma} + \kappa \frac{g_t^{1-1/\varepsilon}}{1-1/\varepsilon} \right], \quad (2)$$

$$\text{such that} \quad c_t = w_t - p_t g_t - s_{t+1} + s_t R_t, \quad (3)$$

where c_t is period t consumption, g_t is period t giving, and p_t is the (after-tax) price of giving. R_t is the gross after-tax rate of return on any savings (i.e., it equals 1 plus the interest rate, r_t , minus the wealth-tax rate, τ_t). w_t is exogenous disposable income.

If r_t is held constant, then $d \log R_t \approx -d\tau_t$ and $dR_t = -d\tau_t$. Hence, for simplicity, we do comparative statics with respect to R_t , but consider this equivalent to the effect of changing (in an opposite direction) the effective wealth tax rate, τ_t .

Proposition 1 *Assume that $R_t = R$ is constant from t to the end of the life-cycle, T . Then the level of giving, g_t , at some point in time, t , is determined by:*

$$g_t^{\sigma/\varepsilon} \left(\frac{p_t}{\kappa} \right)^\sigma \sum_{s=t}^T \beta^{\sigma(s-t)} R^{(\sigma-1)(s-t)} = \sum_{s=t}^T w_s R^{-(s-t)} - g_t p_t^\varepsilon \sum_{s=t}^T p_s^{1-\varepsilon} \beta^{\varepsilon(s-t)} R^{(\varepsilon-1)(s-t)}, \quad (4)$$

where w_t also contains beginning-of-period- t wealth.

Proof: See Appendix A.1.

Discussion: This equation is useful for calibrating the parameters, σ and ε , as we can use it to simulate the effects of changing R (by increasing the wealth tax rate, τ) on g_t , and then determine which parameters produce responses that best resemble the empirical findings.

Proposition 2 *If R_t is constant over time, $p_t = 1$, and $T = \infty$, then the derivative of giving at time t with respect to the future after-tax rate of return, evaluated at $R = \beta^{-1}$, is given by*

$$\frac{dg_t}{dR} = \underbrace{\frac{\varepsilon g_t}{\sigma c_t + \varepsilon g_t}}_{\text{Sensitivity}} \left\{ \underbrace{-(\sigma-1) \sum_{s=t}^{\infty} w_s R^{-(s-t)-1}}_{\text{Income v. substitution}} - \underbrace{\left(1 - \frac{1}{R}\right) \sum_{s=t}^{\infty} (s-t) w_s R^{-(s-t)-1}}_{\text{Human wealth effect}} - \underbrace{\frac{\varepsilon - \sigma}{R^2(1-R^{-1})} g_t}_{\text{Elasticity adjustment}} \right\} \quad (5)$$

Proof: See Appendix A.1.

Discussion: Letting $T \rightarrow \infty$ and evaluating at $R\beta = 1$ allow for simple comparative statics useful for building intuition. The resulting differential equation shows that the key parameters governing responses to tax-induced rate-of-return shocks are σ and ε . The proposition further shows that the effect is governed by familiar sources: the first term in the brackets is the classic income versus substitution trade-off. Inside this term, the agent increases giving if the after-tax rate goes down (due to, e.g., a wealth tax) if and only if $\sigma > 1$. The second term is a human

wealth effect in which lowering R increases the present value of future incomes, and thus giving through a wealth effect. The third term says that if giving is more elastic than consumption ($\varepsilon > \sigma$), then there is an additional intertemporal substitution effect. Empirically, this last term is less important since g_t is small relative to wealth and life-time income. Finally, the *sensitivity* term on the left-hand side says that the giving effect is larger in magnitude whenever the expenditure-weighted giving elasticity (εg_t) is large relative to the consumption elasticity (σc_t).

This proposition demonstrates the theoretical ambiguity in how giving respond to rate-of-return shocks, such as those from wealth taxation. If, for example, $\sigma > \varepsilon > 1$, a reduction in the after-tax rate of return will increase charitable giving. On the other hand, if σ is low and the human wealth effect is weak due to a downward-sloping income path, the effect may instead be negative.

The proposition also shows that one empirical moment is not enough to pin down the structural parameters. Both σ and ε are important, but an empirical estimate of $\frac{dg_t}{dR}$ cannot easily be used to calibrate both. This motivates our empirical strategy, in which we use the estimated own-price elasticity to infer ε (Proposition 3), and then the empirical estimate of $\frac{dg_t}{dR}$ to calibrate σ .

Proposition 3 ε determines both the elasticity of the growth rate of giving with respect to the after-tax gross rate and the compensated own-price elasticity.

$$\frac{d \log(g_{t+1}/g_t)}{d \log(R_t)} = \varepsilon \quad \text{and} \quad \left. \frac{d \log(g_t)}{d \log(p_t)} \right|_{u'(c_t)} = -\varepsilon. \quad (6)$$

Proof: The budget constraints and first-order condition for g_t imply that

$$-\varepsilon \log(g_t) + \log \kappa = \log p_t + \log u'(c_t). \quad (7)$$

When substituting in the first-order condition for s_{t+1} (Euler equation), we obtain $\frac{d \log(g_{t+1}/g_t)}{d \log(R_t)} = \varepsilon$ and $\frac{d \log(g_1)}{d \log(p_1)} = \varepsilon - \frac{d \log u'(c_1)}{d \log(p_1)}$. When considering a *compensated* price elasticity (as obtained in a bunching design), we have that $d \log(u'(c_1)) = 0$, and hence the second term equals $-\varepsilon$.

Discussion: Our simple model and Proposition 2 show that ε is a key parameter in understanding how giving responds to (tax-induced) changes in the rate of return. We further learn that it may be directly estimated either (i) from the effect of rate-of-return changes on the growth of giving or (ii) from estimating the compensated own-price elasticity, for example by using a bunching framework (as we do in section 5).

We also emphasize that $\frac{d \log(g_{t+1}/g_t)}{d \log(R_t)}$ should provide an estimate for ε purely through an intertemporal substitution channel (i.e., Euler equation). Empirically, however, if the shock to R_t is correlated with unexpected shocks to R_{t+1} , then there will be an additional income effect during $t+1$ that drives down g_{t+1} and hence the growth rate. In our quasi-experimental setting,

households treated by more wealth tax exposure in 2013 were even more treated in the subsequent years. Theoretically, this would cause a downward bias in the estimated ε . Accordingly, we do not rely on the Euler-regression approach for estimating ε .⁴

Proposition 4 *The sensitivity of giving to the log gross after-tax rate of return is a constant fraction of the consumption sensitivity, and this fraction equals the ratio of the EIS to compensated own-price elasticity of giving.*

$$\frac{d \log(g_t)}{d \log(R_t)} = \frac{\varepsilon}{\sigma} \frac{d \log(c_t)}{d \log(R_t)}. \quad (8)$$

Proof: This follows from differentiating equation (7) with respect to $\log(R_t)$.

Discussion: Equation (8) shows that the sign of the giving response, $\frac{d \log(g_t)}{d \log(R_t)}$, equals the sign of consumption responses to changes in the after-tax rate of return. This highlights how the theoretical ambiguity regarding consumption responses to rate-of-return changes apply to giving as well. It further emphasizes the role of ε in determining the strength of giving responses relative to consumption responses to rate-of-return shocks.

3 Data and Institutional Setting

3.1 Data

We employ administrative register data on households' income and wealth over the period 2010–2018 (Statistics Norway, 2019). These data contain information on how much a given household pays in wealth taxes each year and the composition of taxable wealth. In particular, estimated housing wealth is one of the key contributors to households' taxable wealth. In assigning housing wealth, the tax authorities distinguish between primary (regular abode) and secondary housing wealth. Importantly for our approach, we observe this distinction in our data.

We combine the administrative data on income and wealth with third-party reported data on charitable giving, recently available from administrative registers for the 2012–2018 period. Since charitable giving is tax deductible in the personal income tax, the tax authorities keep records of how much taxpayers give to charitable organizations. In order to limit the scope for tax evasion and reduce the administrative burden for the taxpayer, the tax authorities require these amounts to be reported directly by the recipient organizations. Hence, we observe the identities of both givers and the recipients. The tax authorities maintain a comprehensive

⁴We provide quasi-experimental evidence on how wealth taxation affects this growth rate in Appendix Table A.6. We find a positive effect of wealth taxation, which implies that increasing R lowers the growth rate, and thus that $\varepsilon < 0$ (although statistically insignificantly so). Hence, it appears that the bias is active. Note that the method is also imprecise, and our preferred estimate of ε from the bunching framework lies within the 95% confidence intervals of the growth-rate estimate.

list of qualified charitable organizations, and all of these report yearly donated amounts at the individual level to the tax authorities. Importantly, data are not truncated at the personal income tax deduction threshold; full amounts are reported. This provides us with rather unique, as well as comprehensive, panel data of charitable giving at the individual and household level. It follows that these data are not affected by issues related to self-reporting. Finally, since the untruncated data are only available from 2012, we supplement with a longer panel of charitable giving deductions from income tax returns. While these are truncated at the deduction threshold, they are used to assess the internal validity of our study by examining pre-trends.⁵

3.2 The Norwegian Wealth Tax

Norway has a long tradition of annually taxing net wealth using a progressive scheme. As of 2009, the wealth tax has taken a relatively simple form, where households pay wealth taxes according to the following formula:

$$wtax_{h,t} = \tau_t \mathbb{1}[TNW_{h,t} > T_t](TNW_{h,t} - T_t), \quad (9)$$

which states that for household h , observed at the end of year t , any taxable net wealth (TNW) in excess of a threshold, T_t , is taxed at a rate of τ_t .⁶ Tax rates and thresholds (2011–2018) are presented in Panel A of Table 1. The threshold increased from NOK 750,000 (USD 125,000) in 2012 to NOK 1,480,000 (USD 250,000) in 2018.⁷ This nonlinear wealth tax schedule may be summarized by the marginal and average wealth tax rates:

$$MWTR_{h,t} = \tau_t \mathbb{1}[TNW_{h,t} > T_t], \quad (10)$$

$$AWTR_{h,t} = \frac{wtax_{h,t}}{Net\ Wealth_{h,t}}, \quad (11)$$

These definitions imply that the marginal *return* on wealth is lowered one-for-one by an increase in $MWTR$ and, similarly, the average return on wealth is lowered one-for-one by increases in $AWTR$. The presence of valuation discounts on some assets imply that taxable wealth (TNW) generally differs from overall net wealth. While financial wealth predominantly enters the tax base, TNW , at third-party reported market values, the estimated market value of housing enters at a discounted fraction:

$$\text{Taxable Value of Housing Wealth}_{h,t} = (1 - d_t^{primary})MVHP_{h,t} + (1 - d_t^{secondary})MVHS_{h,t}, \quad (12)$$

⁵This longer panel is used when we provide event plots in order to evaluate pre-trends. In Appendix Figure A.3 we show that reduced-form effects after 2012 are very close irrespective of which data source is used.

⁶We account for the fact that married households are subject to two times the nominal threshold.

⁷We use the 2012 USD/NOK exchange rates for 2012 of about 6.

where $d_t^{primary}$ and $d_t^{secondary}$ refer to the discount rates for the different types of housing assets. $MVHP$ is the estimated market value for primary housing.⁸ A household's primary home is where households are registered to live according to government registers. $MVHS$ refers to secondary homes, which in government registers appear identical to primary homes. However, they are categorized as secondary rather than primary housing units only because of their registered current use: taxpayers may only own one unit of primary housing, but multiple units of secondary housing.

The evolution of the valuation discount rates, and how they changed for different types of housing, is our central source of identifying variation in wealth tax exposure. The valuation discount on primary housing, $d_t^{primary}$ has been fixed at 75% over the whole period, while the discount on secondary housing, $d_t^{secondary}$, decreased from 60% in 2012 to 10% in 2018. This implies that even if we keep the total value of housing wealth, $MVH = MVHP + MVHS$, constant, households who hold more $MVHS$ will see a clear increase in TNW over time. From equation (9), we see that this may cause both a higher annual wealth tax bill as more wealth is pushed above the wealth tax threshold, as well as higher propensity to face a lower return on any marginal saving (working through $\tau_t \mathbb{1}[TNW_{h,t} > T_t]$).

The presence of a wealth tax threshold in equation (9), as represented by the indicator function, $\mathbb{1}[TNW_{h,t} > T_t]$, is a key ingredient in this institutional setting. The presence of a threshold implies that shocks to TNW (through, e.g., lowering the discount rate on housing wealth) may affect the marginal after-tax return on wealth. If there had been no threshold, an increase in TNW would only increase the average tax rate and hence we would not expect any intertemporal substitution effects since these are driven by marginal tax rate changes.

Finally, we note that it is an easy task for taxpayers to understand how secondary housing wealth affects how much they pay in wealth taxes. Annual tax returns are pre-filled with an individual's assets, and how much they contribute to taxable wealth. From this pre-filled return, it is straightforward to see how different housing assets increase the wealth tax bill. For the average secondary home owner, the annual wealth tax bill more than doubled between 2012 and 2018.⁹ In addition, prior to any given tax year, taxpayers' income tax withholding rates are set by the tax authorities, calculated by predicting future income and wealth taxes.

3.3 Tax Treatment of Charitable Giving

Donations to charitable and religious organizations are tax deductible in the "ordinary income" tax base. The list of exemption-approved organizations is comprehensive (Sivesind, 2015) and

⁸The tax authorities employ a hedonic pricing model to estimate the market value of homes. Ring (2020) provides a detailed description of this methodology.

⁹Appendix Figure A.2 shows that each million NOK (MNOK) of secondary housing wealth increased annual wealth taxes by about NOK 2,800. The summary statistics show that the average secondary home was valued at 1.95 MNOK and that the average secondary home owner paid (after 2012) NOK 10,038 in wealth taxes. Hence, the relative effect is $2800 \cdot 1.95 / (10038 - 2800 \cdot 1.95) = 119\%$.

includes international organizations such as Amnesty International, the Red Cross, and Doctors Without Borders. The ordinary income tax base is taxed at a flat rate, which implies that the after-tax price of giving does not depend on an individual’s taxable income. This differs from other countries, such as the U.S., where charitable giving is also deductible in the tax bases that are subject to progressive taxation. For the 2018 tax year, for example, the government refunded 23% of taxpayers’ charitable giving up to a limit of NOK 40,000 (USD 6,700).

More generally, a taxpayer, i , gets a tax refund of τ_t^g of any charitable giving, $g_{i,t}$, that does not exceed the exemption cap, K_t :

$$\text{Giving Tax Refund}_{i,t} = \min\{\tau_t^g \cdot g_{i,t}, K_t \cdot \tau_t^g\}. \quad (13)$$

This creates a jump in the marginal after-tax price of 1 NOK worth of giving from $1 - \tau_t^g$ to 1 at $g_{i,t} = K_t$. The Norwegian tax treatment of charitable giving thus offers a tax-induced discontinuity in the marginal (after-tax) price. We summarize these tax rates and exemption caps in Panel A of Table 1.

Importantly for identification purposes, while the Norwegian income tax scheme directly incentivizes charitable giving through lowering its after-tax price, the wealth tax does not. This differs from wealth taxation in other settings. For example, in France, charitable giving was partially deductible from the annual wealth tax (Cage and Guillot, 2021): €1 in giving could reduce the wealth tax bill by €0.75. In the U.S. estate tax, charitable donations are deductible from the estate tax base (Bakija, Gale, and Slemrod, 2003), which is subject to high marginal rates. Thus, in these settings, capital taxation has a strong direct effect on incentives to give through altering its after-tax price relative to the price of other consumption goods.

3.4 Summary Statistics

Panel B of Table 1 provides the main summary statistics for our data. Our primary sample consists of households whose $TNW_{h,2012}$ was at most 0.5 million NOK (MNOK) below the threshold and at most 6 MNOK. This restricts the sample to households for whom decreases in housing discount rates may materially affect their wealth tax position. These TNW restrictions are not imposed when studying bunching at the giving deductibility threshold. We further condition on households having strictly positive estimated housing wealth (MVH) as of 2012. The table shows that 43% of the households in our sample paid a wealth tax in any given year, and conditional on paying the tax, paid about NOK 12,734 (USD 2,100) on average.

Approximately 19% of the households in our sample owned a secondary house. We further see that 32% of the households in our sample donate in any given year and, conditional on giving, they give approximately NOK 5,758 (USD 770) on average.

For our quasi-experimental evidence on how wealth taxation affects giving, treated households all own secondary homes. Hence, in Appendix Table A.1, we provide household-level summary

Table 1: INSTITUTIONAL DETAILS AND SUMMARY STATISTICS

Notes: Panel (A) provides information on wealth taxation and the tax-deduction scheme for charitable giving. Panel (B) provides summary statistics for the main sample used to study the effects of wealth taxation on giving. Net wealth equals taxable wealth gross of any valuation discounts. Amounts in Norwegian kroner (NOK) may be divided by 6 to obtain an approximate USD amount as of 2012.

PANEL A: INSTITUTIONAL DETAILS								
	2011	2012	2013	2014	2015	2016	2017	2018
Contribution to TNW, $(1 - d_t^{(\cdot)})$								
Primary, MVHP	25%	25%	25%	25%	25%	25%	25%	25%
Secondary, MVHS	40%	40%	50%	60%	70%	80%	90%	90%
Wealth tax rate, τ_t	1.1%	1.1%	1.1%	1.0%	0.85%	0.85%	0.85%	0.85%
Wealth tax threshold, T_t (MNOK)	0.7	0.75	0.87	1	1.2	1.4	1.48	1.48
Giving tax deduction rate, τ^g	28%	28%	28%	27%	27%	25%	24%	23%
Giving deduction cap, K_t (NOK 1,000)	12	12	12	16.9	20	25	30	40
PANEL B: SUMMARY STATISTICS, MAIN SAMPLE, 2012–2018								
	N	mean	p25	p50	p75			
1[Giving _{<i>h,t</i>} > 0]	4,078,145	0.32						
Giving _{<i>h,t</i>} if > 0	1,069,227	5,758	1,390	3,000	5,800			
1[wtax _{<i>h,t</i>} > 0]	4,078,145	0.43						
wtax _{<i>h,t</i>} if > 0	1,765,842	12,734	3,174	7,835	16,299			
wtax _{<i>h,t</i>} if MVHS _{<i>h,2012</i>} > 0	730,964	10,038	0	2,373	13,735			
MWTR _{<i>h,t</i>}	1,765,842	0.96%						
AWTR _{<i>h,t</i>}	1,765,842	0.20%						
As of 2012:								
MVH _{<i>h,2012</i>} , MNOK	624,969	3.12	1.72	2.55	3.84			
MVHS _{<i>h,2012</i>} if > 0, MNOK	111,469	1.95	0.92	1.56	2.51			
Net wealth _{<i>h,t</i>}	624,969	3,464,991	2,055,887	3,030,239	4,371,869			
Age _{<i>h,2012</i>}	624,969	61	50	62	73			
Gross income _{<i>h,2012</i>}	624,969	640,364	311,318	482,711	800,686			
Number of adults _{<i>h</i>}	624,969	1.39						

statistics by secondary home ownership. Some of the differences are intuitive: secondary home owners are wealthier (28% higher means) and have higher incomes (26%). Importantly, we control for these differences in our regression specifications. It is therefore reassuring that the distributions for income and wealth are substantially overlapping, i.e., the median of one group is between the 25th and 75th percentile of the other group. In terms of their charitable giving behavior, the households are very similar. Secondary home owners are only 1 percentage point more likely to give, and, conditional on giving, the medium amount of giving is identical.

We make a few adjustments to mitigate the impact of outliers. We bound the amount of wealth taxes paid, $wtax_{h,t}$, to 10% of $TNW_{h,2012}$. This adjustment affects only a modest number of households (who saw a cumulative TNW increase of at least 1000%). In addition, to account for moderate level increases from a small initial TNW , we limit $wtax_{h,t}$ to 10% of 1 MNOK if $TNW_{h,2012}$ is below 1 MNOK. We also limit both individual and household-level annual giving to NOK 100,000 (USD 16,700). This also affects very few households.¹⁰ When taking logs of giving, we do not limit the amount, but shift it by an inflation-adjusted NOK 1,000 in order to limit the influence of very small level differences in regressions.¹¹ Gross income, which is used as a control variable, is shifted by an inflation-adjusted NOK 10,000 prior to taking the logarithm.

3.5 Giving and Its Composition Across the Wealth Distribution

This section graphically describes the composition of giving in Norway. For the 2012–2018, we observe the identity of the recipient organizations. Using the organization identifiers and names, we classify all charitable organizations into one of three types. (i) Religious organizations are local or national churches, missionary organizations and non-christian religious organizations. (ii) Internationally-focused organizations includes groups such as the Red Cross, Amnesty International, and climate or environmental organizations such as the World Wildlife Foundation. Finally, (iii) domestic organizations range from groups that provide support for individuals with certain ailments, to coastal rescue organizations, and to local sports clubs.

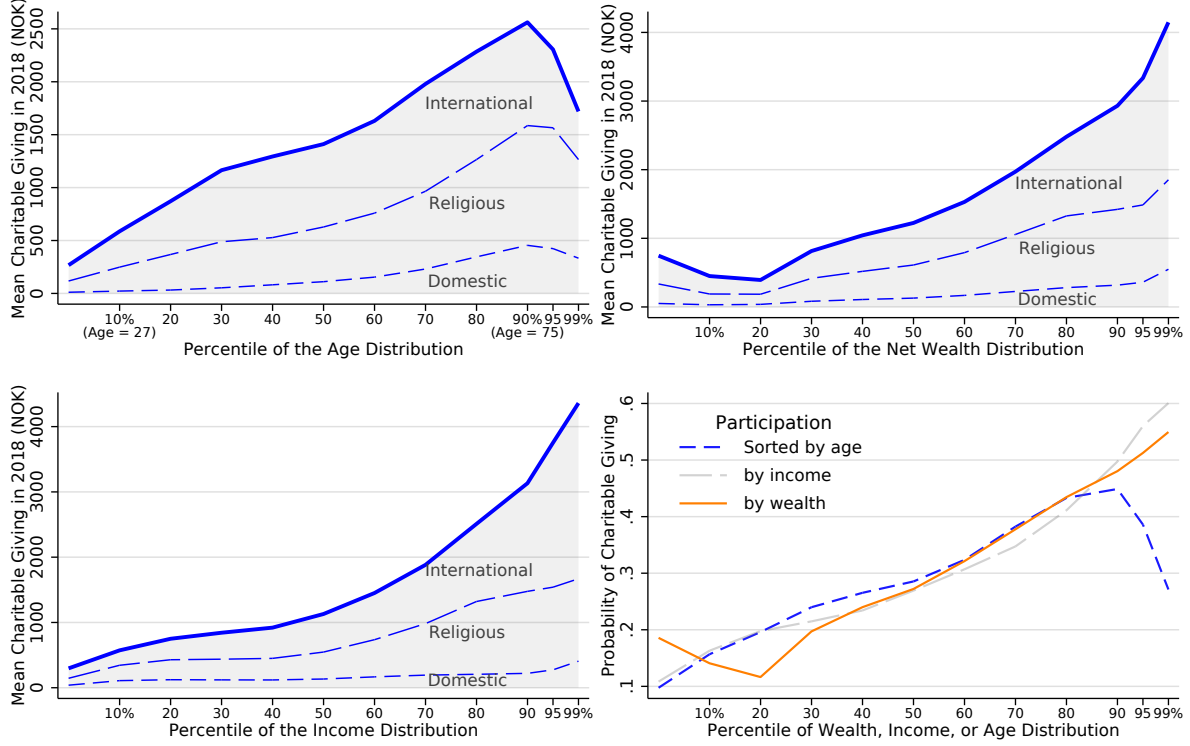
Since we are studying a wealth tax that mostly affects moderately wealthy and older high-income individuals, we plot the amount and decomposition of giving across the wealth, age, and income distributions. We provide our findings in Figure 1. The main distinctive feature of the Norwegian setting is the importance of international giving. While [Meer and Priday \(2021\)](#) find international giving to be modest at around 5% of total giving for moderately wealthy U.S. households, it accounts for almost 50% in our setting. Thus, while it is true that Norwegian households give considerably less on average, they give more than U.S. households to international organizations.¹²

¹⁰Conditional on giving a positive amount, the 99th percentile of household giving is NOK 50,000.

¹¹An alternative would be the inverse hyperbolic sine function (arsinh). While this approach may be appropriate in other settings, it is problematic when considering charitable giving where extensive-margin responses are important. For example, increasing giving from 0 to only NOK 50 is, for most purposes, economically unimportant, but causes an approximated log difference of 4.60 when using arsinh. Using the “log(1+x)” method produces a log difference of 3.93. Essentially, we would interpret this small effect as a 393–460% increase. Our approach, $\log(1000 * 1.02^{(t-2012)} + x)$, for $t = 2012$ instead produces a very small log difference of 0.049 and is equivalent to assuming that all households give an unobserved amount of NOK 1,000 that is unaffected by the treatment.

¹²[Meer and Priday \(2021\)](#) find total giving at the 90th percentile of wealth to be about \$3,300. With an international share of about 5% then equals about \$165. In our setting, mean international giving at the 90th percentile is approximately \$250.

Figure 1: CHARITABLE GIVING ACROSS THE AGE, WEALTH, AND INCOME DISTRIBUTIONS



Notes: The first three panels provide total and decomposed giving across the age, net wealth, and income distributions. The bottom-right panel provides participation rates (i.e., whether total giving exceeds zero). In the first three panels, average giving amounts are stacked to total overall giving. The y-axis value of a given line at, e.g., the 50th (95th) percentile provides unconditional averages for households with wealth weakly above the 50th (95th) percentile and strictly below the 60th (99th) percentile.

4 The Effect of Wealth Taxation on Charitable Giving

Hypotheses. There are at least two ways of producing hypotheses about the effect of wealth taxation on charitable giving. We may consider charitable giving as a form of consumption or assume that it is tightly linked to consumption through intratemporal first-order conditions, as in our model in section 2. In these cases, the sign of the giving response equals the sign of the consumption response, which is theoretically ambiguous (see, e.g., [Straub and Werning 2020](#)), because of countering income and substitution effects. The income effect lowers the amount of charitable giving, as more wealth taxation lowers lifetime after-tax wealth under the assumption that a charitable transfer is a normal good. The substitution effect causes more giving by changing the relative price of consumption across periods. More specifically, future consumption becomes relatively more expensive, which incentivizes households to consume (and give) more today rather than in the future.

The standard assumption is that substitution effects dominate. In structural models, a suffi-

cient condition is generally that the elasticity of intertemporal substitution exceeds unity. Hence, the baseline expectation is that wealth taxation *increases* charitable giving.

4.1 Quasi-Experimental Evidence from Housing Assessments

Our empirical framework exploits the fact that households with the same total housing wealth will see differential wealth taxation from 2013 to 2018, depending on the share of taxable housing wealth due to secondary housing (see section 3.2 for details). This allows us to control for different measures of wealth and thereby minimize the concern that wealthier households were on different counterfactual giving trajectories. This contrasts with the standard approach in the wealth tax literature of comparing households previously below or above tax thresholds when either the threshold or marginal tax rate subsequently changes. By comparing households previously above or below the threshold, empirical findings may be confounded if wealthier households are on different (counterfactual) trajectories. While this issue is not a first-order concern when we reasonably expect confounders to have a minimal impact relative to the treatment effect that one seeks to identify (see, e.g., [Londoño-Vélez and Ávila-Mahecha 2020a](#) and [Londoño-Vélez and Tortarolo 2022](#)), it may be problematic when considering real responses, such as giving, that are small in magnitude.

We use an instrumental variables (IV) approach, where the initial stock of secondary housing wealth is used as an instrument for wealth tax exposure. Importantly, this is done by controlling for overall housing wealth and other measures of wealth and income.¹³ More specifically, we estimate the following system of equations to identify the effect of wealth tax exposure on charitable giving behavior:

$$\begin{aligned} wtax_{h,t} = & \underbrace{f_t(MVHS_{h,12}, TNW_{h,12})\mathbb{1}[t > 2012]}_{\text{Secondary housing wealth instrument}} + \tilde{f}^{FS}(MVHS_{h,12}, TNW_{h,12}) \\ & + g_t^{FS}(MVH_{h,12}, TNW_{h,12}) + \alpha_t^{FS} + \eta_t^{FS}C_{h,t} + \epsilon_{h,t}^{FS}, \end{aligned} \quad (14)$$

$$\begin{aligned} Giving_{h,t} = & \underbrace{\beta x_{i,t}}_{\text{Instrumented variation}} + \underbrace{\tilde{f}^{SS}(MVHS_{h,12}, TNW_{h,12})}_{\text{Differences out 2012 effect}} \\ & + \underbrace{g_t^{SS}(MVH_{h,12}, TNW_{h,12})}_{\text{Controls for total (housing) wealth}} + \alpha_t^{SS} + \eta_t^{SS}C_{h,t} + \epsilon_{h,t}^{SS}, \end{aligned} \quad (15)$$

where MVH includes the total market value of primary ($MVHP$) and secondary housing ($MVHS$), g_t^{FS} and g_t^{SS} are estimated as time-varying polynomial functions, which take the

¹³Since some households are also affected by increases in the wealth-tax thresholds and reductions in the marginal tax rates during 2013–2018, our empirical specification differences out these effects by controlling for ex-ante taxable wealth. Hence, we isolate variation in wealth tax exposure to come from ex-ante ownership in secondary housing.

same functional form as f , \tilde{f}^{FS} , and \tilde{f}_t^{SS} .¹⁴ $C_{h,t}$ is a vector of household-level controls, which includes third-order polynomials in $TNW_{h,2012}$ and age, as well as a second-order polynomial in family size. $C_{h,t}$ further includes a dummy variable for whether there are two adults in the household, and log household labor income in 2012. We also include a 2012-valued indicator variable for ownership in secondary or recreational housing.¹⁵ α_t^{FS} and α_t^{SS} are year fixed effects, and $\epsilon_{h,t}^{FS}$ and $\epsilon_{h,t}^{SS}$ are the error terms.

The instrumental variation comes from $f_t(MVHS_{h,2012}, TNW_{h,2012})$, which provides identifying variation for $t > 2012$, when the post-2012 indicator, $\mathbb{1}[t > 2012]$, turns on. This exploits exogenous variation in wealth tax exposure driven by an increased contribution of $MVHS$ to the wealth tax base over time, which allows us to estimate β , the coefficient of interest. The $\tilde{f}(\cdot)$ terms difference out the baseline (2012) effect from the instrument. In other words, we employ a DiD-IV specification.

In order to estimate differential effects with respect to extensive and intensive margin wealth tax exposure (e.g., using both the marginal and average wealth tax rate as the left-hand-side variable in equation 14), we allow the first-stage effect of $MVHS_{h,12}$ to vary by TNW in 2012. A reasonable assumption is that households with higher initial TNW , and who therefore are further away from the wealth tax threshold, see relatively larger intensive margin (i.e., amount they pay) than extensive margin effects (i.e., whether they pay) of higher assessed secondary housing wealth. We verify this in our first-stage regressions (see Appendix Figure A.1), which means that we are able to separately identify the effects of intensive and extensive margin variation in wealth tax exposure.¹⁶ We parameterize by letting first-stage and reduced-form effects vary non-parametrically with initial TNW .

$$f(MVHS, TNW) = \sum_j \eta_{b_j} \mathbb{1}[b_j \leq TNW - T_{12} \leq b_{j+1}] \cdot MVHS, \quad (16)$$

where T_{12} refers to the wealth tax threshold in 2012,¹⁷ and $b_j = -0.5, -0.4, \dots, 0.3, 0.4, 0.6, 0.8, 1.2, 1.6, 2M, \infty$ MNOK, and the sample is limited to households with $TNW_{h,12} - T_{12} \geq b_1$. We include b_j dummies as control variables in both the first-stage and the reduced-form regressions to isolate the identifying variation to come from increased housing assessment as opposed to differences in ex-ante TNW .

In one application of this approach, we define the intensive margin to be the average wealth tax rate, $AWTR$, and the extensive margin to be the marginal rate, $MWTR$. We thus estimate

¹⁴For each function, g_t^{FS} , g_t^{SS} , f_t , \tilde{f}^{FS} , and \tilde{f}_t^{SS} , we estimate distinct parameters (η_{b_j}).

¹⁵A third category of housing, recreational housing (as cabins and other recreational dwellings) is assessed at historical cost (typically initial construction cost). This indicator variable thus controls for whether the household owns more than one housing unit.

¹⁶See Gruber and Saez (2002) for an implementation of this approach in the context of income taxation and Ring (2020) in the context of saving responses to wealth taxation.

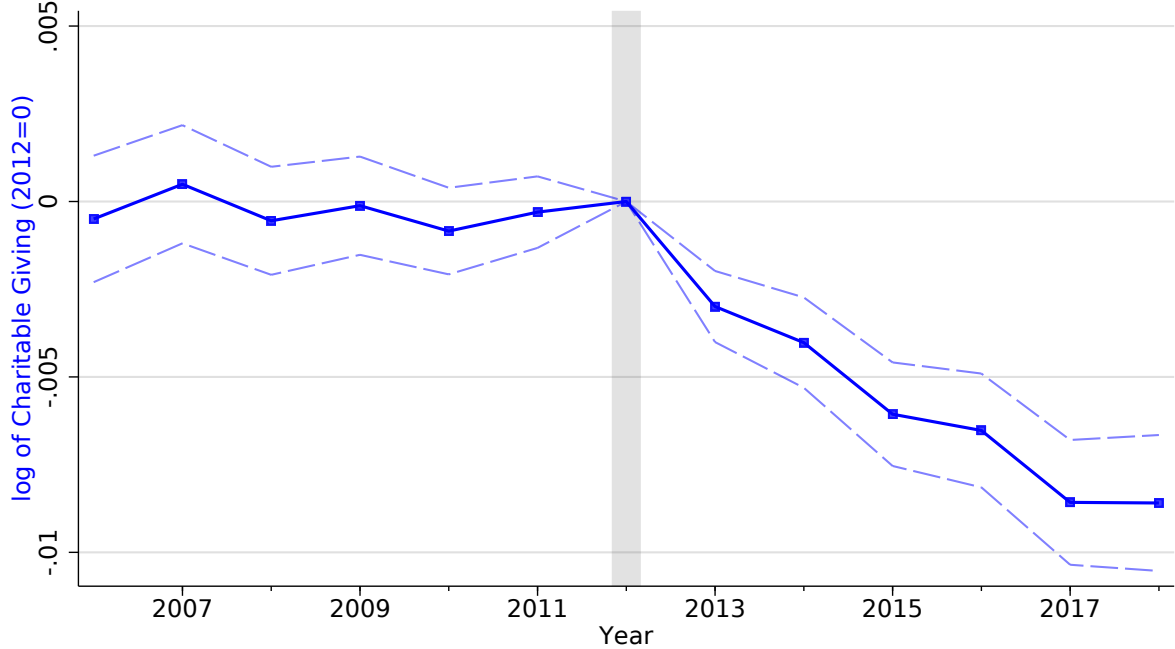
¹⁷While there is no household-level subscript on T_{12} , we account for the fact that married households face a double threshold by using 2012 marital status.

two coefficients, β^{AWTR} and β^{MWTR} . The sum of these two coefficients provide an important statistic: the effect of changing a linear tax rate on capital. That is, the effect of changing the wealth tax rate in an environment in which there is no tax threshold.

4.2 Reduced-form Evidence on How Wealth Taxation Affects Giving

We first provide reduced-form evidence in Figure 2 on how secondary-housing ownership affects wealth tax exposure and charitable giving over time. This is done by estimating a linear, time-varying relationship between initial secondary housing wealth and the log amount of charitable giving.¹⁸ For this analysis, we use the longer panel of giving that is truncated at the deduction limit but allows us to assess pre-trends.

Figure 2: WEALTH TAXATION AND CHARITABLE GIVING: QUASI-EXPERIMENTAL EVIDENCE FROM THE TAX VALUATION OF SECONDARY HOUSING



Notes: This figure shows the reduced-form effect on giving of the 2013-and-onward increase in wealth tax exposure among secondary home owners. Each point estimate shows the effect of owning 1 MNOK more in secondary housing wealth (measured in 2012). The point estimates come from estimating equation (14), simplifying $f(\cdot)$ to be a year-specific linear function of $MVHS_{h,2012}$ alone, and using charitable giving as the left-hand-side variable. Giving is measured using the “long panel” of charitable giving deductions from tax returns. We shift giving by an inflated-adjusted NOK 1,000 to accommodate zeros (and thus entries and exits) and to minimize the impact of small level changes. The dashed lines provide 95% confidence intervals. See Figure 3 for extensive-margin responses.

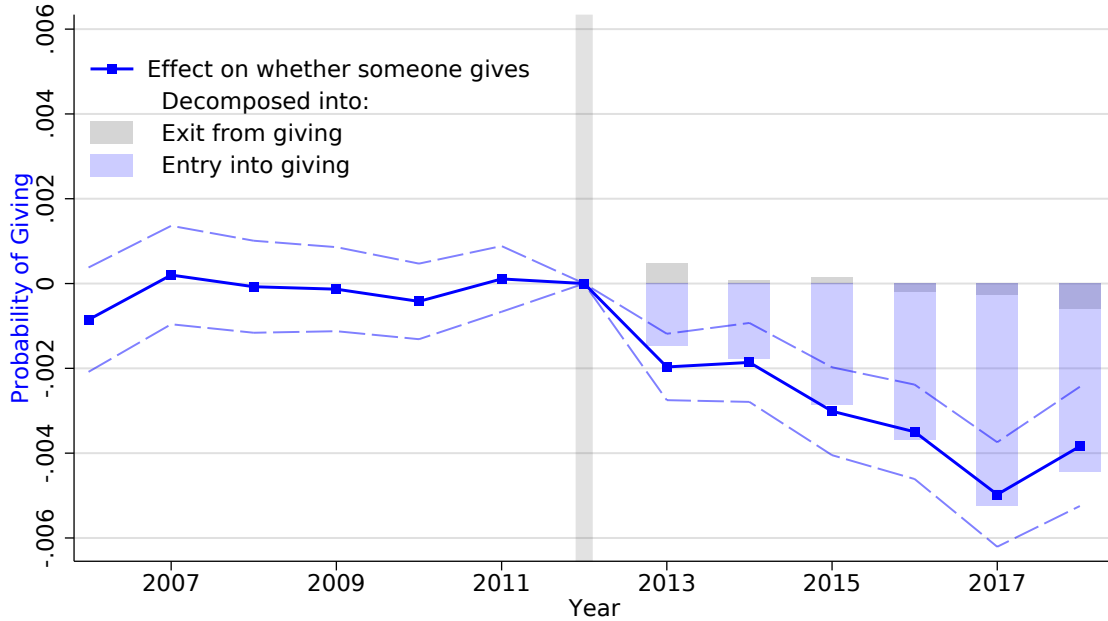
Figure 2 shows that initial ownership of more secondary housing, $MVHS_{h,2012}$, has a marked effect on charitable giving. Each additional MNOK of secondary housing wealth reduces charitable giving by almost 1%. These findings show that households respond gradually. This is

¹⁸We modify f in equations (14)-(15) to be a linear function in $MVHS$, where the slope may differ across years.

reasonable as the valuation discounts were gradually removed, resulting in the first-stage effect on wealth-tax exposure also being gradual (see Appendix Figure A.2).

Extensive-margin responses. The previous results include both intensive and extensive-margin adjustments. We proceed by focusing on extensive-margin effects, that is, whether someone chooses to give. If it is costly for households to participate in charitable giving, due to, e.g., attention or learning costs, households may respond to wealth taxation by stopping to give rather than just reduce the amount they give. Conversely, among households who do not participate in charitable giving, those more exposed to the wealth tax may be less likely to enter.

Figure 3: THE EFFECT OF WEALTH TAXATION ON WHETHER HOUSEHOLDS GIVE:
NO EXITS, BUT REDUCED ENTRY



Notes: This figure shows the reduced-form effect of wealth taxation on *whether* a households give. Each point estimate shows the effect of owning 1 MNOK more in secondary housing wealth (measured in 2012), which is our instrument for wealth tax exposure during 2013–2018. The point estimates come from estimating the reduced-form equation (14), simplifying $f(\cdot)$ to be a year-specific linear function of $MVHS_{h,2012}$ alone. The solid blue line provides the effect on whether a household gives ($\mathbb{1}[G_{h,t} > 0]$). We decompose this effect onto entries (blue) and exits (gray). The blue bars consider the number of times the household has gone from not giving to giving a positive amount, starting in 2012. Similarly, the gray bars consider the effect on the number of exits. The sum of these two variables equal a household fixed effect plus $\mathbb{1}[G_{h,t} > 0]$. Hence, the blue bars minus the gray bars should roughly equal the point-estimate for overall participation (solid blue line). $G_{h,t}$ is measured using the “long panel” of charitable giving deductions from tax returns. The dashed lines provide 95% confidence intervals.

Figure 3 shows clear extensive-margin effects. By 2018, the average treated household ($MVHS = \text{MNOK } 3$) is about 0.8 percentage points less likely to give. We find that this effect is entirely driven by reduced entry into giving. The theoretical implication is that there must exist entry costs. For some households, wealth taxation lowers the optimal amount of giving by enough for the utility of giving to be smaller than the (fixed) utility cost of entering the market for

giving. Similarly, since households who gave during 2006–2012 do not exit in response to wealth taxation, there is unlikely to be sizable per-period participation costs.

The regression equations underlying Figures 2 and 3 do not use the full potential power of the experiment: we are not exploiting the fact that households with different pre-period taxable wealth see very different first-stage effects on wealth tax exposure, which we show in Figure A.1. In addition, the specification does not allow us to distinguish between extensive and intensive-margin wealth tax exposure—which is why we next turn to the richer IV specification.

4.3 IV Results on How Wealth Taxation Affects Giving Behavior

In Table 2, we present results from the full DID-IV estimation, which uses the nontruncated panel on charitable giving that starts in 2012. We present our findings in two ways.

Estimated propensities to give out of higher wealth taxes. Columns (1)-(2) examine the effect on the amount given. In column (1), we find that each additional NOK paid in wealth taxes reduces giving by 0.012. This estimate may also be obtained from considering the dynamic effects on the amount of giving and wealth taxes in Appendix Figure A.2. Importantly, this estimate does not isolate intensive-margin effects. If intertemporal substitution is strong, the estimate of 0.012 is potentially a combination of stronger, negative income effects and slightly weaker positive substitution effects. However, the results in column (2) do not support this. The instrumented variation in *whether* households pay a wealth tax does not cause additional giving. The point estimate of NOK -28.58 (USD 5) is economically small and statistically insignificant.

Semi-elasticity of giving with respect to the wealth tax rate. Columns (3)-(4) consider the log of giving as the outcome variable and tax rates as the explanatory variables.¹⁹ This produces point estimates that correspond to semi-elasticities. Column (3) says that a 1 percentage point reduction in the marginal wealth tax rate reduces charitable giving by about 9%.

Importantly, column (3) implicitly assumes that either marginal and average tax rates are the same or that only marginal tax rates matter. Column (4) instead provides estimates of the effect of both marginal (*MWTR*) and average tax rate (*AWTR*) variation. This shows that the negative effect on giving is due to average tax rate variation: the point estimate on *AWTR* is large, negative, and highly significant. The point estimate on *MWTR* is insignificant. This suggests that while income effects matter, intertemporal substitution effects do not.

Column (4) provides the implied effect of a proportional change in the effective wealth tax rate. By summing the coefficients on *AWTR* and *MWTR*, we get that the implied effect on

¹⁹We shift giving by an inflation-adjusted NOK 1,000 (USD 167) prior to taking logs. This accommodates zeros and reduces the influence of small level differences. This is akin to assuming that all households give an unobserved fixed amount of NOK 1,000 each year.

giving of increasing a proportional (i.e., linear) wealth tax rate by 1 percentage point is -0.26%. We will use this point estimate to calibrate our structural model. This has the benefit of allowing us to be agnostic about the exact decomposition of the effect (i.e., the relative effects of changing *AWTR* and *MWTR*) and to test whether our model calibration corresponds, qualitatively, to the individual coefficients. Since we find an *MWTR* effect of essentially zero, we would expect that a *well-fitted* model produces an elasticity of intertemporal substitution close to zero.

We may use the intensive-margin estimate to compute a back-of-the-envelope *elasticity of giving with respect to after-tax wealth*. Assuming a 2% perpetual interest rate, a horizon of 30 years, the mechanical present value of a 1% tax on wealth is $1\% \times 23.40$. Hence, our findings indicate that a 23.40% reduction in after-tax wealth reduces giving by 22.44%, implying an elasticity of 0.96: As a household’s wealth grows, their charitable giving grows almost as much.

Log-transformations of variables that contain zeros are subject to caveats (Chen and Roth, 2022). However, we note that one obtains a very similar semi-elasticity of giving with respect to *AWTR* by using sample means to transform the level effect in column (2) of 2 to a semi-elasticity.²⁰

Model-based implications for how consumption responds to rate-of-return shocks. Our findings also speak to the theoretical ambiguity of consumption responses to changes in the after-tax rate of return. The frameworks underlying this ambiguity typically consider the effect of a “linear” or “proportional” change in the after-tax rate of return. This implies equally changing average and marginal tax rates. Our approach allows us to estimate the implied effect of such a change, which is the sum of the coefficients on *MWTR* and *AWTR*. We find that this coefficient is statistically significant and negative. Under Proposition 4, our findings imply that income effects dominate substitution effects in how *consumption* responds to rate-of-return shocks.

Additional implications of our findings are discussed when we calibrate our model in section 6. We note that, while we calibrate our model to the implied effect of changing a proportional wealth tax, we do not restrict the model to match the effect decomposition. However, our low calibrated EIS is theoretically consistent with a coefficient on *MWTR* close to zero.

First-stage and reduced-form regressions. We report reduced-form and first-stage coefficients for these IV regressions in Appendix Tables A.3 and A.4. The first-stage results are intuitive: households initially close to the threshold see larger extensive margin (*whether you pay a wealth tax*) effects, and households initially further above the threshold see larger intensive margin (*amount paid in taxes*) effects. It is this heterogeneity that allows us to simultaneously identify the effects of two different margins of wealth tax exposure using only one core instrument (i.e., secondary housing wealth).

Our findings compared to existing estimates on the marginal propensity to give.

²⁰Dividing -0.0121 by the unconditional average amount of giving per household ($5,758 \times 0.32$) and then further by the effect of 1 additional NOK of wealth taxes on the *AWTR* ($1/3464991$) provides a semi-elasticity of -22.74.

Table 2: WEALTH TAXATION AND CHARITABLE GIVING:
MAIN RESULTS FROM DID-IV REGRESSIONS

Notes: The table provides the key coefficients from estimating the system of equations in (14)-(15). *AWTR* is the average wealth tax rate, defined as the amount of wealth taxes paid divided by net wealth (net wealth equals *TNW* absent valuation discounts). *MWTR* is the marginal wealth tax rate and equals the nominal wealth tax rate if the household is above the wealth tax threshold and zero otherwise. See Appendix Tables A.3 and A.4 for the underlying reduced-form and first-stage estimates. In the log-giving specifications, giving is shifted by an inflation-adjusted NOK 1,000 to accommodate zeros and limit influence of outliers. Standard errors are clustered at the household level. One, two, and three stars indicate statistical significance at the 5%, 1%, and 0.1% levels.

	Amount		Adj. log of Giving			$\mathbb{1}[G_{h,t} > 0]$
			(intensive margin)			(extensive margin)
	(1)	(2)	(3)	(4)	(5)	(6)
<u>Instrumented variables</u>						
$wtax_{h,t}$ (NOK)	-0.0124*** (0.0029)	-0.0121*** (0.0034)				
$\mathbb{1}[wtax_{h,t} > 0]$		-28.582 (125.987)				
$AWTR_{h,t}$				-22.44*** (5.28)	-26.03* (11.20)	-23.83*** (3.89)
$MWTR_{h,t}$			-10.29*** (1.68)	-3.73 (2.26)	0.21 (4.71)	-2.72 (1.60)
Implied effect of proportional tax change						
= sum of coefficients on $MWTR$ + $AWTR$				-26.18*** (4.12)	-25.83** (8.42)	-26.55*** (3.12)
Sample restriction					$G_{h,t} > 0$	$G_{h,pre} = 0$
rk- F -statistic	225.69	269.89	291.81	316.54	91.24	189.57
N	4,007,561	4,007,561	4,007,561	4,007,561	1,300,392	2,614,018

Since the effects we find appear to be driven by income effects, it is useful to compare our findings to studies that strictly consider income effects on giving. For this, we rely on [Drouvelis, Isen, and Marx \(2019\)](#) who provide a summary of existing estimates. They show that existing non-experimental estimates on the marginal propensity to give (MPG) out of total income range from 0.024 to 0.093. However, estimates from windfall gains may be more closely aligned with our quasi-random variation in wealth taxation, and these are considerably larger, ranging from 0.16 to 0.74 ([Drouvelis et al. 2019](#)), which is an order of magnitude more than our estimated effect of 0.012. However, this low MPG is likely driven by an overall low level of giving in Norway. In section 6.3, we partially recalibrate our model to a high-giving environment and find an MPG of 0.29, which is still in the lower range of existing windfall-gain estimates.

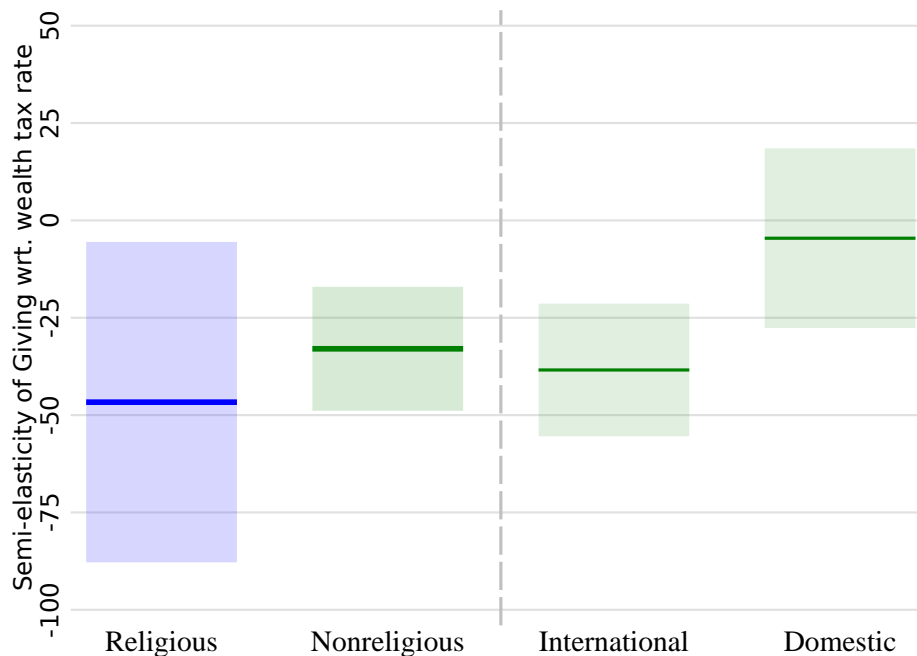
Existing findings on intertemporal substitution. Notably, our findings indicate that intertemporal substitution effects in giving are rather small. This is an interesting finding in accordance with evidence showing that the intertemporal substitution effect with respect to general consumption is low ([Best, Cloyne, Ilzetzki, and Kleven 2020](#); [Ring 2020](#)) and that savings are generally insensitive to tax incentives ([Friedman 2017](#), [Brülhart et al. 2021](#)). For example,

using a different quasi-experimental setting, [Ring \(2020\)](#) shows that income effects dominate substitution effects in household saving responses to wealth taxation. In order to offset the adverse effects of wealth taxes on life-time consumption, households save and work more. However, one may reasonably expect that giving behavior is more intertemporally elastic: while individuals need to consume certain goods and services on a daily basis and are subject to adjustment frictions ([Chetty and Szeidl, 2016](#)), it is not ex-ante obvious that preferences dictate that giving must be smoothed out over time as well.

4.4 Religious Heterogeneity in Responses to Wealth Taxation

Policymakers may place different weights on different kinds of charitable giving when considering the spillover effects of capital taxation. Hence, it may be useful to test whether there are heterogeneous effects across different types of charities. For this purpose, we divide charitable giving into two main types: religious and nonreligious giving. During 2012–2018, we find that only 6% of households gave to religious organizations, but the conditional amount is quite large at NOK 10,758 (USD 1,800) . About 26% gave to nonreligious organizations, but the conditional amount is considerably lower at NOK 3,325 (USD 550).

Figure 4: HOW WEALTH TAXATION AFFECTS DIFFERENT TYPES OF CHARITABLE GIVING



Notes: This figure shows the implied effect of changing a proportional (linear) wealth tax by 1 percentage point on different types of charitable giving. The underlying estimates are based on our IV methodology and are provided in Appendix Table [A.5](#). The lines mark the point estimates and the shaded areas provide 95% confidence intervals.

We provide our main findings in Figure 4. Each effect is estimated on the subset of households

who engaged in that particular type of giving in 2012. We find that religious and nonreligious giving both respond to wealth taxation. While the point estimate for religious giving is larger in magnitude, it is imprecisely estimated since only 6% of households give to religious organizations. In further splitting nonreligious giving into international and domestic, we find that the effect is largely driven by international giving. While the point estimate for domestic giving is small at about -4.56, we cannot rule out more substantial effects (i.e., the lower-bound of the 95% confidence interval is -27.64).

As suggested by the fact that international giving constitutes a substantial share of overall giving, these findings emphasize the possibility that capital taxation may have unintended cross-border spillover effects.

4.5 Discussion of Potential Confounding Factors

The fact that wealth tax payers (by construction) are wealthier poses important challenges in obtaining causal effects of wealth taxation. The standard approach is to exploit wealth tax reforms that change marginal tax rates for households above the wealth tax threshold or reforms that change the threshold, which may cause a new group of households to be subject to the wealth tax. This facilitates the removal of confounders that do not vary with time (i.e., fixed effects). However, identification relies on comparing households who differ in terms of their pre-reform (taxable) wealth (see, e.g., [Jakobsen, Jakobsen, Kleven, and Zucman 2020](#)) who may be on different *trajectories*. To address this, we employ a quasi-experimental approach that allows us to flexibly control for pre-reform wealth and isolate variation coming from changes to the tax authorities’ assessment rules. Since we control for ex-ante wealth interacted with year dummies (which would absorb the identifying variation in other studies), we are not subject to the standard concern that wealthier households are changing their behavior for reasons unrelated to wealth taxation. The causal interpretation of our findings is only limited to the extent that households with larger *MVHS* in 2012—keeping overall wealth and housing wealth fixed—increased or decreased their giving during 2013–2018 for reasons unrelated to wealth taxation.

Importantly, the lack of pre-trends in Figure 2 is reassuring. We are not aware of other reforms or economic shocks occurring in 2012 that would differentially affect secondary home owners. Since we address differential trends that may be driven by differences in initial overall taxable wealth, total housing wealth, income, age, and family size—any confounding factors would be limited to the convex allocation of housing into primary and secondary housing assets—and they would have to only play a role as of 2013. We mitigate this concern by including a control variable that captures whether a household owns either secondary housing or *recreational* housing. Recreational housing units (e.g., cabins) are often qualitatively similar to secondary housing but are treated in a particularly tax-favored way, due to a different valuation system, and does therefore not give much variation in wealth tax exposure.²¹ Thus, this control dummy

²¹Upon construction, houses are typically categorized as either housing (primary or secondary) or recreational

variable may largely take out differential trends arising from owning a second housing unit, while retaining considerable first-stage predictive power on wealth tax exposure. In our opinion, this likely removes potential confounding effects arising due to changes in the housing-unit level economies of scale in homeownership. Unfortunately, the fundamentally different ways in which secondary and recreational housing are valued precludes the use of recreational-housing owners as a (exclusive) control group.

Property taxation. One potential concern is municipality-level property taxation. While the tax authorities’ discount rates are unrelated to local governments’ property tax schemes,²² secondary home owners may benefit from per-house exemption thresholds. These favor a strictly convex allocation of housing wealth into primary *and* secondary housing. It is thus conceivable that we identify effects from households that are paying less in property taxes. However, municipality level property taxation has been in place for decades, and there were no trend-breaking changes in 2012. Hence, the lack of pre-trends in Figure 2 is reassuring, and suggests that our DiD strategy will address this issue by taking out a baseline effect in 2012. Furthermore, if property taxation lowers giving through an income effect, this would lead to an *upward* bias in our estimates (i.e., toward zero). Given our findings, this is not a material concern in terms of the qualitative conclusions.

Price effects. Finally, it is worth noting why potential house price effects of the reform do not affect how we interpret the findings. Most importantly, price effects do not lead to additional income effects that can explain our findings. Any price effect simply decreases the attractiveness of households’ undoing the treatment of the discount-rate change by liquidating their housing position. If they sell, some of the benefits (lower tax bill) is offset by a lower sales price. We find that it is useful to think of the income or wealth effects of the reform in terms of a decision tree. If households never liquidate their secondary-housing position, the income effect is driven only by higher wealth tax payments. If households do liquidate, the wealth-tax income effect ceases and the price effect materializes. Importantly, the two effects are not at play simultaneously. This means that price effects may make the income effects more persistent—but not larger than what is implied by having to pay higher wealth taxes. Secondly, price effects are likely limited. Recall that a given house is categorized as primary or secondary based on its current rather than potential use. If (current) primary and secondary homes are homogenous, there should not be any price effect. However, following the reductions in the valuation discount rates, ownership in secondary housing becomes less attractive for the subset of households who anticipate being

housing by the Norwegian Tax Administration. While tax values for primary and secondary housing are estimated using a hedonic pricing model, recreational housing is assigned a value shortly after construction, which is inflated by using nationwide multiples up until 2010. After 2010, the assigned values have remained fixed, and are typically thought to severely understate the true market value. Older dwellings in areas with high cumulative house price growth since the date of construction are particularly likely to be undervalued.

²²In Norway, effective local property tax rates rarely exceed 0.5% and many municipalities do not collect property taxes. While *some* municipalities began using the tax authorities’ assessment values for property taxation as of 2015 (they were either not allowed or strongly discouraged during 2010–2014, see discussion in Ring 2020), the change in discount rates would not affect the values provided to municipalities, as these are not discounted.

above the wealth tax threshold.²³ Thus, to the extent that secondary homes are more likely to be resold as secondary homes, due to their particular location or features, we would expect a negative house-price capitalization effect.

5 The Own-Price Elasticity of Charitable Giving

We now turn our attention to the *own-price* elasticity of charitable giving by using a bunching design to estimate the compensated own-price elasticity of giving. This estimate both informs our model presented in section 2 and is a key input to determining optimal tax incentives (Saez, 2004). We extend the analysis by studying elasticity heterogeneity in terms of income and wealth and the type of giving, and we study whether households' exposure to wealth taxation affects their own-price elasticity.

5.1 Bunching Methodology

As described in section 3.3, charitable giving is deductible in a specific portion of the income tax base that is subject to a flat tax rate, τ_t^g , of 22–28%. This means that the effective discount on charitable giving is independent of overall income. Importantly, deductions are subject to a cap. The presence of a cap on giving deductions creates a setting in which the marginal after-tax price of giving jumps at a pre-specified threshold.

The bunching methodology exploits the fact that the marginal after-tax price of giving jumps from $1 - \tau_t^g$ to 1 at the exemption cap, K_t and allows us to estimate the giving elasticity,

$$e = \frac{\Delta G^* / K^*}{\Delta \log(1 - \tau^g)}, \quad (17)$$

where ΔG^* is the reduction in giving of the marginal buncher who is at an interior optimum at the exemption cap, K^* (i.e., the kink). In the modeling framework in section 2, this compensated own-price elasticity is a fixed parameter and equal to $-\varepsilon$.

The bunching mass is denoted B . By construction (see Saez 2010 and Kleven 2016 for graphical intuition), B equals $\int_{K^*}^{K^* + \Delta G^*} h_0(G) dG$, where $h_0(G)$ is the counter-factual (absent a kink) probability density function of giving. We apply the standard approximation

$$B = \int_{K^*}^{K^* + \Delta G^*} h_0(G) dG \approx -h_0(K^*) \Delta G^*. \quad (18)$$

Dividing through by K^* , we may write the (approximated) relative change in the giving of the

²³42% of secondary home owners paid a wealth tax as of 2018. Among these wealth-tax-paying secondary home owners, if we consider secondary-housing wealth to be the marginal asset class subject to the wealth tax, approximately 66% of it was subject to the wealth tax. In other words, the presence of the wealth-tax threshold shielded about 72%=1-42%×66% of taxable secondary housing wealth from taxation.

marginal buncher as

$$\frac{\Delta G^*}{K^*} = \frac{-B}{h_0(K^*)K^*} \equiv \frac{-b}{K^*}. \quad (19)$$

This equation represents one of the central insights of the bunching literature, namely that the marginal buncher's response to the kink is proportional to the excess mass at the kink.

We empirically estimate b , which we refer to as the bunching estimate, using the methodology in [Chetty, Friedman, Olsen, and Pistaferri \(2011\)](#). The empirical analog of K^* is the (average) exemption cap for giving denominated in the same units (NOK 100) as the empirical giving bins.²⁴ We write our estimated compensated own-price elasticity as

$$\hat{e} = \frac{-\hat{b}/\bar{K} \cdot 100}{-\log(1 - \tau^g)}, \quad (20)$$

where we have used that $\Delta \log(1 - \tau^g) = \log(1) - \log(1 - \tau^g) = -\log(1 - \tau^g)$. In our main approach, we pool observations across years. We use K_t to sort households into bins and calculate \hat{b} , but use the across-year averages for K_t and τ_t^g , which are denoted without t subscripts, to compute \hat{e} according to equation (20).

5.2 Bunching Results: The Own-Price Elasticity of Giving

Panel A of Figure 5 provides preliminary evidence on bunching by plotting the frequency of different giving amounts. We pool data across years, and see that the spikes in frequencies coincide with the tax-deduction caps. The values of these caps are indicated by the vertical orange lines.

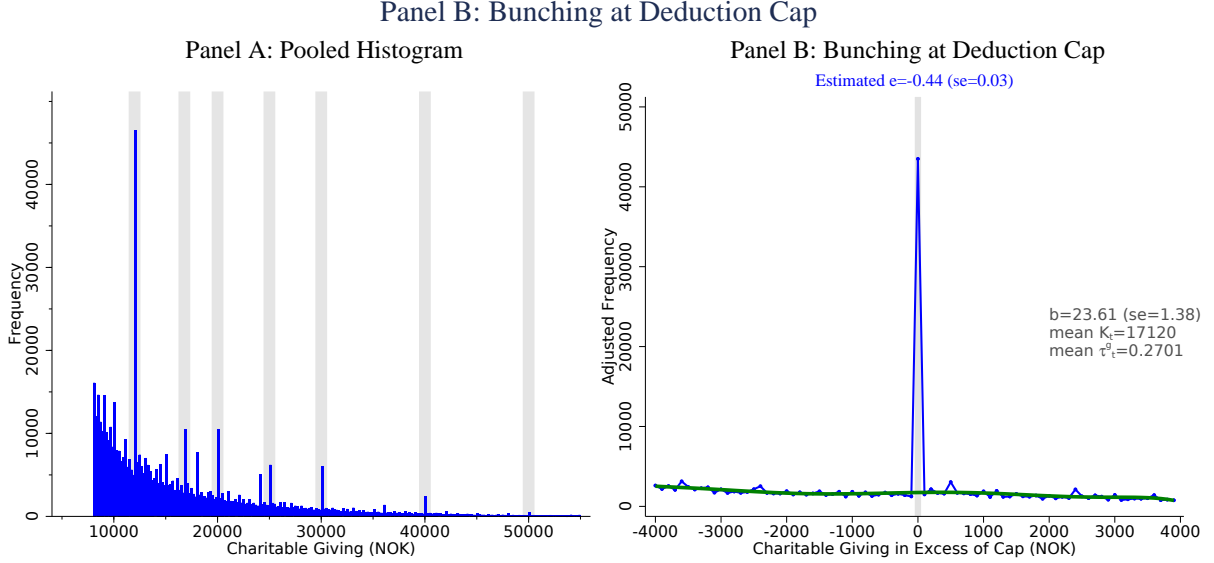
Panel B provides more formal evidence of bunching. We now sort taxpayers based on how much they gave relative to the deduction cap, K_t , applicable for that year. A value of 1000 on the x-axis indicates that the taxpayer donated between NOK 1000 to NOK 1099 more than the cap. Since it is common to donate round-numbered amounts, and donation-exemption caps are round numbers, we adjust the frequencies for round-number bunching that is unrelated to the tax incentive at hand.²⁵ The dotted blue line (at the bottom of Panel B) shows the adjusted frequency, while the solid green line describes the counterfactual adjusted frequency. The counterfactual density function is produced by estimating a 5th order polynomial on all observations

²⁴Thus $K^* = \bar{K}_t/100$. Alternatively, we could multiply \hat{b} by the width of the earnings bins (NOK 100), and let K equal the threshold in NOK.

²⁵This is done by first deducting the mean frequency, at the "giving-bin-year" level, in a leave-me-out fashion. For round-number bins, defined as multiples of NOK 1,000, this involves calculating the mean frequency across other years in which the deduction cap was different. Then to obtain a baseline frequency, absent round-number bunching, the mean frequency of two adjacent bins is added. If the resulting adjusted frequency is below the mean of the two adjacent bins, which occurs for a handful of bins, the value is set to the mean of the adjacent bins. Frequencies are then calculated at the bin level by aggregating across sample years. This adjustment procedure lowers the estimated bunching elasticity by 12%.

outside the bunching region, $BR_t \equiv [K_t - 1000, K_t + 1000]$, measured in NOK.²⁶

Figure 5: BUNCHING AT THE TAX-DEDUCTION CAP FOR CHARITABLE GIVING



Notes: The vertical gray lines, from left to right, in Panel A, represent the deductibility caps for 2012–2013, 2014, 2015, 2016, 2017, and 2017, respectively. The average amount of charitable giving at the kink point in panel B is NOK 20,240, and the bin width is NOK 100. In Panel B, the frequency is adjusted for round-number (thousands) bunching. Standard errors (se) are obtained from a 200-repetition bootstrap procedure.

The estimated \hat{b} is 23.61. This number says that there are 2,361% excess givers at the kink. This large bunching mass translates into a more modest elasticity due to the strong change in tax incentives at the kink: inserting $\hat{b} = 23.61$, $\bar{K}/100 = 15.914$, and $\tau^g = 0.2719$ into equation (20) produces an elasticity estimate (and standard error) of

$$\hat{e} = \frac{-0.44}{(0.03)} . \quad (21)$$

The implied estimate for the parameter ε in our model (see section 2) is thus 0.44.

Own-price elasticity estimate relative to other findings. Our own-price elasticity of -0.44 is considerably smaller than the elasticity of around -1 found in several analyses based on U.S. data. Finding a smaller elasticity in Norway than in the U.S. is reasonable as a U.S. price elasticity reflects both real giving responses as well as itemization responses, where the latter reflects whether taxpayers choose to deduct charitable giving (Meer and Priday, 2020).²⁷

²⁶An alternative approach is to use prior years' empirical density function as a counterfactual (see Londoño-Vélez and Ávila-Mahecha 2020a). This approach is not well suited for our setting since the exemption threshold was always present and increasing over time. Hence, the prior years' distribution around the current year's kink may have been affected by the prior years' kinks.

²⁷The self-reporting aspect is also present in the U.K., where Almunia et al. (2020) document an intensive margin elasticity of only -0.2. However, in the U.K., the after-tax price is lowered by a combination of tax deductibility and governmental matching of private donations. Beyond the different research designs, the possibility that taxpayers value matching less than deductions may explain the different findings (see, e.g., Karlan and List

It is unlikely that the magnitude of our elasticity estimate is low due to the absence of income effects in our bunching design (see discussion below). In section 6.2, we use our calibrated model to calculate an implied uncompensated elasticity of -0.39.

Compensated or uncompensated elasticity. Despite the large change in the relative price of giving at the threshold, intuition suggests that the income effect of the threshold has a limited impact on the elasticity estimates. To see this, consider the households counterfactually located far beyond the exemption cap in Panel B of Figure 5. Absent the exemption cap, these households would pay NOK 4,000 times 27% less in tax. This is about NOK 1,100 (USD 180). Multiplying this by our estimated wealth-tax income effect coefficient (MPG) of 0.012, we would expect the total income effect here to only be NOK 12 (USD 2). This implied income effect is thus much too small to have a meaningful effect on the estimated elasticity. Although the interpretation of implied elasticity estimates from bunching as reflecting compensated effects is subject to caveats (Blomquist et al. 2021; Kleven 2016), we adopt the convention of considering our bunching estimate as a compensated elasticity. This implies that our estimate of $\hat{\epsilon}$ equals (the negative of) the preference parameter, ϵ , from Proposition 3 in section 2.

We also note that $\hat{\epsilon}$ is a *locally* estimated elasticity: it rationalizes observed giving behavior of individuals near the exemption cap. Without structural assumptions, such as those imposed in section 2, it is not directly informative of the price elasticity of individuals who either barely give (and are thus far below the cap) or those who give well in excess of the cap. Below, in section 5.3.3, we discuss how our finding that wealth taxation does not affect the price elasticity supports the constant elasticity assumption. We also note that since the deduction cap has varied over time, our estimate is effectively an average local elasticity across giving levels ranging from NOK 12,000 to 50,000 (USD 2,000 to 8,000).

5.3 Heterogeneity in the Own-Price Elasticity of Giving

5.3.1 Regression-based Methodology for Uncovering Bunching Heterogeneity

Following Bastani and Waldenström (2020a), we estimate a linear probability model to study the correlation between various characteristics, Z , and the probability of bunching at the deduction cap for giving (a measure of price sensitivity), $\mathbb{P}[G_{i,t} \in BR_t]$. This is done by estimating the following regression equation,

$$\mathbb{1}[G_{i,t} \in BR_t] = \alpha_t + \delta \mathbf{Z}_{i,t} + p(G_{i,t}) + \varepsilon_{i,t}, \quad (22)$$

where α_t takes out year fixed effects and $\mathbf{Z}_{i,t}$ is a vector of characteristics of interest. A particularly attractive feature of this regression-based approach for uncovering bunching heterogeneity is that it avoids the omitted variable bias that normally arises when simply doing sample splits by some variable. In our case, we are particularly interested in income and wealth as explanatory

2007).

variable, which are two highly correlated characteristics. We also include age as an explanatory variable since it is highly correlated with income and particularly wealth. An innovation in our paper is to also include a third-order polynomial in the amount of charitable giving, $p(G_{i,t})$, as a control variable to address the issue that whether someone bunches is generally correlated with the amount they donate.²⁸ We run these regressions for observations where $G_{i,t} \in SR_t \equiv [K_t - 10\,000, K_t + 10\,000)$ and report the results in Table 3.

5.3.2 Relating Bunching Heterogeneity to Price-Elasticity Heterogeneity

The estimates in Table 3 show that older, higher income, and wealthier taxpayers are less likely to bunch at the deduction cap for giving. This implies that they are less price elastic, but the economic relevance of this heterogeneity does not follow directly from the regression estimates. To address this, we introduce a simple methodology for relating the estimated coefficients from equation (22) to differences in the own-price elasticity.

First note that the estimated relative excess mass of givers at the threshold, \hat{b} , can be rewritten as the relative excess probability of observing an individual at the threshold,

$$\hat{b} = \frac{\hat{P}^a[G_i \in BR] - \hat{P}^{cf}[G_i \in BR]}{\hat{P}^{cf}[G_i \in BR]}, \quad (23)$$

where \hat{P}^a denotes the actual empirical probability of observing anyone in the bunching region, BR , and \hat{P}^{cf} refers to the estimated counterfactual probability. We thus use equation (20) and rewrite \hat{e} as

$$\hat{e} = -\frac{\hat{P}^a[G_i \in BR] - \hat{P}^{cf}[G_i \in BR]}{\hat{P}^{cf}[G_i \in BR]} \frac{100}{\overline{K}} \frac{1}{-\log(1 - \tau^g)}, \quad (24)$$

where $100/\overline{K} = 1/K^*$. Hence, the effect of a unit increase in some covariate, $Z_{i,t}$, on e is

$$\widehat{\frac{de}{dZ}} = -\hat{\delta} \frac{1}{\hat{P}^{cf}[G_i \in BR]} \frac{100}{\overline{K}} \frac{1}{-\log(1 - \tau^g)}, \quad (25)$$

where $\hat{\delta}$ is a regression estimate of how Z affects the propensity to bunch. We obtain the estimate for the scaling parameter $\hat{P}^{cf}[G_i \in BR]$ by solving equation (23), using the bunching-sample mean $\hat{P}^a[G_i \in BR] = 5.09\%$ and the estimated \hat{b} from Figure 5. This implies that

²⁸Roughly, we may think that this correlation problem occurs when the threshold location differs from the mean amount of donations in the sample or estimation region. Since many characteristics correlate with G_i , we control flexibly for the amount of charitable giving, $G_{i,t}$, in order to minimize the risk of picking up spurious correlations with bunching behavior. Of course, we may not control too flexibly for $G_{i,t}$: for example, granular fixed-effect bins are in danger of completely absorbing the dependent variable, the bunching indicator. A third-order polynomial, on the other hand, seems to be a reasonable way to address correlations between G_i and Z_i without absorbing the correlation between Z_i and $\mathbf{1}[G_{i,t} \in BR_t]$.

$\hat{P}^{cf}[G_i \in BR] = \frac{5.09\%}{23.67-1} = 0.2245\%$. We further use $\tau^g = 27\%$ and $\bar{K} = 17,120$ to get

$$\widehat{\frac{de}{dZ}} \approx -\hat{\delta} \times 8.27. \quad (26)$$

5.3.3 Own-Price Elasticity Heterogeneity Results

Table 3 shows that the propensity to bunch is decreasing in age, income, wealth (see column 4). When simultaneously estimating these correlations, however, the role of wealth is removed in favor of income. Importantly, these correlations, although statistically significant, are economically modest when translated into differences in elasticities.

Table 3: HETEROGENEITY IN THE AFTER-TAX OWN-PRICE ELASTICITY

Notes: The dependent variable is a bunching indicator (multiplied by 100). The unconditional sample mean is 5.09 (%). The implied effect on the tax-price elasticity can be obtained by multiplying the point estimates by the elasticity multiplier near the bottom of the table. To calculate the multipliers, the sample-specific estimates for b are used (see Figure 5). The sample consists of homeowners, but we do not exclude households based on $TNW_{h,2012}$. Stars indicate significance at the 5%, 1%, and 0.1% levels.

	BUNCHING PROBABILITY (%)			
	(1)	(2)	(3)	(4)
Age _{<i>i,t</i>}	-0.0938*** (0.0024)			-0.0973*** (0.0027)
log(Gross Income _{<i>h,t</i>})		-0.5472*** (0.0636)		-1.1892** (0.0695)
log(Net Wealth _{<i>h,t</i>})			-0.2925*** (0.0348)	-0.0627 (0.0346)
f(<i>G_{i,t}</i>)	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	Yes
Unconditional Probability			5.09%	
$\hat{P}^{cf}[G_i \in BR]$			0.2245%	
Multiplier for Elasticity			8.27	
R ²	0.2441	0.2403	0.2399	0.2380
N	792,271	789,820	740,570	739,411

Age heterogeneity. We find that 10 year older households are 1 percentage point less likely to bunch, which translates into a

$$10 \cdot \frac{0.0938}{100} \times 8.27 \approx 8 \text{ percentage point} \quad (27)$$

lower elasticity (see equation 26). This is not a negligible correlation, but it is also not particularly large relative to the common baseline elasticity of -1 . With an average elasticity of -0.44 , it would be a fair characterization that even households 20 years younger than the mean are still modestly elastic at -0.63 .

Income heterogeneity. Controlling for age and wealth, we find that a 10% higher income is associated with a 0.98 percentage point lower elasticity. While statistically significant, this

correlation is zero in terms of its economic impact.

Wealth-based heterogeneity. When controlling for income and age, we find that wealth is a statistically and economically modest predictor of the price elasticity: a 10% increase in *Net Wealth* lowers the elasticity by about 0.05 percentage points.

The finding of no economically important heterogeneity is consistent with the constant elasticity assumption inherent to our model in section 2. However, it may be driven by preference heterogeneity cancelling out a causal effect of higher income and wealth. For example, having higher incomes may cause taxpayers to become less elastic, but at the same time, high-income households may tend to have less elastic preferences. We explore this potential explanation in subsection 5.4.

Literature comparison. Our findings contrast the intensive-margin results reported in Bakija and Heim (2011) and Almunia, Guceri, Lockwood, and Scharf (2020), both pointing to *increasing* responsiveness with respect to income. They also contrast the finding in Fack and Landais (2010) that price elasticities are increasing in the level of giving, which is highly correlated with income and wealth. The fact that we find different results is likely driven by differences in the reporting regimes. Third-party reporting is shown to substantially increase the number of observed givers in tax data (Gillitzer and Skov, 2018), which broadens the set of households that we can study. When focusing on self-reported giving, Meer and Priday (2020) find that the main driver of their estimated price elasticity is households’ beginning to report charitable giving on the extensive margin, i.e., starting to itemize. In our setting, however, individuals cannot choose whether their donations are reported to the tax authorities; thus the extent to which higher-income taxpayers are more diligent in claiming tax deductions from giving does not play a role.

5.4 Quasi-Experimental Evidence on the Relationship between Wealth Taxation and the After-Tax Own-Price Elasticity

In this section, we provide quasi-experimental evidence on how wealth taxation affects the own-price elasticity of giving. As in subsection 5.3.3, we use a regression-based framework to uncover bunching heterogeneity and map this into differences in the own-price elasticity. The innovation in this section is that we employ the quasi-experimental variation in wealth taxation from section 4 as the source of heterogeneity. This allows us to obtain plausibly causal evidence on the effects of wealth taxation on the own-price elasticity.

This exercise is useful for two reasons. First, given our finding that wealth taxation reduces charitable giving, a natural follow-up question is how a government may use direct tax incentives (e.g., deductibility) to undo the effect of introducing a wealth tax. To answer this question, one needs to know the own-price elasticity of giving. However, it is not obvious that elasticity estimates obtained under a non-wealth-tax regime apply when a wealth tax is in effect. On the one hand, wealth taxation may cause negative income-like effects among givers and cause them

to become more price elastic. On the other hand, givers may become less price elastic because the opportunity cost of giving is lower when any savings are subjected to a wealth tax.

Second, studying how wealth taxation affects the own-price elasticity allows us to test the constant compensated-price elasticity assumption underlying our model in section 2. The modeling assumption of constant compensated price elasticities is ubiquitous in public finance and macroeconomics (e.g. when assuming a constant Frisch elasticity in models of labor supply decisions). However, it is challenging to discern whether descriptive heterogeneity in the price-elasticity is due to differing preferences or is causally driven by differences in income or wealth. This matters for optimal taxation since tax policy affects after-tax income and wealth but not necessarily preferences. Since wealth taxation reduces (after-tax) wealth, we may use our quasi-experimental setting to test whether wealth causally affects the own-price elasticity of giving.

Table 4: WEALTH TAXATION AND THE PRICE ELASTICITY OF GIVING:
QUASI-EXPERIMENTAL EVIDENCE

Notes: This table provides the key instrumental variables coefficients from estimating the system of equations in (14)-(15). Standard errors are clustered at the household level. One, two, and three stars indicate statistical significance at the 5%, 1%, and 0.1% levels. Point estimates may be multiplied by 6.74 to obtain the effect on the magnitude of the price elasticity. We consider a wider sample region ($SR_t = [K_t - 10\,000, K_t + 10\,000]$) than in Figure 5 in order to obtain a larger sample to increase power in our IV design.

	(1)	(2)	(3)
Dependent variable =	Bunching Probability (%)		
<u>Instrumented variables</u>			
wtax _{h,t} (NOK 1,000)	0.049 (0.059)	0.036 (0.066)	
1[wtax _{h,t} > 0]		1.499 (2.317)	
AWTR _{h,t} (%)			4.422 (7.423)
MWTR _{h,t} (%)			1.135 (2.518)
<hr/>			
sum of coefficients AWTR+MWTR			5.557 (6.137)
rk- <i>F</i> -statistic	40.49	39.43	39.838
N	299019	299019	299019

We present our findings in Table 4. Focusing on column (1), we see that a NOK 1,000 increase in wealth taxes increases the propensity to bunch by a modest and statistically insignificant 0.05 percentage points. To grasp the potential economic significance of the point estimate, we use the methodology introduced in section 5.3.2 to map these bunching effects into a price-elasticity effect by multiplying point estimates by 8.²⁹ Hence, considering the effect of a fairly large

²⁹Previously, the multiplier was 8.27. We obtain a slightly different multiplier due to differences in sample characteristics. Here, we have that $\hat{P}^a[G_{i,t} \in BR_t] = 5.2837\%$, which implies that $\hat{P}^{cf}[G_{i,t} \in BR_t] = 0.2330\%$. Thus the multiplicative factor becomes $8 \approx (1/0.002330) * (100/17120) * (1/\ln(1-0.27))$.

increase in the annual wealth tax bill of NOK 10,000, we find that

$$\Delta wtax_{h,t} = \text{NOK } 10,000 \Rightarrow \Delta \hat{e} = \frac{0.0392}{(0.0472)} , \quad (28)$$

which is a very modest effect. In column (2), we see that these point estimates are fairly robust to isolating intensive-margin wealth-tax effects by also instrumenting for whether a household pays a wealth tax ($\mathbb{1}[wtax_{h,t} > 0]$).

We also transform our findings into a second-order cross elasticity by using average and marginal wealth tax rates (*AWTR* and *MWTR*) as our instrumented explanatory variables in column (3). Keeping the marginal tax rate constant, we find that a 0.2 percentage point increase in *AWTR* (the mean *AWTR* of wealth tax payers) would increase the probability of bunching by about 0.884 percentage points, and hence the own-price elasticity by a modest 0.07.³⁰

Interestingly, there is no evidence that the compensated own-price elasticity is decreasing in the marginal tax rate on wealth. This is what could happen if after-tax own-price savings are less valuable as the higher *MWTR* causes households to prefer current rather than future consumption and giving through an intertemporal-substitution like effect.

In summary, our empirical findings support the notion that the own-price elasticity is invariant with respect to after-tax wealth. Implicitly, our findings thus support the notion that compensated price elasticities are not causally related to wealth or disposable income. Furthermore, we cannot reject the null hypothesis of a constant compensated price elasticity, inherent in our model in section 2. A natural follow-up question is whether the *uncompensated* price elasticity is sensitive to wealth taxation. In section 6, we use our calibrated model to address this question, and find virtually no sensitivity in our setting.

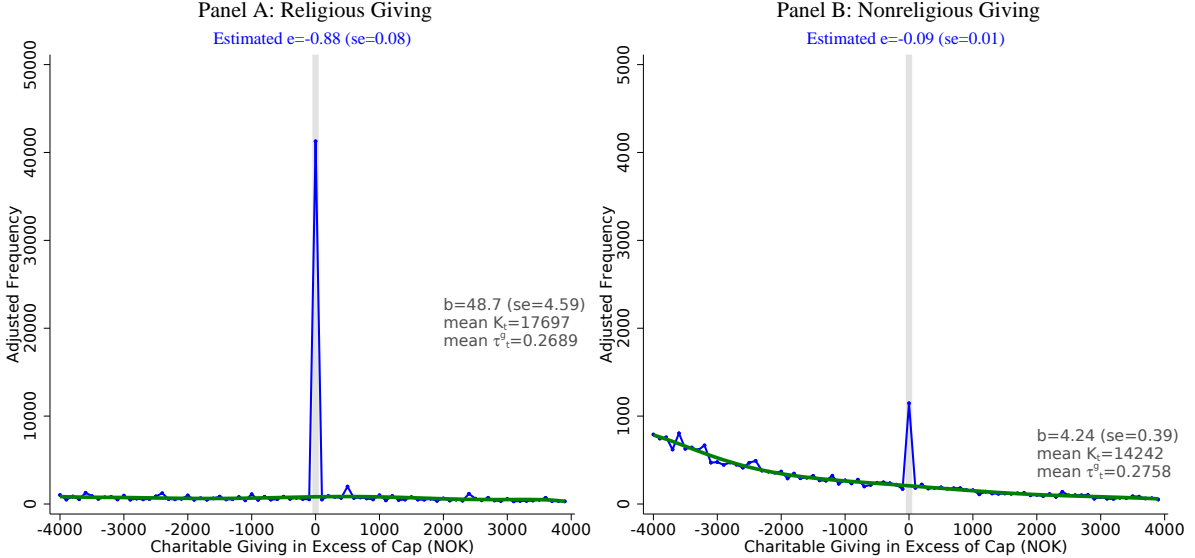
5.5 Religious Price-Elasticity Heterogeneity

Our final heterogeneity exercise considers the *type* of charitable giving. Optimal tax incentives for expenditures, such as giving, depend critically on how elastic these expenditures are (Saez, 2004). If, for example, the government cares equally about all types of giving, but price elasticities are heterogenous, then the government will be undersubsidizing some types, and over-subsidizing others. While there is some work on tax-price heterogeneity,³¹ we benefit from an empirical setting that offers both quasi-experimental individual-level price variation and data that distinguish between different types of giving.

³⁰This is small relative to the knife-edge case where the elasticity takes a value of 1. While it is less small relative to our own estimate of 0.44, our estimate is already quite low. For example, Saez (2004) only considers values of 0.5, 1, and 1.5 in calibrating a model of optimal tax deductions on giving.

³¹See, e.g., Brooks 2007, Yetman and Yetman 2013, and Duquette 2016; and Reinstein 2011 who finds that religious giving is less sensitive than overall giving.

Figure 6: RELIGIOUS OWN-PRICE ELASTICITY HETEROGENEITY



Notes: We repeat the main bunching analysis (Panel B of figure 5) for two subsamples. In Panel A, we consider the subsample of individuals who only gave to religious organizations (e.g., local or national churches, missionary organizations, theological institutes, and non-Christian charitable organizations). In Panel B, we consider households who only gave to nonreligious organizations. Appendix Figure A.5 considers the subset of givers who only give to internationally-focused (nonreligious) organizations.

We present our main findings in Figure 6. We find an astonishing difference in the propensity to bunch. Religious giving displays a bunching elasticity that is more than ten times higher than nonreligious giving. Per equation (20), we translate the bunching estimates into the implied compensated own-price elasticities. We find that religious giving displays an elasticity of 0.88, while nonreligious giving displays an implied elasticity of only 0.09. While religious giving is close to elastic, nonreligious giving is particularly irresponsive to tax incentives. In Appendix Figure A.5, we zoom in on giving to internationally-focused organizations. Here, we find a slightly lower implied elasticity of 0.07.

6 Model Calibration, The EIS, and Applications

6.1 Main calibration results

We calibrate our model presented in section 2 to our empirical setting, and contrast the theoretical and observed treatment effects of wealth taxation in Figure 7. We employ a representative agent model (as in Jakobsen, Jakobsen, Kleven, and Zucman 2020) and endow the agent with the characteristics of the average household from Table 1. We use the precise estimate of ϵ of 0.44 obtained from bunching at the deductibility cap, and plot theoretical treatment effects for different values of the elasticity of intertemporal substitution (σ). We describe the calibration exercise in more detail in Appendix B. For the purpose of this calibration exercise, we leave heterogeneity

with respect to the type of giving aside and focus on overall charitable giving.

Our main finding is that a very low elasticity of intertemporal substitution is needed to replicate our main empirical findings. The empirical finding we try to match is the implied effect on log giving of changing a linear (proportional) tax on wealth by 1 percentage point, which we found to be -0.2618 percentage points in column (4) of Table 2. Figure 7 shows that this requires

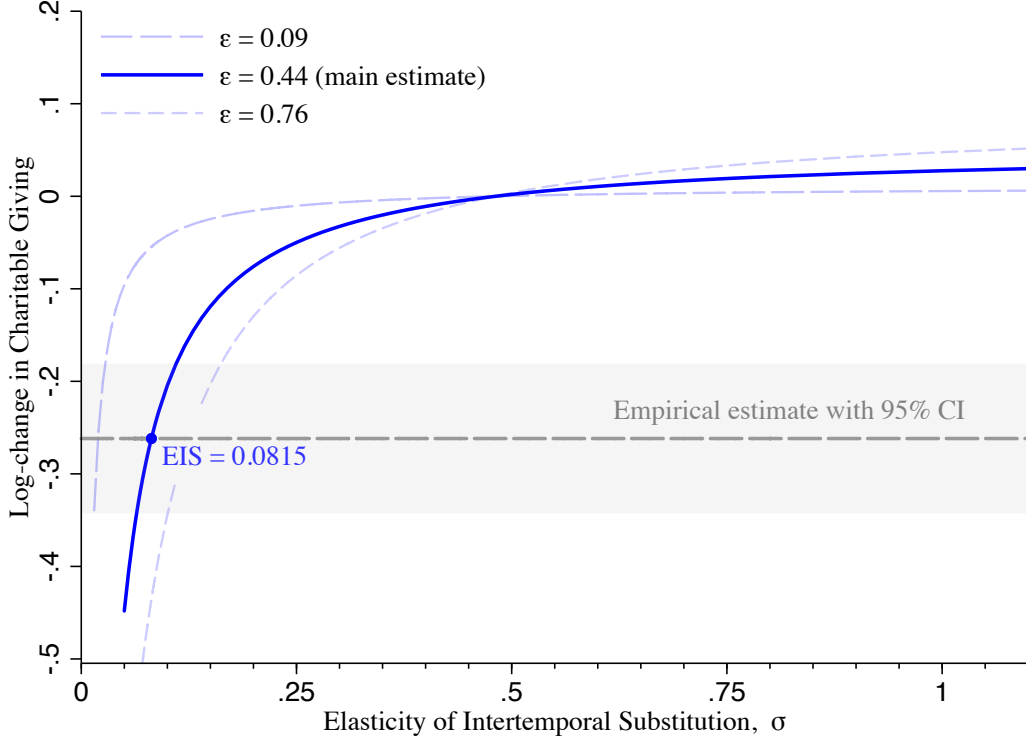
$$\sigma = 0.0815. \tag{29}$$

This low EIS is considerably smaller than most empirical estimates (Havránek, 2015). While Jakobsen et al. (2020) also calibrate the EIS to quasi-experimental moments from wealth taxation, their EIS estimates of 2 to 6 are considerably larger. A possible explanation is that their EIS estimate is driven by changes to evasion or avoidance behavior since they examine changes in taxable wealth where (self-reported) business assets is a major component. When using changes to reported wealth to infer how intertemporally elastic consumption is (via the budget constraint), changes to reported wealth that is driven by evasion or avoidance will make consumption seem more intertemporally elastic than it truly is. Our approach differs by considering only third-party reported behavior on giving, and inferring the consumption EIS indirectly (through the intratemporal first-order conditions). This approach allows us to avoid important empirical issues related to self-reporting of taxable wealth (see Advani and Tarrant 2021 for a review). Reassuringly, our low EIS estimate is very close to that from other quasi-experimental settings where self-reporting is not a factor (see, e.g., Best, Cloyne, Ilzetzi, and Kleven 2020 who find an EIS of 0.1).

An interesting observation is that the structural parameter, ε , which governs the own-price elasticity, is almost six times larger than the consumption elasticity, σ . As Proposition 2 illustrates, this means that giving effects can be large relative to consumption effects even the initial level of giving is relatively low. In terms of the relative economic importance of giving vis-a-vis consumption (measured in terms of expenditures), the high ratio of ε to σ may substantially make up for the fact that the ex-ante level of giving is small relative to consumption. This can be inferred from the *Sensitivity* term in Proposition 2, where the effect on giving is multiplied by $\varepsilon g_t / (\varepsilon g_t + \sigma c_t)$.

Figure 7: COMPARING MODEL-BASED AND EMPIRICAL EFFECTS OF WEALTH TAXATION ON CHARITABLE GIVING

Notes: This figure shows model-based treatment effects on charitable giving for combinations of the elasticity of intertemporal substitution (σ) and the absolute value of the compensated own-price elasticity (ε). Using a life-cycle model with endogenous giving that is calibrated to our empirical setting, we plot the effects of a 1 percentage point unexpected reduction in the after-tax rate (i.e., a 1% wealth tax). Under our baseline $\varepsilon = 0.44$, we find that $\sigma = 0.0815$ replicates our empirical findings. Our calibration is described in more detail in Appendix B and the main elements are presented in Proposition 1. The dashed gray line provides the empirical estimate of the implied effect of changing a proportional wealth tax rate, which is presented in column (4) of Table 2. The middle, blue line provides the theoretical treatment effect assuming that $\varepsilon = 0.44$, corresponding to the bunching evidence in section 5.2. The dashed lines provide the theoretical effects assuming ε equals the estimates from the religious and nonreligious subsamples (see Figure 6)



Unmodeled frictions and potential biases in the inferred EIS. We find that our findings are best rationalized by intertemporally inelastic consumption preferences. An alternative explanation is that inattention or other frictions prevent even high-EIS households from responding more. In our setting, the negative effect on disposable incomes may be more salient relative to the intertemporal tax incentive that should accelerate giving. Hence, the fact that households may be quite passive in response to intertemporal incentives (Chetty, Friedman, Leth-Petersen, Nielsen, and Olsen 2014) may lead to low-EIS mimicking behavior.

In general, the presence of financial frictions may cause low-EIS-like behavior since households who are financially constrained may be unable to optimally increase their giving in response to wealth taxation. However, we believe financial frictions play a very limited role in our setting. Firstly, Norwegian wealth tax payers generally have considerable amounts of liquid wealth (Ring, 2020). Furthermore, secondary home owners who pay a wealth tax will, by design, tend to have

considerable home equity, implying that they have collateral available for borrowing. Finally, intertemporal substitution effects should be driven by those affected on the extensive margin (i.e., those who see a reduction in the marginal after-tax rate of return) and these are the households for whom the overall liquidity effect should be the smallest.

Another potential caveat is that biases arising from estimating ε using a bunching design may influence the calibrated EIS. For example, it may be costly for households to pay attention to the tax rules. Hence, they may not know the exact deduction threshold and thus be less likely to bunch. Even if households know the location of the threshold, bunching may be costly (Mavrokonstantis and Seibold, 2022). These frictions may downward bias the estimated ε . Recent evidence also suggests that households may keep bunching at a prior year’s tax threshold to avoid revealing tax evasion (Londoño-Vélez and Ávila-Mahecha, 2020a). However, in our setting, charitable giving bunching is not indicative of evasion. In addition, any downward bias in our estimate of ε only strengthens our key finding of a low EIS. This is apparent from Proposition 2 and Figure 7 that show that a lower ε requires a lower σ to replicate our empirical findings.

A related concern is that some of the households in our sample may be irresponsive to wealth taxation since they are bunching at the tax-deduction cap. This implies that our model may exaggerate giving responses and thus allow us to replicate a given negative effect with a higher EIS. To the extent that this biases our EIS, the bias works in the opposite direction of our finding of a low EIS.

In terms of external validity, it is possible that our sample is skewed in favor of nonreligious giving, which we find to be less price elastic (see Figure 6). However, even if we calibrate the EIS to a considerably higher ε of 0.76 (from the religious-giving subsample), the required EIS is still very small at about 0.12. This can be seen in Figure 7, where we provide simulated treatment effects when $\varepsilon = 0.76$.

6.2 Model Application I: Uncompensated Own-Price Elasticity

In section 5.2, we estimated an own-price elasticity of -0.44. Importantly, this elasticity identifies the response of marginal bunchers and thus provides a *compensated* price elasticity. This elasticity does not give the full answer to the question of what would happen to charitable giving if, for example, the tax deductibility was removed. This is because income effects would also play a role: as the tax-incentive is removed, households become poorer and thus give less.

Our calibrated model (with $\sigma = 0.08$ and $\varepsilon = 0.44$) allows us to calculate the implied uncompensated price elasticity. We do this by calculating the log-change in giving caused by changing the after-tax price p_s from 0.73 to 1 for $s = t, \dots, T$. This results in a

$$\text{Model-implied } \hat{e}^{\text{uncompensated}} = -0.4914, \quad (30)$$

which is quite close to our compensated elasticity estimate. This implies rather weak income effects of price changes, consistent with the low *level* of giving in our data. If we calibrate our model to a 10 times higher level of giving (through increasing the utility weight κ , see discussion in next subsection), the uncompensated elasticity remains modest at -0.6626. Hence, our model implies that our low elasticity estimate relative to the existing literature is unlikely to be driven by the fact that we estimate a compensated as opposed to uncompensated elasticity.

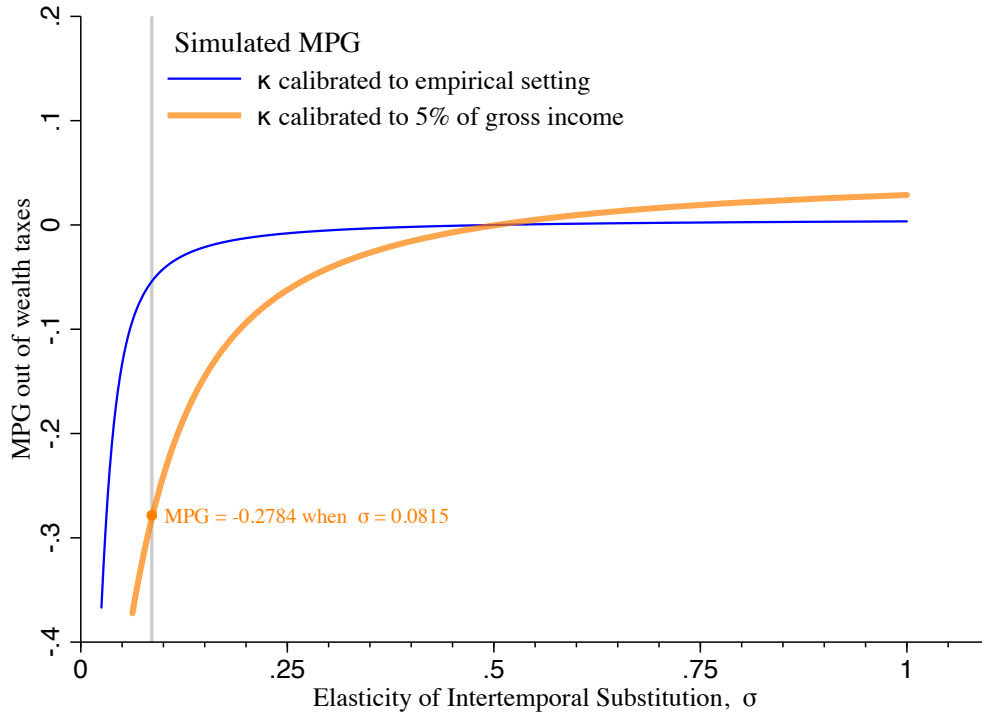
Our low implied uncompensated elasticity says that giving is inelastic, and thus that increasing the deduction rate (τ^g) would likely be an inefficient way to undo the negative effects of wealth taxes on giving.

Wealth taxation and the *uncompensated* own-price elasticity. We also investigate whether our model predicts different elasticities depending on the wealth tax regime. We investigate this by calculating the uncompensated elasticity in the presence of a 1% proportional (applied to all wealth) wealth tax. We find virtually no effect: the own-price elasticity decreases modestly in magnitude to -0.4810.

6.3 Model Application II: Wealth Taxation in a High-Giving Environment

Figure 8: CROWD-OUT FROM WEALTH TAXATION

Notes: This figure shows model-based treatment effects on charitable giving for combinations of the elasticity of intertemporal substitution (σ). The outcome variable (y-axis) the marginal propensity to give (MPG) out of wealth taxes. We plot simulated treatment effects for two scenarios. The first (thin blue line) shows the effect when the level of giving is calibrated to our empirical setting. The second (thick orange) line provides the effect when the level of giving is calibrated to equal 5% of gross income.



We found empirically that each additional NOK of wealth tax reduces charitable giving by about 0.012. This is rather small in terms of a crowd-out effect. If governments care about the wealth tax revenue *plus* the amount of charitable giving, the net effect of raising another NOK of wealth taxes is only reduced to 0.988 due to this crowd-out. However, this small effect is in part driven by the low *level* of giving in Norway. While intuition suggests that the large relative effect on giving implies that the crowd-out would be larger if the level of giving were larger, we proceed by quantifying this by using our model.

We recalibrate our model to an agent that gives an amount equal to 5% of gross income, the average fraction reported by List (2011) for the U.S. This is done through increasing the utility weight on giving, κ (see equation 56), and keeping $\sigma = 0.0815$ and $\varepsilon = 0.44$.³² We then calculate the implied level effect on g_t and divide this by the amount raised in wealth taxes, $1\% \cdot s_{t+1}$, where s_{t+1} includes behavioral responses to the tax increase. We illustrate our findings in Figure 8.

We find that the

$$\text{Model-implied } MPG \text{ out of wealth taxes in a high-giving environment} = -0.2784, \quad (31)$$

which represents a substantial effect. If governments want to compensate charitable organizations for such a large crowd-out effect, they effectively face a 28% “tax” on any wealth tax revenues.

We further numerically explore the importance of the EIS for this large crowd-out effect. Setting the EIS to be 1, which corresponds to log utility of consumption, we instead find that the

$$\text{Model-implied } MPG \text{ out of wealth taxes in a high-giving environment} \big|_{\sigma=1} = 0.0335 > 0. \quad (32)$$

This stark change in both magnitude and sign highlights the importance of pinning down the EIS to in order to model how giving responds to tax reforms that affect the after-tax rate-of-return.

While our rudimentary extrapolation to a high-giving setting by no means provides a perfect forecast of how a wealth tax would affect giving in a high-giving country such as the U.S., it nevertheless emphasizes the point that the link between charitable giving and wealth taxation may be a first order concern for policymakers.

³²One justification for assuming that the elasticity parameters, σ and ε stay fixed but κ changes is that we may think of κ as not being a deep structural parameter but rather something that depends on the extent of overall government services and redistribution.

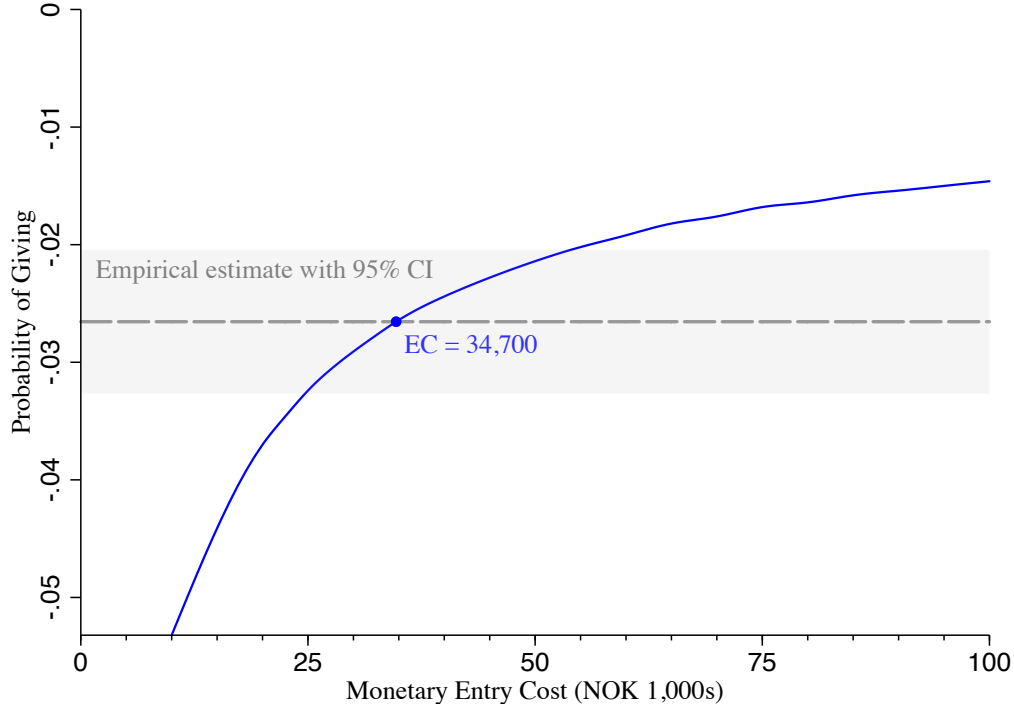
6.4 Calibrating Fixed Costs of Giving

Our baseline calibration exercise ignores fixed costs of giving. However, since we find an effect of wealth taxation on whether households give, we use this to infer the implied fixed cost. Since our empirical findings do not indicate any effect on exits, we focus solely on entry costs (as opposed to per-period participation costs that would also induce exits).

Acknowledging the presence of fixed costs in giving, we recalibrate the EIS, σ , to the effect on conditional giving, since conditional on giving, the optimal giving amount is not affected by fixed costs. We find that for the purpose of calibrating the EIS, it does not matter if we model extensive-margin variation directly. Our new estimate is 0.0785 and thus very similar to the baseline estimate of 0.0815.

Figure 9: INFERRING ENTRY COSTS INTO GIVING FROM QUASI-EXPERIMENTAL EVIDENCE ON PARTICIPATION

Notes: This figure shows model-based treatment effects on entry into charitable giving. Using a life-cycle model with endogenous giving that is calibrated to our empirical setting, we plot the effects of a 0.1 percentage point unexpected reduction in the after-tax rate (i.e., a 0.1% wealth tax). We find that a monetary entry cost (*MEC*) of NOK 34,700 replicates our empirical findings. Our calibration is described in more detail in Appendix B. The dashed gray line provides the empirical estimate of the implied effect of changing a proportional wealth tax rate, which is presented in column (6) of Table 2. The blue solid provides the simulated treatment effect assuming that $\varepsilon = 0.44$, corresponding to the bunching evidence in section 5.2, and an EIS (σ) of 0.083 (calibrated to the empirical effect on conditional giving).



Using an EIS estimate of 0.0785 and our existing estimate of $\varepsilon = 0.44$, we model how households respond on the extensive margin in the presence of entry costs. We describe the exercise in detail in Appendix B. For a given entry cost, EC , a household chooses to give if the utility of

giving exceeds the utility of not giving—and if not giving, receiving an endowment shock equal to EC . This implies that households only choose to give whenever the utility gain from giving exceeds the utility gain from optimally consuming EC over their life-cycle. This approach of modeling entry costs imply that agents behave as if the entry cost entered as a one-time utility cost in their utility function, but it allows us to model the entry cost as a monetary amount.

One issue with the standard approach of modeling per-period utility from giving as $\kappa \frac{g^{1-\epsilon}}{1-\epsilon}$ is that the marginal utility of giving tends to infinity as $g \rightarrow 0$. We circumvent this issue by using a first-order Taylor-series approximation around the optimal conditional giving amount to pin down the per-period utility losses from seizing to give.

The main complication with incorporating fixed costs of giving is that we need heterogeneity in order to simulate non-binary effects on participation. Following [Almunia, Guceri, Lockwood, and Scharf \(2020\)](#), we obtain this heterogeneity by modeling dispersion in the strength of the warm-glow of giving, i.e., κ . We first pick some monetary entry cost, EC , and then calibrate a log-normal κ distribution. This first-step calibration ensures that the correct fraction of households chooses to start giving (absent a wealth tax shock), and that the mean κ corresponds to the κ that reproduces the sample-average conditional amount of giving.

We graphically illustrate our second-step calibration in Figure 9. We find that a one-time entry cost of

$$EC = \text{NOK } 34,700 \text{ (USD } 5,800) \tag{33}$$

is needed to replicate our findings. This equals 43.06% of the present-value life-time giving of the marginal untreated entrant.

We know of no comparable entry cost estimates. While [Almunia, Guceri, Lockwood, and Scharf \(2020\)](#) estimate fixed costs, they estimate the per-period cost of *declaring* charitable giving on tax returns—not cost of participating in or entering into charitable giving.

We note that the fact that positive relationship between the probability of giving and entry costs in Figure 9 is not a result by itself. It is rather caused by the first-step calibration: keeping the fraction of households that chooses to give constant, a larger EC makes it less likely that the wealth-tax shock is large enough to push someone’s optimal giving amount above or below the threshold for participation.

Counterfactual experiment of removing entry costs. Our calibrated model with heterogeneity in κ allows us to investigate the effect of removing the cost of entering into giving. While this is a purely hypothetical exercise (since we do not know exactly what creates the entry costs), it is useful to quantify the overall importance of fixed costs in charitable giving. We find that the average ex-ante nonparticipant would enter and give NOK 575 (\$96) if entry costs were removed. This is about 52% of simulated unconditional mean amount of giving. Assuming a giving participation rate of around 30%, which we see in our data, removing entry costs would

increase total giving by around 58%.³³

7 Summary

The public finance literature has devoted significant attention to why and how the personal income tax system can be used to encourage charitable giving (see, e.g., [Saez 2004](#), [Diamond 2006](#), and [List 2011](#)). To what extent donation behavior could be influenced by capital taxation, in the form of a wealth, capital income or capital gains tax, has been largely neglected. New evidence thus seems prudent in light of the surging interest for using capital taxation, and wealth taxes in particular, as a policy instrument to address economic inequality (see, e.g., [Saez and Zucman 2019b](#)).

To our knowledge, this paper is the first to present empirical evidence on how capital taxation, or wealth taxation in particular, affects charitable giving. Our finding of a negative effect is surprising in the sense that standard preference parameters (e.g., an elasticity of intertemporal substitution of 1) imply a positive effect. Our results thus indicate that the redistributive effects of wealth taxation may be offset by reduced giving.

We exploit our empirical setting to provide novel quasi-experimental evidence on the presence of fixed costs in giving. Our findings suggest that removing entry barriers may substantially increase giving.

We also provide new evidence on how responsive giving is to its after-tax own price. Our findings suggest that actual giving may be considerably less responsive to tax incentives than reported giving. Our heterogeneity analyses show that the sensitivity of giving is not higher for higher-income and wealthier individuals. However, our data demonstrates substantially heterogeneity with respect to the type of charitable giving. The presence of elasticity heterogeneity with respect to the type of giving implies that, even if a social planner values different types of giving equally, efficiency losses may arise from using uniform tax incentives.

In order to more concretely guide optimal taxation, we provide the key preference parameters that can explain our empirical findings. These may be more useful than our reduced-form findings as they do not explicitly depend on the peculiarities of the tax code (e.g., the progressivity) or the characteristics of the taxpayers. Importantly, they may be used to model the effects of a wide range of tax reforms, such as changing the direct tax incentives for giving, changing the tax rate on capital income, or introducing a progressive wealth tax.

While we find that the crowd-out effect of giving is modest in a NOK for NOK sense, our structural parameters imply that this crowd-out may be economically large in a setting where the ex-ante level of giving is higher, such as the U.S. or the U.K. While crowding out giving by itself reduces the potential benefits of wealth taxation, this crowding-out cannot, structurally, occur in a vacuum. This is because the crowd-out is driven by a low elasticity of intertemporal

³³The unconditional mean (and thus also the total amount of giving) would increase by $0.5222 \times (1 - 0.3) = 36.55\%$.

substitution. A low EIS in standard frameworks generally imply low efficiency costs of capital taxation. Our findings suggest that optimal tax models that ignore giving may overstate the attractiveness of more comprehensive capital taxation when the EIS is low and the ex-ante level of giving is high.

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A Model and Proofs

Consider the following optimization problem:

$$\max_{\{c_t, s_{t+1}, g_t\}} \sum_{t=0}^T \beta^t \left[\frac{c_t^{1-1/\sigma}}{1-1/\sigma} + \kappa \frac{g_t^{1-1/\varepsilon}}{1-1/\varepsilon} \right], \quad (34)$$

$$\text{such that} \quad c_t = w_t - p_t g_t - s_{t+1} + s_t R_t, \quad (35)$$

where c_t , s_{t+1} , and g_t is the consumption, saving, and charitable giving undertaken in year t . β is the temporal discount factor, σ is the elasticity of intertemporal substitution (the EIS, for consumption) and ε is the giving elasticity. w_t is exogenous income. w_0 may be considered the initial wealth of the agent. p_t is the price of a unit of giving. R_t is the gross after-tax rate of return, net of any wealth taxes. Agents start their optimization problem at $t = 0$ and end their life cycle at $t = T$.

In applying this model to our empirical setting, it will be convenient to assume that agents start their optimization problem at some other time $t > 0$, such as when they face a wealth tax shock. In this case, we can replace time 0 with t and replace T with $T - t$.

A.1 Solving the life-cycle optimization problem.

s_{t+1} -FOC may be written the following ways:

$$c_t^{-1/\sigma} = \beta R_{t+1} c_{t+1}^{-1/\sigma} \Leftrightarrow c_{t+1} = (\beta R_{t+1})^\sigma c_t \quad (36)$$

s_t -FOC iteration yields:

$$c_t = \beta^{\sigma t} \Pi_{s=0}^t R_s^\sigma c_0 \quad (37)$$

g_t -FOC may be written the following ways:

$$c_t^{-1/\sigma} p_t = \kappa g_t^{-1/\varepsilon} \Leftrightarrow c_t = \left(\frac{p_t}{\kappa} \right)^\sigma g_t^{\sigma/\varepsilon} \Leftrightarrow g_t = \left(\frac{\kappa}{p_t} \right)^\varepsilon c_t^{\varepsilon/\sigma} \quad (38)$$

Substitute in the s_t -FOC iteration into the g_t -FOC, and then the g_0 -FOC:

$$g_t = \left(\frac{\kappa}{p_t} \right)^\varepsilon c_t^{\varepsilon/\sigma} = \left(\frac{\kappa}{p_t} \right)^\varepsilon \beta^{\varepsilon t} \Pi_{s=0}^t R_s^\varepsilon c_0^{\varepsilon/\sigma} = \left(\frac{\kappa}{p_t} \right)^\varepsilon \beta^{\varepsilon t} \Pi_{s=0}^t R_s^\varepsilon \left(\frac{p_0}{\kappa} \right)^\varepsilon g_0 \quad (39)$$

$$= \left(\frac{p_0}{p_t} \right)^\varepsilon \beta^{\varepsilon t} \Pi_{s=0}^t R_s^\varepsilon g_0 \quad (40)$$

Budget constraint iteration yields the present-value condition:

$$\sum_{t=0}^T c_t \Pi_{s=0}^t R_s^{-1} = \sum_{t=0}^T w_t \Pi_{s=0}^t R_s^{-1} - \sum_{t=0}^T p_t g_t \Pi_{s=0}^t R_s^{-1}. \quad (41)$$

Now we substitute in the iterated FOC expressions for g_t and c_t .

$$c_0 \sum_{t=0}^T \beta^{\sigma t} \Pi_{s=0}^t R_s^{\sigma-1} = \sum_{t=0}^T w_t \Pi_{s=0}^t R_s^{-1} - g_0 \sum_{t=0}^T p_t^{1-\varepsilon} p_0^\varepsilon \beta^{\varepsilon t} \Pi_{s=0}^t R_s^{\varepsilon-1}. \quad (42)$$

We then substitute in the g_0 -FOC:

$$g_0^{\sigma/\varepsilon} \left(\frac{p_0}{\kappa} \right)^\sigma \sum_{t=0}^T \beta^{\sigma t} \Pi_{s=0}^t R_s^{\sigma-1} = \sum_{t=0}^T w_t \Pi_{s=0}^t R_s^{-1} - g_0 \sum_{t=0}^T p_t^{1-\varepsilon} p_0^\varepsilon \beta^{\varepsilon t} \Pi_{s=0}^t R_s^{\varepsilon-1}. \quad (43)$$

This expression provides the initial level of giving, g_0 , that solves the optimization problem. All other endogenous variables then follow from the FOCs.

Simplified problem. Now, we simplify and set $R_t = R$ and $p_t = 1 \forall t$.

$$g_0^{\sigma/\varepsilon} \left(\frac{1}{\kappa} \right)^\sigma \frac{(\beta^\sigma R^{\sigma-1})^{T+1} - 1}{\beta^\sigma R^{\sigma-1} - 1} = \sum_{t=0}^T w_t R^{-t} - g_0 \frac{(\beta^\varepsilon R^{\varepsilon-1})^{T+1} - 1}{\beta^\varepsilon R^{\varepsilon-1} - 1}. \quad (44)$$

Note that we may substitute in s for 0, and also replace T with $T - s$ to determine the level of giving for some point in time s . The in-text proposition is similar to the simplified expression above but does not assume that p_t is constant over time.

Simplified, infinite-horizon, when $T = \infty$, assuming $\beta^\varepsilon R^{\varepsilon-1} < 1$ and $\beta^\sigma R^{\sigma-1} < 1$, yields condition:

$$g_0^{\sigma/\varepsilon} \frac{1}{1 - \beta^\sigma R^{\sigma-1}} \left(\frac{1}{\kappa} \right)^\sigma = \sum_{t=0}^T w_t R^{-t} - g_0 \frac{1}{1 - \beta^\varepsilon R^{\varepsilon-1}} \quad (45)$$

$$g_0^{\sigma/\varepsilon} \left(\frac{1}{\kappa} \right)^\sigma = (1 - \beta^\sigma R^{\sigma-1}) \sum_{t=0}^T w_t R^{-t} - g_0 \frac{1 - \beta^\sigma R^{\sigma-1}}{1 - \beta^\varepsilon R^{\varepsilon-1}} \quad (46)$$

Differentiate

$$\frac{\sigma}{\varepsilon} g_0^{\sigma/\varepsilon-1} \left(\frac{1}{\kappa} \right)^\sigma dg_0 = -(\sigma - 1) \beta^\sigma R^{\sigma-2} \sum_{t=0}^T w_t R^{-t} dR - (1 - \beta^\sigma R^{\sigma-1}) \sum_{t=0}^T t w_t R^{-t-1} dR \quad (47)$$

$$- dg_0 \frac{1 - \beta^\sigma R^{\sigma-1}}{1 - \beta^\varepsilon R^{\varepsilon-1}} \quad (48)$$

$$+ g_0 \frac{(\varepsilon - \sigma) \beta^{\sigma+\varepsilon} - (\varepsilon - 1) \beta^\varepsilon R^{\sigma+1} + (\sigma - 1) \beta^\sigma R^{\sigma+1}}{R (\beta^\varepsilon R^\varepsilon - R)^2} dR \quad (49)$$

Reorder terms, particularly the last term, and get

$$\left(\frac{\sigma}{\varepsilon} g_0^{\sigma/\varepsilon-1} \left(\frac{1}{\kappa} \right)^\sigma + \frac{1 - \beta^\sigma R^{\sigma-1}}{1 - \beta^\varepsilon R^{\varepsilon-1}} \right) \frac{dg_0}{dR} = -(\sigma - 1) \beta^\sigma R^{\sigma-1} \sum_{t=0}^T w_t R^{-t} \quad (50)$$

$$\begin{aligned} & - (1 - \beta^\sigma R^{\sigma-1}) \sum_{t=0}^T t w_t R^{-t-1} \\ & - g_0 \left[\frac{(\varepsilon - 1) \beta^\varepsilon R^{\varepsilon-2} (1 - \beta^\sigma R^{\sigma-1})}{(1 - \beta^\varepsilon R^{\varepsilon-1})^2} - \frac{(\sigma - 1) \beta^\sigma R^{\sigma-2}}{1 - \beta^\varepsilon R^{\varepsilon-1}} \right] \end{aligned} \quad (51)$$

Now consider this differential equation when $\beta R = 1$:

$$\left(\frac{\sigma}{\varepsilon} g_0^{\sigma/\varepsilon-1} \left(\frac{1}{\kappa} \right)^\sigma + 1 \right) \frac{dg_0}{dR} = -(\sigma - 1) R^{-1} \sum_{t=0}^T w_t R^{-t} \quad (52)$$

$$\begin{aligned} & - (1 - R^{-1}) \sum_{t=0}^T t w_t R^{-t-1} \\ & - g_0 \left[\frac{(\varepsilon - 1) R^{-2} (1 - R^{-1})}{(1 - R^{-1})^2} - \frac{(\sigma - 1) R^{-2}}{1 - R^{-1}} \right] \end{aligned} \quad (53)$$

After some reordering and cancelling out, we get

$$\left(\frac{\sigma}{\varepsilon} \kappa^{-\sigma} g_0^{\sigma/\varepsilon-1} + 1 \right) \frac{dg_0}{dR} = \underbrace{-(\sigma - 1) \sum_{t=0}^T w_t R^{-t-1}}_{\text{Income v. substitution}} - \underbrace{(1 - R^{-1}) \sum_{t=0}^T t w_t R^{-t-1}}_{\text{Human wealth effect}} - \underbrace{\frac{\varepsilon - \sigma}{R^2 (1 - R^{-1})} g_0}_{\text{Elasticity adjustment}}. \quad (54)$$

The first two terms on the right-hand side are driven by consumption adjustments across periods. The third term says that if giving is more elastic than consumption ($\varepsilon > \sigma$), then current giving drops by even more than what is implied by the first two consumption-related terms. Note that from the intratemporal g_0 -FOC, we get that $\kappa^{-\sigma} g_0^{\sigma/\varepsilon} = c_0$. Hence, we can also write

$$\left(\frac{\sigma}{\varepsilon} \frac{c_0}{g_0} + 1 \right) \frac{dg_0}{dR} = \underbrace{-(\sigma - 1) \sum_{t=0}^T w_t R^{-t-1}}_{\text{Income v. substitution}} - \underbrace{(1 - R^{-1}) \sum_{t=0}^T t w_t R^{-t-1}}_{\text{Human wealth effect}} - \underbrace{\frac{\varepsilon - \sigma}{R^2 (1 - R^{-1})} g_0}_{\text{Elasticity adjustment}}. \quad (55)$$

Note that left-hand-side term in parentheses may also be rewritten as $\frac{\sigma c_t + \varepsilon g_t}{\varepsilon g_t}$.

B Calibration

B.1 Baseline calibration, without entry costs

We use Proposition 1 for calibration. We first set $\varepsilon = 0.44$ in line with our bunching estimate and set $\beta = 0.98$. We set $R^{base} = \beta^{-1}$. This implies $R \approx 1.02$, which is in line with the low interest rates in Norway during this time period. We then select a σ , and compute a value for κ (see below). We enter empirical values for $\{w_s\}_{s=t}^T$ and T (also see below), and solve for the g_t

that would arise under two values of R : R^{base} and $R^{shocked}$, and compute the log difference. We do this for many σ values, and search for the one that gives the closest match to our empirical estimates on the implied effect of a proportional wealth tax rate change.

We choose κ from a level calibration. Given, σ and ε , we use the first-order condition that $\kappa = g_t^{1/\varepsilon} c_t^{-1/\sigma} p_t$ to assign κ . We assume that c_t equals 93% of after-tax income, consistent with the net saving rate of 6–8% found by [Fagereng, Holm, Moll, and Natvik \(2019\)](#) for wealthier households. Gross income is set to the mean amount of gross income, 0.77 MNOK. We assume a tax rate of 40%. Hence, $\hat{c}_t = 0.93 \cdot 0.77 \cdot (1 - 0.40) = 0.43$. The mean amount of giving in our sample is 5,225 conditional on giving (see Table [A.1](#)). Multiplying by the probability of giving, 0.3, we set $\hat{g}_t = 1,568$. We set $p_t = 0.72$, which is the after-tax price in 2012. Note these values, \hat{g}_t and \hat{c}_t , are only used to calibrate κ , and do not otherwise affect the calibration exercise.

$$\kappa \equiv \hat{c}^{-1/\sigma} \hat{g}^{1/\varepsilon} = 0.43^{-1/\sigma} \cdot 0.001568^{1/\varepsilon} \cdot 0.72. \quad (56)$$

w_t is set to $0.77 \cdot (1 - 0.40) + 4.23$. This is the average yearly income and wealth for households in our sample as of 2012 (Table [A.1](#)). The average person in our sample is 59. we set $t = 58$ and assume that starts declining linearly at $t = 62$ and reaches 50% of the original income at age 66 based an assumed retirement at 66 and a 50% income replacement rate. We set the duration of the life cycle to be $T = 85$.

When calculating simulated log-treatment effects, we shift giving amounts by NOK 1,000 prior to taking logs in order to match the empirical specification, which employs this shifting argument to accommodate zeros and avoid a large influence from trivial level changes.

B.2 Calibration, with entry costs

Intensive-margin calibration. We still use our empirical bunching-design estimate to inform ε . We calibrate σ to intensive-margin giving responses, using the same procedure as when there were no entry costs.

We calibrate the warm-glow strength as before:

$$\kappa^c \equiv \hat{c}^{-1/\sigma} \hat{g}^{1/\varepsilon} \cdot p_t, \quad (57)$$

where $\hat{c} = 0.93 \cdot 0.77 \cdot (1 - 0.4)$, $\hat{g} = 0.005225$ (both denominated in MNOK), and p_t equals the 2012 after-tax price of giving of 0.72. Our calibration exercise thus differs from the baseline (no entry cost) by using a higher (conditional) \hat{g} . We obtain an EIS estimate of 0.0696.

Extensive-margin calibration

In order to calibrate entry costs, we first choose some entry cost, EC , and iterate until we find the EC that produces similar treatment effects on extensive-margin giving as we find empirically.

Households will choose between giving (as if there were no entry cost) or not giving, in which case they receive EC as compensation and can use this to optimally increase consumption.

One issue that arises is that the utility form yields negative infinity utility once $g_t \rightarrow 0$. We circumvent this INADA condition by assuming that the utility loss from not giving the optimal amount (conditional on giving) is captured by a *first-order Taylor series approximation* around the optimal conditional giving amount.

Our exercise then defines two identical households, where one is the control and one is the treated household who sees a reduction in R . We then examine the effect on participation, $\mathbb{1}[g_t > 0]$.

In order to simulate a non-binary effect on participation, we need heterogeneity (i.e., only a fraction of shocked households should start giving). We choose to model this as heterogeneity in the strength of the warm-glow utility of giving, that is, κ . This approach has precedent in [Almunia et al. \(2020\)](#). We assume that κ is distributed according to the log-normal distribution, $\ln(\psi_{mean}, \psi_{sdev})$, with CDF $\Psi(\kappa)$. We choose μ and σ such that (i) the probability of giving matches the share of ex-ante non-givers who choose to give during 2013–2018. In the model, we solve for κ^* , above which a household chooses to start giving in the presence of entry costs. (ii) We also require the mean κ to correspond to the κ^c that produces the sample-mean conditional amount of giving. This means that we have two equations in two unknowns that we can use to solve for the log-normal distribution parameters, ψ_{mean}, ψ_{sdev} :

$$\Psi(\kappa^*; \psi_{mean}; \psi_{sdev}) \equiv 12\%, \quad (58)$$

$$\mathbb{E}[\kappa] = \kappa^c. \quad (59)$$

Since these cannot be solved analytically, we obtain our estimates for ψ_{mean}, ψ_{sdev} by minimizing the sum of squared differences between the simulated moments and the moments we want to match.

We discretize the heterogeneity by modeling the behavior of $N_h = 4,999$ households. Each household draws a κ from a unique quantile of the κ distribution. We denote this as $h = 1, 2, \dots, N_h$. A given household's κ is then

$$\kappa_h = \Psi^{-1} \left(\frac{h}{N_h + 1} \right). \quad (60)$$

Our calibration exercise then considers the mean treatment effect (by comparing treatment and control) across households, h . We graphically show which EC provide a simulated mean treatment that replicate our empirical findings.

This approach provides a EC and κ distribution that replicates the observed probability of giving during 2013–2018 among those who did give during 2006–2012. However, it does not replicate the fact that these households did not give ex-ante. To replicate this dynamic one

would need some kind of shock to either preferences, endowment or other state variables. For example, we could (trivially) replicate ex-ante non-participation by assuming that prior to year t , all the households had very low κ_h , but in year t everyone draws a new κ_h according to our calibrated log-normal distribution.

C Supplementary Figures and Tables

Table A.1: SUMMARY STATISTICS BY SECONDARY HOME OWNERSHIP

Notes: This table provides household-level summary statistics for two subsamples of our main analysis sample as of 2012. Net wealth equals taxable wealth gross of any valuation discounts. Amounts in Norwegian kroner (NOK) may be divided by 6 to obtain an approximate USD amount as of 2012. Note that the statistics for charitable giving are done at the household level in this table. In the main summary statistics table, summary statistics for giving are computed at the individual level since this is the unit of analysis when using the bunching framework.

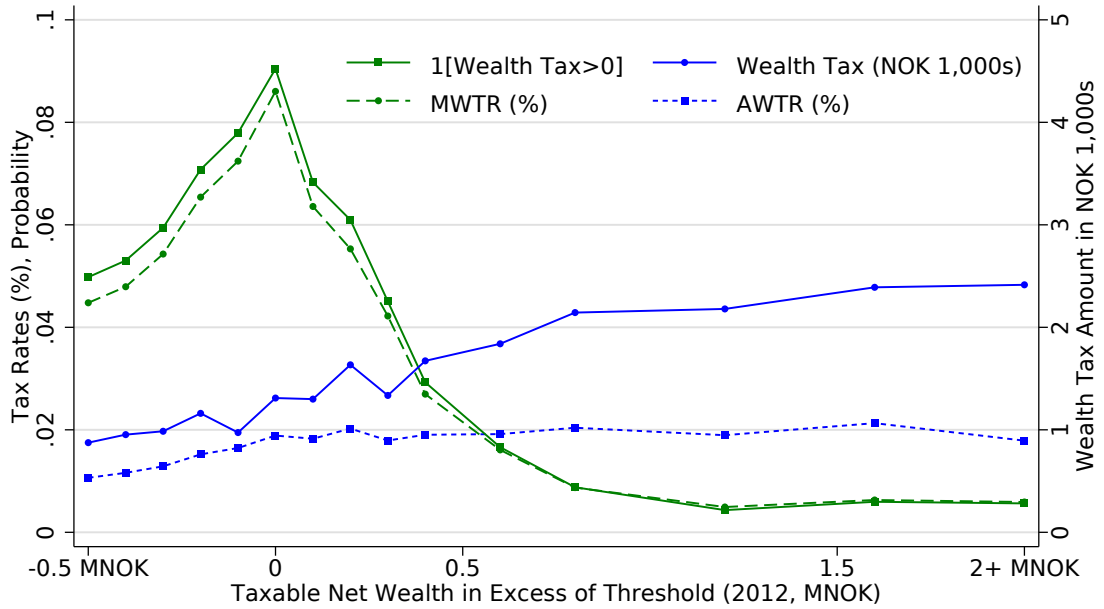
	$MVHS_{h,2012} = 0$				$MVHS_{h,2012} > 0$			
	mean	p25	p50	p75	mean	p25	p50	p75
$1[\text{Giving}_{h,2012} > 0]$	0.29				0.30			
$\text{Giving}_{h,2012}$ if > 0	4,906	1,200	2,880	5,020	5,225	1,400	2,880	5,400
$1[\text{wtax}_{h,2012} > 0]$	0.52				0.61			
$\text{wtax}_{h,2012}$ if > 0	9,298	2,621	6,416	13,089	11,914	3,644	8,746	17,110
$MVH_{h,2012}$, MNOK	2.87	1.70	2.44	3.52	4.23	1.95	3.52	5.70
$MVHS_{h,2012}$, MNOK	0.00	0.00	0.00	0.00	1.95	0.92	1.56	2.51
Net Wealth $_{h,2012}$, MNOK	3.30	2.03	2.93	4.14	4.23	2.23	3.74	5.62
Age of household head $_{h,2012}$	62	50	63	74	58	48	59	68
Gross income $_{h,2012}$, MNOK	0.61	0.30	0.46	0.75	0.77	0.36	0.60	1.01
Number of adults $_h$	1.37				1.49			

Table A.2: SUMMARY STATISTICS FOR SECONDARY HOME OWNERS
BY WHETHER THEY GAVE DURING 2006–2012

Notes: This table provides household-level summary statistics for two subsamples of treated households (those with $MVHS_{h,2012} > 0$) as of 2012. Nonparticipants gave zero each year during 2006–2012. Participants gave a strictly positive amount each year. Net wealth equals taxable wealth gross of any valuation discounts. Amounts in Norwegian kroner (NOK) may be divided by 6 to obtain an approximate USD amount as of 2012.

	Nonparticipants ($G_{h,pre} = 0$)				Participants ($G_{h,pre} > 0$)			
	mean	p25	p50	p75	mean	p25	p50	p75
$1[Giving_{h,t} > 0]$ if $t > 2012$	0.12				0.94			
$Giving_{h,2012}$					6,496	2,640	3,440	6,660
$1[wtax_{h,2012} > 0]$	0.58				0.67			
$wtax_{h,2012}$ if > 0	10,949	3,273	7,957	15,692	14,066	4,758	10,722	20,387
$MVH_{h,2012}$, MNOK	3.77	1.69	3.06	5.05	5.27	2.97	4.67	6.92
$MVHS_{h,2012}$, MNOK	1.85	0.88	1.47	2.37	2.12	0.99	1.75	2.72
Net Wealth $_{h,2012}$, MNOK	3.83	1.96	3.32	5.08	5.24	3.26	4.81	6.77
Age $_{h,2012}$	57	47	58	68	61	53	61	68
Gross income $_{h,2012}$, MNOK	0.66	0.32	0.51	0.85	1.05	0.56	0.90	1.33
Number of adults $_h$	1.40	1.00	1.00	2.00	1.73	1.00	2.00	2.00

Figure A.1: FIRST-STAGE HETEROGENEITY



This figure shows our first-stage regression coefficients: that is, how the effect of owning more secondary housing wealth (i.e., more $MVHS_{h,2012}$) affected wealth tax exposure for households with different levels of taxable wealth in 2012 ($TNW_{h,2012}$). The point estimates correspond to the coefficients in Appendix Table A.4.

Figure A.2: DYNAMIC REDUCED-FORM EFFECTS ON LEVEL OF GIVING AND WEALTH TAXES

Notes: The point estimates come from estimating the equation (14), with giving as the left-hand-side variable and simplifying $f(\cdot)$ to be a year-specific linear function of $MVHS_{h,2012}$ alone. Giving is measured using the “long panel” of charitable giving deductions from tax returns. The horizontal lines provide 95% confidence intervals

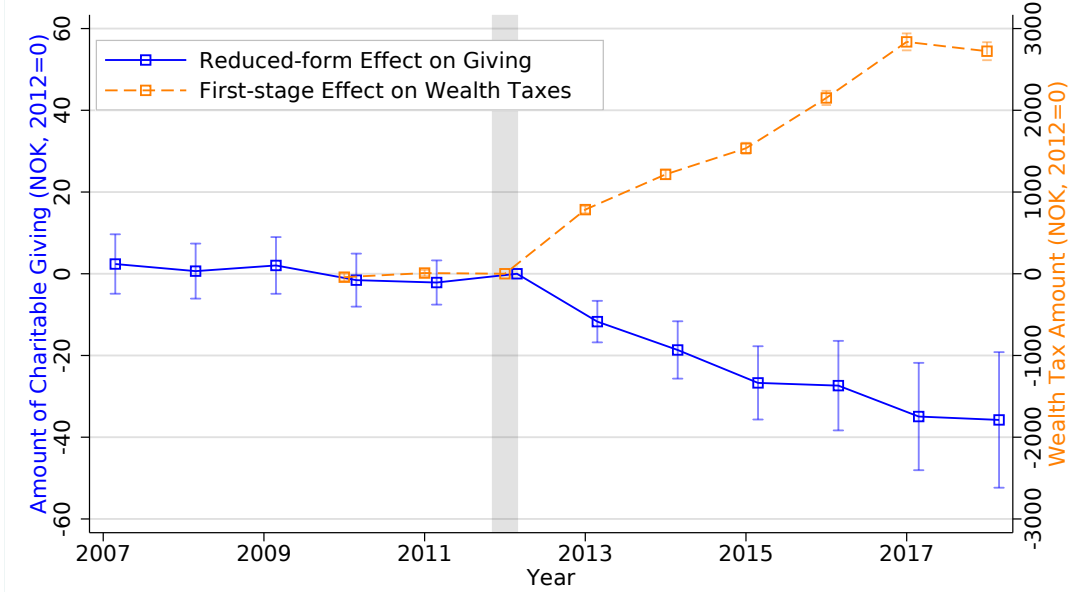


Figure A.3: REDUCED-FORM EFFECTS ON GIVING USING BOTH DATA SOURCES ON GIVING

Notes: The long panel comes from tax returns and reflects actual amount deducted from the income tax base. The short panel is the total amount of giving that was directly reported by the charitable organizations. The point estimates come from estimating equation (14), simplifying $f(\cdot)$ to be a year-specific linear function of $MVHS_{h,2012}$ alone.

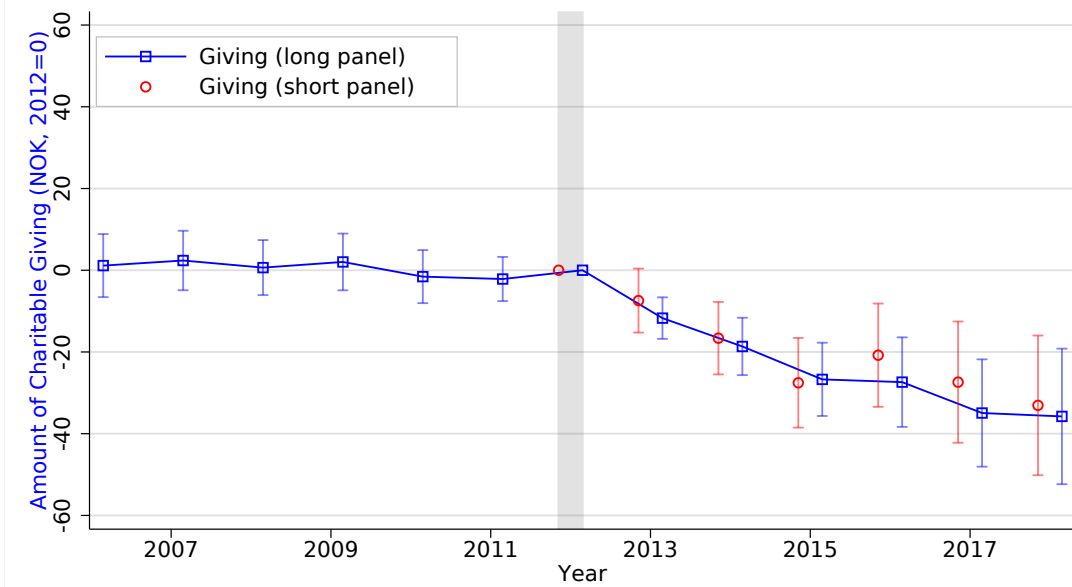


Table A.3: WEALTH TAXATION AND CHARITABLE GIVING: REDUCED-FORM REGRESSION ESTIMATES

Notes: The table provides the underlying **reduced-form** coefficients on instruments from estimating the system of equations in (14)-(15). Each b corresponds to a binned range of $TNW_{h,2012} - T_{12}$. Note that the underlying regression specification also includes controls for the b_j indicator variables, and thus only uses the interaction between ex-ante taxable wealth and initial $MVHS$ for identification. Standard errors are clustered at the household level. One, two, and three stars indicate statistical significance at the 10%, 5%, and 1% levels.

	(1)	(2)	(3)	(4)	(5)	(6)
	Amount		Adj. log of Giving		$\mathbb{1}[G_{h,t} > 0]$	
					(int. marg)	(ext. marg)
$1[b=-0.5M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-3.974 (21.403)	-3.974 (21.403)	-0.008** (0.003)	-0.008** (0.003)	-0.001 (0.006)	-0.008*** (0.002)
$1[b=-0.4M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-30.571 (23.319)	-30.571 (23.319)	-0.008** (0.003)	-0.008** (0.003)	-0.010 (0.006)	-0.002 (0.002)
$1[b=-0.3M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-16.097 (16.603)	-16.097 (16.603)	-0.004 (0.003)	-0.004 (0.003)	-0.006 (0.006)	-0.004* (0.002)
$1[b=-0.2M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-10.766 (17.829)	-10.766 (17.829)	-0.006* (0.003)	-0.006* (0.003)	0.004 (0.006)	-0.008*** (0.002)
$1[b=-0.1M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	5.816 (17.954)	5.816 (17.954)	-0.001 (0.003)	-0.001 (0.003)	0.004 (0.006)	-0.006** (0.002)
$1[b= 0.0M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-14.155 (18.562)	-14.155 (18.562)	-0.006* (0.003)	-0.006* (0.003)	-0.007 (0.006)	-0.003 (0.002)
$1[b= 0.1M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-59.328** (22.199)	-59.328** (22.199)	-0.010*** (0.003)	-0.010*** (0.003)	-0.008 (0.007)	-0.011*** (0.002)
$1[b= 0.2M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-18.478 (21.550)	-18.478 (21.550)	-0.007* (0.003)	-0.007* (0.003)	-0.000 (0.007)	-0.006** (0.002)
$1[b= 0.3M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-7.962 (19.955)	-7.962 (19.955)	-0.007* (0.003)	-0.007* (0.003)	-0.011 (0.007)	-0.007** (0.002)
$1[b= 0.4M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-14.326 (16.643)	-14.326 (16.643)	-0.005 (0.002)	-0.005 (0.002)	-0.008 (0.005)	-0.006*** (0.002)
$1[b= 0.6M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-22.321 (18.710)	-22.321 (18.710)	-0.005 (0.003)	-0.005 (0.003)	0.001 (0.005)	-0.005* (0.002)
$1[b= 0.8M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-19.500 (12.999)	-19.500 (12.999)	-0.006** (0.002)	-0.006** (0.002)	-0.005 (0.004)	-0.004* (0.002)
$1[b= 1.2M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-48.403** (16.653)	-48.403** (16.653)	-0.007*** (0.002)	-0.007*** (0.002)	-0.009 (0.005)	-0.005** (0.002)
$1[b= 1.6M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-30.594 (18.392)	-30.594 (18.392)	-0.002 (0.002)	-0.002 (0.002)	-0.013** (0.005)	-0.003 (0.002)
$1[b= 2.0M] \times MVHS_{h,12} \times \mathbb{1}[t > 2012]$	-22.053 (14.605)	-22.053 (14.605)	-0.004* (0.002)	-0.004* (0.002)	-0.001 (0.003)	-0.006*** (0.001)

Table A.4: WEALTH TAXATION AND CHARITABLE GIVING: FIRST-STAGE REGRESSION ESTIMATES

Notes: The table provides the underlying **first-stage** coefficients on instruments from estimating the system of equations in (14)-(15) corresponding to columns (1)-(4) in Table 2. Each b corresponds to a binned range of $TNW_{h,2012} - T_{12}$. Note that the underlying regression specification also includes controls for the b_j indicator variables, and thus only obtains identifying first-stage variation from the interaction between ex-ante taxable wealth and initial $MVHS$. Standard errors are clustered at the household level. One, two, and three stars indicate statistical significance at the 10%, 5%, and 1% levels.

	(1) $1[\text{wtax}_{h,t} > 0]$	(2) $\text{wtax}_{h,t}$	(3) $\text{MWTR}_{h,t}$	(4) $\text{AWTR}_{h,t}$
$1[b=-0.5M] \times MVHS_{h,12} \times 1[t > 2012]$	0.04980*** (0.00204)	874.73955*** (65.51051)	0.00045*** (0.00002)	0.00011*** (0.00001)
$1[b=-0.4M] \times MVHS_{h,12} \times 1[t > 2012]$	0.05303*** (0.00219)	953.16318*** (67.78993)	0.00048*** (0.00002)	0.00012*** (0.00001)
$1[b=-0.3M] \times MVHS_{h,12} \times 1[t > 2012]$	0.05944*** (0.00231)	985.71233*** (80.83699)	0.00054*** (0.00002)	0.00013*** (0.00001)
$1[b=-0.2M] \times MVHS_{h,12} \times 1[t > 2012]$	0.07080*** (0.00256)	1160.72179*** (89.39980)	0.00065*** (0.00002)	0.00015*** (0.00001)
$1[b=-0.1M] \times MVHS_{h,12} \times 1[t > 2012]$	0.07793*** (0.00281)	973.36174*** (78.10608)	0.00072*** (0.00003)	0.00016*** (0.00001)
$1[b= 0.0M] \times MVHS_{h,12} \times 1[t > 2012]$	0.09045*** (0.00340)	1310.09234*** (92.70186)	0.00086*** (0.00003)	0.00019*** (0.00001)
$1[b= 0.1M] \times MVHS_{h,12} \times 1[t > 2012]$	0.06830*** (0.00344)	1299.60026*** (88.07988)	0.00064*** (0.00003)	0.00018*** (0.00001)
$1[b= 0.2M] \times MVHS_{h,12} \times 1[t > 2012]$	0.06096*** (0.00317)	1634.19799*** (94.93024)	0.00055*** (0.00003)	0.00020*** (0.00001)
$1[b= 0.3M] \times MVHS_{h,12} \times 1[t > 2012]$	0.04507*** (0.00320)	1335.55172*** (94.91412)	0.00042*** (0.00003)	0.00018*** (0.00001)
$1[b= 0.4M] \times MVHS_{h,12} \times 1[t > 2012]$	0.02926*** (0.00225)	1672.77469*** (78.45694)	0.00027*** (0.00002)	0.00019*** (0.00001)
$1[b= 0.6M] \times MVHS_{h,12} \times 1[t > 2012]$	0.01661*** (0.00227)	1839.41726*** (94.96646)	0.00016*** (0.00002)	0.00019*** (0.00001)
$1[b= 0.8M] \times MVHS_{h,12} \times 1[t > 2012]$	0.00879*** (0.00160)	2143.87031*** (87.35073)	0.00009*** (0.00002)	0.00020*** (0.00001)
$1[b= 1.2M] \times MVHS_{h,12} \times 1[t > 2012]$	0.00433** (0.00162)	2179.56331*** (112.62815)	0.00005** (0.00002)	0.00019*** (0.00001)
$1[b= 1.6M] \times MVHS_{h,12} \times 1[t > 2012]$	0.00596*** (0.00161)	2390.55274*** (131.17084)	0.00006*** (0.00002)	0.00021*** (0.00001)
$1[b= 2.0M] \times MVHS_{h,12} \times 1[t > 2012]$	0.00564*** (0.00088)	2414.78097*** (120.22911)	0.00006*** (0.00001)	0.00018*** (0.00000)

Table A.5: WEALTH TAXATION AND (NON)RELIGIOUS GIVING

Notes: This table considers different types of charitable giving. Religious giving consists of donations to religious organizations such as local or national churches, theological institutions, and missionary organizations. International giving consists of donations to organizations with an international focus, such as humanitarian, aid, human rights, and environmental. The table provides the IV (second-stage) coefficients from estimating the system of equations in (14)-(15). *AWTR* is the average wealth tax rate, defined as the amount of wealth taxes paid divided by net wealth (net wealth equals *TNW* absent valuation discounts). *MWTR* is the marginal wealth tax rate and equals the nominal wealth tax rate if the household is above the wealth tax threshold and zero otherwise. In the log-giving specifications, giving is shifted by an inflation-adjusted NOK 1,000 to accommodate zeros (and thus entries or exits) and limit influence of outliers. Standard errors are clustered at the household level. One, two, and three stars indicate statistical significance at the 5%, 1%, and 0.1% levels.

	Religious Adj. log($G_{h,t}^{rel}$) (1)	Nonreligious Adj. log($G_{h,t}^{nonrel}$) (2)	International Adj. log($G_{h,t}^{intl}$) (3)	Domestic Adj. log($G_{h,t}^{dom}$)
<u>Instrumented variables</u>				
AWTR	-29.91 (25.10)	-19.55 (10.76)	-24.29* (11.67)	2.52 (14.68)
MWTR	-16.77 (9.92)	-13.42** (4.74)	-14.11** (5.22)	-7.09 (6.02)
Implied effect of proportional tax change	-46.68* (20.99)	-32.97*** (8.12)	-38.41*** (8.69)	-4.57 (11.77)
Sample restriction (based on giving in 2012)	Religious > 0	Nonreligious > 0	International > 0	Domestic > 0
rk- <i>F</i> -statistic	24.06	69.67	58.27	23.06
N	265,903	1,046,692	890,086	302,470
Additional summary statistics (full sample, 2012–18)				
Share who give to type of charity	0.06	0.26	0.22	0.07
Conditional mean giving	10,758	3,325	3,449	1,394

Table A.6: WEALTH TAXATION AND CHARITABLE GIVING:
IV GROWTH-RATE REGRESSIONS

Notes: The table provides the key coefficients from estimating the system of equations in (14)-(15). *AWTR* is the average wealth tax rate, defined as the amount of wealth taxes paid divided by *TNW*. *MWTR* is the marginal wealth tax rate and equals the nominal wealth tax rate if the household is above the wealth tax threshold and zero otherwise. See Appendix Tables A.3 and A.4 for the underlying reduced-form and first-stage estimates. Standard errors are clustered at the household level. One, two, and three stars indicate statistical significance at the 10%, 5%, and 1% levels.

	(1)	(2)
	Growth rate	
	$\Delta \log G_{h,t+1}$	$\Delta \log G_{h,t+1}$
MWTR	1.6156 (1.3491)	0.9066 (1.7649)
AWTR		2.7062 (4.3689)
sum of coefficients MWTR + AWTR		3.613 (3.401)
rk- <i>F</i> -statistic	291.81	286.47
N	3425665	3425665

Figure A.4: THE EFFECT OF WEALTH TAXATION ON GIVING
FOR HOUSEHOLDS WHO ALREADY GAVE DURING 2006–2012

Notes: We re-estimate the baseline effect on charitable giving on the subset of households who gave during 2006–2012, that is, $\{h : G_{h,t} > 0, t=2007, \dots, 2012\}$.

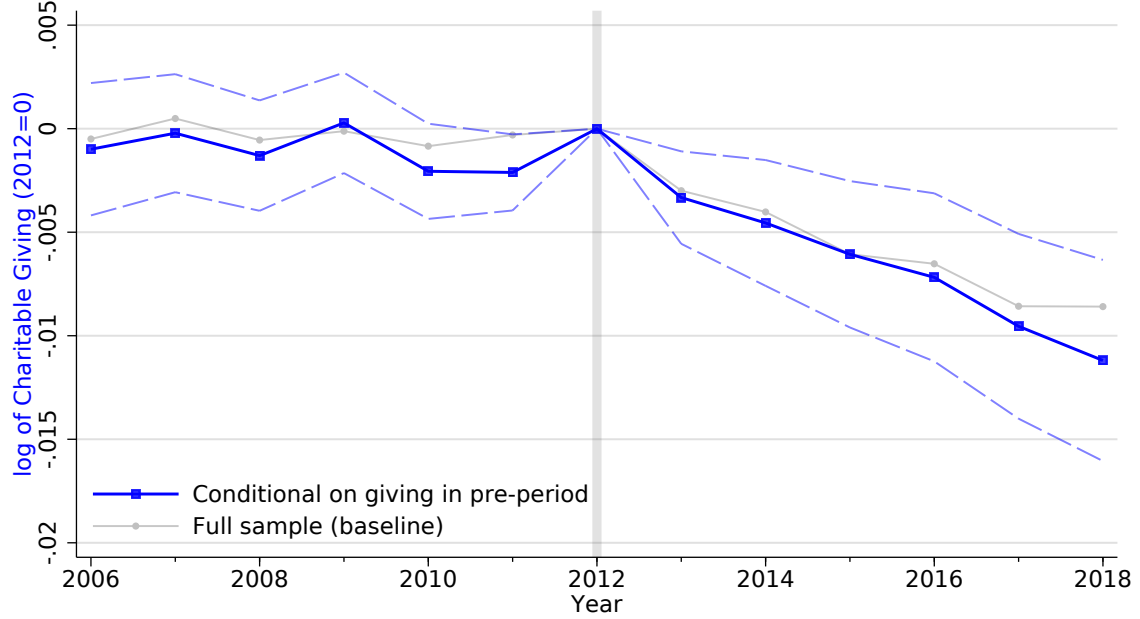


Figure A.5: BUNCHING IN GIVING TO INTERNATIONAL ORGANIZATIONS

Notes: We repeat the bunching analysis done in Panel B of figure 5 for the subsample of givers who only give to organizations with an international focus (excluding missionary organizations), such as the Red Cross, Amnesty, and the World Wildlife Foundation.

