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Sin Taxes and Self-Control†

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According to theory, "sin taxes" are welfare improving if consumers with low self-control respond at least as much to the tax as consumers with high self-control. We investigate empirically if demand response to soft drink and fat tax variations in Denmark depends on consumers' self-control. We use a unique home-scan panel that includes a survey measure of self-control. When taxes increase, consumers with low self-control reduce purchases less strongly than consumers with high self-control. When taxes decrease, both groups increase their purchases similarly. The results show an asymmetry in price elasticities by self-control that is more pronounced when taxes increase. (*JEL* D12, D91, H25, H31, I12, I18, L66)

The "global obesity epidemic" is a major public health challenge (WHO 2000) and a leading risk factor for many non-communicable diseases like type-2 diabetes and coronary heart disease (Smith Jr. 2007). Poor diets that contain high levels of sugar and fat are among the main culprits of this phenomenon (Finkelstein, Ruhm, and Kosa 2005). Hence, the World Health Organization advises governments to consider the introduction of so-called "sin taxes" on unhealthy foods—for example, taxes on sugar-sweetened beverages. A number of countries have already implemented taxes on sugary beverages and other unhealthy foods—for instance, France, Mexico, the United Kingdom, and, until 2014, Denmark.

There are two rationales for sin taxes: externalities and internalities. Externalities mean that sugar consumers do not take the social costs of adverse health behavior into account and the tax is meant to internalize these costs. Internalities in the form of self-control problems imply that people underappreciate the long-term health costs that an unhealthy diet has on themselves. In this paper, we focus on

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the internality argument since it dominates the public debate about sin taxes on foods.[1] The idea is that a sin tax could help consumers with low self-control to follow their long-run utility by increasing the instantaneous costs. Such a tax can even be welfare improving if the corrective gains for individuals with low self-control outweigh the distortionary costs for those without self-control problems. However, to ensure that this is the case, individuals with low self-control must reduce their purchases at least as much as those with high self-control (O'Donoghue and Rabin 2006; Haavio and Kotakorpi 2011).

In this paper, we investigate empirically the effect of self-control on responsiveness to sin tax changes. To do this, we leverage a unique dataset which contains (1) longitudinal scanner data of around 1,300 consumers, (2) a psychologically validated self-control scale answered by the same consumers, and (3) a time period that features multiple exogenous sin tax changes. Namely, we investigate the increase of the Danish soft drink tax in 2012 and its step-wise repeal in 2014. We corroborate our findings by studying the fat tax on saturated fat, which was introduced in 2011 and repealed in 2013.

As a first step, we study the reduced-form demand response to sin tax changes by self-control. Therefore, we use the survey to stratify the sample into high- and low-self-control consumers. In response to the soft drink tax hike, we find that consumers with low self-control reduce their purchases by only 4 percent, significantly less than those with high self-control, who reduce their purchases by 19 percent. In contrast, in response to the soft drink tax repeal, consumers with high and low self-control increase their purchases to a similar extent, between 26 and 28 percent. To study the fat tax, we look at demand responses for butter since it experienced substantial tax variation due to its high content of saturated fat. We find a similar, but less pronounced, pattern for the introduction and repeal of the fat tax.

We find parallel pretax trends between the self-control groups and no difference between self-control groups in placebo tests with untaxed products. We assess a range of alternative explanations for the differential response by self-control. We find that the coefficient of self-control remains stable when including measures for education, nutritional knowledge, income, and preferences for unhealthy food. These findings suggest that the differential response is not driven by correlations with either of these variables. Employing the bounding approach of Oster (2019), we find little evidence that selection on unobservables can explain the results. Moreover, as border-shopping across the German border is popular in Denmark, we also test for whether the differential response by self-control is driven by the proximity to the German border. The differential response differs only mildly by distance to the German border, suggesting that border-shopping does not drive the effect.

[1] First, soft drink taxes are often advocated based on the premise that, in particular, children, who are among the heaviest soft drinks consumers, ignore the long-run consequences of high sugar intake (Dubois, Griffith, and O'Connell 2020). Second, the effectiveness of these taxes is usually assessed by the reduction in consumption and not by tax revenue raised (for externality correction this distinction would be less relevant). For example, on March 13, 2018, the former British finance minister and initiator of the British soft drink tax, George Osborne, tweeted: "In OBR [*Office for Budget Responsibility*] report today is news that our Sugar Tax is even more effective than hoped. Expected receipts halved […]." (https://twitter.com/George_Osborne/status/973647500551827456, retrieved 09/23/19).

Next, we investigate if consumers with high and low self-control experienced different price changes. This could explain the differential response if low-self-control consumers tend to purchase products that passed the tax through to a lesser extent. Therefore, we construct individual level price indices, but find no evidence for differential pass-through by self-control.

Ultimately, we estimate the price elasticities of consumers with high and low self-control, exploiting the tax changes to instrument prices. Using the total sample period, we estimate a price elasticity of -0.8 for consumers with low self-control and -1.1 for consumers with high self-control. We find a similar asymmetry when considering only the pretax period and using nontax price instruments. While the difference in overall price elasticities is relatively small, we uncover a pronounced asymmetry when looking at tax hikes and cuts separately: we find that the elasticities of low- and high-self-control consumers differ significantly for tax hikes, but not for tax cuts.

Our study is motivated by the theoretical literature on taxation of behavioral internalities like imperfect self-control. The idea is that a lack of self-control can lead consumers to overconsume goods with long-run costs that are not fully taken into account at the moment of consumption. A sin tax increases the instantaneous and future costs of consumption and reduces overconsumption. Gruber and Köszegi (2001) show that optimal taxes on cigarettes are substantially higher if individuals are addicted due to present bias. O'Donoghue and Rabin (2006) and Haavio and Kotakorpi (2011) argue that an internality-correcting tax can be welfare improving if individuals with low self-control are at least as responsive to a sin tax as those with high self-control. Further, the comprehensive model of Allcott, Lockwood, and Taubinsky (2019a), which studies the welfare effects and the distributional implications of sin taxes, takes the correction of internalities into account. However, these papers do not make predictions regarding whether consumers with low self-control actually respond to sin taxes, thus leaving this question to empirical research.

We contribute to the burgeoning empirical literature that assesses targeting properties of sin taxes by estimating heterogeneous tax responsiveness by self-control. Allcott, Lockwood, and Taubinsky (2019a) estimate, in their empirical section, the share of soda consumption that is due to a self-reported lack of self-control.[2] They find that the share of consumption that is due to low self-control decreases in income, which means that poor consumers can benefit more from the corrective effects of the tax. However, due to their focus on the regressivity of sin taxes, they do not consider if the price elasticity varies with the level of self-control. In contrast, we use actual tax variation and investigate if the tax successfully targets consumers with low self-control. The targeting properties of a soft drink tax are also investigated by Dubois, Griffith, and O'Connell (2020) in a structural demand model. They estimate price elasticities of different consumer groups and hypothesize that the high soda preference of certain groups (e.g., young people and high sugar consumers) is more likely due to biases. They find that young people are more price responsive, but

[2] They use the Nielsen household panel and classify panelists as having imperfect self-control if they do not respond "Not at all" to the statement "I drink soda pop or other sugar-sweetened beverages more often than I should."

that high sugar consumers are less price responsive than the average consumer. We complement these findings by employing an established measure of self-control (Tangney, Baumeister, and Boone 2004) and by exploiting exogenous variation in prices to explicitly test the impact of self-control on price responsiveness.

Furthermore, we contribute to the empirical literature that estimates price elasticities of sin goods like soft drinks. One strand of this literature uses naturally occurring price variation combined with exogenous instruments or structural approaches to identify price elasticities. For sugary soft drinks, these studies often find demand to be elastic (e.g., elasticity of -1.4 in Allcott, Lockwood, and Taubinsky (2019a); -1.6 in O'Connell and Smith (2020); and -1.2 in the survey by Powell et al. (2013)), with less elastic demand when focusing on on-the-go purchases (-0.9 in Dubois, Griffith, and O'Connell 2020) or explicitly modeling stockpiling (-0.6 in Wang 2015a). Another strand of research exploits quasi-experimental variation in taxes to estimate the impact on purchases (see the surveys in Cawley et al. 2019a; Allcott, Lockwood, and Taubinsky 2019b).[3] While these case studies depend on the specific context and implementation of the considered tax, they are valuable, as consumer reaction to tax-induced price variation may be different compared to the reaction to naturally occurring price variation. For one, these price changes are likely perceived as permanent and therefore allow for the studying of long-term responses. Moreover, a simultaneous price change of all products of a category may lead to distinctive substitution along various margins, including product switching and cross-border shopping. Using the soda tax in Philadelphia, Cawley, Frisvold, and Jones (2019) estimate a price elasticity of -1.3, but do not find a statistically significant reaction in purchases to the soda taxes in San Francisco, Seattle, and Oakland. Seiler, Tuchman, and Yao (2020) also study the soda tax in Philadelphia and estimate that the price elasticity decreases from -1.4 to -0.6 when taking into account cross-border shopping in neighboring jurisdictions. While cross-border shopping is a concern for city-wide taxes, it is arguably less important for national taxes. Colchero et al. (2017) study the Mexican soft drink tax using a pre-post design and estimate a decline in purchase volume of -7.6 percent over two years, which implies a price elasticity of -0.8 using the pass-through rates from Aguilar, Gutierrez, and Seira (2021).[4] We estimate an average price elasticity of -0.9 that is in the range of previous quasi-experimental evidence. However, due to our unique setting and data, we can uncover important asymmetries: when the tax is increased, we estimate that consumers with low self-control are inelastic (elasticity of -0.6), while those with high self-control are elastic (elasticity of -1.8). In contrast, when the tax is repealed, we find only a mild asymmetry by self-control (elasticities of -0.8 and -1.2). Hence, our findings also contribute to a nascent literature that

[3] In earlier work, we analyze the tax pass-through and average purchase response to the 2012 increase and 2014 repeal of the Danish tax on soft drinks using a pre-post design (Schmacker and Smed 2020). Jensen and Smed (2013) and Smed et al. (2016) analyze the effects of the fat tax in Denmark in a pre-post design and document a significant drop in average purchases of saturated fat from, for example, butter and margarine. There is also a long-standing literature that uses sin tax variation on tobacco and alcohol to estimate price elasticities (see the surveys in Chaloupka, Yurekli, and Fong (2012) for tobacco and in Wagenaar, Salois, and Komro (2009) for alcohol).

[4] Aguilar, Gutierrez, and Seira (2021) find a smaller reduction of 3 percent in calories from taxed beverages, but they consider the response immediately after the tax change, while Colchero et al. (2017) show that the response becomes larger over time.

studies asymmetric demand responses to increasing and decreasing prices (Iizuka and Shigeoka forthcoming).

The remainder of the paper proceeds as follows. In Section I, we present the conceptual framework that motivates our empirical analysis. Section II describes the institutional setting and the dataset that we use. Section III presents reduced-form results for differential tax responsiveness by self-control. Section IV analyzes differential pass-through to soft drink prices by self-control. Section V estimates price elasticities by self-control. Section VI concludes.

I. Conceptual Framework

In this section, we briefly summarize a key result of the sin tax literature in behavioral public economics that motivates our research question. Theoretical models of sin taxes typically assume that consumption is associated with costs that are not fully considered by the consumer (O'Donoghue and Rabin 2006; Haavio and Kotakorpi 2011; Allcott, Lockwood, and Taubinsky 2019a).

In a two-good model with sin good x_i and numeraire z_i, consumers maximize $U_i = u_i(x_i, z_i)$, while their long-run utility is $U_i^* = u_i(x_i, z_i) - I_i x_i$, which includes the internality cost, I_i, associated with sin good consumption (Griffith, O'Connell, and Smith 2018). The internality can arise, for example, due to self-control problems in a beta-delta model (Laibson 1997). In the present, consumers underweigh the future health costs relative to their long-run plans. As a result, they overconsume the sin good.

A social planner may decide to impose a tax t on the sin good to help consumers follow their long-run preferences. The corrective tax is lump-sum redistributed back to consumers. Following Haavio and Kotakorpi (2011), the optimal tax can be written as

$$t = \bar{\tilde{I}} + \frac{\operatorname{cov}\left(\tilde{I}_i, \frac{\partial x_i}{\partial t}\right)}{\partial \bar{x} / \partial t}, \tag{1}$$

where $\bar{\tilde{I}}$ is the marginal internality in money-metric terms averaged over all consumers. The first term is adjusted by the covariance between the marginal internality and the demand responsiveness (weighted by the average responsiveness of sin good consumption to tax changes).[5] Intuitively, the optimal tax is larger if those with high internalities (e.g., low self-control) react more to the tax. In that case, the tax is relatively effective in correcting the internality. However, the optimal tax is smaller if consumers with high internalities respond less to the tax.

It is important to note that the optimal tax depends on the price responsiveness in absolute terms (cf. Haavio and Kotakorpi 2011). Hence, even if consumers with low self-control have inelastic demand, the covariance can be positive if their baseline consumption is substantially higher and translates into a greater absolute change in consumption of the sin good. Consequently, in the analysis we will consider both

[5] This formulation resembles the classical result by Diamond (1973) that the optimal externality-correcting tax is a weighted average of marginal externalities, with weights proportional to each individual's demand responsiveness.

relative and absolute changes in purchases, acknowledging that with different baseline levels one does not imply the other.

According to the existing literature, it is an empirical question whether the relationship between self-control and price responsiveness is positive or negative (O'Donoghue and Rabin 2006). Hence, this relationship is what we aim to investigate using the institutional setting described in the next section.

II. Data

A. *Institutional Background*

For identification, we exploit variation in two different sin taxes: the soft drink tax and the fat tax. The tax variation is illustrated in Figure 1. The first tax that we study is the tax on sugary soft drinks. It is a volumetric excise tax that is imposed on all soft drinks that contain more than 0.5 grams of sugar per 100 milliliters. The soft drink tax in Denmark has a longstanding tradition. Both its introduction and subsequent tax reforms have mainly been motivated by the goal of raising tax revenues (Bergman and Hansen 2019). However, the increase of the tax in January 2012 from 1.08 DKK to 1.58 DKK per liter aimed to improve public health.[6] This is also illustrated by the fact that the tax on diet soft drinks remained constant at a lower level. In previous work, we document an average price increase of 1.01 DKK (10.7 percent) in reaction to the tax hike (Schmacker and Smed 2020). Hence, the tax hike is substantially overshifted, which is consistent with the study of Bergman and Hansen (2019) for earlier soft drink tax increases in Denmark. In April 2013, the Danish government announced it would repeal the tax on soft drinks in order to secure jobs in the retail sector in the Danish-German border region and to make up for tax revenue losses due to cross-border trade. The tax was first decreased to 0.82 DKK in July 2013 and completely eliminated in January 2014. In Schmacker and Smed (2020), we estimate an average price drop of 1.88 DKK (23.1 percent) in response to the tax repeal—that is, a bit more than full pass-through (see Figure B.1 in the online Appendix).

The second tax variation that we investigate is the October 2011 introduction and January 2013 repeal of the fat tax. The fat tax was applied to all products that contain more than 2.3g saturated fats per 100g. It amounts to 1.60 DKK per 100g saturated fat. Vallgårda, Holm, and Jensen (2015) analyze the political debate around the introduction and repeal of the fat tax. They conclude that a public health framing dominated the public debate around the tax introduction, while economic arguments (cross-border shopping, administrative burden, and regressive effects on the poor) became more prominent over time, ultimately leading to the repeal of the tax. Since the tax was proportional to the amount of saturated fat, the tax affected product groups very differently. We analyze the demand response for butter since it contains a high amount of saturated fats (approximately 50 percent) and has, therefore, experienced substantial tax variation. In online Appendix C.2, we show that the tax

[6] Note that in Denmark a value-added tax of 25 percent is levied. Hence, if the soft drink tax hike of 0.50 DKK per liter and the VAT were fully passed through, the consumer price increased by 0.63 DKK per liter.

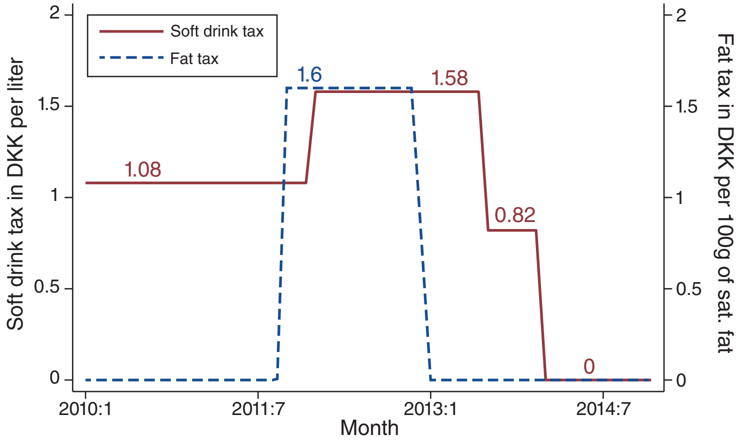

Figure 1. Soft Drink and Fat Taxes in Denmark (in DKK)

Note: Graph shows soft drink and fat tax variation over time. Note that there is a 25 percent VAT on the excise tax-inclusive price.

introduction is associated with an increase in butter prices by 0.76 DKK per 100g and the repeal with a decrease by 0.61 DKK per 100g.

The increasing taxes on sugar and saturated fat were part of the Danish tax reform of 2010. The overall goal was to reduce income taxes and instead increase taxes on consumption goods that have detrimental effects on public health or the environment (The Danish Ministry of Taxation 2009). In addition to the taxes on soft drinks and fat, taxes on sweets, chocolate, ice cream, and tobacco were increased.[7] Moreover, a tax on the content of sugar in many goods was planned but never realized.

B. *Dataset*

To investigate the response in purchases due to the tax variation, we use household panel data from GfK for the years 2009 to 2014 (GfK Denmark A/S 2015). Panelists are asked to scan the barcodes and report prices and quantities of all the grocery purchases that they bring into the home. GfK aims for a representative panel with respect to geography, age, education, and family size. Once a year, panelists fill out a questionnaire on demographic and socioeconomic characteristics. In 2013 and 2015, an additional survey containing a broad range of questions about self-control and dietary habits was sent to panelists. The responses to this questionnaire are matched with the purchase records using the panel identifier.

[7] While the tax on tobacco was increased in April 2012, the taxes on sweets, chocolate, and ice cream were increased in January 2012; that is, simultaneously with the soft drink tax. Harding, Leibtag, and Lovenheim (2012) find cross-price elasticities between soda and candy/snacks close to zero; hence, we do not expect these simultaneous tax changes to affect our analysis of the soft drink tax. We investigated differential responses in sugar consumption by self-control to candy, chocolate, and ice cream taxes and find qualitatively similar results compared to our main results. To determine the sugar content, we use the FRIDA food nutrient database from 2018 by the Danish Technical University (DTU Food Institute 2018). However, since these goods tend to be underreported according to the data provider, we did not include them in the main analysis. The results are available upon request.

In the analysis, we only include those households that report at least one purchase of the product in question per year and that responded to the self-control questionnaire. These restrictions leave us with 1,278 panelists for the soft drink tax estimations and 1,324 for the fat tax estimations.

When looking at quantity purchased, we aggregate the purchases to monthly observations to account for potential stockpiling. Hence, in the soft drink estimations, the dependent variable is monthly purchases of taxed soft drinks in centiliters (including colas, lemonades, ice tea, and juices with added sugar). In the fat tax estimations, the dependent variable is monthly purchases of butter in grams. We assign months a zero if any purchases are observed but none of these purchases belong to the product category in question (soft drinks or butter). If no purchase is observed in a given month, it is considered a missing observation for that month.

As is often the case with household-level purchase data, the distribution of purchases is characterized by a right-skewed distribution with some extreme outliers. To prevent outliers from having an undue influence on the parameter estimates, we employ multiple measures. First, we clean the data from anomalous values (as described in online Appendix B.1 and C.1). Second, we winsorize the monthly quantities at the 99 percent level; i.e., the largest 1 percent of quantities is set to the quantity at the 99th percentile. Third, in the analysis, we employ Poisson QMLE and two-part models, which are more robust to outliers compared to OLS.

C. *Measuring Self-Control*

Self-control is measured using the scale developed and validated by Tangney, Baumeister, and Boone (2004), which consists of 36 statements concerning different domains of trait self-control (the complete scale is provided in Table A.2 in the online Appendix). The respondents indicate their approval to each of these statements on a 5-point Likert scale. Whenever possible, we use the 2013 data and, if the panelist did not fill in the survey in 2013, we impute the missing data with data from 2015. Hence, we assume that self-control is a time-constant trait, which is supported by a high retest reliability: among the 1,234 panelists, who answered the self-control scale in both years, the scores from 2013 and 2015 correlate with $r = 0.783$.[8]

Tangney, Baumeister, and Boone (2004) perform an exploratory factor analysis using the self-control scale and identify five latent factors of self-control. Apart from a factor related to "healthy habits," they find, for example, factors for "deliberate/nonimpulsive actions" or "work ethic." Thus, computing an unweighted average over all 36 items, including those that are unrelated to health-related self-control, can lead to measurement error and, consequently, attenuation bias (Heckman, Pinto, and Savelyev 2013; Piatek and Pinger 2016).

[8] In line with our assumption, Gottfredson and Hirschi (1990) argue that trait self-control is developed during childhood and that relative self-control within-cohort remains stable afterwards. Hay and Forrest (2006) find that absolute and relative measures of self-control remain stable for 84 percent of the sample through ages 7 to 15, and Coyne and Wright (2014) show in a twin study that 76 percent of the variation in self-control is explained by genetic variation. Among others, Tangney, Baumeister, and Boone (2004) and Fullerton et al. (2018) report a high retest reliability of the self-control scale.

TABLE 1—HIGHEST LOADING ITEMS ON THE SELECTED SELF-CONTROL FACTOR

		(1)	(2)
1.	I am good at resisting temptations	0.644	0.695
2.	I have a hard time breaking bad habits (R)	0.608	0.694
3.	I am lazy (R)	0.439	0.489
11.	I refuse things that are bad for me	0.347	0.419
12.	I spend too much money (R)	0.367	0.420
15.	I wish I had more self-discipline (R)	0.459	0.623
25.	I am able to work effectively toward long-term goals	0.305	0.378
26.	Sometimes I can't stop myself from doing something, even if I know it is wrong (R)	0.316	0.402
33.	People would say that I have an iron self-discipline	0.397	0.435
34.	I have many healthy habits	0.708	0.538
35.	(I eat healthy food)	0.712	.

Notes: Table shows the items with a rotated factor loading of at least 0.3 on the selected self-control factor after principal component factor analysis (varimax rotation) with five factors, using GfK data. Column 1 shows factor loadings when using all 36 items of the self-control scale and column 2 when excluding item 35 ("I eat healthy food"). (R) indicates that the item is reverse coded.

In order to reduce the large number of items and to find the latent dimensions of self-control, we perform a principal component factor analysis using all 2,387 panelists who filled in the self-control scale. We follow Tangney, Baumeister, and Boone (2004) and extract five factors.[9] The resulting factor structure is described in online Appendix A.1.

To find the self-control factor that matters for food choices, we correlate the five extracted self-control factors with body mass index (BMI) as well as stated intentions to reduce weight and to eat healthier. In Table A.3 in the online Appendix, we see that panelists who have low self-control according to the second factor have higher BMI and are more likely to be obese. They are more likely to respond that they would like to reduce their weight and agree more often that they should eat less sugar and animal fat. All of these correlations are substantially weaker or nonexistent for the other four self-control factors. Hence, in the main analysis we focus on the second of the five self-control factors.

Table 1 shows the highest loading items on the selected self-control factor. The factor can be described as temptation tolerance and is associated with health-related habits. The factor also loads high on the item "I eat healthy foods." To make sure that the inclusion of this item does not drive the results, we rerun the factor analysis excluding this item (results of this factor analysis are shown in the second column of Table 1 and in online Appendix Table A.4). The resulting factor loadings are qualitatively similar, suggesting that the results are not driven by the respective item. In the robustness section, we conduct robustness checks using this alternative measure of self-control and find very similar results (Table B.9). We perform further sensitivity analyses using only the highest loading items as well as aggregated, unweighted indices of the entire self-control scale (Table B.8).

[9] The scree plot of our factor analysis suggests extracting between three and six factors.

D. *Descriptive Statistics*

In Table A.1 in the online Appendix, we show descriptive statistics of the overall sample used in the analysis, as well as descriptives of the sample split by self-control. Moreover, in the last column, we show descriptives for the unrestricted sample, which also includes panelists who report at least one purchase in every sample year but for whom we have no information on self-control.

The demographic characteristics appear quite similar across the different sample restrictions. However, there is an intuitive association between self-control and education, with high-self-control respondents having higher education. In the robustness section, we address whether the differential response by self-control is affected if we also control for heterogenous responses by education.

III. Do Purchase Responses to Sin Taxes Differ by Self-Control?

A. *Empirical Strategy*

As a first step, we test whether purchases by consumers with high and low self-control react differently to the tax changes. Therefore, we exploit within-household variation in soft drink purchases the year before and after the tax changes.

We are interested in both absolute and relative responses in purchases. While the former is the relevant statistic in models of sin taxes, the latter allows us to estimate demand elasticities that we can compare to estimates from the literature.

The empirical model for estimating the absolute change in purchase quantity Y_{it} in month t by consumer i is

$$(2)\quad \mathrm{Y}_{it} \;=\; \alpha_0 + \alpha_1 \,\mathrm{tax}_t + \alpha_2\Big(\mathrm{tax}_t \times \mathbf{1}\big(\beta_i \;=\; \beta^{high}\big)\Big) + \gamma_i + \eta_t + \alpha_3 Z_{it} + \epsilon_{it},$$

where Y_{it} is the observed quantity. The variable tax_t is a dummy variable that is 1 after the tax change and 0 before. We interact the tax dummy with indicators that specify whether consumer i is characterized by low or high levels of self-control as defined in the previous section. Hence, α_2 estimates the differential effect of the tax change on purchase quantity for consumers with high self-control compared to those with low self-control. γ_i denotes household fixed effects, which are included to control for time-invariant unobserved heterogeneity, and η_t denotes month-of-the-year fixed effects. Z_{it} is a set of household-specific controls that includes the household size, income group, and labor market status of the main shopper. In the analysis of the soft drink tax, Z_{it} also includes the average monthly temperature in Denmark.[10]

To estimate the relative change in purchases, we use an exponential model to estimate the conditional expectation for purchase quantity Y_{it} by consumer i in month t:

$$(3)\quad E\big[Y_{it}|X_{it}\big] \;=\; \gamma_i exp\Big(\alpha_0 + \alpha_1 \,\mathrm{tax}_t + \alpha_2\Big(\mathrm{tax}_t \times \mathbf{1}\big(\beta_i \;=\; \beta^{high}\big)\Big) + \eta_t + \alpha_3 Z_{it},\Big)$$

[10] The temperature data were retrieved from Danmarks Meteorologiske Institut (DMI 2017).

which is estimated by a Poisson quasi-maximum likelihood estimator (QMLE) with consumer fixed effects, γ_i (Wooldridge 2010).[11] As we use an exponential model, the coefficients can be interepreted as semi-elasticities. Hence, α_1 estimates the relative purchase response by low-self-control consumers, while $\alpha_1 + \alpha_2$ estimates the relative response by consumers with high self-control. η_t and Z_{it} are defined as above.

In both estimations, the main coefficient of interest is the interaction effect of the tax dummy and the self-control indicator, α_2. In order to identify whether the differential responsiveness is due to self-control, we must make two assumptions.

First, we assume that the demand for the taxed goods by consumers with high and low self-control would have evolved similarly absent the tax changes. While we cannot observe the counterfactual demand, we can assess the plausibility of the assumption by checking whether the trends are parallel in the years preceding the tax changes.[12] Therefore, we regress purchases on “placebo tax” indicators in the pretax period. These results of these placebo tests are discussed in Section IIIB for the soft drink tax and in Section IIIC for the fat tax. Moreover, we can analyze purchases of other untaxed products, which are not substitutes or complements of the taxed products, to check whether demand in general evolved similarly around the time of the tax changes. These results are presented in Figure 3.

Second, we assume that differences in price responsiveness are due to self-control and not due to other correlated characteristics, like income and education. Therefore, we investigate if the differential response by self-control remains stable when additionally controlling for differential changes by income, education, nutritional knowledge, and taste for unhealthy food. The results are presented in Section IIID.

In order to assess whether the purchase response is driven by consumers stopping to consume or consumers reducing the purchased amount, we estimate the extensive margin and the intensive margin of consumption separately. First, we estimate equation (2) as a linear probability model, using the likelihood to consume any soft drink in a given month as dependent variable (extensive margin). Second, we estimate equation (2) using the log-transformed quantity, conditional on observing a purchase, as dependent variable (intensive margin).

As is often the case with household-level purchase data, the distribution of purchases is characterized by a mass at 0 and a right-skewed distribution with some extreme outliers. To assess the robustness of our estimations, we use the estimates from the extensive and intensive margin to additionally estimate a two-part model that explicitly takes the mass at 0 into account (Duan et al. 1983; Mullahy 1998).

For each tax event, we consider one year before the first tax change and one year after the final tax change. For the soft drink tax estimations, we employ a “donut”

[11] Since the purchase data contain many zeros, we cannot use OLS with a log-transformed dependent variable. In such cases, several authors recommend using Poisson QMLE due to its robustness to distributional misspecification (e.g., Santos Silva and Tenreyro 2006; Wooldridge 2010). The estimator is consistent and asymptotically normal, as long as the conditional mean is correctly specified and it is not restricted to count data or Poisson distributed data (Wooldridge 1999). Moreover, by using cluster-robust standard errors, the Poisson QMLE is robust to overdispersion.

[12] As noted above, we are interested in both absolute and relative responses to sin tax changes. Since parallel pretax trends in absolute and relative terms only coincide in the case of flat pretax trends, we check if the yearly absolute and relative trends are indeed flat.

approach and omit the months January and December of each year. Otherwise, we might overestimate the effect of the tax hike in January 2012 due to customers stockpiling soft drinks in December 2011 and living off stock in January 2012. To make years comparable, we omit the months January and December in all soft drink tax estimations. For the fat tax, where stockpiling is less of a concern since butter is a perishable product, we compare the years before and after the introduction in October 2011 and before and after the repeal in January 2013.

In the empirical analysis, we investigate the differential responsiveness by self-control, first for soft drink tax changes and, second, for fat tax changes. In both cases, we provide graphical evidence on the development of purchases surrounding the tax changes before we present the regression results.

B. *Soft Drink Tax Variation*

Figure 2, panel A shows monthly averages of reported soft drink purchases by self-control, controlling for month and consumer fixed effects. In Figure 2, panel B we plot the relative difference between consumers with high and low self-control.[13]

First, in the years before the tax hike in 2012, we observe that the difference between consumers with low and high self-control is relatively stable. However, as can be seen in Figure 2, panel A, when the tax increased in 2012, soft drink purchases by consumers with low self-control did not change strongly, while we observe a marked drop for consumers with high self-control. In Figure 2, panel B this shows up as an increasing difference between purchase quantities by consumers with low and high self-control. In contrast, when the tax was cut in July 2013 and ultimately repealed in January 2014, we observe a marked increase in purchases by both consumer groups. Figure 2, panel B shows that the difference between consumer groups remains at a similar level.

In order to quantify the purchase response to the tax variation, we show estimation results of the empirical model in Table 2 for the two parameters of interest (α_1 and α_2). The coefficient α_1 is the tax indicator variable, which gives the change in purchases by low-self-control consumers (the reference category), and α_2 is the interaction of the tax dummy with the high-self-control indicator, which gives the differential change in purchases by high-self-control consumers. Panel A shows results for the tax hike and panel B for the tax repeal. In columns 1 and 2, we use OLS to estimate the absolute change in centiliters. In columns 3 and 4, we use the Poisson QMLE model, such that coefficients approximate the relative changes in percent. In column 5, we use OLS with purchase incidence as the dependent variable (extensive margin), and in column 6 we use OLS and the log-transformed quantity given a purchase as the dependent variable (intensive margin).

The results in panel A of Table 2 reveal that consumers with high self-control decreased their purchases significantly more than consumers with low self-control in response to the tax hike, both in absolute and relative terms. Column 4 shows that consumers with low self-control reduced their purchases by 3.6 percent (not statistically

[13] In Figure B.2 in the online Appendix, we plot the raw data without adjusting for consumer and month fixed effects. The figure shows a strong seasonal component, but gives a similar picture overall.

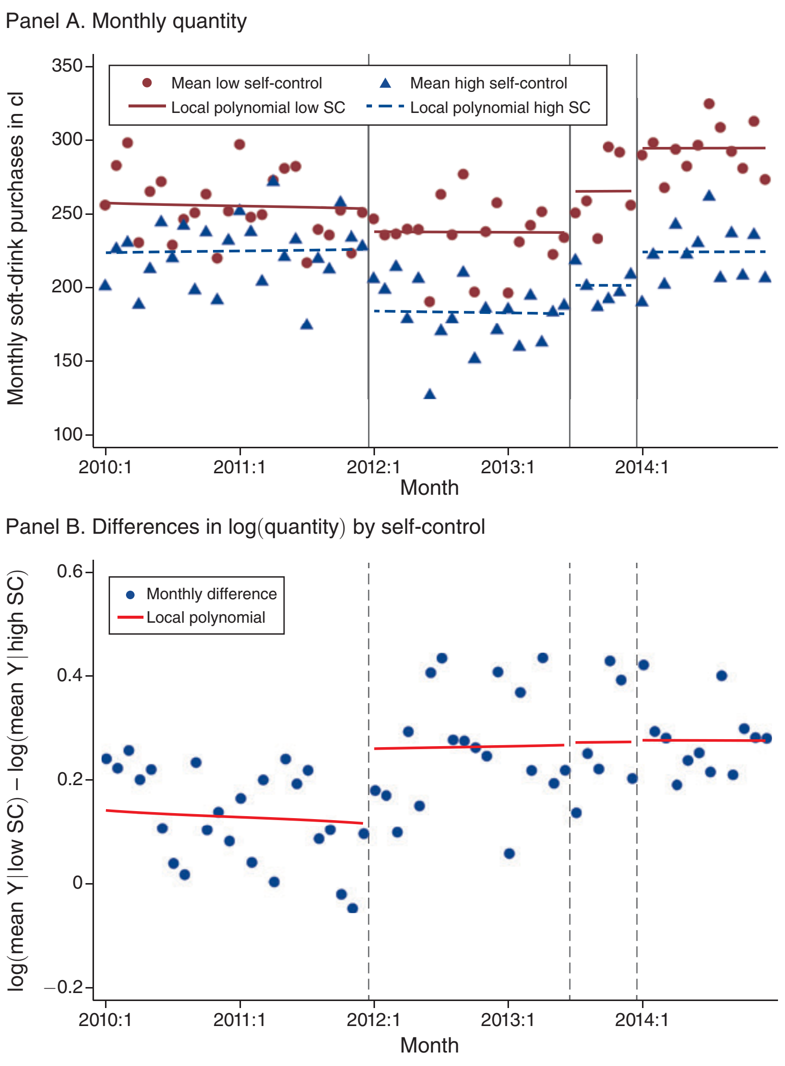

FIGURE 2. MONTHLY SOFT DRINK PURCHASES BY SELF-CONTROL (DESEASONALIZED)

Notes: Panel A shows monthly average residuals from a regression of soft drink purchase quantity on consumer and month fixed effects. The residuals are added to the sample mean. Panel B shows the monthly differences in log(quantity) from panel A between high- and low-self-control consumers. Local polynomials use degree 0 and 12 month bandwidth. The vertical lines indicate the timing of soft drink tax changes.

Source: GfK

significant), while consumers with high self-control reduced their purchases significantly more. Adding coefficients reveals that consumers with high self-control reduce their purchases by 19.1 percent. Columns 5 and 6 show that consumers with high self-control responded more strongly to the tax hike both in terms of purchase probability (extensive margin) and purchase quantity (intensive margin), but the differential response is only significant at conventional levels on the extensive margin.

Table 2—Soft Drink Purchases in Response to Soft Drink Tax Changes by Self-Control

	Absolute change (1)	Absolute change (2)	Relative change (3)	Relative change (4)	Extensive margin (5)	Intensive margin (6)
Panel A. Tax hike						
Tax hike 01/12	−10.504	−7.200	−0.039	−0.036	−0.012	0.004
	(12.485)	(12.796)	(0.046)	(0.047)	(0.011)	(0.037)
Tax hike 01/12 × High SC	−38.810	−36.738	−0.169	−0.155	−0.036	−0.090
	(15.859)	(15.994)	(0.062)	(0.061)	(0.014)	(0.046)
Households	1,104	1,104	1,104	1,104	1,104	1,104
Household months	19,543	19,543	19,543	19,543	19,543	7,466
Panel B. Tax repeal						
Tax cut 07/13	32.282	29.206	0.116	0.102	0.023	0.076
	(12.500)	(13.891)	(0.044)	(0.049)	(0.012)	(0.040)
Tax cut 01/14	31.446	36.723	0.100	0.125	0.028	0.083
	(14.330)	(16.216)	(0.046)	(0.054)	(0.012)	(0.040)
Tax cut 07/13 × High SC	−6.887	−3.974	0.001	0.014	−0.006	0.005
	(17.450)	(17.347)	(0.071)	(0.070)	(0.015)	(0.051)
Tax cut 01/14 × High SC	−2.519	−3.264	0.021	0.019	0.009	−0.031
	(17.729)	(17.747)	(0.063)	(0.063)	(0.015)	(0.050)
Households	1,141	1,141	1,141	1,141	1,141	1,141
Household months	25,904	25,904	25,904	25,904	25,904	9,782
Controls		✓		✓	✓	✓
Household fixed effects	✓	✓	✓	✓	✓	✓

Notes: Table shows regression results with standard errors clustered on household level, using GfK data. In columns 1 and 2, estimations are conducted using OLS; that is, coefficients can be interpreted as absolute changes. In columns 3 and 4 the estimation uses Poisson quasi-maximum likelihood estimation; that is, coefficients can be interpreted as relative changes. Column 5 uses OLS with purchase incidence in a given month as dependent variable; column 6 uses OLS with log-transformed quantity as dependent variable. Controls include household size, income, labor market status, temperature, and month-of-the-year fixed effects. The estimations only include observations that exhibit within-household variation in purchases.

In panel B of Table 2, we conduct the same exercise for the tax repeal. Here, we estimate separate coefficients for the two tax cuts in July 2013 and January 2014. We see a significant increase in purchases in response to both tax cuts, but there is no significant difference between high- and low-self-control consumers. In response to the first tax cut, column 4 shows that low-self-control consumers increased their purchases by 10.2 percent and those with high self-control by 11.6 percent. In response to the second tax cut, the former increased their purchases by 12.5 and the latter by 14.4 percent.

Our analysis assumes that, absent the tax changes, soft drink demand by consumers with low and high levels of self-control would have evolved in the same way. While we cannot observe the counterfactual situation, we can check the plausibility of this assumption by running a series of placebo tests.

First, we run the same estimations for placebo tax changes preceding the actual tax changes. Thus, we can assess the pretax trends using the same model as in our main specification. In Table B.1 in the online Appendix, we complete this exercise for placebo tax changes on January 1, 2010, and January 1, 2011. We observe no differential change in purchases by high-self-control consumers, neither in absolute nor in relative terms. There are no significant positive or negative pretax trends by either group.

Second, we check whether "placebo products," for which the tax was not changed, exhibit a differential change in purchases by self-control. We consider two groups of untaxed products: milk and household supplies.[14] The former captures common shocks to consumer demand for beverages, but is a potential substitute and thus may not be completely unaffected by the tax change. The latter group consists of products completely unrelated to soft drinks. If the asymmetric response to soft drink taxes were driven by other idiosyncratic factors—for example, a change in reporting behavior or the economic condition of the household—we would expect to see a difference by self-control for these placebo goods as well.

Figure 3 reproduces Figure 2, panel B using the placebo product groups on the same scale. Figure 3, panel A shows the differences in log(quantity) of milk purchases by self-control. Figure 3, panels B to D show the differences in log(expenditure) for the three subgroups of household supplies: toiletries, cleaning supplies, and tissues.[15] Figure 3 does not show peculiarities around the timing of the soft drink tax changes, suggesting that the asymmetric pattern for soft drinks is not generated by instantaneous factors that coincide with the timing of the tax change. In Table B.2, we show regression results for the placebo product groups, but we do not observe a differential response by self-control that is statistically or economically significant.

Robustness.—We provide further robustness checks in the online Appendix. We calculate predicted values of a two-part model using the estimates from the extensive and intensive margin, along with Duan smearing factors (Duan 1983). The two-part model is an alternative estimation method for non-negative data with many zeros (Duan et al. 1983; Mullahy 1998; Egger et al. 2011). The results in Table B.3 in the online Appendix corroborate the findings from the QMLE exponential model. In response to the tax hike, consumers with high self-control reduced their purchases by 21.2 percent, while consumers with low self-control reduced their purchases by 3.8 percent, with the latter response not statistically significant. In contrast, when the tax is reduced, both consumer groups increase their purchases to a similar extent.

In Figures B.4 and B.5, we show results from a permutation test with 2,500 iterations, in which we randomly reshuffle if consumers are classified as high or low self-control. For the tax hike, only 0.7 percent of coefficients are more negative than the actually estimated interaction coefficient, which corroborates its significance. For the tax cuts, 61.0 and 57.2 percent of randomly reshuffled iterations produce a more negative effect, suggesting that there is no significant difference for the tax repeal. While we cluster the standard errors on the consumer level in the main specification, in online Appendix Table B.4, we follow the suggestion of Bertrand, Duflo,

[14] Household supplies consist of three subgroups: toiletries (toothpaste, deodorant, shampoo, soap, shower gel, hair styling), cleaning supplies (cleaning supplies, dishwashing and laundry detergent, rinse aid), and paper and tissue (toilet and kitchen paper).

[15] We aggregate to average monthly expenditure instead of purchase quantity, as there is no natural way to aggregate the purchased quantity for most of these goods (e.g., deodorant and soap). Moreover, if the asymmetric response was driven by a change in economic conditions, we would expect that consumers buy similar quantities of household supplies (e.g., toiletries), but choose cheaper products and, hence, have a lower expenditure.

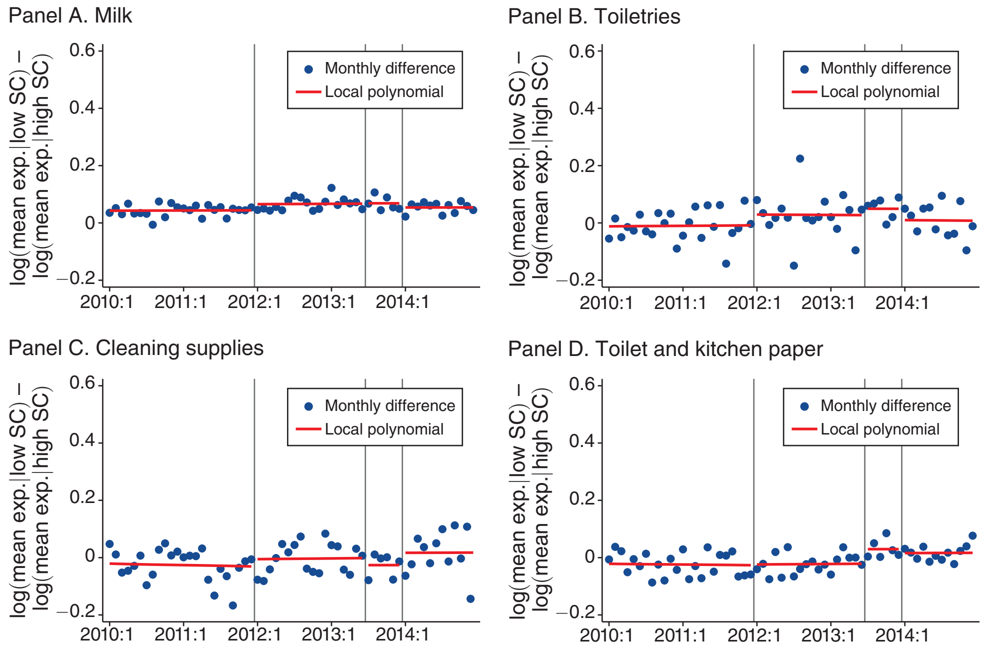

FIGURE 3. MONTHLY PURCHASES OF PLACEBO PRODUCTS BY SELF-CONTROL (DESEASONALIZED)

Notes: This figure shows monthly differences between high- and low-self-control consumers in log(quantity) and log(expenditure), based on residuals from a regression on consumer and month fixed effects. Panel A shows differences in log(quantity) for milk, panels B to D monthly average expenditures for toiletries (toothpaste, deodorant, shampoo, soap, shower gel, hair styling), cleaning products (cleaning supplies, dishwashing and laundry detergent, rinse aid), and paper and tissues (toilet and kitchen paper), respectively. Local polynomials use degree 0 and 12 month bandwidth. The vertical lines indicate the timing of soft drink tax changes. The residualized means are shown in Figure B.3.

Source: GfK

and Mullainathan (2004) and collapse the months into one pretax and one post-tax change period. The results show that the standard errors are very similar compared to the main specification.

In our setup, we observe purchases by many consumers that we assign to two large groups (low and high self-control). Donald and Lang (2007) argue that with a small number of groups the standard errors can be underestimated and propose a two-step procedure to obtain more accurate standard errors. We follow the exposition of the two-step procedure in Bedard and Kuhn (2015). First, we calculate the monthly average purchase quantity in the low- and high-self-control group; that is, for the tax hike, we collapse the sample from 1,278 consumers $\times$ 20 months to 2 groups $\times$ 20 months. In the second step, we regress the time series of monthly differences between consumer groups on the tax dummy. This method results in correct inference under the assumption that shocks to the difference between the low- and high-self-control groups are independent and identically distributed. This assumption would be violated if there were consumer-specific autocorrelated shocks. To allow shocks to have some persistency, we follow Bedard and Kuhn (2015) and additionally calculate Newey-West standard errors with a lag of three months. Online Appendic Table B.5 shows the results, but the results do not alter our conclusions.

As further robustness tests, we re-estimate our main specification on the subsample of single households. The reason is that there is likely heterogeneity in soft drink preferences within households, and the main shopper (whose self-control we elicit) may not be the person demanding to buy the soft drinks. By restricting the analysis to single households, we can be sure that measured self-control coincides with the self-control of the individual who actually buys and consumes the soft drinks. Online Appendix Table B.6 presents the results, which reiterate the previous findings: high-self-control individuals reduce their purchases significantly more than low-self-control consumers when the tax goes up, and the interaction coefficient is even larger than in the full sample. However, when taxes go down there is no differential change that is statistically significant at conventional levels.

Finally, we scrutinize the robustness of the results regarding the measure of self-control. In online Appendix Table B.7, we split the sample into terciles according to self-control. The results show that in response to the tax hike, consumers with low self-control reduced their purchases the least and those with high self-control the most, while for the tax repeal there is no systematic difference in responsiveness.

In column (1) of Table B.8 in the online Appendix, we use the continuous self-control factor instead of a median split, but the results are very similar to our main results. In columns (2) and (3), we split the sample according to the items that load highest on the self-control factor, but which do not ask directly for consumption of healthy food: "I am good at resisting temptations" and "I have many healthy habits." The differential responsiveness to the tax hike is stronger for the former item, suggesting that temptation resistance is an important factor behind the observed results. In columns (4) and (5), we form an unweighted index of the three and four highest loading items (on top of the items mentioned before, these are "I have a hard time breaking bad habits" and " I wish I had more self-discipline"). The results turn out very similar to the self-control factor in column (1) and hardly change when adding the fourth item. In column (6), we form an index of the complete self-control scale (i.e., computing the unweighted average of all 36 items of the self-control scale). Although we still observe a differential responsiveness to the tax hike, the magnitude is smaller, which would be expected if weighting all items equally increases measurement error and attenuation bias (Heckman, Pinto, and Savelyev 2013). In online Appendix Table B.9, we use the alternative measure of self-control that excludes the revealed preference item about healthy food consumption contained in the original scale (cf. Section IIC). However, the results are similar to the main specification.

To account for measurement error in elicited self-control, we can take advantage of the fact that self-control was elicited in 2013 and 2015, giving us multiple self-control measurements for many panelists. Hence, in online Appendix Table B.10, we use the continuous self-control factor elicited in 2015 to instrument the continuous self-control factor elicited in 2013 (cf. Gillen, Snowberg, and Yariv 2019).[16] Using the 2013 and 2015 self-control factors separately, as well as the IV

[16] Since the second stage of the IV is a nonlinear exponential model, it is implemented using the control function approach by Lin and Wooldridge (2019). That means the predicted residuals from the first stage are included as regressors in the second stage (see Section V). The t-test for $\hat{\nu}$ tests the null hypothesis that the instrumented

estimation, corroborates our findings. Moreover, the statistical insignificance of the predicted residuals $\hat{\nu}$ indicates that we cannot reject the null hypothesis that the 2013 self-control measure is exogenous.

C. *Fat Tax Variation*

In the previous section, we show that consumers with low self-control respond less to increasing soft drink taxes than consumers with high self-control. In contrast, when soft drink taxes are cut, there is no systematically different response. In this section, we check whether this pattern is particular to soft drink tax changes or whether it also emerges for the introduction and repeal of the fat tax.

In the following, we look at butter as it is one of the goods that contains the most saturated fat and is frequently purchased. By extending the analysis to butter, we can assess whether the differential responsiveness to tax changes is only observed for soft drinks or also observed for other products.

We run the same estimations as described in Section IIIA on the data for butter.[17] Figure 4, panel A shows monthly averages for butter purchases by self-control, controlling for consumer and month fixed effects. In Figure 4, panel B we plot the relative differences in purchases by self-control. When the tax is introduced, we find again that consumers with high self-control reduce their purchases more than those with low self-control. However, the differential response is less pronounced compared to the soft drink tax hike. When the tax is repealed, both consumer groups increase their purchases to a similar extent.

We show estimation results of the coefficients of interest from the empirical model in Table 3. Panel A illustrates that, in response to the fat tax introduction, consumers with high self-control reduce their purchases significantly more than consumers with low self-control. Column 4 shows that consumers with low self-control reduce their purchases by around 5.2 percent, while those with high self-control reduce their purchases by around 10.1 percent.

In panel B of Table 3, we run the estimation for the tax repeal. As before, we find little evidence for a differential response to the tax repeal. If anything, consumers with low self-control increase their purchases more than those with high self-control, but the differential response is not significant at conventional levels.

In Table C.3, the results of placebo tax changes in January 2010 and October 2010 are shown. The interaction coefficients, which measure differential changes in response to the placebo tax changes, are insignificant and close to 0.

In sum, we find evidence supporting the findings of the soft drink tax analysis. In response to the fat tax, consumers with low self-control respond less to increasing prices. While the general pattern persists, the results appear slightly noisier and less strong than in the case of soft drinks.

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variable is exogenous. For simplicity, when looking at the tax repeal, we exclude the months between the tax cuts and compare the year before the first tax cut to the year after the final tax cut.

[17] The estimations mirror those for soft drinks. The only notable differences are, first, that we restrict the sample to households that report a butter purchase in the years 2010 through 2013. Second, we do not include the average temperature as a control variable since temperature is arguably less relevant for butter demand than it is for soft drink demand.

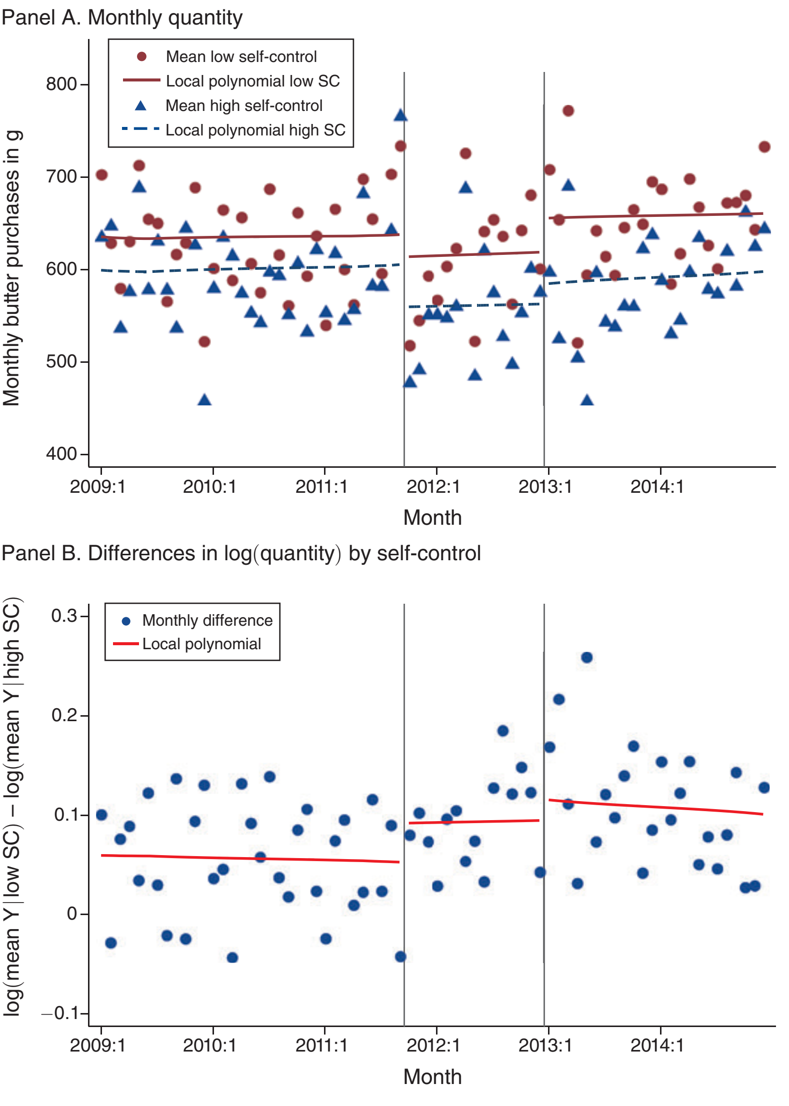

Figure 4. Monthly Butter Purchases by Self-Control (Deseasonalized)

Notes: Panel A shows monthly average residuals from a regression of butter purchase quantity on consumer and month fixed effects. The residuals are added to the sample mean. Panel B shows the monthly differences in log(quantity) from panel A between high- and low -self-control consumers. Local polynomials use degree 0 and 12 month bandwidth. The vertical lines indicate the timing of fat tax changes.

Source: GfK

Robustness.—In online Appendix C.3, we conduct the main robustness checks for the fat tax variation. Figure C.2 shows coefficients from a permutation test with 2,500 iterations. While for the tax introductions, only 1.7 percent of estimates are more negative than the actual interaction coefficient, this is true for 15.0 percent of coefficients for the tax repeal. Hence, while the differential response for the tax introduction is unlikely to be purely random, this cannot be rejected for the tax

Table 3—Butter Purchases in Response to Fat Tax by Self-Control

	Absolute change (1)	Absolute change (2)	Relative change (3)	Relative change (4)	Extensive margin (5)	Intensive margin (6)
Panel A. Tax introduction						
Tax hike	−36.400	−32.997	−0.057	−0.052	−0.014	−0.051
	(10.210)	(10.276)	(0.016)	(0.016)	(0.008)	(0.013)
Tax hike × High SC	−27.565	−27.991	−0.050	−0.049	−0.018	−0.003
	(14.096)	(14.065)	(0.023)	(0.023)	(0.011)	(0.017)
Households	1,296	1,296	1,296	1,296	1,296	1,296
Household Months	27,692	27,692	27,692	27,692	27,692	17,941
Panel B. Tax repeal						
Tax repeal	24.373	25.617	0.037	0.040	0.012	0.045
	(10.006)	(11.291)	(0.015)	(0.017)	(0.008)	(0.014)
Tax repeal × High SC	−16.049	−16.286	−0.023	−0.023	0.005	−0.026
	(13.793)	(13.875)	(0.022)	(0.022)	(0.010)	(0.018)
Households	1,302	1,302	1,302	1,302	1,302	1,302
Household months	28,483	28,483	28,483	28,483	28,483	18,818
Controls		✓		✓	✓	✓
Household fixed effects	✓	✓	✓	✓	✓	✓

Notes: Table shows regression results with standard errors clustered on household level, using GfK data. In columns 1 and 2, estimations are conducted using OLS; that is, coefficients can be interpreted as absolute changes. In columns 3 and 4 the estimation uses Poisson quasi-maximum likelihood estimation; that is, coefficients can be interpreted as relative changes. Column 5 uses OLS with purchase incidence in a given month as dependent variable; column 6 uses OLS with log-transformed quantity as dependent variable. Controls include household size, income, labor market status, and month-of-the-year fixed effects. The estimations only include observations that exhibit within-household variation in purchases.

repeal. In online Appendix Table C.4, we collapse the pre- and post-tax month and find the standard errors to be very similar to those for the main specification.

In Table C.5, we assess the sensitivity to the sample split in the main specification. Instead of a median split, we split the sample into three quantiles and observe that consumers with the lowest level of self-control respond the least to the tax hike. After the tax repeal, we again do not observe a systematic differential response by self-control. In Table C.7, we use the alternative measure of self-control described in Section IIC. The results for the main specification in the first column are similar to those obtained for the original self-control scale.

D. *Alternative Explanations*

In the following, we investigate other potential explanations for the differential response by self-control. Therefore, we interact the tax indicator with other observable characteristics, such as income, education, taste for unhealthy food, and nutritional knowledge. Moreover, we employ a bounding approach to assess the importance of selection on unobservables, and we show that cross-border shopping is unlikely to explain the differential response by self-control.

Education and Nutritional Knowledge.—First, we address the concern that self-control is correlated with education and that education is responsible for the differential response. In the second column of Table 4, we interact the soft drink tax dummy additionally with an indicator for high education. High education means that the panelist has attended at least one year of tertiary education, whereas the reference category is no tertiary education. The interaction coefficients of self-control and the tax change indicators are almost unaffected compared to the main specification. In the second column of Table C.6 in the online Appendix, we conduct the same exercise for the fat tax. Similarly, including education does not change the interaction coefficient of self-control and tax indicators.

Second, it is conceivable that our measure of self-control is associated with knowledge about the healthiness of food and that this drives the differential response. To account for this, we add the interaction of the tax change dummy with an indicator if consumers agree with the statement "I believe I would make healthier food choices if I had more information on how to eat healthy." In the third column of Table 4, we show the results for the soft drink tax variation. For both the tax hike and the tax repeal, the interaction with self-control remains of similar magnitude. Further, in online Appendix Table C.6, we observe a similar pattern for the fat tax variation.

Income.—As self-control is typically positively correlated with income, it could be that tighter budget constraints are the reason for the differential response. However, if that were the case, we would expect consumers with low self-control (and low income) to reduce purchases *more* than consumers with high self-control (and high income). Hence, the finding that low-self-control consumers respond *less* to the tax hike already suggests that budget constraints do not drive the differential response.

In the fourth column of Table 4, we rerun the main specification for the soft drink tax variation, but add an interaction with a dummy indicating whether a panelist is in the top half of the distribution of equivalized incomes. We observe that the coefficients for the interaction of the soft drink tax hike and repeal with self-control are of a similar magnitude compared to our main specification. In online Appendix Table C.6, we conduct the same exercise for the fat tax and observe that the coefficients of interest also move very little when including interactions with income.

Taste for Unhealthy Food.—It is conceivable that measured self-control is correlated with taste for unhealthy food. To check if the differential response by self-control can be attributed to differences in taste, we add the interaction with a dummy variable that indicates if consumers agree with the statement "I believe I would make healthier food choices if unhealthy food was less tasty." In the fifth column of Table 4, we observe in panel A that the interaction of the tax hike with self-control becomes slightly smaller, but remains sizeable and significant. Consumers who like unhealthy food seem to be less likely to reduce their purchases in response to the tax hike, but the interaction is only significant at the 10 percent level. In panel B, there is—as in the main specification—not much evidence for a differential effect by self-control.

Table 4—Soft Drink Purchases in Response to Soft Drink Tax Changes, Alternative Explanations

	Main (1)	Education (2)	Knowledge (3)	Income (4)	Taste (5)	All (6)
Panel A. Tax hike						
Tax hike	−0.036	−0.051	−0.034	−0.039	−0.096	−0.098
	(0.047)	(0.053)	(0.054)	(0.057)	(0.064)	(0.073)
Tax hike						
× High SC	−0.155	−0.160	−0.157	−0.155	−0.137	−0.146
	(0.061)	(0.061)	(0.065)	(0.063)	(0.065)	(0.069)
× Interaction term		0.050	−0.053	0.010	0.096	
		(0.060)	(0.086)	(0.070)	(0.066)	
Households	1,104	1,104	1,033	1,104	1,033	1,033
Household months	19,543	19,543	18,297	19,543	18,297	18,297
Panel B. Tax repeal						
Tax cut 07/2013	0.102	0.120	0.108	0.081	0.139	0.105
	(0.049)	(0.053)	(0.055)	(0.056)	(0.064)	(0.076)
Tax cut 01/2014	0.125	0.099	0.108	0.166	0.116	0.150
	(0.054)	(0.056)	(0.061)	(0.062)	(0.071)	(0.081)
Tax cut 07/2013						
× High SC	0.014	0.021	−0.008	0.009	−0.018	−0.025
	(0.070)	(0.072)	(0.074)	(0.073)	(0.073)	(0.077)
× Interaction term		−0.064	−0.016	0.047	−0.064	
		(0.072)	(0.093)	(0.072)	(0.072)	
Tax cut 01/2014						
× High SC	0.019	0.008	0.041	0.031	0.036	0.048
	(0.063)	(0.064)	(0.068)	(0.065)	(0.069)	(0.071)
× Interaction term		0.093	0.042	−0.084	0.004	
		(0.070)	(0.092)	(0.067)	(0.069)	
Households	1,141	1,141	1,068	1,141	1,068	1,068
Household months	25,904	25,904	24,307	25,904	24,307	24,307
Interaction term	None	High education	Lacks knowledge	High income	Unhealthy taste	All
Controls	✓	✓	✓	✓	✓	✓
Household fixed effects	✓	✓	✓	✓	✓	✓

Notes: Table shows Poisson QMLE regression results with standard errors clustered on household level, using GfK data. The dependent variable is monthly quantity in centiliters. "High education" means tertiary education (ref.: vocational education), "Lacks knowledge" indicates consumers who agree that they would make healthier food choices if they had more information, "High income" indicates consumers in the top half of the distribution of equivalized incomes, "Unhealthy taste" indicates consumers who agree that they would make healthier food choices if unhealthy food were less tasty. Controls include household size, income, labor market status, temperature, and month-of-the-year fixed effects. The estimations only include observations that exhibit within-household variation in purchases.

In online Appendix Table C.6, we add the interaction with a preference for unhealthy food to the fat tax estimation. We observe, in contrast to the case of soft drinks, that consumers with a preference for unhealthy food decrease their purchases more in reaction to the fat tax introduction. Nevertheless, controlling for taste leaves the differential response by self-control almost unaffected. In panel B, we observe that there is no differential response to the tax repeal by taste differences.

Selection on Unobservables.—While we cannot directly test for the influence of selection on unobservables, we can draw some inferences based on the movement of coefficients and explained variance when controlling for observables. We adapt the approach suggested by Oster (2019), which builds on Altonji, Elder, and Taber (2019). The idea is to bound the estimates by making assumptions about the relative importance of unobserved (relative to observed) variables and about the highest explainable variance.

We aim to determine whether within-household changes in purchases vary due to differences in self-control or due to unobserved differences between self-control groups. Hence, the baseline estimate is a fixed-effects OLS regression of purchases on only a tax dummy and the tax dummy interacted with the self-control indicator. In the controlled specification, we additionally control for time-varying controls and interactions of the tax dummy with education, nutritional knowledge, income, and unhealthy taste (i.e., the specification in the last columns of Tables 4 and C.6). We assume that selection on unobservables is as important as selection on observables and that it can either go in the same or in the opposite direction. The argument is that if controlling for informative observables does not change the coefficients much, controlling for unobservables would not do so either. In online Appendix D, we describe this approach in detail.

Online Appendix Table D.1 presents the results of the bounding exercise. It can be seen that the coefficients do not move much when including controls and the bounds for the soft drink tax hike and the fat tax introduction do not contain 0, suggesting that proportional selection on unobservables is unlikely to drive the results. However, due to the relatively low R^2, we cannot make a definite statement regarding the importance of unobservables.

Cross-Border Shopping.—As mentioned above, the tax on soft drinks was principally repealed to reduce cross-border shopping in Germany. In general, this should not be a concern for our analysis since in the GfK data, consumers also report purchases abroad. However, one may be concerned that cross-border purchases are underreported and consumers engage differently in border-shopping depending on self-control. To assess the importance of this channel, we distinguish if consumers have access to the German border without using a toll bridge or ferry.[18] Thus, the "No Toll" indicator is a proxy for how easy and economic it is to buy groceries in Germany.

In online Appendix Table B.11, we separately estimate the model on "Toll" and "No toll" households to assess heterogenous effects by distance to border. For the soft drink tax hike, we observe in panel A that among consumers in the "Toll" region (i.e., where the border is not easily accessible) the difference between low and high self-control is somewhat stronger compared to consumers in the "No toll" region. However, even in the "No toll" region, the interaction is not substantially smaller compared to the main specification and remains significant at the 10 percent level despite the reduced sample size. This seems to suggest that consumers with high

[18] While households in Jutland and Funen do not have to use a toll bridge or ferry to reach the German border, households in Sealand, Copenhagen, and Bornholm must. The costs of using the ferry or bridge for a standard car start at 30 euros each way. In Schmacker and Smed (2020), we provide descriptive evidence that this distinction is informative about the propensity to engage in cross-border shopping.

self-control do not reduce their purchases as much when there are close-by opportunities to avoid the tax. In panel B, we see that purchases increased more in the "Toll" region after the tax repeal, but there is again not much evidence for a differential response by self-control.

Cross-border shopping is arguably less important for butter purchases since butter is not as storable as soft drinks. In line with this argument, Table C.8 in the online Appendix shows that the interaction of self-control and the fat tax dummy is similar in magnitude for "Toll" and "No toll" consumers. The magnitude of the interaction coefficient in panel A is almost the same compared to the main specification, but is estimated more noisily due to the reduced sample size.

IV. Did Consumers with Low Self-Control Face Different Price Changes?

In the previous section, we found that consumers with low self-control responded less to a tax hike than high-self-control consumers, suggesting that they have a lower price elasticity. However, it could be that the asymmetry is due to low-self-control consumers buying products for which the tax changed the prices to a lesser extent, even though their price elasticity is the same. This could happen, for example, if low-self-control consumers buy primarily discount brands and high -self-control consumers buy branded products. If there is lower pass-through of the tax for discount brands, we could see an asymmetric demand response despite a similar demand elasticity.

In this section, we investigate whether the experienced pass-through of the tax differs by self-control. As a first step, Figure 5 shows the average reported prices of soft drinks by self-control. The figure shows that both consumer groups report higher prices after the tax hike and lower prices after the tax cuts, but we do not find strong differences between consumer groups.

It is important to note that the average prices reported by the panelists have to be interpreted with caution, as consumers do not only choose the quantity, but also the quality of products (Deaton 1988). For example, a tax hike may lead consumers to switch to cheaper products, leading to a smaller increase in reported prices. The propensity to substitute to cheaper products when prices rise may differ between consumers with high and low self-control. Moreover, if some consumers purchase less frequently due to the tax hike, the composition of consumers who report prices in the dataset changes. For example, if the tax hike induced all consumers of expensive branded products to stop consuming, this would lead to a lower average price observed in the data.

In the following, we control for differential substitution by constructing individual price indices. Our disaggregated purchase data allow us to construct consumer-specific product baskets and to calculate the price change if the consumer had consumed the same basket after the tax changes. Previous literature has used similar techniques to study heterogeneity in inflation rates (e.g., Kaplan and Schulhofer-Wohl 2017; Jaravel and O'Connell 2020).

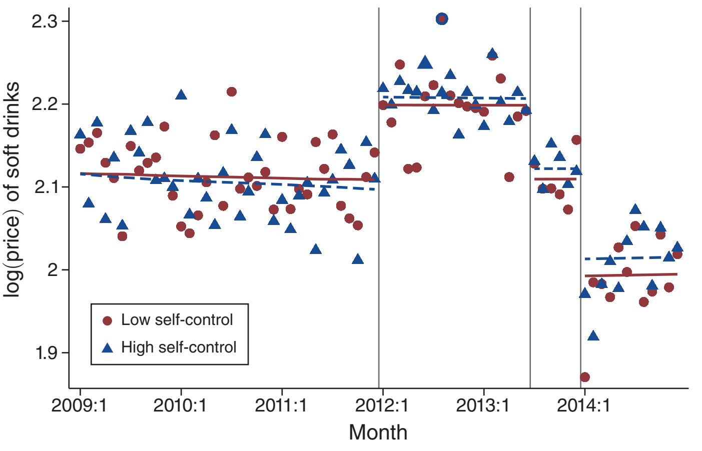

FIGURE 5. MONTHLY AVERAGES OF REPORTED SOFT DRINK PRICES BY SELF-CONTROL

Note: This figure shows monthly average log(prices) of purchased soft drinks by consumers with low and high self-control, respectively, using GfK data.

We calculate both Laspeyres- and Paasche-type price indices on the individual level. While the Laspeyres index π^L takes the pretax period as the base period, the Paasche index π^P considers the consumption bundle from the post-tax period:

$$(4) \qquad 1 + \pi_{i,t+1}^{Lasp} = \frac{\sum_j q_{ij,t} \cdot p_{j,t+1}}{\sum_j q_{ij,t} \cdot p_{j,t}}.$$

$$(5) \qquad 1 + \pi_{i,t+1}^{Paas} = \frac{\sum_j q_{ij,t+1} \cdot p_{j,t+1}}{\sum_j q_{ij,t+1} \cdot p_{j,t}}.$$

where $q_{ij,t}$ is the quantity of product j that consumer i purchases in period t. A product j is defined by its brand, package type (e.g., 2-liter bottle or can), and shop type (e.g., discounter or supermarket).[19] p_{jt} are national average prices on the quarterly level.[20] For each consumer, we compute the expenditure weights of each purchased product in the base year. These observations are projected forward (for the Laspeyres index) or backwards (for the Paasche index) and matched with the national average

[19] When deciding how to aggregate products, we face a tradeoff between ensuring similarity of products and retaining products that are observed in as many periods as possible. We decided to aggregate by brand, package type, and shop type to capture the most important margins of tax-induced substitution (switching to cheaper brands, larger container sizes, and cheaper stores). We report summary statistics for the products in online Appendix Tables B.12 to B.14.

[20] We follow Jaravel and O'Connell (2020) and use national average prices across all households. The alternative—using the price that the consumer pays for the same product in the comparison period—would significantly reduce the sample due to missing values. Similarly, we choose the quarterly instead of the monthly level to reduce the number of missing prices (e.g., a product that was bought in November 2011 may have no reported price in November 2012). We also experimented with monthly, biannual, and annual price averages, but the results turned out very similar.

prices in the comparison period to calculate the individual price index. If consumers substitute to cheaper products after a tax hike, we expect the Laspeyres index to be larger than the Paasche index. Conversely, if consumers upgrade quality after tax cuts, we expect the Paasche index to be larger.

Figure 6 shows the distribution of individual price indices for the tax hike and the tax repeal. For both tax changes, we calculate the price indices comparing the year before the tax change to the year after the tax change (i.e., for the tax repeal the year before the first tax cut to the year after the second tax cut).

We observe heterogeneity in the pass-through of the tax changes across consumers. For example, the interquartile range of the Laspeyres index is 13.1 percentage points for the tax hike and 14.6 for the tax repeal. However, graphically we see no systematic differences by self-control since the histograms largely overlap.

To analyze whether there are significant differences in the pass-through of the tax by self-control, we run the following fixed-effects regression:

$$\log(p_{it}) = \omega_0 + \omega_1 \, tax_t \times \mathbf{1}\left(\beta_i = \beta^{high}\right)) + \delta_i + \epsilon_{it}. \tag{6}$$

where p_{it} is the average price that a consumer would pay in a given period given the consumer-specific product share in the base period:

$$p_{it} = \sum_j \frac{p_{j,0} \, q_{ij,0}}{\sum_{j'} p_{j',0} \, q_{ij',0}} p_{jt}. \tag{7}$$

That means that for each consumer there is one individual price in every period that reflects the prices in the comparison period, weighted with the consumer's product basket in the base period.

In online Appendix Table B.15, we regress the price index on the tax dummies, interacted with an indicator whether a consumer has high or low self-control. In column (1), the dependent variable is the average price of the pretax consumption bundle (Laspeyres) and in column (2) the price of the post-tax consumption bundle (Paasche). For the tax hike, we estimate an average price change of 12.1 percent for the Laspeyres index and 9.3 percent for the Paasche index. For the tax repeal, we estimate a price decrease of 6.8 percent for the first tax cut and 15.1 to 15.4 for the second tax cut, with little difference between price indices. If anything, prices decreased slightly less for consumers with high self-control, but the difference is small in magnitude. In total, we do not find evidence for differential price changes by self-control.

V. Do Price Elasticities Differ by Self-Control?

Ultimately, we are interested in estimating the price elasticities of consumers with high and low self-control. In our setting, we are in the unusual position of having multiple sin tax changes within a short period of time that lead to sizeable exogenous shifts in prices. In contrast, previous literature had to rely on naturally

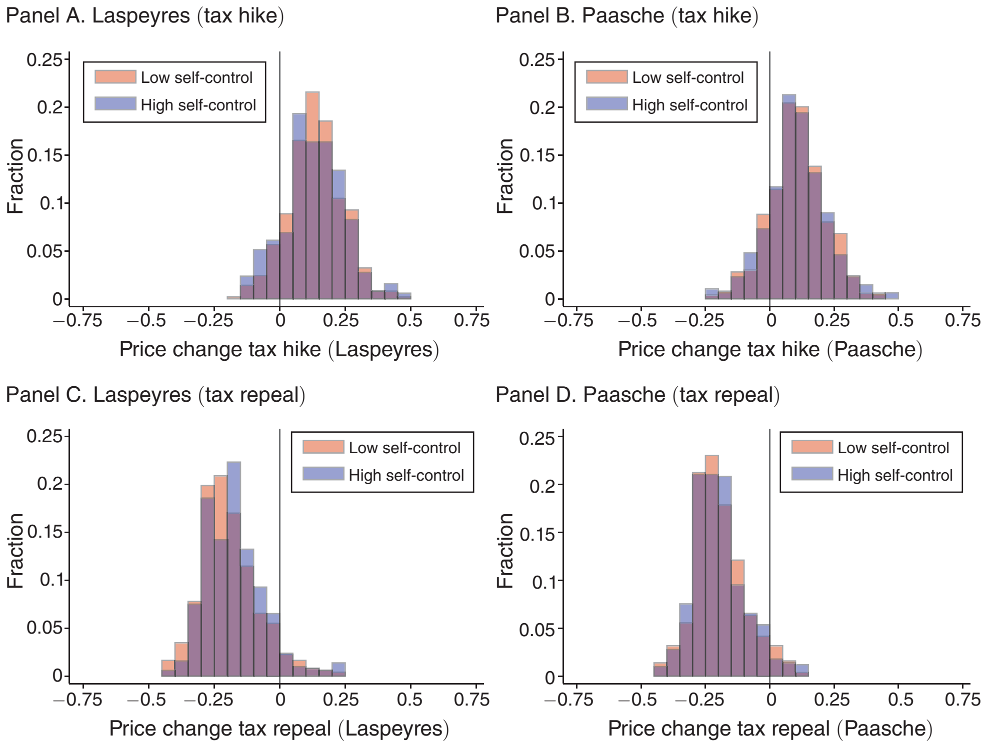

Figure 6. Annual Changes of Individual Price Indices for Soft Drinks by Self-Control

Notes: The figure shows the distribution of individual price indices comparing the year before and after the respective tax changes. The Laspeyres index uses pretax product weights and the Paasche index post-tax product weights. We omit the lowest and highest percentile from the figure.

Source: GfK

occurring price variation and instrumental variable strategies to overcome the problem of simultaneity.[21]

There are several reasons why studying tax-induced price variation is of particular relevance. For example, price variation induced by taxes is likely to be perceived as permanent, while naturally occurring price variation often comes in the form of temporary sales and price discounts. If consumers stockpile during sales, this can lead to a wedge between short-run price elasticities, which are typically estimated by demand models, and long-run price elasticities, which are the policy relevant statistic (Wang 2015a). Moreover, a tax that is passed through leads to simultaneous price changes of all taxed products, whereas natural price variation typically affects only a subset of products, which can lead to potentially different substitution patterns (including product substitution and cross-border shopping). Hence, in our

[21] For example, structural demand models reconstruct the choice set that consumers face and exploit price variation over time that differs across brands and retailers (e.g., Dubois, Griffith, and O'Connell 2020; O'Connell and Smith 2020; Wang 2015a). In contrast, Allcott, Lockwood, and Taubinsky (2019a) instrument prices with those at the same chain in other regions. In the robustness section, we apply the latter strategy to our pretax-change data.

case the source of price variation is exactly that which is relevant for policy: tax variation.

However, as in our dataset prices are reported by consumers, we cannot take them at face value. As discussed above, consumers do not only choose the quantity, but also the quality of purchased products. For example, a tax hike may lead consumers to downgrade on quality and choose cheaper products. Moreover, we only observe prices and quantities if consumers actually make a purchase. For consumers who stopped consuming after the tax hike, post-tax prices are not observed, such that regressing instantaneous quantities on prices would mean excluding these individuals from the analysis.

Therefore, we regress the reported purchase quantity on the individual price index using the product basket from the previous year, resembling a Laspeyres-type price index (see equation (7)). Hence, the product choice is predetermined and we can also include consumers who stop consuming soft drinks after the tax hike. As can be seen in Figure 6, there is substantial variation in the Laspeyres price indices on the individual level that we can exploit. The price elasticities are, thus, identified from variation in individual price indices, as the excise tax changes the prices of different product baskets to different extents.[22]

A. *Estimation Strategy*

In the estimations, we use an instrumental variable strategy that instruments the prices with the tax changes. We run the estimations both on the overall sample and separately by self-control.

As a first stage of the IV, we use OLS to regress the log-transformed price index on the tax dummy:

$$\log(p_{it}) = \omega_0 + \omega_1 tax_t + \xi_t + \delta_i + u_{it}, \tag{8}$$

where p_{it} is the price index of the pretax change product basket, defined in equation (7). ξ_t are month-of-the-year dummies and δ_i are individual fixed effects.

Since the dependent variable (monthly purchases) contains many zeros, we estimate a nonlinear exponential model in the second stage. Hence, to implement the IV, we employ the control function approach proposed by Lin and Wooldridge (2019) and obtain the fixed-effects residuals from the first stage $\hat{u}_{it} = \log(p_{it}) - \big(\hat{\omega}_0 + \hat{\omega}_1 \,\text{tax}_t + \hat{\xi}_t\big)$. The residuals are then included as regressors in the second stage regression:

$$E\big[Y_{it} | X_{it}\big] = \gamma_i exp(\alpha_0 + \alpha_1 \log(p_{it}) + \eta_t + \nu u_{it}), \tag{9}$$

[22] Mechanically, the same excise tax hike leads to a smaller relative price change for an expensive brand than for a cheap brand. Moreover, pass-through may differ by product and retailer depending on various factors like, among others, market power or demand elasticities. In Tables B.12 to B.14 in the online Appendix, we report the average prices by brand, packaging, and shop type in the years before and after the tax changes. Comparing the average prices over the years shows that the price changes, around the tax changes, differ strongly by brand, packaging, and shop type.

which is estimated by Poisson QMLE, such that α_1 can be interpreted as a price elasticity. Due to the control function approach, we use the predicted values $\hat{u}_{it}$ for u_{it}, which ensures that we only use the tax-induced price variation by controlling for the endogenous part. Moreover, by testing the H0: $\nu = 0$, we can check whether the price is indeed endogenous. To account for the fact that one regressor is estimated, we use the bootstrap with 2,000 replications to obtain standard errors.

The exclusion restriction requires that the tax does not have a direct effect on the purchase quantity. We cannot fully exclude that the tax hike or the surrounding media discussion had a signaling effect that discouraged consumption (Rees-Jones and Rozema 2020; Taylor et al. 2019); hence, the elasticity estimates have to be interpreted with caution. However, in the robustness section we will use the IV strategy proposed by Allcott, Lockwood, and Taubinsky (2019a) that uses pretax data and is, therefore, not subject to the same concerns.

B. *Results*

Table 5 shows the estimated price elasticities for soft drinks when we instrument prices with tax changes. We estimate price elasticities for the period 2011 to 2014, which comprises all tax changes, as well as the tax hike and tax repeal separately.

Using all tax changes without differentiating into tax hikes and cuts, we estimate an overall price elasticity of -0.93. However, when running these estimations separately by self-control, we observe that consumers with low self-control tend to be somewhat inelastic (elasticity of -0.82), while consumers with high self-control tend to be unitary elastic (elasticity of -1.07). However, the difference in elasticities is not statistically significant.

Next, we look at tax hike and tax repeal separately—that is, we use the year before the tax change as base period and compare it to the year after the tax change.[23] When looking only at the tax hike, we uncover strong heterogeneity in price elasticities by self-control. While we estimate consumers with low self-control to be inelastic with a price elasticity of -0.62, consumers with high self-control have elastic demand with an elasticity of -1.78. This asymmetry is weaker when using tax cuts as instruments. Here, consumers with low self-control have an elasticity of -0.82 and those with high self-control of -1.20. Using bootstrapped standard errors, we find that the elasticities are significantly different for the tax hike ($p = 0.022$), but not for the tax repeal ($p = 0.236$).

In most estimations, we reject that $\hat{\nu} = 0$, indicating that prices are indeed endogenous. Moreover, the first stage coefficients show that the tax is highly predictive of prices and that pass-through is similar for consumers with high and low self-control, as discussed in detail in Section IV.

Robustness.—In online Appendix E, we follow the estimation strategy proposed by Allcott, Lockwood, and Taubinsky (2019a) to see if the results also hold when

[23] Note that this changes the base period for the tax repeal. While the product basket for "All tax changes" and "Tax hike" is determined in a period of low prices before the tax hike, the product basket for "Tax repeal" is determined after the tax hike when prices are high.

Table 5—Soft Drink Price Elasticities By Self-Control (Laspeyres)

	Total	Low self-control	High self-control
Panel A. All tax changes			
Price elasticity	−0.929	−0.823	−1.068
	(0.142)[b]	(0.181)[b]	(0.230)[b]
$\hat{\nu}$	0.539	0.180	0.948
	(0.193)	(0.258)	(0.283)
First stage			
Tax hike (01/2012)	0.111	0.111	0.112
	(0.004)	(0.006)	(0.005)
Tax cut (07/2013)	−0.089	−0.085	−0.092
	(0.005)	(0.007)	(0.007)
Tax cut (01/2014)	−0.134	−0.146	−0.122
	(0.006)	(0.008)	(0.010)
First stage F	870.5	465.1	426.0
Households	1,021	503	518
Household months	36,431	17,964	18,467
Panel B. Only tax hike			
Price elasticity	−1.165	−0.615	−1.780
	(0.258)[b]	(0.376)[b]	(0.355)[b]
$\hat{\nu}$	0.453	−0.371	1.395
	(0.325)	(0.478)	(0.442)
First stage			
Tax hike (01/2012)	0.121	0.121	0.122
	(0.004)	(0.005)	(0.005)
First stage F	1,152.1	619.0	540.6
Households	1,010	498	512
Household months	17,989	8,883	9,106
Panel C. Only tax repeal			
Price elasticity	−0.986	−0.818	−1.201
	(0.152)[b]	(0.197)[b]	(0.255)[b]
$\hat{\nu}$	0.702	0.180	1.262
	(0.215)	(0.270)	(0.322)
First stage			
Tax cut (07/2013)	−0.083	−0.078	−0.087
	(0.004)	(0.006)	(0.006)
Tax cut (01/2014)	−0.130	−0.145	−0.115
	(0.006)	(0.007)	(0.010)
First stage F	1,136.0	635.5	542.5
Households	976	487	489
Household months	22,107	11,020	11,087
Household fixed effects	✓	✓	✓
Month-of-year fixed effects	✓	✓	✓

Notes: The table shows regression results from an IV Poisson QMLE regression, using the control function approach proposed by Lin and Wooldridge (2019). In the first stage, we use FE OLS to regress the log price index on the tax dummy and control variables. In the second stage, we use FE Poisson QMLE to regress the purchase quantity on the price index and the predicted residuals. Hence, the price coefficient can be interpreted as a price elasticity. $\hat{\nu}$ is the coefficient of the predicted residuals. In panel A, we use the years 2011 to 2014 and 2011 as the base year to determine the product weights. In panels B and C, the sample spans one year before and after the respective tax change, and the pretax-change year is the base period. Standard errors marked with [b] are bootstrapped with 2,000 replications and clustered on the household level, while the unmarked standard errors are clustered on the household level.

Source: GfK

using nontax price variation. Allcott, Lockwood, and Taubinsky (2019a) instrument prices with prices at retailers of the same chain outside a consumer's region to control for local demand shocks. Moreover, they subtract the national average price for the same product to control for product-specific national demand shocks like brand advertising campaigns. We use only the pretax years 2010 and 2011 to run this analysis.

In online Appendix Table E.1, we observe an asymmetry in line with our main results: for low-self-control consumers we estimate a price elasticity of -0.78, and for high-self-control consumers an elasticity of -1.17. However, it has to be noted that unlike Allcott, Lockwood, and Taubinsky (2019a), we do not observe marketing variables (e.g., whether a product is featured or on display) that often coincide with sales. Hence, these price elasticities may be biased. We discuss further differences to our main results in online Appendix E.

VI. Conclusion

In both policy debates and in the economic literature, it is often argued that sin taxes can help consumers with low self-control to act more in accordance with their own long-run interest. However, this requires that consumers with low self-control respond to tax changes by reducing their consumption. This paper presents evidence that consumers with low self-control respond systematically less to increasing sin taxes than do high-self-control consumers. However, we find no difference between the groups when the tax is reduced, indicating an asymmetry in the responsiveness to tax hikes and cuts.

We estimate that consumers with high self-control tend to have elastic demand for soft drinks and consumers with low self-control tend to have inelastic demand. We find that the difference in elasticities is more pronounced and statistically significant when taxes increase.

A potential reason for the asymmetry is the habituating nature of soft drink consumption (Zhen et al. 2011). While reducing consumption in response to a tax hike may require self-control, increasing consumption after a tax cut may not require self-control. Consistent with this idea, Tangney, Baumeister, and Boone (2004) describe trait self-control as related to the capability of breaking habits and resisting temptation. In their meta-analysis on trait self-control, de Ridder et al. (2012) distinguish between automatic behavior (e.g., habitual smoking) and deliberate behavior (e.g., quitting smoking) and argue that self-control is particularly relevant for the latter. In this context, reducing consumption after a tax hike may require breaking with a habit (where self-control matters), while increasing consumption after a tax cut means indulging in a habit more intensively (where self-control does not matter). In line with this explanation, we find a stronger asymmetry in price responsiveness between high- and low-self-control consumers for soft drinks than for butter—with sugary beverages more likely to trigger cravings than butter.

Even if consumers with low self-control are less price elastic, they could still react more strongly in absolute terms if their baseline consumption is sufficiently higher. Since the optimal internality-correcting tax depends on the *absolute* price responsiveness (O'Donoghue and Rabin 2006; Haavio and Kotakorpi 2011), it is important to consider whether differences in price elasticities translate to differences in absolute price responsiveness. We find that both in relative and absolute terms, consumers with low self-control react less strongly to increasing soft drink and fat taxes.

These results suggest that the correction of overconsumption due to low self-control does not suffice to justify the introduction of a sin tax on welfare grounds. However, it must be noted that the optimal sin tax is additive in externalities and internalities (Diamond 1973; Allcott, Lockwood, and Taubinsky 2019a). Hence, a positive sin tax may still be welfare-optimal if it corrects externalities on public health or internalities due to other factors than a lack of self-control (e.g., lack of nutritional knowledge).

Furthermore, while consumers with low self-control may not respond to the price incentives themselves, smart sin tax design can still improve the diets of individuals with low self-control (Grummon et al. 2019). If taxes on the harmful ingredient increase (e.g., sugar in soft drinks), producers are incentivized to make their product less unhealthy, as documented for the tiered soft drink tax in the UK (Dickson, Gehrsitz, and Kemp 2021). Since the Danish soft drink tax was volumetric, this incentive was not given.

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