Incidence and Distributional Effects of Value Added Taxes*

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Abstract: In this paper, we examine the incidence and distributional effects of VAT in a setting with plausibly exogenous variation in tax rates. The context of our study is a sharp change in the VAT policy on food items in Norway. Using a regression discontinuity design, we examine the direct impact of the policy change on the consumer prices of food items as well as any cross-price effects on other goods. Our estimates suggest that taxes levied on food items are completely shifted to consumer prices, whereas the pricing of most other goods does not seem to be materially affected. To understand the distributional effects of the VAT reform, we use expenditure data and estimate the compensating variation of the tax induced price changes. We find that lowering the VAT on food attenuates inequality in consumer welfare, in part because households adjust their spending patterns in response to the price changes. By comparison, the usual first-order approximation of the distributional effects, which ignores behavioral responses, seriously understates the redistributive nature of the VAT reform.

Keywords: Value added taxes; incidence; distributional effects; pass-through

Jel Codes: H20, H22, H23, H31, H32.

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1 Introduction

Are taxes levied on commodities completely shifted to consumer prices, or does the incidence also fall on firms? What are the welfare implications of commodity taxes for poor and rich households? These questions are important for both policy and scientific research, as commodity taxes make up a large part of fiscal revenue in most developed countries. In Europe, much of the controversy surrounding recent policy proposals to broaden the base for value added taxes (VAT) revolves around who ultimately bears the burden of these taxes. In the United States, recent debates on whether to increase reliance on consumption-based taxes have raised concerns over the distributional effects of such policy changes. The typical assumption is that consumer prices fully reflect taxes, so that the main empirical question is how the tax induced price changes affect members of different income groups. However, the evidence base is scarce (Crawford, Keen, and Smith, 2010) and market imperfections could generate both over and under-shifting of commodity taxes to consumer prices (Seade, 1985; Delipalla and Keen, 1992; Anderson, De Palma, and Kreider, 2001b).

The aim of this paper is to investigate the incidence and distributional effects of commodity taxes in a setting with plausibly exogenous variation in tax rates. The context of our study is an abrupt change in the VAT policy on food in Norway. As in most European countries, food retailing in Norway is highly concentrated with a few chains commanding most of the market. On July 1st, 2001 the Norwegian government reduced the VAT on all food items from 24 to 12 percent, while the VAT on non-food items remained at 24 percent. This sharp change in VAT policy provides an attractive setting to analyze the pass-through of commodity taxes using a regression discontinuity (RD) design that compares consumer prices just before (i.e. the control group) and after (i.e. the treatment group) the reform date. We apply this design to rich data on retail prices for a representative sample of consumer goods. This allows us to estimate the direct impact of the policy change on the consumer prices of food items as well as any cross-price effects on other goods. We challenge the identifying assumptions of the RD design through a number of robustness checks, finding little cause for worry.

The RD estimates tell us whether the gains from the VAT reform ultimately fall on consumers or producers. However, the distributional effects also depend on the extent to which poor and rich households are affected by the pass-through to consumer prices. Using survey data on consumer expenditure, we perform a first-order approximation of the distributional effects. An advantage of this approach is that it simply requires information on the price changes and the pre-reform expenditure patterns of the households. However, the VAT reform generated substantial rather than marginal changes in food prices. In such cases, substitution effects can be non-trivial, as consumers substitute towards relatively cheaper goods. The first-order approximations ignore these effects, and therefore, can be seriously biased (see e.g. Banks, Blundell, and Lewbel, 1996). To address this concern, we use expenditure data to estimate the Almost Ideal (AI) demand system. This allows us to incorporate behavioral responses in estimating the compensating variation of the changes in prices associated with the VAT reform.

The insights from our empirical results may be summarized with two broad conclusions. First, the VAT on food items is completely shifted to consumer prices, implying that producers bear none of the tax burden. By comparison, there appears to be little spillover effects of the VAT reform to the consumer prices of most non-food items. Second, lowering the VAT on food substantially attenuates inequality in consumer welfare. This reduction in inequality is partly because poor households have a higher expenditure share on food prior to the reform, but also because households adjust their spending patterns in response to prices changes. By comparison, the usual first-order approximation of the distributional effects, which ignores behavioral responses, seriously understates the redistributive nature of the VAT reform.

Our findings have implications for recent proposals for tax reforms. For example, the Mirrlees Review (2012) sets out a comprehensive proposal for tax reform in the United Kingdom. A key element of the reform package is to broaden the base for VAT, in part by removing the zero rating for food. Arguing that there is little credible evidence to draw on, the Mirrlees Review assumes the incidence of VAT is fully on consumer prices. Atkinson (2013) questions the reform proposal, stressing that until direct evidence is available, "we should remain agnostic about the strength of the optimal tax argument for extending VAT to food" (p. 6). Our paper helps to address this issue by providing transparent and credible identification of the incidence of VAT taxes on food.

Our paper is primarily related to two recent empirical studies on the incidence of VAT. Carbonnier (2007) studies two VAT reforms in France which reduced the rates on new car sales and on housing repair services. He uses the variation in consumer prices across goods and over time to estimate the pass-through of these VAT reforms. His estimates suggest a majority of the tax burden is paid by consumers, especially in the competitive market for housing repair services. Kosonen (2015) analyze the incidence of VAT in the context of hairdressing services in Finland. He uses a difference-in-differences strategy where the control group consists of beauty salons, day spas and massage services. His estimates suggest the tax burden on hairdressing services is shared between consumers and producers.

Our paper expands on this research in several important ways. First, our study provides novel evidence on the incidence of a VAT tax system with lower rates on perceived necessities such as food. Second, our data allows us to look at cross-price effects on other goods, and therefore, capture the entire change in the price structure. Third, we quantify the extent to which a lower VAT rate on food redistributes resources from better-off households to less well-off households. To the best of our knowledge, substitution effects have not been incorporated in distributional analysis of VAT reforms. We do so here, and investigate the accuracy of the usual first-order approximation to welfare implications of tax reforms.

Our paper is also related to an empirical literature on the pass-through of sales taxes. Unlike VAT, a sales tax is imposed only at the retail level, which could have important implications for how the tax burden is shared between consumers and producers (see e.g. Anderson, De Palma, and Kreider, 2001b). Poterba (1996) and Besley and Rosen (1999) examined tax shifting in the United States by comparing local sales taxes and consumer prices across areas and over time. Their estimates suggest that consumers tend to pay for sales taxes. In addition, researchers have examined the incidence of per unit (excise) taxes on goods such as tobacco, alcoholic beverages, and gasoline.[1] Economic theory predicts that in markets with imperfect competition, the consumer share of excise taxes could differ from that of VAT (see e.g. Delipalla and Keen, 1992, Anderson, De Palma, and Kreider, 2001a, and Carbonnier, 2014). This theoretical prediction is supported by the empirical evidence in Delipalla and O'Donnell (2001) and Carbonnier (2013).

The remainder of the paper proceeds as follows. In Section 2, we describe our data and discuss the VAT reform and its expected impact. In Section 3, we discuss the RD design, present our main findings on VAT incidence, and report robustness checks. In Section 4, we present the demand system and analyze the distributional effects of VAT. Section 5 concludes.

[1]For example, Doyle and Samphantharak (2008) and Marion and Muehlegger (2011) study tax incidence on gasoline. Their findings point to a complete pass-through to consumer prices. Other studies have examined the consumer share of excise taxes on tobacco and alcohol beverages (see e.g. Young and Bielińska-Kwapisz, 2002, Kenkel, 2005 and DeCicca, Kenkel, and Liu, 2013).

2 Data and Background

2.1 Data sources and summary statistics

Our analysis uses two data sources. The first is a rich data set on retail prices for a representative sample of consumer goods. The data is collected by Statistics Norway and forms the basis for calculating the Norwegian Consumer Price Index (CPI). Every month, Statistics Norway collects information about the consumer prices on a variety of items. In 2001, we observe 250 different food items, there are 545 different retailers reporting prices, adding up to more that 180,000 recorded prices. Because of the detailed nature of the data, it is possible to follow prices on a given item and retailer over time.

Table 1 displays summary statistics for the major consumer goods in 2001. This table shows that the food category consist of 250 different items. In each category, the average consumer price is computed as the weighted average of the retail prices on the items that belong to this category. We follow the procedure used to construct the CPI in the choice of weights and classification of items. For food, the average consumer price is 38 NOK per item. In general, we see that the consumer prices vary considerably within and between the different types of goods.

The other data source we will be using is the Norwegian Consumer Expenditure Survey for the years 1991–2001. In addition to detailed information on each household's expenditure, there is also a rich set of household characteristics, including information on household size, age of household members, gender, marital status, region, labor status, occupation, and household disposable income. We use the same classification of goods for the expenditure data as for the price data. Our sample consists of households in which the household head is between 20 and 70 years old and not self-employed; the sample is top and bottom coded at the 99th and 1st percentile level of the distribution of household income.[2] Throughout the paper, we use sampling weights to produce representative estimates for the corresponding population of households.

Table 2 summarizes the expenditure shares for non-durable goods. As expected, food purchases form the largest share of household expenditure and the expenditure share declines in household income. For example, food purchases make up 28.3 % of household expenditure in the bottom quartile of the household income distributions,

[2] The top and bottom coding reduces the likelihood that outliers create nonlinearities in the budget-share equations.

Table 1: Summary Statistics: Consumer Price Data

	Consumer price		Number of		
	Average	St. dev	Items	Retailers	Obs.
Food	38	50	250	545	180,510
Clothing	398	602	104	522	43,380
Services	142	354	33	365	22,253
HH fuels	1680	2365	25	2602	19,959
Alcohol	45	33	11	388	19,700
Transport	454	1034	28	377	33,493
Other non-durables	265	982	242	1341	126,301

Notes: This table displays summary statistics for non-durable goods in 2001. In each category, the average consumer price is computed as the weighted average of the retail prices on the items that belong to this category. We follow the procedure used to construct the Consumer Price Index in the choice of weights and classification of items.
Data source: Retail prices collected to calculate the Consumer Price Index, Statistics Norway.

whereas the expenditure share on food is only 23.9 % in the top quartile. We see the same pattern for other perceived necessities such as fuel, while the share of household expenditure on goods like clothing and transport increases in household income.

Figure 1 looks closer at the relationship between household income and food expenditure by graphing the Engel curve. This figure provides a nonparametric description of the Engel curve and suggests that a log-linear specification approximates well the food share curve. This result aligns well with previous evidence from developed countries (see e.g. Banks, Blundell, and Lewbel, 1997).

2.2 The VAT reform and institutional details

In Norway, VAT are levied on the sale of goods and services on registered businesses with annual turnover above NOK 50,000 (approximately USD 7,000). VAT applies to all sales, whether to private consumers or other businesses. Under the "invoice-credit" form of the VAT, registered businesses offset the VAT they have been charged on purchases against the liability on their sales, remitting only the net amount due. The result is that no net revenue is collected from the taxation of intermediate goods sales, so that the ultimate base of the tax is final consumption.[3]

Before the VAT reform in 2001, Norway had a VAT rate of 24 % on most goods. Notable exceptions were certain transport services which had reduced rates and newspapers and books with zero ratings. On July 1st, 2001 the Norwegian government reduced

[3]We refer to Crawford, Keen, and Smith (2010) for a detailed discussion of the structure of VAT in OECD countries.

Table 2: Summary statistics: Expenditure Data on Non-Durable Goods

			Household income					
	Full Sample		Bottom quartile		Middle quartile		Top quartile	
Expenditure shares:	Mean	St.dev	Mean	St.dev	Mean	St.dev	Mean	St.dev
Food	0.268	0.12	0.283	0.14	0.276	0.12	0.239	0.11
Clothing	0.103	0.10	0.091	0.10	0.103	0.09	0.117	0.09
Services	0.160	0.12	0.159	0.14	0.154	0.12	0.173	0.12
HH Fuel	0.087	0.06	0.111	0.08	0.084	0.05	0.070	0.04
Alcohol	0.011	0.02	0.012	0.03	0.011	0.02	0.011	0.02
Transport	0.144	0.13	0.129	0.14	0.146	0.12	0.154	0.13
Other non-d.	0.227	0.14	0.214	0.15	0.227	0.14	0.238	0.14
No. of obs	11107		2777		5554		2776	

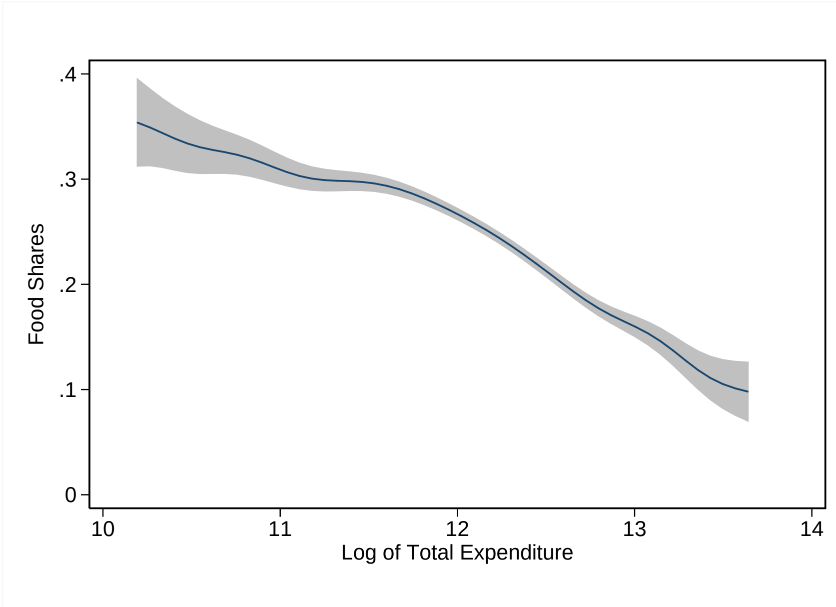
Notes: Column 1-2 shows means and st. deviations in expenditure shares on various goods in the full sample, while columns 3-8 report the same statistics for different income groups. The sample comprises of households in which the household head is between 20 and 70 years old and not self-employed. The sample is top and bottom coded at the the 99th and 1st percentile level of the distribution of household income.
Data source: Norwegian Consumer Expenditure Survey, Statistics Norway.

the VAT on all food items from 24 to 12 percent, while the VAT on other goods did not change. The reduction in the VAT on food items was announced in December 2000.

The key motivation for the reform was that the broad-based VAT system with a uniform rate on most goods places a too large tax burden on poor households. The view that reduced VAT on food items would reduce the tax burden of the poor was based on the standard assumption that the tax would be shifted forward to consumers through price decreases. However, the market structure may affect the incidence of the tax so that consumers may not necessarily get the gains from the reduction in the VAT.

Like most European countries, food retailing in Norway is highly concentrated. The largest chain had in 2004 a market share of 34.6 %, whereas the three largest chains commanded 82 % of the market (Einarsson, 2007). In total there are about 20 different food retailers in Norway that are all linked to one out of the four biggest chains (see Konkurransetilsynet, 2009). Another widely used measure of the level of competition in a marked is the Herfindahl-Hirschman Index (HHI). The US regulatory authorities, like other competition authorities, use HHI in their assessment of mergers. If there is only one firm in the market, the HHI will equal 10,000; if the market is divided equally between a large number of firms the HHI will approach 0; and if the figure is higher than 1800, US law states there is a risk of significant concentration and any potential merger under such circumstances is subjected to careful scrutiny. By this standard, the retail market in Norway was highly concentrated with a HHI of 2600.

Figure 1: Nonparametric Engel Curve for Food Shares



Notes: The solid line shows the estimated relationship between the expenditure share on food and total household expenditure. The relationship is estimated using a nonparametric kernel regression with Gaussian kernel and a mean integrated squared-error optimal smoothing parameter. Household expenditure is defined as yearly total expenditure on non-durables. The shaded area shows the 95 % confidence bands. Sample: households in which the household head is between 20 and 70 years old and not self-employed. The sample is top and bottom coded at the the 99th and 1st percentile level of the distribution of household income.
Data source: Norwegian Consumer Expenditure Survey, Statistics Norway.

The Norwegian food retail market is also highly concentrated in comparison with other countries: Einarsson (2007) reports HHI figures of 1600 in France and Germany, 1800 in the United Kingdom, and as low as 300–500 in Spain.

2.3 Expected reform effects

As in the Mirrlees Review, it is customary to think of the burden of VAT as being borne by consumers in the form of higher after-tax prices, but in theory there is considerable scope for shifting of the tax burden. Indeed, there are plausible circumstances in which consumers bear more than 100 % of the burden or pay little if anything of the VAT. Below, we describe the measures of tax shifting we will use and briefly discuss the challenges to making informative theoretical predictions of the pass-through of the VAT reform.

Tax shifting measures. Let τ denote the VAT rate. The producer (or pre-tax) price is $\frac{p}{1+\tau}$ where p is the consumer (or after-tax) price. The amount of taxes paid per unit sold is $\frac{\tau p}{1+\tau}$. After a change in the VAT rate, the consumer price variation is $\frac{dp}{d\tau}$ and the tax variation is $\frac{d}{d\tau}\left(\frac{\tau p}{1+\tau}\right)$. The consumer share of the change in VAT is then given by

$$cs = \frac{(1+\tau)}{p}\frac{dp}{d\tau}\frac{(1+\tau)}{1+\tau\frac{(1+\tau)}{p}\frac{dp}{d\tau}}. \tag{1}$$

A consumer share of more than one means the tax is over-shifted. If the consumer share is equal to one the tax is fully forward shifted, and consumers bear the full cost of the VAT change. By comparison a consumer share less than one implies the tax is under-shifted, and producers bear some of the cost.

Equation (1) tells us the tax is fully forward shifted when $\frac{dp}{d\tau} = \frac{p}{1+\tau}$, which is equivalent to $\frac{d}{d\tau}\left(\frac{p}{1+\tau}\right) = 0$. This means the tax is fully shifted whenever the producer price does not change as a result of the VAT change. In our setting, this implies that the VAT reduction from 24 % to 12 % on food items is fully forward shifted if consumer prices on food decreases by 9.7 %.

Theory predictions. A priori, it is challenging to credibly predict the pass-through of a VAT reform, as it requires detailed knowledge of the market structure and reliable estimates of demand and supply.

Consider first the benchmark of perfect competition, in which case the consumer price variation is given by

$$\frac{dp}{d\tau} = \frac{p}{(1+\tau)}\frac{1}{1+\frac{\eta_D}{\eta_S}}, \tag{2}$$

where $\eta_S = \frac{p}{(1+\tau)}\frac{1}{S}\frac{\partial S}{\partial p}$ is the supply elasticity evaluated at the producer price and $\eta_D = -\frac{p}{D}\frac{\partial D}{\partial p}$ is the demand elasticity evaluated at the consumer price. Equation (2) shows that even with perfect competition, it is difficult to predict the pass-through of a VAT reform: While over-shifting is not possible, the consumer share can range from 0 to a 100 %. If demand for the taxed good is relatively elastic compared to supply then producers bear most of the tax burden, whereas the consumer share is larger if demand is less elastic than supply.

In our setting, there is strong market concentration and imperfect competition is likely, which make it even more difficult to credibly predict the pass-through of a VAT reform.[4] To see this, suppose there are n firms and each firm produces a variant of a

[4]See Anderson, De Palma, and Kreider (2001b) for a theoretical analysis of the incidence of VAT in an oligopolistic industry with differentiated products and price-setting (Bertrand) firms. Seade (1985) and Delipalla and Keen (1992) provide a theoretical analysis of incidence in the case of an oligopolistic industry with homogenous demand and quantity-setting (Cournot) firms. Weyl and Fabinger (2013) extends the analysis of incidence to a general model of imperfect competition.

differentiated product. Firm i's profit is given by

$$\pi_i = \frac{p_i}{1+\tau_i} D_i\left(p_i; \mathbf{p}_{-i}\right) - c\left(D_i\right)$$

where $c(\cdot)$ is the cost function common for each firm, $D_i(p_i; \mathbf{p}_{-i})$ is the demand for firm i's product as a function of firm i's own consumer price, p_i, and a vector consisting of the other firms' consumer prices, $\mathbf{p}_{-i}$. Further, the function $D_i(p_i; \mathbf{p}_{-i})$ is continuously differentiable, decreasing in p_i and increasing in all elements of $\mathbf{p}_{-i}$. At a Bertrand-Nash equilibrium, assuming an interior solution, each firm will set a price p_i, given $\mathbf{p}_{-i}$, such that the first-order condition is satisfied:

$$(p_i - \tilde{c}_i)\frac{\partial D_i(p_i; \mathbf{p}_{-i})}{\partial p_i} + D_i(p_i; \mathbf{p}_{-i}) = 0, \tag{3}$$

where $\tilde{c}_i = (1+\tau_i)c_i$ denotes effective cost.

The effects of an increase in the VAT on own producer prices are given by total differentiating the first-order conditions given in (3)

$$\frac{dp_i}{d\tau_i} = \frac{p_i}{1+\tau_i}\frac{1+\varepsilon_{ii}}{2\varepsilon_{ii}-E_{ii}} - \sum_{j\neq i}\frac{p_i}{p_j}\frac{\varepsilon_{ii}E_{ij}+E_{ij}-\varepsilon_{ij}}{2\varepsilon_{ii}-E_{ii}}\frac{dp_j}{d\tau_i} \tag{4}$$

where

$$\frac{dp_j}{d\tau_i} = -\sum_{k\neq j}\frac{p_j}{p_k}\frac{\varepsilon_{jj}E_{jk}+E_{jk}-\varepsilon_{jk}}{2\varepsilon_{jj}-E_{jj}}\frac{dp_k}{d\tau_i} \quad \text{for } j \neq i,$$

where we have substituted c_i from the first-order condition (3), $\varepsilon_{ij} = \frac{\partial D_i}{\partial p_j}\frac{p_j}{D_i}$ is the own or cross price elasticity of demand, and $E_{ij} = \frac{\partial^2 D_i}{\partial p_i \partial p_j}\frac{p_j}{\partial D_i/\partial p_i}$ is the elasticity of the slope of the demand curve.

While under perfect competition the pass-through rate is entirely determined by the elasticity of supply and demand, the predictions are more complicated under imperfect competition. Equation (4) shows that in particular, the curvature of demand also plays a role. Consider, for example, the case in which an increase in the VAT of good i does not lead to a price change for good j, $\frac{dp_j}{ds_i} = 0 \; \forall j \neq i$. In this case, equation (4) is equal to

$$\frac{dp_i}{d\tau_i} = \frac{p_i}{(1+\tau_i)}\frac{1+\varepsilon_{ii}}{2\varepsilon_{ii}-E_{ii}}.$$

From the consumer share equation (1), it follows that the consumer share exceeds a 100 % if the curvature of the demand function is such that $E_{ii} > \varepsilon_{ii} - 1$. Because standard demand forms restrict this curvature in ways that have little empirical or theoretical foundation (see e.g. Fabinger and Weyl, 2015), imperfect competition makes it particularly difficult to credibly predict the pass-through rate.

Taken together, the challenges of making informative theoretical predictions motivate our empirical analysis of the incidence of VAT in a setting with plausibly exogenous variation in tax rates.

3 Incidence of the VAT reform

3.1 Research design

On July 1st, 2001 the VAT on all food items was reduced from 24 to 12 percent, while the VAT on non-food items remained at 24 percent. This sharp change in the VAT policy provides an attractive setting to analyze the pass-through of commodity taxes using a RD design that compares consumer prices just before (i.e. the control group) and after (i.e. the treatment group) the reform date.

Our RD design can be described by the following regression model:[5]

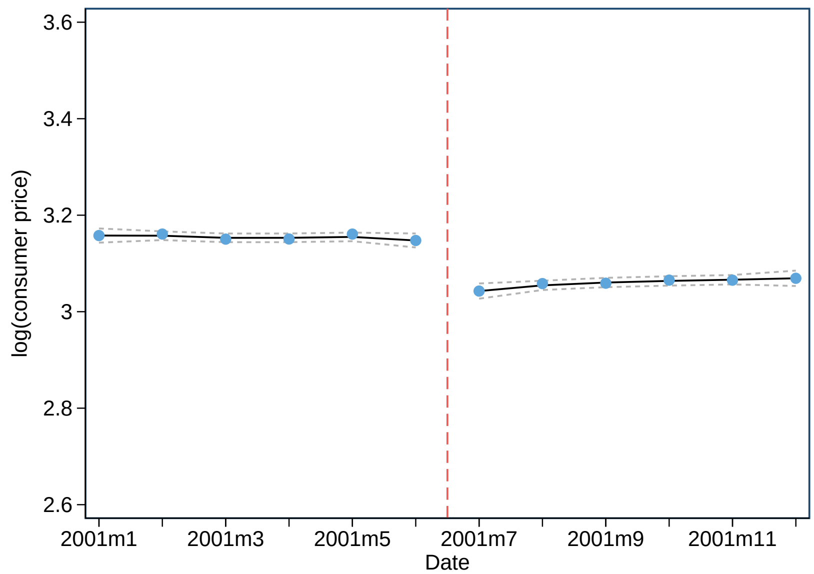
$$y_{it} = \alpha + \mathbf{1}\left\{t \geq c\right\}\left[g_l\left(t-c\right) + \lambda\right] + \mathbf{1}\left\{t < c\right\} g_r\left(c-t\right) + e_{it} \tag{5}$$

where y_{it} denotes log consumer price on good i in month t, c is the reform date (July 1st, 2001), e_{it} is an error term, and g_l, and g_r are unknown functions. The key identifying assumptions are that prices do not change in anticipation of the VAT reform and that other factors determining consumer prices evolve smoothly around the reform date. Under these assumptions, we can consistently estimate the parameter λ, which gives the impact of the VAT reform on the consumer price of good i. Below, we challenge the identifying assumptions of the RD design, finding little cause for worry.

To implement the RD design, we need to specify g_l and g_r and decide on the window on each side of the reform date. Our first specification uses a local linear regression with triangular kernel density and 2 months of bandwidth on each side of the reform date. Our second specification uses a window of just one month on each side of the reform date. Because we have monthly data on consumer prices, the RD model is then equivalent to a first-difference (FD) model: the average consumer prices in June 2001

[5]See e.g. Lee and Lemieux (2009) for a detailed discussion of the RD design.

Figure 2: Evolution of Consumer Price on Food over Time



Notes: Each observation is the average consumer price for food reported on the 15th each month. The dashed vertical line denote the reform date. The solid lines are from a local linear regression with triangular weights on monthly consumer price data. The dashed lines represent the 95% confidence intervals. The y-axes are scaled to $+/-$ .5 st.dev. of the mean consumer price.

is compared to the average consumer prices in July 2001.[6]

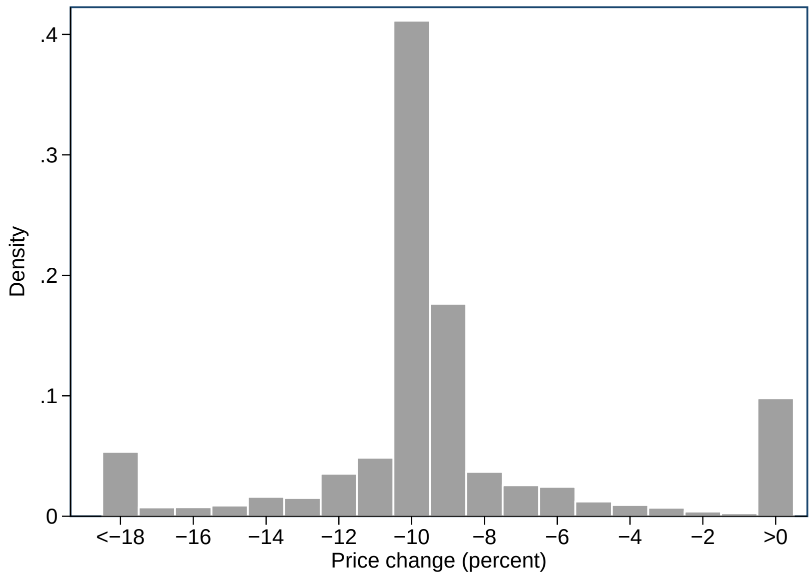
3.2 Graphical evidence

A virtue of the RD design is that it provides a transparent way of showing how the reform impact is identified. To this end, we begin with a graphical depiction before turning to a more detailed regression-based analysis.

Figure 2 shows both the unrestricted and the estimated monthly means of consumer prices for food items during 2001. The estimated monthly means come from a local linear regression with a triangular kernel applied to each side of the reform date: While the regression lines better illustrate the trends in the data and the size of the jump at the reform date, the unrestricted means indicate the underlying noise in the data. The figure shows evidence of a sharp decline in the average food price at the time of the reform, suggesting that the tax is heavily shifted to consumer prices. To further zoom in on how individual prices reacted around the reform date, Figure 3 shows a histogram

[6]When following the procedure of Imbens and Kalyanaraman (2012), we estimate an optimal bandwidth of about 1 month. This result motivates the choice of bandwidths in our setting. The results barely move if we widen the bandwidth to 4 or 6 months.

Figure 3: Histogram over Percentage Change in Consumer Price of Food from June 2001 to July 2001



Notes: Histogram of consumer price changes in percent from June 15th 2001 to July 15th 2001. The width of each bin is equal to 1 percent.

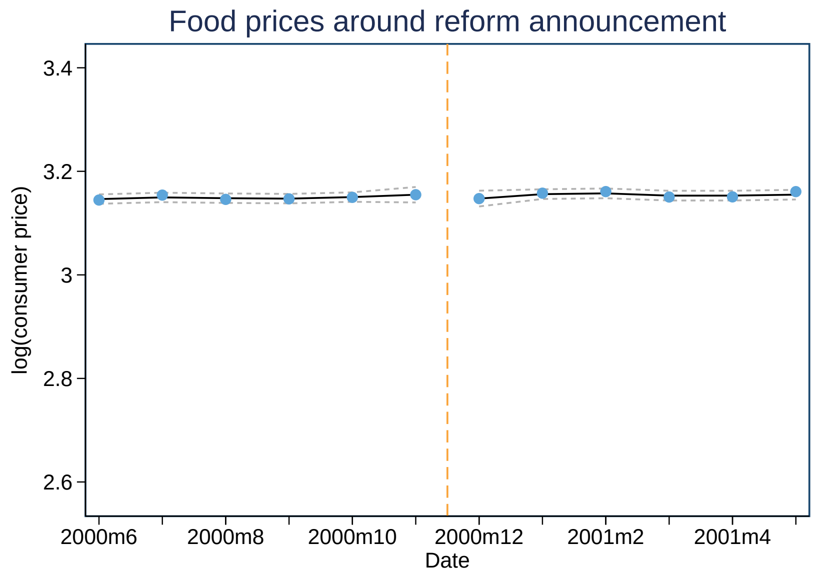
of the percent price change in consumer prices on food items from June 2001 to July 2001. The histogram shows that 90 % of food prices are lower in July 2001 compared to June 2001, illustrating that there was a downward shift in the price for most food items. Moreover, 80 % of the consumer prices were lowered with 8 percent or more.

3.3 Threats to identification

The validity of our RD design requires that prices do not change in anticipation of the VAT reform. Figure 2 shows no evidence of changes in food prices prior to the VAT reform, suggesting that firms did not change food prices in anticipation of the reform. Additionally, food prices barely move in the months following the reform suggesting that firms respond swiftly to the change in VAT. This suggests that the estimated short run effect from June to July is representative for the long run effect of the reform.

To further challenge the assumption on anticipation effects, Figure 4 shows the unrestricted and estimated monthly means of consumer prices for food items during the months before and after December 2000, wehn the reform was first announced. The figure shows no evidence of a discontinuous change in prices around this date. Moreover, there is no evidence of discontinuous changes in prices either in the months

Figure 4: Evolution of Consumer Price on Food over Time



Notes: Each observation is the average consumer price for food reported on the 15th each month. The dashed vertical line denote the date the reform was announced. The solid lines are from a local linear regression with triangular weights on monthly consumer price data. The dashed lines represent the 95% confidence intervals. The y-axes are scaled to +/− .5 st.dev. of the mean consumer price.

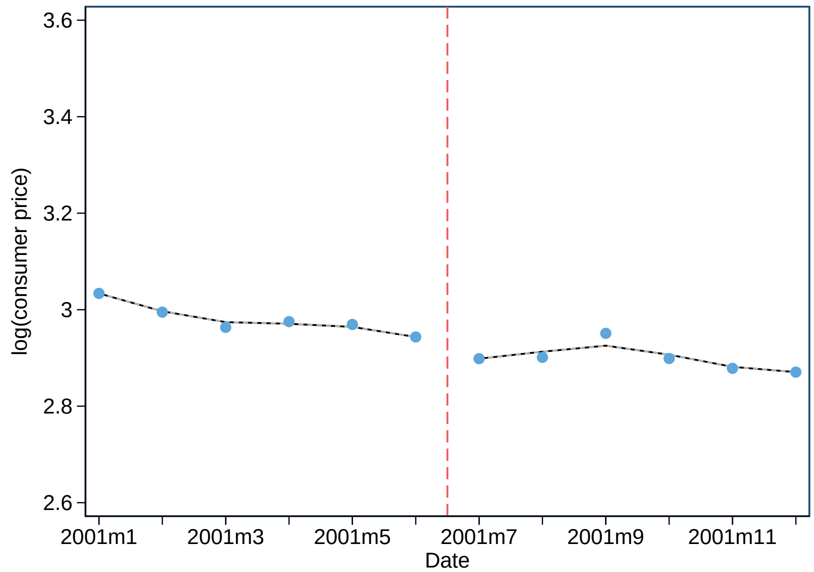
before or after the announcement date, further suggesting that prices did not change in anticipation of the reform.

A second assumption behind our RD design is that consumer prices would have evolved smoothly around the reform date in the absence of the policy change. This continuity condition implies that other observable determinants of consumer prices should have the same distribution just before and after the reform. For simplicity, we consider a scalar representation of the observable determinants, given by the predictions from a regression of food prices on a flexible set of lagged values of oil prices and exchange rates.[7] The covariates are jointly predictive of food prices (with an F-statistic of 34). Figure 5 displays the predicted price in each month, showing no evidence of discontinuous changes in observables around the time of the reform. This implies that the discontinuity in consumer prices of food observed around the reform date is not driven by discontinuities in the covariates.[8] Indeed, when looking at each covariate

[7]We control for oil prices to proxy for energy and transportation prices. Since the price of imported food is likely to depend on the exchange rate, we control for the import weighted exchange rate as well as the exchange rates of Norway's key trading partners.

[8]Figure 5 shows the predicted consumer price with a four week lags for the covariates. The results are robust to using shorter and longer lags.

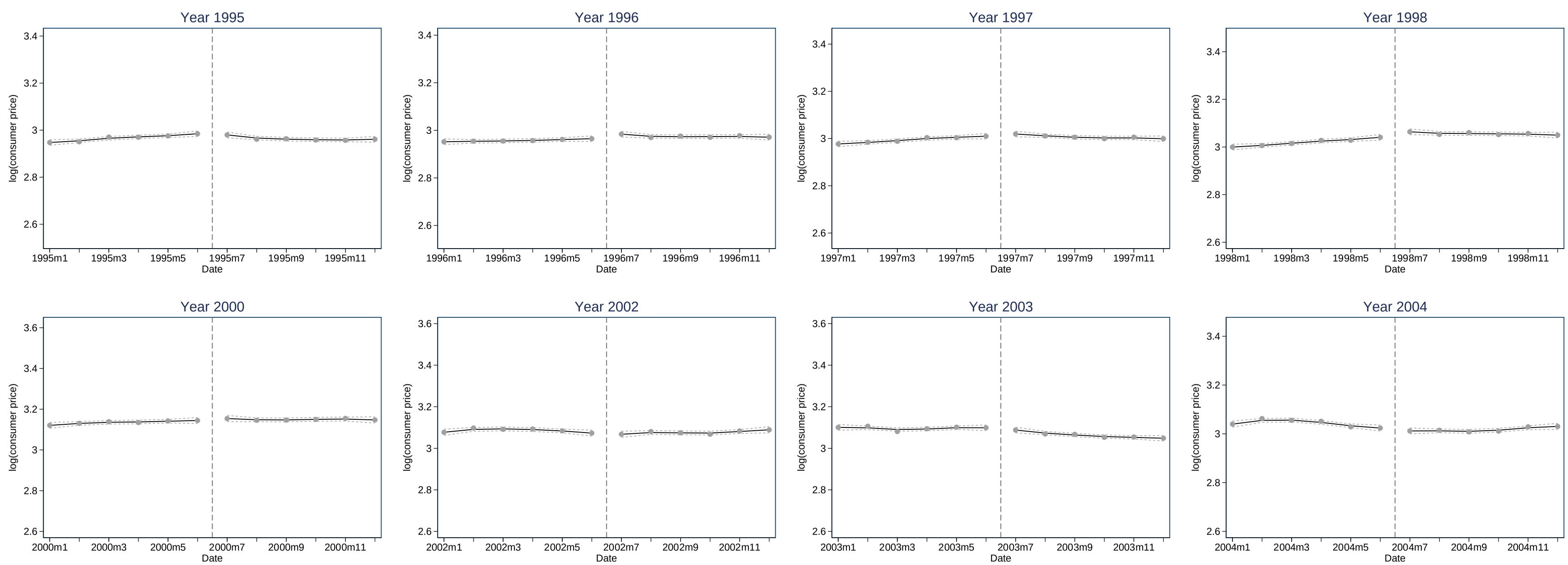
Figure 5: Evolution of Predicted Consumer Price on Food over Time



Notes: Predicted food consumer prices is given by a regression of food consumer prices on oil prices (brent), euro/NOK, SEK/NOK, GDP/NOK and an import weighted exchange rate. The covariates are lagged four weeks and each observation is the average predicted consumer price for food reported on the 15th each month. The dashed vertical line denote the reform date. The solid lines are from a local linear regression with triangular weights on the predicted food consumer prices. The y-axes are scaled to +/− .5 st.dev. of the mean predicted consumer price.

separately, we find no evidence of any discontinuous changes around the time of the reform.

Figure 6: Evolution of Consumer Prices on Food, Non-Reform Years



Notes: Each observation is the average consumer price for food reported on the 15th each month. The dashed vertical line denote July 1st, 2000. The solid lines are from a local linear regression with triangular weights on monthly consumer price data. The dashed lines represent the 95% confidence intervals. The y-axes are scaled to +/− .5 st.dev. of the mean consumer price. Year 1999 is omitted since Statistics Norway did a restructuring of the classification of goods in August 1999 (see Official Statistics of Norway, 2001)

Another threat to the continuity condition is that seasonality or month effects could generate discontinuous changes in food prices around the time of the reform. Although Figure 2 shows no sign of changes in food prices in the months before and after the reform, we cannot rule out that there would have been a discontinuous change in July 2001 in the absence of the policy change. To investigate the possibility of a July-specific month effect, Figure 6 shows the unrestricted and the estimated monthly means of consumer prices for food items during the 6 years prior to the reform and 3 years after the reform.[9] Overall, there seem to be no systematic month-of-July discontinuity in the data; this finding is reassuring because there were no reform in the VAT system from June to July during these years.

3.4 Regression estimates

Having shown the raw patterns on the variables of interests around the reform date we now turn to regression-based estimates.

Table 3 shows the point estimate and standard error of the impact of the VAT reform on consumer prices for food. The first column reports the result from the FD model, comparing consumer prices in June and July 2001. The point estimate suggests the reform reduced food prices by 10.5 percent. By way of comparison, full shifting would imply a reduction in food prices of 9.7 percent. This suggests that VAT on food items are completely shifted – or even slightly over-shifted – to consumer prices. Indeed, the FD estimate is sufficiently close to -9.7 percent that we cannot reject the null hypothesis of full shifting.

The second column of Table 3 reports the RD estimates with 2 months of bandwidth on each side of the reform date. The key difference between the FD model and the RD model is their assumptions regarding how the prices would have changed over time in the absence of the reform. The FD specification takes the consumer price on a good in June 2001 as a counterfactual for the price on the same good in July 2001. If there were secular changes in prices over this time period, the FD model would produce biased estimates of the effect of the VAT reform, because the price in June 2001 would be an inappropriate counterfactual for the price in July 2001. In this type of "smoothly contaminated" experiment, the RD specification uses the observed trends in prices on each side of the reform date to construct an appropriate counterfactual. As is evident

[9] Year 1999 is omitted since Statistics Norway did a restructuring of the classification of goods in August 1999 (see Official Statistics of Norway, 2001)

Table 3: Reform Effects on Consumer Prices of Food

	FD	RD		
Dep. Variable: Log Consumer Price	(1)	(2)	(3)	(4)
Food and non-alcoholic beverages	-0.105***	-0.106***	-0.103***	-0.106***
	(0.008)	(0.014)	(0.005)	(0.006)
Item F.E.	No	No	Yes	Yes
Month Effects	No	No	No	Yes
Hypothesis tests:		P-value		
H_0: $\lambda = -0.097$ vs H_1: $\lambda \neq -0.097$	0.315	0.533	0.226	0.142
H_0: $\lambda \geq -0.097$ vs H_1: $\lambda < -0.097$	0.158	0.267	0.113	0.071

*** p<0.01, ** p<0.05, * p<0.1

Notes: The coefficient in column (1) is estimated using the FD model with log consumer prices as dependent variable. The coefficients in column (2) – (4) are estimated using a RD model with log consumer prices as dependent variable, triangular weights and two months bandwidth. Column (2) report the results with no controls, column (3) includes item fixed effects, and column (4) also control for possible month effects using a DiD strategy. The standard errors are clustered at the firm level and robust to heteroskedasticity. We report p-values for the two-sided test that the VAT for food items is fully shifted and from one-sided tests of the null hypothesis that the VAT for food items is undershifted to consumer prices.

from the second column of Table 3, the RD estimates are very similar to the FD estimates for food. The point estimate suggests the reform reduced food prices by 10.6 percent, which supports the conclusion that the VAT is completely shifted to consumer prices.

3.5 Specification checks

To increase the confidence in our identification strategy, we now show that our regression estimates are robust to several specification checks.

We begin by adding a full set of item fixed effects to the regression model. The third column of Table 3 report the results. We find that the estimates change little when including fixed effects, suggesting the estimated reform effects are not driven by changes in the composition of commodities over time. However, including the fixed effects reduces the residual variance and is thus a useful way to gain precision.

Next, we estimate a difference-in-differences (DiD) specification of the RD model. The main motivation for this robustness check is that seasonality or month effects could generate discontinuous changes in consumer prices. The DiD specification exploits the fact that there was no change in the VAT rates in 2000: significant changes in the food prices in July 2000 would therefore be unrelated to the VAT system and should instead capture month effects. The DiD estimate is obtained by separately estimating equation (5) using data from 2000 and subtract it from the RD estimate of the VAT reform.

The fourth column of Table 3 reports estimates from the DiD specification. The point estimate barely move from column (3) to column (4), but in this specification we are able to rule out under-shifting at a 10 % level of significance. Additional robustness checks on the DiD specification of the RD model is provided in Appendix A. Table A1 report results where the potential month of July effect is captured by estimating equation (5) separately for multiple years between 1994 and 2004. The results further suggest that month effects do not confound the conclusions drawn about the pass-through of the VAT reform.

We further explore the assumption that consumer prices would have evolved smoothly around the reform date in the absence of the policy change by including the covariates oil prices and exchange rates (lagged with 1 or 4 weeks) to equation (5). Table A2 in Appendix A shows that the estimates of λ does not change appreciatively by adding these observable determinants of consumer prices.

Lastly, we examine whether prices change in anticipation of the VAT reform. The change in tax rates was announced in December 2000, and it is conceivable that firms or consumers adjust their behavior prior to the reform date. However, the graphical evidence presented in Figure 2 and 4 showed no sign of changes in food prices outside the reform window nor around the announcement date. Further evidence against anticipation effects is provided by splitting the set of food items into fresh and storable food. The idea is that any anticipation effect should be stronger for storable food than for fresh food, and as a result, put downward pressure on the estimated pass-through of the VAT reform for storable food. However, when estimating the RD model separately for storable and fresh food, we find very similar reform effects. The RD estimates of the reform effect is -0.109 (s.e.=0.0047) for storable goods and -0.098 (s.e.=0.0067) for fresh goods.

3.6 Heterogeneity and cross price effects

Appendix Table A3 explores heterogeneity across different food items as well as cross-price effects on non-food items.

Panel A in Table A3 present estimates of the impact of the VAT reform for different types of food items, based on separate regressions for 6 subcategories.[10] The point

[10] We base our categories of food on the classifications used by United Nations Statistics Division and categorized food items using the COICOP classification.COICOP is an abbreviation of Classification of Individual Consumption According to Purpose and the subgroups corresponds to COICOP classes 111-119 and 121-122.

estimates are all negative and statistically different from zero at conventional levels. Further, the estimated pass-through is broadly similar across the different categories of food items. Indeed, for all but one category of food, the point estimates are consistent with full-shifting or slight over-shifting.

Panel B in Table A3 explores whether the VAT reform had an impact on consumer prices for goods that where not directly affected by the reform. The table shows the point estimate and standard error of the impact of the VAT reform from separate regressions on 6 non-food categories of non-durable goods. As before, the dependent variable is the log consumer price in all of these regressions. For most of the categories, we find no evidence of cross-price effects of the VAT reform. In terms of magnitudes, the most notable change is for the category services, although the estimated price increase is not significant at conventional levels.

4 Distributional effects of VAT reform

The RD estimates demonstrated that the gains from the VAT reform ultimately fell on consumers rather than producers. In this section, we investigate how the pass-through to consumer prices affected the welfare of poor and rich households.

4.1 Model and estimates of demand system

Demand system. To study the welfare effects of the VAT reform, we apply the AI demand system first proposed by Deaton and Muellbauer (1980). In the AI demand system, preferences belong to the Price-Independent Generalized Logarithmic (PIGLOG) class (Muellbauer, 1976) and they are defined by the expenditure function

$$\log c\left(u, \mathbf{p}\right) = \left(1-u\right)\log a\left(\mathbf{p}\right) + u\log b\left(\mathbf{p}\right) \tag{6}$$

where u is the indirect utility and $\mathbf{p}$ is a vector of prices of n goods. The functions $a\left(\mathbf{p}\right)$ and $b\left(\mathbf{p}\right)$ are specified by the following functional forms:

$$\log a\left(\mathbf{p}\right) = \alpha_0 + \sum_{i=1}^{n}\alpha_i \log p_i + \frac{1}{2}\sum_{i=1}^{n}\sum_{j=1}^{n}\gamma_{ij}^{*}\log p_i \log p_j, \tag{7}$$

and

$$\log b\left(\mathbf{p}\right) = \log a\left(\mathbf{p}\right) + \prod_{i=1}^{n} p_i^{\beta_i}. \tag{8}$$

In line with Browning and Meghir (1991) and Blundell, Pashardes, and Weber (1993), we use a two-stage budgeting framework. Preferences are characterized such that, in each period t, household h makes decisions on how much to consume of a set of non-durable commodities, conditional on household characteristics and the consumption level of a second group of commodities with possibly less flexible demand. The commodities we model directly ($\mathbf{q}$) are food, clothing, services, household fuel, alcohol, transport, and other non-durable goods. The second group contains housing, some durables, and labor-market decisions which together with household characteristics, is represented by $\mathbf{z}$. Household utility is defined over $\mathbf{q}_t^h$ for household h in period t conditional on the set of demographics and other conditioning variables $\mathbf{z}_t^h$.

The first stage of the budgeting framework is to allocate expenditures to commodities $\mathbf{q}_t^h$, denoted by m_t^h. In the second stage of the budgeting framework, households decide on how much to spend on food, clothing, services, household fuel, alcohol, transport and other non-durable goods conditional on m_t^h. More specifically, inserting for (7) and (8) in (6) and applying Roy's identity gives the second stage budget shares

$$w_{it}^h = \alpha_{it}^h + \sum_j \gamma_{ij} \ln p_{jt} + \beta_{it}^h \ln \left[\frac{m_t^h}{a(\mathbf{p})} \right] \tag{9}$$

where w_{it}^h is household h's budget share of good i, and p_{jt} is the price of good j at time t. The term $\left[m_t^h / a(\mathbf{p}) \right]$ represents relative income with $a(\mathbf{p})$ being a price index. Household preferences are incorporated by allowing the constant α_{it}^h to depend on household characteristics, z_{kt}^h,

$$\alpha_{it}^h = \alpha_i + \sum_k \alpha_{ik} z_{kt}^h + \sum_k \delta_k T_{kt},$$

in which we have also added a full set of indicator variables for year and season T_{kt}.

Both the indirect utility function and the demand functions for each good that arise from Equations (6) – (8) are linear in the log of total expenditure. Figure (1) in Section 2 examined this assumption for our main commodity of interest: food items. This figure provided a nonparametric description of the Engel curve and shows that the linear model seems to be a reasonable approximation for the food share curve.[11]

Estimation procedure. To consistently estimate (γ_{ij}, β_i) for every commodity i,

[11] This is consistent with what Banks, Blundell, and Lewbel (1997) find using British household data.

we use the two-step estimation method of Browning and Meghir (1991) and Blundell, Pashardes, and Weber (1993). This estimation method incorporates a set of theoretical within-equation and cross-equation restrictions. Furthermore, it accounts for the endogeneity in m_t^h in the budget share equations.

The first step imposes the within-equation restrictions of adding-up and (zero-degree) homogeneity on (9) by expressing all prices relative to the price of "other" goods together with excluding this equation from the system. Each equation is estimated separately, allowing for endogeneity in m_t^h as well as heteroscedasticity in the error terms. We use GMM to obtain unrestricted consistent estimates for each equation where we instrument m_t^h in each budget share equation with total household income. Additionally, an iterative method is applied where one takes advantage of the conditional linearity of equation (9) given $a(\mathbf{p})$. That is, given $a(\mathbf{p})$, the system is linear in parameters, and this suggests a natural iterative procedure conditioning on an update of $a(\mathbf{p})$ at each iteration.[12]

The second step imposes the cross-equation restriction of symmetry. Let ϕ (ϕ^*) denote the vector of unrestricted (restricted) parameters obtained in the step outlined above. The cross-equation restrictions on ϕ can then be expressed as

$$\phi = \mathbf{K}\phi^*, \tag{10}$$

where $\mathbf{K}$ is a matrix of rank $l - m(m-1)/2$ and l is the number of unrestricted parameters in the demand equation system. To impose these restrictions the MCS method chooses an estimator $\hat{\phi}^*$ so as to minimize the quadratic form

$$\hat{\phi}^* = \arg\min \left[\hat{\phi} - \mathbf{K}\phi^*\right]' \mathbf{\Sigma}_\phi^{-\mathbf{1}} \left[\hat{\phi} - \mathbf{K}\phi^*\right] \tag{11}$$

where $\hat{\phi}$ is the vector of unrestricted parameter estimates and $\mathbf{\Sigma}_\phi$ is its estimated covariance matrix. The estimated covariance matrix of the symmetry constrained estimator is given by $(\mathbf{K}'\mathbf{\Sigma}_\phi^{-\mathbf{1}}\mathbf{K})^{-1}$.

Parameter estimates and elasticities. When estimating the individual household expenditure functions given by Equation (9), we rely on cross sectional data on household expenditures. These data does, however, not contain information on prices. To construct prices we use the consumer price data and aggregate up to commodity-

[12] As a first approximation to $a(\mathbf{p})$, we compute household-specific Stone price indices.

quarter-year specific indicies.[13]

The price and income coefficients that correspond to the γ_{ij} and β_i parameters in Equation (9) are given in Table A4 in the Appendix. The estimated income parameters suggest that food and fuel are necessities, while clothing, services and transport are luxury goods. Consistent with the findings in Blundell, Pashardes, and Weber (1993), we find a positive relationship between the price of food and the expenditure share on food. However, this and some of the other parameters are too noisily estimated to draw firm conclusions about the values in the population of Norwegian households.

To interpret the parameters, it is useful to consider the implied elasticities. The budget elasticity at reference price is defined as

$$\varepsilon_i^h = \frac{\beta_i}{w_i^h} + 1.$$

The budget elasticities will vary with family composition since the predicted expenditure share w_i^h varies across households. The uncompensated demand elasticity of good i w.r.t. the price of good j at reference prices is given by

$$\varepsilon_{ij}^u = \frac{1}{w_i^h}\left[\gamma_{ij} - \beta_i\left(\alpha_j + \sum_{k=1}^{N}\gamma_{jk}\ln p_k\right)\right] - \delta_{ij}$$

where δ_{ij} is the Kronecker delta. Again, we see that the elasticities vary across households due to different budget shares. The compensated price elasticity is

$$\varepsilon_{ij}^c = \varepsilon_{ij}^u + \varepsilon_i^h w_j^h,$$

where the compensated price elasticity allows the consumers to revise their expenditure decision made in stage one of the budgeting framework when the price of good j changes. The elasticities are reported in Table 4. These elasticities are calculated for each household individually, and then a weighted average is constructed, with the weights being equal to the household's share of total sample expenditure of the relevant good. As expected, the uncompensated and compensated own price elasticities are negative for all goods. In terms of magnitudes, the elasticity estimates are in line

[13]The indices are constructed using the same aggregation method as the CPI (see Official Statistics of Norway, 2001). The demand analysis exploits both the cross-sectional variation in prices as well as the temporal variation in prices within goods (conditional on covariates). While this is standard in demand analysis (see e.g Blundell, Pashardes, and Weber 1993), one may worry about correlated unobservables.

with the findings of Banks, Blundell, and Lewbel (1997)

Table 4: Implied Elasticities using the Parameters of the Demand System

A: Budget Elasticities:

	Commodity						
	Food	Clothing	Services	Fuel	Alcohol	Transport	Other
	0.676	1.279	1.194	0.256	0.980	1.231	1.099

B: Uncompensated Cross Price Elasticities:

	Commodity						
Price	Food	Clothing	Services	HH fuel	Alcohol	Transport	Other
Food	-0.736	0.103	-0.121	-0.014	-0.001	0.104	-0.011
Clothing	0.067	-1.289	0.129	0.047	0.111	-0.412	0.067
Service	-0.293	0.096	-0.928	-0.024	-0.009	-0.018	-0.017
HH fuel	0.055	0.197	0.099	-1.105	0.013	0.236	0.250
Alcohol	-0.054	0.768	-0.065	0.023	-1.274	-0.168	-0.210
Transport	0.0194	-0.281	-0.026	0.034	-0.020	-1.093	0.137
Other	-0.107	0.049	0.003	0.013	-0.016	0.111	-1.153

C: Compensated Cross Price Elasticities:

	Commodity						
Price	Food	Clothing	Services	HH fuel	Alcohol	Transport	Other
Food	-0.520	0.172	-0.025	0.040	0.007	0.192	0.135
Clothing	0.357	-1.063	0.328	0.128	0.123	-0.235	0.362
Services	-0.033	0.220	-0.637	0.052	0.003	0.154	0.241
HH fuel	0.143	0.217	0.130	-1.061	0.016	0.262	0.293
Alcohol	0.200	0.867	0.098	0.092	-1.232	-0.030	0.009
Transport	0.282	-0.161	0.161	0.110	-0.009	-0.776	0.392
Other	0.132	0.160	0.158	0.082	-0.005	0.252	-0.779

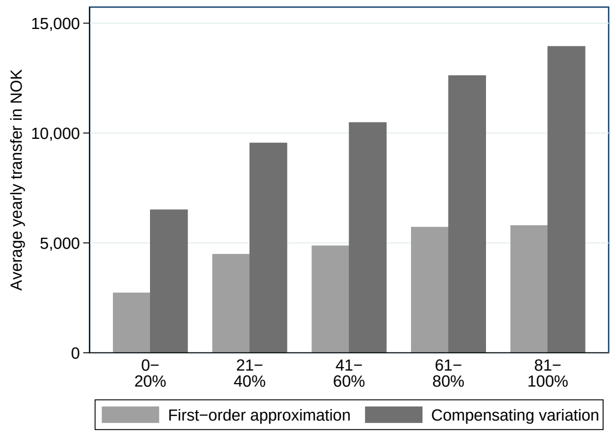
Notes: Calculated elasticities using the γ-symmetry constrained Almost Ideal estimates reported in Appendix Table A4. Sample: households in which the household head is between 20 and 70 years old and not self-employed. The sample is top and bottom coded at the the 99th and 1st percentile level of the distribution of household income.

4.2 Distributional effects

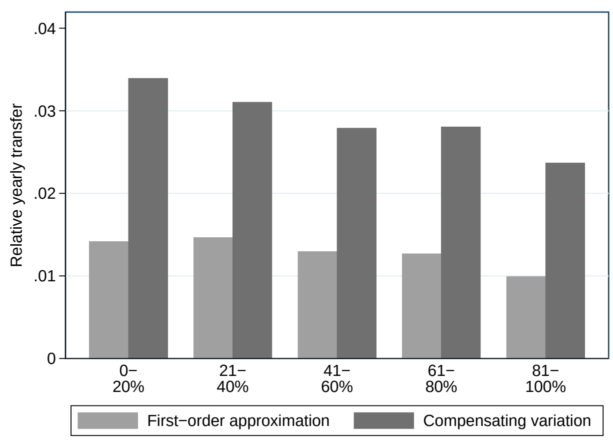
First-order approximation. We begin with a first-order approximation of the distributional effects, ignoring any behavioral responses to the reduction in prices. To this end, we multiple the pre-reform expenditure levels on the various goods with the RD estimates of the prices changes. After the VAT reform, households can buy the same bundle of goods at a lower price. Assuming no behavioral responses and ignoring cross-price effects, the VAT reform amounts to a cash transfer equal to 10.6 % of a

Figure 7: Size of Transfer over Household Income Quintiles

(a) Average transfer



(b) Relative transfer



Notes: The transfers are calculated using the direct price response to the VAT reform reported in column 4 of Table 3. First-order approximation is defined as 10.6 % of the household's pre-reform expenditure on food. Compensating variation is defined as the difference in the cost functionsc $\left(\mathbf{p}^0, z, u^0\right) - c\left(\mathbf{p}^1, z, u^0\right)$, where the post-reform cost function is evaluated at the pre-reform indirect utility level. Relative transfer is defined as transfer/household income. Sample: households in which the household head is between 20 and 70 years old and not self-employed. The sample is top and bottom coded at the the 99th and 1st percentile level of the distribution of household income.

household's expenditure on food.

Figure 7 illustrates the variation in the size of this transfer across households over income quintiles. The lighter bars in panel (a) shows the average size of the transfer at each quintile. The average transfer to the poorest 20 percent in our sample is 2,725 NOK/year. By comparison, the average transfer to the 20 percent richest households in our sample is 5,792 NOK/year, which is more than twice as large as the transfer to the

20 percent poorest households. Similarly, the lighter bars in panel (b) shows the size of the transfer as a fraction of income. The relative transfer decreases with household income. Taken together, the evidence from the first-order approximation suggest that richer households receives a larger absolute transfer from the reform, whereas poorer households received more relative to their income.

Allowing for behavioral responses. There is an obvious attraction to simply using information on observed expenditure patterns to assess the welfare implications of the VAT reform. No response parameters are required, and therefore the analysis is not subject to estimation error in own- or cross-price demand elasticities.[14] However, the VAT reform generated substantial rather than marginal changes in food prices. In such cases, substitution effects can be non-trivial, as consumers substitute towards relatively cheaper goods. The first-order approximations ignore these effects, and therefore, can be seriously biased (see e.g. Banks, Blundell, and Lewbel, 1996).

To allow for behavioral responses, we use the parameter estimates of the AI model to calculate the indirect utility of the households from Equations (6) - (8). The transfer to a given household with characteristics z is measured as the compensating variation, given by the difference in the cost functions $c\left(\mathbf{p}^0, z, u^0\right) - c\left(\mathbf{p}^1, z, u^0\right)$, where the post-reform cost function is evaluated at the pre-reform indirect utility level. This welfare measure tells us the maximum amount of income a household is willing to pay for the VAT reform.

Distributional effects

We begin with a graphical depiction of the distributional effects of the VAT reform, before quantifying its impact on inequality. The darker bars in panel (a) in Figure 7 shows that the magnitude of the compensating variation increases with household income. As rich households consume more food, the willingness to pay increases with total expenditure. Panel (b) in Figure 7 complements by showing the relative size of the compensating variation. This figure reveals that richer households are willing to pay a smaller fraction of their total income for the VAT reform.

To summarize the impact of the VAT reform on inequality, we employ the much used Gini coefficient. In 2000 (before the VAT reform), the distribution of household income in our sample gives a Gini coefficient of 0.210. To assess the impact of the VAT

[14] In our example, the own- and cross-price parameters are noisily estimated implying that we should be cautious in generalizing the welfare effects and extrapolating from the sample to the population.

Table 5: Percentage Change in Gini Coefficient

	(1)	(2)
A: First-order approximation		
Δ Gini coefficient	-0.44 %	-0.32 %
B: Behavioral response		
Δ Gini coefficient	-0.82 %	-0.88 %
Indirect price responses	No	Yes

Notes: Column (1) shows the percentage change in the Gini coefficent when only allowing for direct price response to the VAT reform. Column (2)2 shows the percentage change in the Gini coefficient allowing for both direct and indirect price responses to the VAT reform. The direct price response are reported in column 4 of Table 3 and the indirect price responses are reported in column (4) of Appendix Table A3.

reform, we add the size of the transfer to each household income in 2000, and then compute the Gini coefficient in this counterfactual distribution of household income.

Panel A of Table 5 reports the change in the Gini coefficient when the size of the transfer is computed by the first-order approximation. If we abstract from cross-price effects, the Gini coefficient is reduced by 0.44 percent when we include these transfers. Allowing for cross-price effects, results in a smaller reduction of the Gini coefficient. Panel B of Table 5 reports the change in the Gini coefficient when the size of the transfer is computed by the compensating variation. These transfers are substantially larger in absolute amounts and they have a larger impact on the distribution of household income. Column 1 abstracts from cross-price effects and shows that the reduction in the Gini coefficient is equal to 0.82 percent. In column (2), we allow for cross-price effects and the result is that the Gini coefficient is reduced further to 0.88 percent. Put into perspective, this reduction in the Gini coefficient corresponds to introducing a 0.88 percent proportional tax on earnings and then redistributing the derived tax revenue as equal sized amounts to the individuals (Aaberge, 1997).

One caveat with the analysis in Table 5 is that we do not hold the total tax burden constant. Ideally, one should incorporate that the tax revenue may finance public expenses reducing inequalities or the tax revenue may be replaced by other taxes with different progressivity. In practice, however, it is difficult to tell exactly how the government uses a particular tax revenue, and a balanced budget analysis is beyond the scope of this paper.

5 Conclusion

Much of the controversy surrounding recent policy proposals to broaden the base for VAT revolves around who ultimately bears the burden of these taxes. The typical assumption is that consumer prices fully reflect taxes, so that the main empirical question is how the tax induced price changes affect members of different income groups. For example, the Mirrlees Review assumes the incidence is fully on consumer prices in their proposal to broaden the base for VAT by removing the zero rating for food. However, the evidence base is scarce, and as critics of such policy changes point out (see e.g. Atkinson, 2013), market imperfections could generate both over and under-shifting of VAT to consumer prices.

In this paper, we examined the incidence and distributional effects of VAT in a setting with plausibly exogenous variation in tax rates. The context of our study was a sharp change in the VAT policy on food items in Norway. Using a RD design, we examined the direct impact of the policy change on the consumer prices of food items as well as any cross-price effects on other goods. Our estimates suggested that taxes levied on food items are completely shifted to consumer prices, whereas the pricing of most other goods is not materially affected. To understand the distributional effects of the VAT reform, we used expenditure data and estimated the compensating variation of the tax induced price changes. We found that lowering the VAT on food attenuates inequality in consumer welfare, in part because households adjust their spending patterns in response to the prices changes. By comparison, the usual first-order approximation of the distributional effects, which ignores behavioral responses, seriously understates the redistributive nature of the VAT reform.

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Appendix: Additional Tables and Figures

Table A1: Reform Effects on Consumer Prices of Food, DiD Specification of the RD Model, Multiple Years

Dep. variable: Log Consumer Price	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Reform effect	-0.100***	-0.102***	-0.125***	-0.106***	-0.111***	-0.108***	-0.106***	-0.104***
	(0.005)	(0.006)	(0.005)	(0.005)	(0.005)	(0.007)	(0.006)	(0.007)
Year of month of July effect	1994	1995	1996	1997	1998	2002	2003	2004

Notes: All coefficients are estimated using the DiD specification of RD model with log consumer prices as dependent variable, item fixed effects, triangular weights and two months bandwidth. The first column report the results where the month of July effect is from 1994, in column (2) the month of July effect is from 1995, etc. The standard errors are clustered at the firm level and robust to heteroskedasticity.

Table A2: Reform Effects on Food, Controlling for Observable Determinants of Food Prices

Dep. variable: Log Consumer Price	(1)	(2)	(3)
Reform effect	-0.108***	-0.113***	-0.100***
	(0.004)	(0.006)	(0.007)
Covariates lagged 1 week	No	Yes	No
Covariates lagged 4 weeks	No	No	Yes

Notes: All coefficients are estimated using a RD model with log consumer prices as dependent variable, item fixed effects, triangular weights and two months bandwidth. The first column report the results controlling for no covariates, the second column includes covariates lagged on week, and the third column includes covariates lagged 4 weeks. The included covariates are oil prices (brent), euro/NOK, SEK/NOK, GDP/NOK and an import weighted exchange rate. The standard errors are clustered at the firm level and robust to heteroskedasticity.

Table A3: Reform Effects on Consumer Prices.

Dep. Variable: Log Consumer Price						
A: Heterogeneity in Reform Effects across Food items						
	Bread and cereals	Meat, fish and seafood	Dairy, eggs and fats	Fruit and vegetables	Sugar, chocolate and other food products	Non-alcoholic beverages
Reform effect	-0.135***	-0.085***	-0.096***	-0.131***	-0.096***	-0.121***
	(0.009)	(0.012)	(0.007)	(0.015)	(0.007)	(0.010)
B: Reform Effects on Consumer Prices of Non-food Items						
	Clothing	Services	Household fuel	Alcohol	Transport	Other non-durables
Reform effect	-0.0123	0.0427	-0.000436	-0.0141**	-0.00941	-0.00324
	(0.0273)	(0.0592)	(0.0097)	(0.0069)	(0.0109)	(0.0084)

*** p<0.01, ** p<0.05, * p<0.1

Notes: All coefficients are estimated using a RD model with log consumer prices as dependent variable, triangular weights and two months bandwidth. We control for item fixed effects and possible month effects using a DiD strategy. The standard errors are clustered at the firm level and robust to heteroskedasticity.

Table A4: Estimated Parameters from the Demand System

	Share equations					
	Food	Clothing	Services	HH fuel	Alcohol	Transport
Expenditure	-0.0768***	0.0319***	0.0329***	-0.0507***	-0.0003	0.0382***
	(0.0048)	(0.0040)	(0.0051)	(0.0026)	(0.0010)	(0.0058)
Price food	0.0374	0.0182	-0.0389**	-0.0129***	-0.0010	0.0158
	(0.0280)	(0.0182)	(0.0142)	(0.0022)	(0.0079)	(0.0272)
Price clothing	0.0182	-0.0304	0.0190	0.0093***	0.0130**	-0.0434*
	(0.0182)	(0.0191)	(0.0109)	(0.0020)	(0.0049)	(0.0213)
Price services	-0.0389**	0.0190	0.0165	-0.0000	-0.0012	0.0007
	(0.0142)	(0.0109)	(0.0144)	(0.0023)	(0.0038)	(0.0225)
Price hh fuel	-0.0129***	0.0093***	-0.0000	-0.0135***	0.0004	0.0103***
	(0.0022)	(0.0020)	(0.0023)	(0.0011)	(0.0005)	(0.0028)
Price alcohol	-0.0010	0.0130**	-0.0012	0.0004	-0.0047	-0.0029
	(0.0079)	(0.0049)	(0.0038)	(0.0005)	(0.0056)	(0.0084)
Price transport	0.0158	-0.0434*	0.0007	0.0103***	-0.0029	-0.0111
	(0.0272)	(0.0213)	(0.0225)	(0.0028)	(0.0084)	(0.0477)

Notes: γ-symmetry constrained Almost Ideal estimates using the Norwegian Consumer Expenditure Survey from 1991–2001. Control variables include: indicator variable for year and season; age and gender of the head of household; an indicator equal to one if the head of household is a single parent or retired; number of cars in the household; working status of wife; number of children by age 0-7 years, 7-16 years and 16-20 years; indicator variable for smoking, in addition to a full set of indicators for region and population size. Total expenditures on non-durables are treated as endogenous and household income is used as the excluded instrument. Sample: households in which the household head is between 20 and 70 years old and not self-employed. The sample is top and bottom coded at the the 99th and 1st percentile level of the distribution of household income. Standard errors are reported in parenthesis. $*$p$<$ 0.1, $**$p$<$ 0.5, $***$p$<$ 0.01.