Estimating taxpayer responses to top tax reforms*

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Abstract

We propose a new geographic grouped-data estimator of the elasticity of taxable income in response to top-bracket tax reforms, based on changes in reported income and marginal tax rates at the census tract level. Our estimator shares features of both the cross-section and panel data estimators common in the existing literature, and it addresses potential sources of inconsistency in both those estimators. We estimate medium-term and short-term retiming responses to recent top tax increases in Canada. Estimated avoidance responses are large and persistent over time, suggesting that revenue gains to the reforms have been extremely small.

1 Introduction

We estimate the response to top bracket tax rate increases of high-income taxpayers through changes in economic activities, changes in tax planning strategies, and other avoidance behaviour. Understanding response to tax reforms is highly policy relevant. If taxpayers respond to increased tax rates by reducing reported income, then the incremental revenue from the increase will be smaller than would otherwise be anticipated and, arguably, the economic costs of the reform will be larger. For this reason, an extensive literature has emerged to estimate taxpayer responses to tax changes since Feldstein (1995). But identifying taxpayer response to top bracket tax reforms is notoriously difficult, because treatment by top tax reforms is inherently endogenous, and treated and untreated taxpayers may be affected differently by other, contemporaneous economic changes. For this reason,

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the range of estimated responses is large, and there is little consensus in the literature on the desirability of increasing top tax rates (Saez et al., 2012).

To learn the causal effects of taxes on reported incomes through observational data, we must estimate how top-bracket taxpayers' incomes would have evolved over time in the absence of the tax changes; the difference between actual and counterfactual incomes then represents the causal effect of tax reform. To do so, we compare incomes of top bracket taxpayers to other high-income taxpayers, whose incomes are likely to evolve similarly over time, but who were largely unaffected by the recent tax reforms. Difference-in-difference estimators of this kind are of course commonly used to answer such questions, but the previous literature has pursued one of two very different strategies for constructing treatment and control groups for tax reforms: (i) cross-section methods, and (ii) panel microdata methods. While both methods have similar structures and the same objectives, they address endogeneity very differently, and they often yield very different estimates of the causal effects of tax reforms (Saez et al., 2012). For this reason, the nature and magnitude of high-income taxpayers' response to tax changes is far from a settled matter.

In this paper, we present a new approach to estimating responses to top tax reforms from administrative tax return data, based on a grouped-data instrumental variables estimator that compares changes in reported incomes and predicted changes in marginal tax rates around reforms, within census tracts over time. Because taxpayers choose residence assortatively by income, the average tax change within a census tract is informative about the predicted impact of the reform on an individual taxpayer there, but it is are not affected by the problems of mean reversion in individual incomes that may bias traditional panel-data estimates, nor by the problems of income distribution rank reversals that may bias traditional cross-section estimates. We document these problems in both traditional approaches with our data, and we compare our estimates to those obtained from the traditional approaches.

In one recent example of the cross-section method, Saez (2017) examines responses to the 2013 top tax rate increases in the United States using aggregate data, by comparing changes in the share of income accruing to the top percentile and the rest of the top decile from 2011 to 2014. He concludes that most of the observed change in top incomes represented retiming of income from 2013 into 2012, and the medium term tax avoidance response in 2013 and beyond was small. On the basis of this research, one would conclude that recent top tax increases have led to a large but transitory retiming of income due to anticipation effects, but a fairly small fundamental behavioural response.

But these conclusions cannot be supported by the data. The problem comes from using taxpayers' observed incomes in each year to define the notional treatment and control groups for the reform. The rich in one year are not necessarily the rich in the next. From year to year, people shift between the two groups, affecting average income in both groups and making changes in average incomes over time non-comparable. If these changes are correlated with tax changes, then difference-in-difference estimates of tax responses using this approach will be biased to some degree. What is worse, the magnitude and even direction of the bias in unknown in general.[1]

The group-shifting problem is especially severe in the presence of short-term retiming responses to tax changes. As we document below, mobility in the top one percent is high in Canada – and it was especially high in 2015 and 2016. A business owner whose income typically was below the top one percent threshold might declare a special dividend in 2015 for purposes of intertemporal arbitrage. So we cannot compare aggregate income changes from year to year and conclude that they are offsetting. Thus the shifting nature of the treatment and control group renders aggregate analysis (and its cousin, pooled cross-section microdata analysis) inadequate to detecting the effect of taxes on income reporting behaviour – and particularly, but not solely, for detecting short-run retiming responses.

To deal with the group shifting problem, in this paper we use panel microdata methods, following individual taxpayers through reform periods to measure responses. This allows us to identify unchanging treatment and control groups to compare around reforms – based on the level of individual income in a year prior to the reform. It also allows us to difference out individual heterogeneity that, as discussed above, could bias the aggregate data estimates. As is recognized in the existing literature, panel data methods are also potentially biased, due to mean reversion in individual incomes. For a top tax rate increase, this bias is likely positive, as high income in the pre-reform year predicts both a larger future tax increase, and lower future income due to mean reversion. We introduce a novel strategy to deal with mean reversion, instrumenting an individual's predicted tax change around tax reforms with the average predicted change for the individual's neighbours in the same census tract and year.

We follow the existing literature in summarizing taxpayer responses to taxation by estimates of the elasticity of taxable income (ETI): the percentage change in reported income, given a one percent change in the taxpayer's take-home share of income (one minus their marginal tax rate). The estimated ETI is highly policy relevant. It is informative about the deadweight cost of taxation and in forecasting the revenue impact of tax changes (Feldstein, 1999; Chetty, 2009). The ETI is also usefully in designing policies to increase tax compliance (Slemrod and Kopczuk, 2002) and in coordinating tax policies across different tax bases and levels of governments (Milligan and Smart, 2019).

Our research uses data from a 20 percent sample of tax returns from Canada, using longitudinal data on individual taxpayers observed over the 1982-2016 period. Top-bracket taxpayers in Canada have experienced rather large increases in top tax rates in recent years. In 2016, the federal government introduced a new top tax rate of 33 percent on incomes over $200,000, an increase of four percentage points on the previous top marginal tax rate. Together with increases in top provincial tax rates implemented in the past decade in most provinces, top marginal rates have increased about 7.5 percentage points on average since 2009.

The initial evidence suggests that taxpayer responses to the 2016 federal tax increase in particular have been large. As noted by other observers[2] and as we document in de-

[1] This has long been recognized in the literature. See for example Slemrod (1989) and Goolsbee (1999).
[2] See Laurin (2018) and Parliamentary Budget Office (2019).

tail below, the average income of top-bracket taxpayers in fact *rose* in 2015, suggesting that some taxpayers anticipated the 2016 reform and retimed income realizations, moving income into the 2015 tax year in order to take advantage of the lower, pre-reform tax rate. In 2016, average top-bracket incomes then fell substantially, which could reflect the combined impact of retiming behaviour, plus a more persistent increase in tax avoidance behaviour that will continue to decrease the tax base in 2017 and subsequent tax years.[3]

There is a large literature on estimating ETI using both quasi-experimental methods based on tax reforms, but the ETI is known to vary across countries and over time based on differences in tax rules, and there is relatively little evidence on the ETI for Canada.[4] A few other papers have examined grouped data instrumental variables estimates of response to tax changes (Blundell et al., 1998; Burns and Ziliak, 2016). But our use of fine geographic detail to construct groups is new, as is our comparison of estimates from the grouped data IV and the traditional cross-section and individual panel data estimators.

We find a large elasticity of reported incomes in response to top-bracket tax increases. Based on our preferred approach, the elasticity of reported market income excluding capital gains is 0.5 for top-bracket taxpayers. Notably, this elasticity appears to represent a fundamental behavioural response to tax increases over the medium term, rather than merely the effects of retiming of income in anticipation of future tax increases. Indeed, examining the 2016 federal reform alone using our panel data methods, we do not find any evidence that taxpayers predicted to be in the top bracket increased reported in incomes in 2015 in anticipation of the future tax increase. In this sense, the observed income response in 2016 is likely to be a persistent consequence of higher taxes on the rich. As we show below, the responses are large enough that it is likely the 2016 tax increase generated no additional tax revenues for the federal and provincial governments taken together.

The plan of the paper is as follows. Section 2 describes recent tax changes and discusses apparent responses in incomes based on aggregate data. Section 3 describes our estimation strategies and our data. Section 4 presents results based on the pooled cross-section and panel data estimators in turn. On the basis of the estimated elasticity, in Section 5, we simulate the medium-term revenue impact of the recent federal top-bracket tax increase. Section 6 concludes.

2 Recent income tax changes in Canada

Figure 1 shows the evolution of the combined federal and provincial top marginal tax rates among provinces since 1988. The pattern is clear. Top rates in Canada had fallen with the tax reforms of 1981 and 1987, but then started to rise again in the 1990s, largely due to the increases in federal surtaxes during an era of deficit-fighting under the Mulroney

[3]Income data for 2017 and subsequent years are not yet publicly available.

[4]Sillamaa and Veall (2001) examined the 1987 federal tax reform; Department of Finance (2010) and Milligan and Smart (2015) examined provincial tax reforms of the 1990s and early 2000s. But both the methodologies and results of these studies varied widely, and it is safe to say there is no consensus on the elasticity of taxable income for Canada.

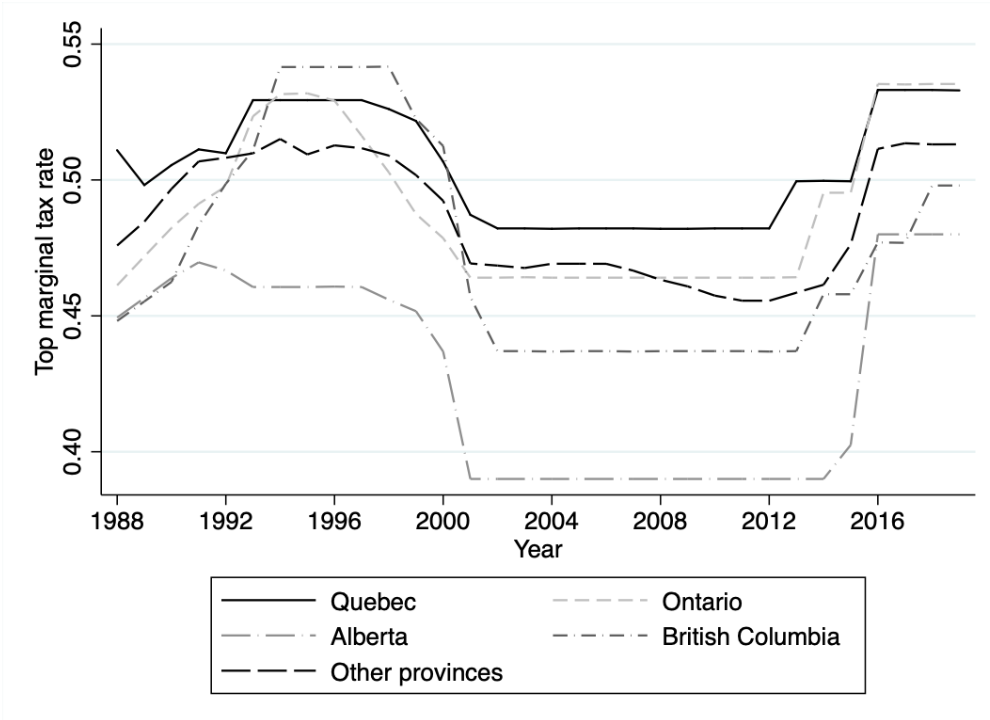

Figure 1: The evolution of provincial top marginal tax rates

and Chretien Governments, including a high-income surtax that applied to the top two percent of taxpayers. The early 1990s also saw increases in some provincial tax rates. By the end of the decade, the tide had turned again towards tax flattening. Federal surtaxes were abolished in 1999 and 2000. As well, the 2000 tax year saw the introduction of the "Tax on Income" (TonI) system, under which provincial governments were permitted under Tax Collection Agreements to levy taxes as a function of taxable income rather than as a percentage of basic federal tax. This new provincial flexibility led to a major tax-flattening reform in Alberta in 2000, followed by less dramatic reductions in top rates in other provinces subsequently. With these changes, the weighted average top marginal rate fell to 44.8 percent in 2001, compared to 51.3 percent in 1998.

Following the financial crisis and recession of 2008-9, top rates began to increase again. The changes in federal and provincial top tax brackets since 2011 are summarized in Table 1. New top brackets were created in six provinces over the period, beginning with Ontario in 2012. The table shows the taxable income threshold at which the new bracket begins, and the increase in the marginal tax rate at that income, relative to the rate applying in 2011. Although top rates have increased, there have also been marked increases in the threshold income at which these rates apply. The weighted average of provincial top bracket thresholds is now over $210,000, compared to just over $125,000 in 2009. The federal top bracket threshold has increased commensurately.

In 2016, a new top federal rate of 33 per cent on incomes over $200,000 was intro-

Table 1: Top tax changes, 2012-19

Province	Top bracket threshold ($)		Change in MTR since 2011 (%)
Newfoundland (2015-16)[e,f]	175,000	a,b	3.5
New Brunswick (2015)[d]	250,000	a,b	7.91
New Brunswick (2016-)	150,000	a,b	3.3
Quebec (2013-)	100,000	a	1.75
Ontario (2012-13)[c]	500,000	a	3.12
Ontario (2014-)	220,000	a,b	3.12
Alberta (2015-)[e]	300,000	a,b	5.0
British Columbia (2014-15, 2018-)[d]	150,000	a	2.1
Federal	200,000	a	4.0

Notes:

a Bracket threshold indexed for inflation

b Additional new bracket created below top threshold

c Commenced July 1, 2012

d Abolished in 2016

e Rate changed in 2016

f Rate changed in 2017

duced, replacing the previous top federal rate of 29 per cent. The average of top rates, weighted by provincial shares of top income taxpayers, is now 52 percent, up from 44 percent in 2011. More relevant for taxpayers' decisions is probably the take home rate (one minus the marginal tax rate), which has decreased by 13.5 percent on average since 2012. Such large changes in tax rates are apt to have a discernible effect on tax avoidance and reported taxable income. The new federal bracket was proposed during the federal election in October 2015, and it appears to have been highly salient to voters and taxpayers. As such, the 2016 rate change might have been anticipated, leading to changes in the timing of income and other tax avoidance strategies in the fall of 2015.

The 2016 changes that we study were accompanied by other tax reforms at the federal level. The middle bracket tax rate dropped from 22 percent to 20.5 percent. This applies to taxable incomes in approximately the range from $50,000 to $90,000. As well, the tax environment for privately-controlled companies has been changing in recent years. Canadian-controlled private corporations (CCPCs) with income and assets below certain thresholds are subject to corporate-level taxes at much lower than either general corporate tax rates or top individual rates.[5] Dividends paid to Canadian residents are generally taxable as individual income, with an imputation system in place that is intended to ensure that combined corporate and personal taxes on earnings are about the same as ordinary

[5] The federal rate for active business income of CCPCs below $500,000 is currently 9 per cent, compared to the 15 per cent general corporate rate. Provincial rate reductions are proportionately larger.

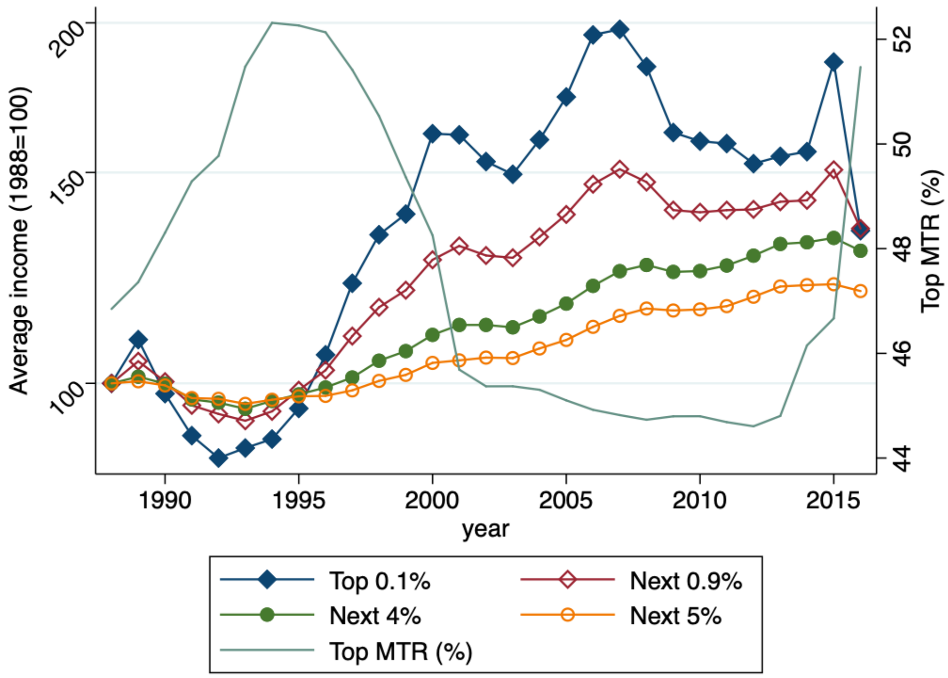

Figure 2: Top percentile market incomes, 1988-2016

income. But this individual-level tax may be deferred indefinitely by retaining earnings inside the corporation, and some tax planning strategies are available to convert corporate earnings into lightly-taxed capital gains in the hands of shareholders (Wolfson et al., 2016). The use of CCPCs has been expanding rapidly in recent years. As such, CCPCs are an increasingly important part of the tax planning strategies available to many high-income taxpayers. The government began a review of tax expenditures and of the taxation of CCPCs in 2017, with certain reforms being introduced late in the year. These changes, if anticipated, might also have had an effect on tax planning strategies in 2016 and 2017.

2.1 Taxpayer responses: Evidence from aggregate data

As a first look at the data and a preview of our identification strategy, we first look at the patterns visible in the aggregate data on high-income taxpayers.[6] Figure 2 shows the evolution since 1988 of average market incomes excluding capital gains for various groups of high-income taxpayers who are more or less affected by top bracket tax reforms. The lines in the figure show average real incomes of parts of the top decile of income: the top 0.1 per cent of taxpayers by market income in each year, the rest of the top one percent, the rest of the top five per cent, and so on. To focus on the differences in income growth

[6]The source for data in this section is Statistics Canada's High income tax filers in Canada data table 11-10-0055-01. It is derived from the microdata used in our empirical analysis below.

between groups, the series have been rebased so that all are equal to 100 in 1988. In 2016, the top 0.1 per cent had individual market incomes above $673,000, the top one per cent had incomes above $224,000, the top five per cent had incomes above $118,000, and the top decile had incomes above $91,000. For the most part, P99 and above are treated by top tax reforms, while P90 and P95 are control groups that are much less affected by top marginal tax rate changes but which arguably face similar aggregate income dynamics as the treatment groups.

Figure 2 also shows the top marginal tax rate, averaged across provinces using the shares of each province in the population of P99.9 taxpayers as weights. The top 0.1 per cent and next 0.9 per cent incomes in aggregate do appear to respond to tax changes in the aggregate time series, relative to contemporaneous changes in the lower income groups. There is a discernable dip in reported income in the early 1990s when tax rates were increasing, and an increase as tax rates fell again in the early 2000s. Also discernable is the spike in incomes in 2015, followed by a sharp drop coincident with the 2016 federal rate increase. Of course, these broad correlations are merely suggestive of a causal effect of taxes.

The data are also notable for the lack of a strong trend in top tail inequality in recent years, even before the recent downturn in top incomes. Between 2004 and 2014, the average market income excluding capital gains of the top one per cent rose just 3.2 per cent in real terms, from $461,500 to $476,300. In the preceding decade, average income rose 53.8 per cent, from $299,900 to $461,500. Some of this decline in trends growth reflects the increased use of incorporation by top income taxpayers, but that effect appears to be small (Wolfson et al., 2016). The slowing of top income growth is in contrast to the United States, where top one per cent real income excluding capital gains grew by 69.2 per cent and 29.3 per cent over the same two decennial periods.[7] The lack of a trend in top tail inequality in recent years in principle makes it easier to estimate the causal effect of top tax rate changes using difference in difference methods, since the parallel trends assumption is more plausible with stable patterns of top tail inequality.

Figure 3 looks more closely at top one per cent incomes in the more recent years that we examine in detail in the analysis below. Notable in the figure is the sharp changes in 2015 and 2016. Average real market incomes of the top one percent rose 12.2 percent in 2015 compared to the previous year, and then fell 9.5 per cent (relative to the 2014 base level) in 2016. It seems plausible that these sharp changes reflect at least to some degree retiming effects from the anticipated increase in tax rates in 2016.

How did these changes come about? Figure 4 shows the components of income since 2009 for taxpayers with taxable income over $250,000, taken from public aggregates reported by the Canada Revenue Agency.[8] The data in this case are for taxable income, except that we have adjusted the data reported for capital gains to add back the 50 per

[7]Based on the data in Piketty and Saez (2003) and updated data downloaded from Emmanuel Saez's website.

[8]CRA, T1 Final Statistics, Table 2, ALl returns by total income class, `https://www.canada.ca/en/revenue-agency/programs/about-canada-revenue-agency-cra/income-statistics-gst-hst-statistics/t1-final-statistics.html`.

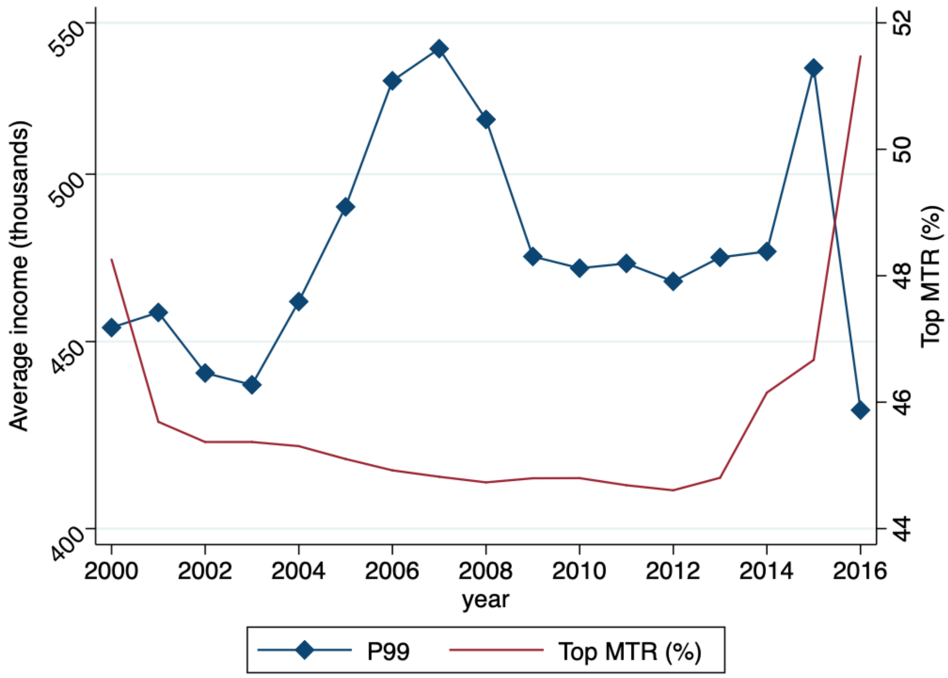

Figure 3: Top one per cent market incomes, 2000-2016

cent of gains that is excluded from taxation under current tax rules. The decline in reported income reflects declines in employment income and in self-employment income, particularly the "professional income" category that includes unincorporated business income of doctors and lawyers, among others. Employment income might be subject to retiming responses, as for example if bonuses were paid to employees at the end of 2015 to beat the tax increase. Business income including professional income could be subject to the same phenomenon, if proprietors are able to delay deductible expenses to 2016 (when they were deductible at a higher rate), or accelerate realizations of gross income into 2015. Capital gains are variable from year to year, but there is no obvious downward trend over time, nor short-run increase in 2015. Most notable in Figure 4, there was a sharp increase in taxable dividends in 2015, followed by a sharp decline in 2016. It seems likely that this reflects a retiming response by owners of small business corporations who paid out retained earnings in 2015 in order to avoid some portion of the 2016 tax increase. In short, many components of the income of top-bracket taxpayers may be influenced by tax panning responses, some of which are likely transitory, while others are of a more long-run character.

Since the 2016 changes caused the take-home share of top bracket taxpayers to decline by about 11.5 per cent on average, these income changes are consistent with rather high short-run elasticities of taxable income – equal to -1.1 for the 2015 anticipation effect, and 0.8 for the short-run effect in 2016. What does this imply for longer-run responses? It

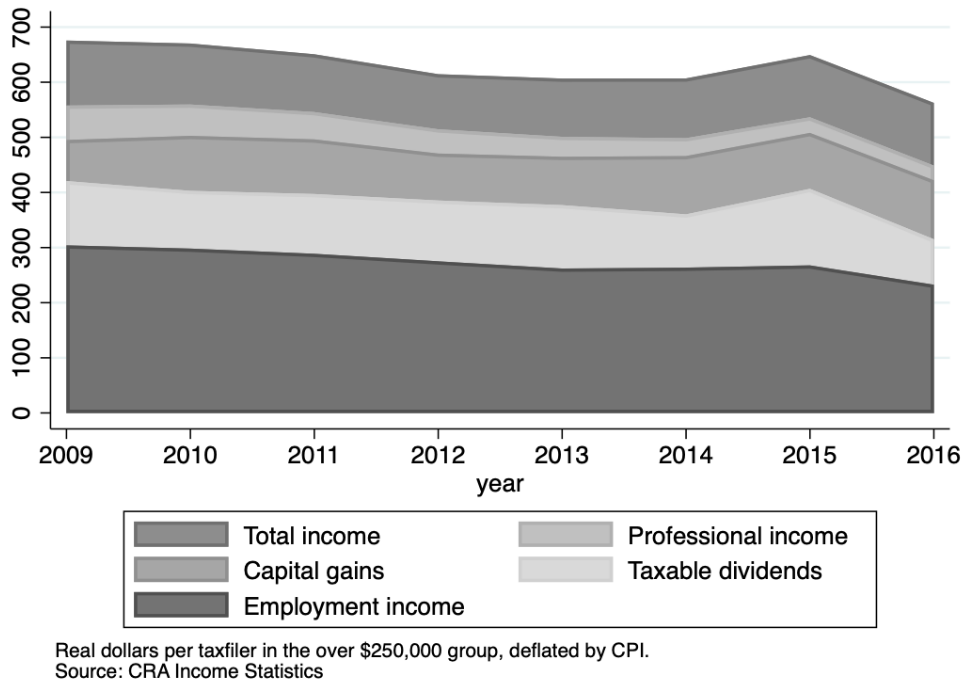

Figure 4: Components of taxable incomes over $250,000

is tempting to assume that all of the 2015 increase represents retiming of income which otherwise would have been reported in 2016. If so, then the two figures could be added together to estimate the expected medium-term response to the tax change, after the transitory impact of retiming responses has dissipated. Since the 2015 increase was larger than the 2016 decrease, this would imply that the medium-term response to the tax increase was in fact negligible. But, as we argue below, this rosy conclusion can't really be suported by the data and what we know about tax avoidance responses.

3 Microdata evidence

Our starting point is a standard simple structural model in which the log taxable income of individual i in year t is a linear function of the log of one minus her marginal tax rate:

$$\log y_{it} = \alpha_i + at + b\log\theta_{it}(y_{it}) + u_{it} \tag{1}$$

where y is income, $\theta = (1-\tau)$ is the take-home rate, α_i is log permanent income, at is a trend term, and u_{it} is a shock to current income that follows some dynamic process. We are interested in recovering an estimate of the ETI b, which we assume to be homogeneous in the population of taxpayers affected by tax reforms.

Suppose we have a series of tax reforms that change over time the take home rate for top-bracket taxpayers, who are identified by the indicator variable:

$$d_{it} = \mathbb{1}(y_{it} > k_t) \tag{2}$$

where k_t is the top bracket threshold. The tax reform changes only top-bracket tax rates $\Delta\theta_{it}$, i.e.

$$\theta_{it}(y_{it}) = \theta_{i,t-1} + \Delta\theta_{it} d_{it}$$

Notice that this treatment $\Delta\theta_{it}$ may be heterogeneous across the population, but let's assume it is non-stochastic (and observable). For example, the magnitude of the reform might depend on family structure, or on the province of residence of the taxpayer.

The chief challenge to identification is that an individual's tax parameter θ_{it} is clearly endogenous: it is actually a mechanical function of the post-treatment outcome of interest y_{it}. An individual's tax rate at each date is therefore positively correlated with her unobservable permanent income y_i^p. So we need an instrument z_{it} for treatment. Consistent with the existing literature (Saez et al., 2012), we explore two approaches to identifying b through instrumental variables, based on (i) pooled cross-section analysis, and (ii) panel data analysis.

Pooled cross-section analysis simply identifies the treatment group for the tax reform in each year as the group of taxpayers in a top fractile of the income distribution in that year, who are by definition most affected by top bracket tax reforms. We define this group to be the top one per cent of market incomes, although we also experiment with other smaller treatment groups. Thus the instrument is

$$z_{it}^{CS} = \mathbb{1}(F_t(y_{it}) > 0.99) \tag{3}$$

This is a difference in difference estimator, which identifies ETI from the average change in income in the top percentile around tax reforms, compared to the contemporaneous change in the next income fractile down.

The pooled cross-section estimator is the microdata analogue of the aggregate analysis of the preceding section. The two estimators are similar, but the microdata approach permits us to control for the effects of taxpayers' individual characteristics over time. In particular, in what follows, we control for the effects of taxpayer gender, language, immigration status, and province of residence, as well as linear trend terms in all these controls. The pooled cross-section estimator identifies the ETI under the assumption that there are no rank reversals in the income distribution (Goolsbee, 1999), so that the top fractile of the distribution identifies the same set of people in each year and the estimator can difference out the average values of unobservable permanent income in the treatment and control groups over time. Below, use the longitudinal component of our data to investigate whether the assumption of no rank reversals is a reasonable one. (It is probably not.) Of course, the difference in difference approach also requires a parallel trends assumption, i.e. the counterfactual changes in income should be the same on average over time, rather than inequality increasing or decreasing over time. But this assumption can at least be tested by examining income trends in pre-reform periods.

Panel data estimators instead look at responses to tax reforms at the individual taxpayer level. Differencing in (1) to eliminate individual fixed effects,

$$\Delta \log y_{it} = a + b\Delta \log\theta_{it}\mathbb{1}(y_{it} > k_t) + \Delta u_{it} \tag{4}$$

Again here, $\Delta\log\theta_{it}$ is mechanically endogenous, as it is high-income taxpayers who are treated by the tax reforms given (2). To deal with this, we follow Gruber and Saez (2002) and most of the subsequent literature by constructing the simulated tax rate change the taxpayer would experience due to tax reform if her income remained at the pre-reform level, i.e. the instrument is

$$z_{it}^{PD} = \delta\log\theta_{it}\mathbb{1}(y_{i,t-1} > k_t) \tag{5}$$

Following Chetty (2009), we construct our sample by differencing income over three-year intervals and stacking the resulting first differences in a single estimation sample; i.e. Δy_{it} is the log difference in income between an individual's income in 2011 and 2014, and 2012 and 2015, and so on.

This "simulated tax change" instrument may nevertheless be endogenous is (4), if lagged income predicts future innovations to income, i.e.

$$E(\Delta u_{it}|y_{i,t-1}) \neq 0$$

In the presence of mean reversion in individual incomes, this correlation is typically negative. For progressivity-increasing reforms such as the 2016 tax change, this would result in an upward bias in the DD estimate of b (at least in the absence of other covariates), as high transitory income shocks in the pre-reform period predict treatment with a tax increase, and also predict reversion to the mean. Exploring this issue, Weber (2014) suggest constructing the simulated tax change instrument with further lags of the taxpayer's income, i.e. $y_{i,t-k}$ for $k > 1$. But if the stochastic process for individual income has an autoregressive component, then all finite lags of income predict income innovations at t, even if this effect dies out asymptotically over time. We therefore adopt a new approach to constructing the instrument for $\Delta\theta_{it}$.

To deal with mean reversion in (4), we instrument for z_{it}^{PD} of an individual with its value for her neighbours residing in the same Census Tract in year $t-1$. Census Tracts are contiguous geographic areas of between 2500 and 8000 residents urban areas of the country, which are defined by Statistics Canada with the goal of identifying relatively homogeneous populations in terms of socioeconomic characteristics, and which have boundaries that follow natural features like rivers, or major roads and municipal boundaries that change rarely over time.[9]

The Census Tract instrument deals with the mean reversion problem by averaging the treatment indicator over a group of similar individuals who are subject to similar economic

[9]Individuals residing outside urban areas (defined as population agglomerations of at least 50,000 people) are not assigned to Census Tracts. We construct a single residual group in each province consisting of all these rural residents.

conditions. We implement the estimator using grouped data IV, i.e. simply estimating the grouped data mean version of (4),

$$\overline{\Delta \log y}_{nt} = a + b\overline{\Delta \log\theta}_{nt} P_{nt} + \overline{\Delta u}_{nt} \tag{6}$$

where bars indicate averages over individuals in census tract T_n, and P_{nt} is the treatment propensity there, i.e.

$$P_{nt} = \Pr(y_{it} > k_t | i \in T_n)$$

To compute the efficient two-stage least squares estimate of (6) in the manner proposed by Angrist (1990), we weight observations by census tract populations, and we include only census tracts with at least 25 taxpayers in the top decile of incomes to avoid cases where the taxpayer herself has a large mechanical impact on the neighbourhood average.

Our grouped data estimator is a natural "halfway point" between the traditional panel data and pooled cross-section approaches to estimating ETI. Like the panel data approach, we identify a fixed set of taxpayers most likely to be treated by tax reform based on their pre-reform observable characteristics, instead of allowing the treatment group to change in unobservable ways over time, as with the pooled cross-section approach. However, like the pooled cross-section approach, we deal with mean reversion in individual incomes by averaging income changes over a large set of similar taxpayers, unlike the traditional panel data estimator which relies only on individual taxpayer information to construct the instrument and which is therefore prone to bias, even in large sample, due to mean reversion.

3.1 Anticipation effects

The discussion of aggregate data trends suggested that retiming of income around anticipated tax reforms may be an important part of tax responsiveness for high-income taxpayers. Disentangling retiming from medium-term responses is in turn important for estimating the effects of tax reforms on government revenues and deadweight loss form taxation. To do so, we estimate an extended version of the standard equation (1) for reported income that incorporates anticipation effects:

$$\log y_{it} = y_i^p + at + b_1 \log\theta_{it}(y_{it}) + b_2 E \log\theta_{i,t+1}(y_{i,t+1}) + u_{it} \tag{7}$$

where $E\log\theta_{i,t+1}(y_{i,t+1})$ is the taxpayer's expected log take-home rate in the following year.[10] We proxy the expected future tax term by the taxpayer's actual take-home share one year forward; for 2016, since we do not observe 2017 incomes we calculate this based on actual 2016 incomes. In the panel data analysis, we instrument for the change in the future tax term by the simulated change in the tax term one year forward, using the taxpayer's neighbours' incomes four years previously – i.e. the same instrument as for the current year tax change, except that the tax policy parameters are for the following year. In this specification, b_2 is anticipation effect and b_1 the immediate response once the

[10] A similar but more elaborate approach to anticipation effects is implemented in Kawano et al. (2016).

tax change has been implemented; the sum of the two coefficient therefore captures the medium-term elasticity of reported income, netting out the transitional effect on income in the year of the reform and immediately before it.

3.2 Data

The data for our analysis come from Statistics Canada's Longitudinal Administrative Databank (LAD), a 20 per cent longitudinal sample of personal tax returns covering the 1982-2016 period. The LAD is longitudinal in the sense that individual taxpayers are added each year to replace panel attrition, and a taxpayer once sampled in a particular year is included in the LAD in all subsequent years that he or she files a tax return. Thus, attrition from the LAD is extremely rare, and occurs only due to death, cessation of residence in Canada, or in extremely rare cases the failure of the individual to file returns.

The full LAD over the 1982-2016 period consists of approximately 150 million taxpayer-year observations. In our work, since we are concerned with recent top tax reforms, we confine attention to taxpayers in the top 10 percent of the distribution of market income over the 1998-2016 period. As well, we exclude taxpayers who are age 65 or older on filing their tax returns, since older taxpayers face different tax rates and rules for a variety of reasons. We also exclude taxpayers who changed province of residence in the previous three years, since different provinces have different tax rates. We also exclude taxpayers whose income is imputed to the record by Statistics Canada, which occurs in a small fraction of the sample when no tax return was filed for the relevant year. Taken together, these sample selection rules result in approximately 8.5 million observations in the analysis sample. For the shorter 2008-2016 period covering recent tax reforms, the regression sample is approximately 3.8 million observations.

Our interest is in how reported income responds to tax changes, but of course there are many components to income. In most of this paper, our key variable of interest is "market income," defined in the LAD as the sum of reported individual income from all sources, excluding government transfer payments and realized capital gains. Capital gains are generally taxable at 50 per cent of ordinary income tax rates (and in some cases excluded altogether), so that the implications of marginal tax rate changes for capital gains income are ambiguous. We do look at market income including capital gains in one case below. Market income in the LAD corresponds roughly to broad income which is used to estimate tax responses in much of the US literature. Taxpayers are permitted certain deductions from total income in computing taxable income, but these have been fairly small (much smaller than in the US) since the 1987 tax reform that converted most deductions into non-refundable tax credits, the value of which are independent of the taxpayer's marginal tax rate. The chief remaining deductions are for contributions to tax-prepaid savings plans known as individual Registered Retirement Savings Plans and employer-provided Registered Pension Plans, and for certain tax losses carried forward from previous years. In one case below, we examine impacts on Net Income for Tax Purposes, adjusted for (what did we adjust for), which is Taxable Income before certain final timing adjustments are applied. Regardless of which measure of reported income is examined, for consistency with the literature we use the term "elasticity of taxable income" to describe the parameter measuring the responsiveness of income to taxation. (Consistent with the discussion in this paragraph, we do find that most of the response to taxation is through changes to income included in market income, rather than through changes to deductions.)

Lastly, our work requires a measure of the marginal tax rate on reported income facing each taxpayer in each province and year. We simulate marginal tax rates using the Canadian Tax and Credit Simulator (CTACS), described in Milligan (2016), based on taxpayers' actual taxable incomes as reported in the LAD.

4 Results

4.1 Pooled cross-section estimates

We begin with estimates of the ETI for market ("broad") income based on the pooled cross-section method using top quantile indicators as the instrument for treatment defining treatment and control groups. Estimates from these specifications, corresponding to equations (1)–(3), are presented in Table 2. In the first two columns the quantile is the top one per cent. Note that we have pooled data for the 2008-16 period, incorporating all the recent provincial and federal top tax reforms, plus a short pre-reform period useful in identifying trends in top incomes, which we allow the differ on the basis of gender, language, family status, and province of residence.

The coefficient on the log take-home rate in column (1) of the table is therefore the elasticity of reported income ignoring anticipation effects, which we estimate to be 0.565. In parentheses are estimated standard errors of the stimates, robust to heteroskedasticity clustered in province-year groups. This estimate is in the range of other estimates, for Canada and internationally. In particular, Milligan and Smart (2015) estimate ETI to be in the range of 0.6-0.8, using aggregate quantile income share data and provincial-level variation in top marginal tax rates in the preceding decades (1988-2011). Department of Finance (2010) reports similar estimates, based on both aggregate data and panel microdata approaches. The next column adds the future year tax term to capture the anticipation effect. In this case, the immediate elasticity in response to the tax change is much larger at 1.156, but the anticipation term is -0.597 (although this is not significantly different from zero), so that the combined estimated medium-term elasticity is virtually the same at 0.559.

The remaining columns of Table 2 consider different ways of identifying the treatment group, in order to get at heterogeneity in responses. In columns (3)-(4), the treatment group are taxpayers in the top 0.1 per cent of the income distribution. The estimated ETI is much higher here at 1.687. The next two columns take the treatment group to be top 0.1 per cent taxpayers who own shares in Canadian-controlled private corporations, the closely held business entities that appear to be related to several tax planning opportunities, including those related to the short-term retiming of income. The elasticity for this group

Table 2: Pooled cross-section estimates of the elasticity of taxable income

	(1)	(2)	(3)	(4)	(5)	(6)
					P99.9	P99.9
	P99	P99	P99.9	P99.9	CCPC	CCPC
log take-home rate	0.565*	1.156*	1.687***	3.400***	1.121***	1.883***
	(0.271)	(0.541)	(0.332)	(0.709)	(0.230)	(0.451)
F.log take-home rate		-0.597		-1.906**		-0.764*
		(0.466)		(0.666)		(0.387)
N	3753200	3647730	3753200	3647730	3753200	3647730
First stage partial R^2	0.023	0.011	0.002	0.001	0.007	0.003

Standard errors in parentheses, clustered by province and year

All regressions include year fixed effects plus fixed effects and linear trends in gender, language, family status, and province of residence.

* $p<0.05$, ** $p<0.01$, *** $p<0.001$

is somewhat smaller at 1.121, but there is again ample evidence of a retiming response around the reforms. That said, these estimates are biased in the presence of heterogeneity in the ETI among top percentile taxpayers, since we are assigning the remainder of the top one per cent to the control group to the control group as we change specifications in the table. This raises the average change in tax rates in the treatment group, so that the parallel trends assumption is violated in the presence of ETI heterogeneity (Saez et al., 2012).

Another and potentially more serious bias comes from the changing identity of taxpayers in the treatment group in all the pooled cross-section specifications, as taxpayers switch between treatment and control groups over time in response to transitory changes in individual income – i.e. the rank reversal problem. Since taxpayers switching in and out of the top percentile might have higher or lower incomes than those persistently in the top percentile – and since this difference may be correlated with tax changes – there is a potential bias to the pooled cross-section estimates that is unknown direction and magnitude.

To get a sense of the importance of income mobility in and out of the top quantiles, we calculate Markov transition probabilities for the P99 and P99.9 groups, i.e. the proportion of taxpayers in each group were also in the same group in the previous year. These are presented graphically in Figure 5. There is indeed considerable mobility. About 70-75 per cent of the top percentile were also in the top percentile in the previous year; for the top 0.1 per cent, this probability is lower at 60-65 per cent in most years. These relatively high probabilities of transition correspond to the first-stage partial R^2 of the quantile dummies reported in Table 2. The P99 dummy explains only 2.3 percent of the variation in individual taxpayers' log take-home rates over time, indicating that there is a

Figure 5: Markov transition probabilities for top quantiles

lot more going on with tax rates than the quantile dummy captures. The partial R^2 falls to 1.1 per cent when it is used to explain movements in the future tax rate. These diagnostic statistics are in turn much lower when the instrument in membership in the P99.9 group, consistent with the much higher transition probability in that group.

Observe that these probabilities dip substantially lower in 2015 and 2016 around the federal tax reform, and also in 1999-2001, during the earlier period when federal and provincial top tax rates were changing sharply. (They are also lower during the 2008-09 financial crisis.) This suggests that mobility in and out of the top quantiles is indeed correlated with tax reforms. This is problematic for our ability to estimate the ETI using the conventional cross-section methods – and particularly for disentangling anticipation effects from medium-term behavioural responses. The problem is that the taxpayers we are comparing over time are not identical, and the changes in the composition of the treatment group are correlated with the very event we are trying to study. In particular, the evidence in Figure 5 strongly suggests that some of the increase in incomes in 2015 came from taxpayers who were not typically in the top percentile, but who entered it that year as a result of extraordinary realizations of income. That is still of course an anticipation effect of the reform, but it is not possible to estimate its magnitude by comparing average income changes in the top one percent of time. Slemrod (1989), examining the problem of estimating tax responsiveness of capital gains realization using this method, put the problem succinctly: “the message conveyed by comparing the behavior of percentile

Table 3: Panel data estimates of elasticity of taxable income

	(1)	(2)	(3)	(4)	(5)	(6)
log take-home rate	0.506**	0.593**	0.421**	0.532*	0.791***	0.624
	(0.154)	(0.211)	(0.149)	(0.209)	(0.237)	(0.323)
F.log take-home rate		-0.116		-0.151		0.0133
		(0.192)		(0.194)		(0.248)
N	19795	19795	19795	19795	2321	4620
First stage slope	0.855		0.865		0.989	
	(0.015)		(0.016)		(0.027)	
First stage partial R^2	0.276	0.150	0.279	0.149	0.459	0.350
Years	2008-16	2008-16	2008-16	2008-16	2016	2015-16
Linear trends	No	No	Yes	Yes	No	No

Standard errors in parentheses, clustered by census tract and year

All regressions include fixed effects in gender, language, family status, and province of residence.

* $p<0.05$, ** $p<0.01$, *** $p<0.001$

groups across years may give a very misleading picture of the actual differential behavioral response across groups." Instead we need panel data to compare the same taxpayers' behaviour over time.

4.2 Panel data estimates

Next we turn to estimates of the panel data model in first differences, corresponding to equations (4)–(5), which we implement through the grouped-data instrumental variables estimator based on census tract averages, as in (6). Table 3 presents the basic version of this specification. Here, our estimates of ETI without anticipation effects is 0.506, about the same as for the pooled cross-section estimator based on the P99 dummy instrument. In this case, we report additional first-stage diagnostics on the instrument. The partial R^2 for the simulated tax rate change is 0.276, indicating it is a strong instrument for the actual tax rate change at the neighbourhood level. Even more reassuring, the slope of the first stage relationship is 0.855, indicating that differences across neighbourhoods really do explain much of the propensity for taxpayers to be treated by the reform. Turning to the extended model (7) with anticipation effects, the results are now quite different than for the pooled cross-section estimator. The estimated anticipation elasticity is an -0.116, compared to -0.597 for the pooled cross-section estimate (however neither is significantly different from zero). This again suggests that retiming effects in advance of tax changes, while real and somewhat measurable, are poorly identified from the cross-section estimator because they are correlated with a taxpayer's propensity to be treated by the reform that leads to rank reversals.

The remaining columns of the table explore the sensitivity of results to our strategy of estimating responses by pooling first differences across the several tax changes of the 2008-2016 period. This approach would be biased if there were a trend in top tail inequality over the period that were correlated with the generally increasing trend in top tax rates. But, as we noted in Section 2.1, there does not appear to be any such trend in the aggregate data. That said, the estimates reported in columns (3) and (4) are from a specification that includes a linear spline in incomes above the P99 threshold of market incomes. The estimated ETI drops slightly from 0.506 to 0.423, and the pattern of short-run and anticipation effect elasticities in column (4) matches that of column (2) as well. Columns (5) and (6) report results when the sample is restricted to just the 2015 and 2016 years (or, more accurately, the income and tax changes for 2012-15 and 2013-16). Here the estimated ETI is substantially higher at 0.791, which may reflect the larger magnitude and salience of the 2016 tax change. In column (6), the estimated anticipation elasticity now has the wrong sign (although it is again insignificant). This highlights the difference between our panel data approach compared to the pooled cross-section approach and the patterns visible in the aggregate data. Although top incomes did rise on average in 2015 in advance of the reform, there is no evidence that taxpayers who could be expected to be in the treatment group for the reform were those who increased their incomes in that year. For this reason, it is simply not possible to disentangle retiming responses from medium-term elasticities by comparing average income changes from year to year in the top quantile.

Next, we examine observable heterogeneity in the ETI. Based on previous research and simple introspection, we know that some taxpayers have more access to tax planning strategies than others, which may permit then to respond more and more quickly to tax increases, whether this is the result of short-term retiming of income across tax years, or shifting of income between corporate and personal tax bases, or other strategies. As we noted previously, chief among these groups with higher predicted ETI is probably owners of closely-held small business corporations (CCPCs), who may easily retime realizations of individual income across tax years by paying or deferring dividends paid to themselves as shareholders, or salaries paid to themselves as managers. It is also possible, although more conjectural, that CCPC owners can avail themselves of long-term tax avoidance strategies by permanently reinvesting earnings in their corporations and ultimately realizing the income at the individual level as lightly taxed capital gains (Wolfson et al., 2016). Another group that is likely to have access to greater tax planning opportunities is corporate executives who receive a portion of their compensation in the form of stock options, which give additional opportunities both for income retiming and for long-term tax avoidance. Under Canadian tax law, the value of stock and stock options granted to an employee is taxable, but only one-half of the value is included in taxable income. This means that stock-based compensation is taxed in the same way as capital gains income received by investors, but it is tax-favoured relative to cash-based forms of employee compensation

Table 4: Heterogeneity in the elasticity of taxable income

	(1)	(2)	(3)	(4)
	Base	CCPC owners	Stock option deduction	Finance industry
log take-home rate*Heterogeneity dummy:				
No	0.506**	0.539***	0.633***	0.643***
	(0.154)	(0.155)	(0.150)	(0.159)
Yes		1.624***	0.787**	0.214
		(0.285)	(0.300)	(0.221)
N	19795	28316	28647	28856
First stage slope	0.855	0.787	0.823	0.831
	(0.015)	(0.015)	(0.013)	(0.013)
First stage partial R^2: No	0.276	0.253	0.240	0.247
First stage partial R^2: Yes		0.330	0.486	0.461

Standard errors in parentheses, clustered by census tract and year

All regressions include fixed effects in gender, language, family status, and province of residence. Controls are fully interacted with heterogeneity dummies.
* p<0.05, ** p<0.01, *** p<0.001

at the personal level.[11] In the medium or long term, we might therefore expect that an increase in marginal tax rates on ordinary income would lead to an increase in the use of stock-based compensation for top employees. In the short run, we also expect to see greater retiming of income in anticipation of tax increases, since these employees likely have substantial unrealized capital gains that could be realized and brought into taxable income in advance of the increase. Previous research for the US suggests that income retiming by top managers of large companies explains a substantial portion of the aggregate short-term response to tax increases (Goolsbee, 2000).

In Table 4, we report estimated ETIs for subpopulations of taxpayers of these various kinds. To do so with our census tract grouped data estimator, we compute average income changes and tax changes separately for each heterogeneity group in each census tract and year, and we interact all control variables with indicator variables for the heterogeneity groups. (This is equivalent to computing two-stage least squares estimates separately for each heterogeneity group, but it allows hypothesis testing between the groups.) We define CCPC owners as those who owned a CCPC in the base year prior to the tax change being studies, i.e. in 2012 for the 2012-15 change, and so on. We define those with stock option

[11] There is generally no deduction from corporate taxable income for the value of stock options granted.

Table 5: Alternative definitions of income

	(1)	(2)	(3)
	Dependent variable: log of		
	Market income	Adjusted net income	Market income w/CGs
log take-home rate	0.506**	0.637***	0.296*
	(0.154)	(0.128)	(0.151)
N	19795	19795	19795

Standard errors in parentheses, clustered by census tract and year
All regressions include fixed effects in gender, language, family status, and province of residence.
* p<0.05, ** p<0.01, *** p<0.001

income as those who ever reported a stock option deduction on their tax form in any year up to the base year, i.e. in the 1984-2012 period for the 2012-15 tax change, and so on. As a final cut of the data, we estimate ETI separately for taxpayers employed in the Finance, Insurance, and Management of Companies industries, based on NAICS codes recorded in our data. (This is perhaps a placebo test.) For purposes of reference, column (1) of table 5 reports the estimated ETI of 0.506 for the full sample, taken from Table 3. The next column reports ETI for CCPC owners and others. The estimate for owners is 1.624, compared to 0.529 for others affected by the tax changes. In this sense, small business owners do respond more to tax changes in the short and medium term. Column (3) reports ETI for stock grant recipients 0.787, compared 0.633 for others, estimates which are not significantly different. The final column reports the estimated ETI for employees in the financial sector as an insignificant 0.214, compared to 0.643 for others.

The results highlight the importance of CCPC owners in explaining the response to taxation among high-income taxpayers. The estimated ETI could reflect retiming effects or persistent medium-term effects. Given the relatively short time period studied and the importance of the tax change in 2016, the final year of our data, it is difficult to say which. Disentangling the two effects in more detail would evidently be easier as well if we observed the corporate income and balance sheets of the corporations owned by these CCPC owners.

In Table 5, we report estimates of ETI for different measures of income. Again, the first column reports the reference specification for market income without capital gains, from Table 3. In the next column, the dependent variable is the (log change in) our adjusted net income measure, which is close to the actual taxable income reported by taxpayers. The estimated ETI rises slightly to 0.637 from 0.506. By way of comparison, Gruber and Saez (2002) report an ETI for broad income of 0.171 for high-income taxpayers in the US, much

Table 6: Simulated revenue impacts of the 2016 reform

	$ millions
Total top bracket income in 2016	133,074.0
Mechanical increase of tax increase	2,693.2
Behavioural effect on:	
- Federal tax revenues	-1,699.2
- Provincial tax revenues	-1,021.7
Simulated change in total tax revenues	-27.7

Note: Simulated change in tax revenues from a four percentage point increase in marginal rate, based on 2016 data and an elasticity of taxable income equal to 0.5

smaller than the estimated ETI of 0.567 for taxable income. This suggests that deductions from income serve a smaller role in tax responsiveness than income inclusion in Canada. This is as expected, given the relatively small role of deductions in the Canadian system, where tax allowances for taxpayer expenditures have since 1988 mainly been delivered through non-refundable credits whose value is independent of a taxpayer's marginal tax rate. The final column reports an ETI for market income including capital gains, which is smaller at 0.296. This may reflect the ability of taxpayers to substitute towards lightly taxed capital gains in response to tax increases.

5 Revenue impacts of 2016 reform

Was the 2016 federal tax increase a good policy? The answer presumably lies in comparing the additional revenues raised through the reform to the economic costs of the tax avoidance behaviour induced by the higher tax rate. The federal rate increase induced a "mechanical" increase in federal tax revenues that is easily estimated, given data on total taxable income over $200,000 that is subject to the rate increase. But we must also consider the "behavioural" effect through which rate increase induce additional tax avoidance behaviour, shrinking the tax base and reducing the marginal revenue gains associated with the rate increase. Furthermore, federal and provincial governments tax the same income tax base, so that federal rate changes induce reductions in provincial as well as federal revenues through the behavioural effect. Our estimates suggest that tax responsiveness among top bracket taxpayers is high, and revenue gains from the change considerably smaller than a naive mechanical calculation would suggest. But how much smaller?

To answer this question, we performed tax revenue simulations based on the estimated elasticity of 0.5 for market income. (Although the estimated response for taxable income is in fact larger, this might reflect in part retiming responses, as for example through contributions to registered savings vehicles, which do not change the long-run present value of tax revenues.) The tax base for these calculations is actual taxable income of

Table 7: Simulated impacts on provincial revenues

	Top tax bracket:		Simulated change in revenue:	
Province	Taxable income	MTR	Total	Per capita
	($ millions)	(%)	($ millions)	($)
NL	1,193.0	16.8	-8.0	-15
PE	152.2	18.4	-1.2	-8
NS	1,593.5	21.0	-14.5	-15
NB	1,020.1	20.3	-8.9	-12
QC	20,872.9	20.3	-181.6	-22
ON	59,402.5	20.5	-524.9	-38
MB	2,905.8	17.4	-20.4	-16
SK	2,548.7	15.0	-14.7	-13
AB	24,190.2	15.0	-139.6	-33
BC	19,195.1	14.7	-107.9	-22
Federal	133,074.0	33.0	994.0	28
Total			-27.7	-1

taxpayers above $200,000 in each province in 2016. Beginning from the 2016 actual top rates, we simulate the removal of four percentage point from the federal top rate. Table 6 shows that aggregate taxable income of top-bracket taxpayers was $133.1 billion, but only $67.3 billion of this amount was subject to the top rate; the mechanical effect of a four percentage point change in the rate is therefore $2.7 billion. At the same time, our estimates suggest the rate increase results in a tax base that is $5.5 billion smaller than it would otherwise be. This results in a loss in federal tax revenues of about $1.7 billion through the behavioural effect, leaving at net increase in federal revenues of just under $1 billion. At the same time, however, shrinkage in the shared tax base reduced provincial tax revenues by an additional $1.0 billion, leaving combined federal and provincial revenue essentially unchanged from a four percentage point tax increase. Put somewhat differently, we conclude that top income tax rates are currently very close to their revenue-maximizing levels, so that rate changes have only negligible impacts on revenues.

Table 7 shows the simulated revenue losses in individual provinces, which range from $1.2 million in Prince Edward Island, to $524.9 million in Ontario, or between $8 and $38 per capita. The losses are proportionately larger in provinces that impose high top tax rates (like Ontario) and that have greater top tail income inequality, so that the proportion of the tax base in the top bracket is larger (like Ontario). The fiscal externalities among federal and provincial governments and the importance of heterogeneity among provinces is analyzed further in Milligan and Smart (2019).

6 Concluding remarks

Our research shows that high-income taxpayers have responded to recent top bracket tax increases by reducing taxable income, and these changes are likely to be persistent rather than merely transitory. High-income taxpayers, and high-income business owners in particular, appear to have access to substantial tax avoidance opportunities. As long as these tax avoidance opportunities remain available, it is unlikely that further increases in top marginal tax rates can increase the tax actually paid by the rich.

In this paper, we have explored an estimator of the ETI that uses panel data methods to deal with the rank reversal problem that makes inferences about tax responsiveness difficult with aggregate and pooled cross-section data methods. We also introduce a neighbourhood grouped-data IV estimator to deal with the mean reversion problem that arises with the traditional panel data estimator which relies only on individual taxpayer information to construct the instrument. But ore needs to be in done. In future work, we plan to explore further the neighbourhood IV, including comparisons to traditional panel data estimates, overidentification tests, heterogeneity in ETI among taxpayers, and so on.

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