Corporate tax avoidance and sales: micro evidence and aggregate implications *

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March 15, 2024

Abstract

This paper investigates the influence of corporate tax avoidance (CTA) on US firm-level sales, and its aggregate implications. In theory, CTA gives a competitive advantage to avoiding firms, which affects the distribution of sales in the economy. In practice, we find a causal impact of CTA on firm-level sales using a broad set of measures of tax avoidance. A quantitative exercise reveals that the strength of CTA in shaping changes in sales distribution varies across industries. In several key industries, the increase in CTA of large firms has reinforced their dominant position, leading to more concentration. Overall, changes in tax planning across firms explains about 15% of the changes in concentration across US industries between 1994 and 2017. Further analysis shows the impact of CTA-induced distortions on industrial output is relevant at a macroeconomic scale.

JEL codes: D22, H26, L11, D4, F23

Keywords: Tax Avoidance, Distorted Sales, Industry Concentration, IRS Audit Probability.

*This paper was first published as a CEPR DP in 2020 under the title "Corporate tax avoidance and industry concentration". We wish to thank Kimberly Clausing, Ron Davies, Michael Devereux, Peter Egger, James Hines, Martin Jacob, Marko Köthenbürger, Philippe Martin, Florian Mayneris, Sophie Osotimehin, Nadine Riedel, Leslie Robinson, Bridget Stomberg and seminar participants at CEPII, ENS Paris-Saclay, Boston University, ENS-GATE, CREST, KU-Leuven, Münster university, Kiel University, Econometric Society World Congress, ECARES-ULB, Université Paris-Dauphine, Sciences-Po, ETH Zürich, LMU Munich, and the UQAM "Corporate Taxation in a Globalized World" virtual conference meetings for useful comments and discussions. We are particularly thankful to Baptiste Souillard for sharing tax-haven information with us, and to Felipe Saffie for helping us with the CRP data. We wish to thank the ENS Paris-Saclay Booster program for funding. Martin is supported in part by funding from the Social Sciences and Humanities Research Council, and by the UQAM research chair on the local impact of multinationals.

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1 Introduction

The prevalence of corporate tax avoidance (CTA) and its implications for governments' tax revenues have received considerable attention.[1] This paper shows that, beyond its adverse impact on government revenues, CTA distorts the distribution of sales across firms within industries.

Many U.S. firms have experienced a decreasing tax burden over the last decades (Dyreng et al., 2017). This is partly due to weaker IRS enforcement (Hoopes et al., 2012; Nessa et al., 2020) and changes in U.S. regulations encouraging profit shifting to low-tax jurisdictions (Blouin and Krull, 2014; Congress, 2010; Sullivan, 2004; Hope et al., 2013; Herrmann and Thomas, 2000). We exploit such changes in audit risks and regulations across firms to show that CTA gives a competitive edge to avoiding firms, resulting in higher relative sales. The positive impact of tax planning on sales has important implication at the macro level. A quantitative exercise reveals that the strength of tax planning in shaping changes in sales distribution varies across industries. In several key industries, the increase in CTA of large firms has reinforced their dominant position, leading to more concentration. Overall, changes in tax planning across firms explains about 15% of concentration dynamics across US industries between 1994 and 2017. These tax planning practices have thus contributed to the much discussed trends in concentration in the US.[2]

We organize our analysis around three main building blocks. In the first block (section 2), we sketch the main theoretical mechanisms which guide our subsequent empirical analysis on the impacts of tax avoidance on firm-level sales. We develop a model of imperfect competition in which firms are heterogeneous with respect to their productivity and with respect to their ability to reduce reported pre-tax profits. Profit-maximizing sales are distorted when firms can inflate production costs for calculating their income tax liability. In turn, lower *effective*

[1] See, among many others, Tørsløv et al. (2022), Zucman (2014), Clausing (2016), Clausing (2020).

[2] The trend in U.S. industry concentration is discussed in Grullon et al. (2019), Furman and Orszag (2015), Philippon (2019), and Shambaugh et al. (2018). See Affeldt et al. (2021), or Bighelli et al. (2022) for evidence of increased concentration in European markets.

taxation allows avoiding firms to expand their activity. In the model, this mechanism is captured by the negative impact of tax avoidance on *effective* marginal costs.

The model highlights two key aspects regarding how CTA may eventually lead to changes in industry concentration. First, to observe industry concentration, the strength of tax avoidance must vary among firms. A uniform shift in the level of avoidance within an industry has no effect on the sales distribution, assuming all else equal. If instead larger firms adopt more aggressive tax avoidance strategies than smaller ones, they expand, thereby raising industry concentration. Second, the model allows us to draw a link between production, firms' productivity, and CTA. We show that heterogeneous CTA across firms distorts production allocation within industries. Importantly, this mechanism is operative even in a context where firms have the same elasticity of sales to tax avoidance but differ in their tax avoidance intensity.

In the second block (sections 3 to 4), we present the empirical analysis. We use firm-level data from Compustat covering the period from 1994 to 2017 to assess the causal relationship between corporate tax avoidance (CTA) and individual sales. This analysis encounters two primary challenges: defining and measuring CTA, and addressing potential endogeneity between CTA and sales. To address the first challenge, we draw upon the extensive literature which defines CTA as firm-level strategies that involves actions taken by the firm to lower taxes and outcomes of regulations taken by governments (Dyreng and Hanlon, 2021). Such definitions involve both aggressive tax plannings and also statutorily intended avoidance (accelerated depreciation deductions, research tax credits, etc.) instituted by governments. To quantify CTA, we employ four measures: a measure of tax avoidance robust to data truncation and firm loss (Henry and Sansing, 2018), permanent book-tax differences (Wilson, 2009), discretionary permanent differences (Frank et al., 2009), and tax-sheltering residuals (Desai and Dharmapala, 2006). We show in our empirical analysis that these measures of CTA capture firm tax planning, in line with the above-mentioned papers. All these measures yield consistent results regarding the impact of CTA on sales.

To address the second challenge regarding the endogeneity of CTA, we develop two strategies. The first strategy uses an instrumental variable approach that is inspired by the literature that finds a negative correlation between IRS audit probability and corporate tax avoidance (Hoopes et al., 2012; Nessa et al., 2020). Our instrumental variable is based on the residual of a regression of raw audit probabilities on asset classes and year fixed effects. In words, our exclusion restriction posits that any fluctuations in audit probabilities *within asset classes over time* will solely affect total sales through the readjustment of companies' tax strategies. Working with these audit-probability shocks instead of raw audit probability levels allows us to evacuate concerns stemming from the fact that large firms may be more likely to be audited than smaller firms. To strengthen causal interpretation, we incorporate controls for sector-specific, year-specific, and firm-specific fixed effects, along with variables such as productivity levels, the proportion of intangible assets, the likelihood of acquisition, multinational status, R&D intensity, and lobbying efforts related to taxation issues. The results are robust to the inclusion of these specific effects and controls, bolstering the interpretation of the key coefficient as the causal effect of CTA on firm-level sales.

Our second strategy takes advantage of changes in U.S. regulations that provided greater incentives for multinational corporations to engage in tax avoidance. We focus on the check-the-box (CTB) regulations enacted in 1997 and the SFAS-131 enacted in December 1998. Both regulations have been shown to facilitate the avoidance of treated firms. This alternative strategy confirms the causal positive impact of CTA on individual sales.

In the third block (sections 5 and 6), we quantify the macro-implications of tax avoidance on sales. We build upon our theoretical framework and estimates of the impact of avoidance on sales to gauge the contribution of tax avoidance to the changes in the dominance of large firms. We provide a systematic quantification of the role of CTA on the patterns of industry concentration across US industries. We show that CTA has important positive effects on industry concentration, but the impact is quite heterogeneous across industries. In several key sectors including nonstore retail, telecommunication, computer and electronic product manufacturing, or health and personal care stores, CTA accounts for a 3pp to 10pp increase in the concentration ratio of the 4 largest firms. In other industries, the tax avoidance of large relative to smaller firms has changed little or decreased, which has a weak or a negative effect on industry concentration. While CTA has an heterogeneous impact on industry concentration across industries, it has to be noted that the dynamics of concentration are themselves quite heterogeneous across industries. We show that changes in tax avoidance of large relative to small firms explains about 15% of the variation of concentration across US industries between 1994 and 2017.

Finally, we show that distortions in firm-level sales caused by CTA significantly affect real production in various industries. Using our model, we establish a connection between production, firm productivity, and CTA. The model allows us to assess whether avoiding firms in 1994 were producing more or less than expected based on their productivity levels. If they were producing more, CTA would exacerbate distortions, leading to decreased output by favoring these firms. Conversely, if avoiding firms in 1994 were producing relatively less than predicted, tax avoidance by these firms would alleviate distortions, resulting in increased real output. We find movements in CTA can induce swings in industries' output ranging from a drop of 6% to an increase of 5%. We find important effects in many industries, such as chemical manufacturing, electronic products, or non-store retailers, where CTA-induced concentration was prominent. Our quantitative exercise points to a sizable impact of corporate tax policy on real output.

In addition, to the works cited above, we contribute to different strands of the literature. The tax-planning channel has been overlooked so far by the literature on the drivers of industry concentration, which has discussed the role of technology (Autor et al., 2020), intangible capital (Crouzet and Eberly, 2019; Bajgar et al., 2021), increasing barriers to entry, lax or ineffective antitrust enforcement (Gutiérrez and Philippon, 2018; Philippon, 2019).[3] We show that when large firms increase tax avoidance more than their competitors,

[3] Cai and Liu (2009) explore the link between competition and tax avoidance using Chinese data. They show that an increase in competition increases tax avoidance.

it reinforces their dominant position. And we document that this mechanism has been at play in several sectors including nonstore retail, telecommunication, computer and electronic product manufacturing, or health and personal care stores.

We also add new insights to recent papers that analyze the distortive impact of tax avoidance. For example, Baugh et al. (2018) show how avoiding sales taxes has allowed online retailers to maintain a price advantage over brick-and-mortar retailers, which has reshaped the U.S. retail industry. Gauss et al. (2022) show that transfer pricing regulations in the E.U. increase the effective taxation of multinationals and foster the profits and sales of domestic firms. Both examples of one industry or one type of avoidance align with our findings that tax avoidance distorts individual sales. We further show that such distortions have contributed to industry concentration dynamics across many U.S. industries. Arayavechkit et al. (2018) show that heterogeneity in firms' lobbying for capital-based tax benefits generates capital misallocation.[4] We instead discuss potential misallocation driven by a distortion in sales. A strand of the literature has also focused on the real impact of one specific type of tax avoidance, namely profit shifting. Guvenen et al. (2022) show the consequences of tax avoidance on measuring various U.S. macroeconomic indicators. Egger and Wamser (2015), Mooij and Liu (2020), and de Mooij and Liu (2021) also show the adverse impact of anti-profit shifting regulations on firms' investment. Serrato (2018) shows that the repeal of section 936 of the Internal Revenue Code, which prevents U.S. MNEs from shifting profits to affiliates in Puerto Rico, has substantial real effects on the U.S. economy. Souillard (2020b) and Alstadsæter et al. (2022) show that profit shifting impacts wage inequality. Complementary to these papers, we show that tax avoidance affects the distribution of sales in the economy.

A recent study by Gallemore et al. (2023) confirms that tax planning activities of industry leaders relative to contenders is associated with changes in concentration across US industries.[5] We adopt a different approach by analyzing the causal impact of tax plan-

[4] In a related vein, Kaymak and Schott (2019) examine the capital misallocation induced by loss-offset provisions in the tax code. Kopczuk et al. (2016) emphasize how the incidence of a tax is influenced by the heterogeneity in the opportunity of evasion across economic agents.

[5] In the majority of specifications in Tables 6 to 8, Gallemore et al. (2023) show that an increase in the

ning on sales *at the firm-level*, which is new to the literature.[6] The quantification based on these micro-estimates implies that CTA explains about 15% of the variation in concentration trends across industries. Consistent with the nuanced interpretation of Gallemore et al. (2023), we document the large heterogeneity in the role of CTA across industries.

Last, we contribute to the literature on tax avoidance and firm size. Several papers in the literature argue that large firms (Gumpert et al., 2016; Davies et al., 2018) or firms with a greater product market power (Kubick et al., 2014) are likely to engage in greater tax avoidance. Our results do not dismiss the idea that large firms are more likely to avoid taxes. Empirically, we account for this by developing an IV strategy to show that causality also goes the other way. More broadly, the counterfactual experiments show that both effects might be at work, reinforcing each other. In several US industries, large firms have been more likely (or had more incentives) to avoid taxes, which reinforced their dominant position and increased industry concentration.

2 Theoretical framework

We incorporate tax avoidance in an otherwise standard model to examine how tax avoidance affects firm-level sales and under which conditions it may increase industry concentration.

Model set-up. We consider a simple economy with a distribution of heterogeneous firms that produce horizontally differentiated goods. Firm $i \in [1; N]$ produces $q_i = \varphi_i k_i$ units under constant returns from a single input k_i, for instance, capital or labor available at a unitary price, and φ_i denotes its productivity. Firms also differ in their ability $\theta_i \geq 1$ to reduce the pre-tax profits declared to tax authorities. As we clarify below, the joint distribution of (φ_i, θ_i) plays no role in the subsequent analysis, so we leave it unspecified.

relative tax planning of industry leaders – either due to their own increased tax planning efforts or a more pronounced decrease in tax planning among contenders – is associated with a rise in the Herfindahl index.

[6] From an empirical perspective, working with firm-level data allows us to have an IV strategy and include firm and industry-year fixed effects thus addressing potential confounding factors such as regulatory or technological variations and unobserved firm effects.

Denoting the statutory tax rate by t^s, firm i's income taxes are given by $t^s(p_i - \theta_i \varphi_i^{-1})q_i$. In other words, a tax-avoiding firm can increase the production costs used in calculating its income tax liability by a factor $\theta_i > 1$.

After-tax profits of firm i can be written as follows:

$$\pi_i = (1 - t^s)\left(p_i - \frac{1 - t^s\theta_i}{1 - t^s}\varphi_i^{-1}\right) q_i \tag{1}$$

Equation (1) shows that tax avoidance, that is $\theta_i > 1$, gives a competitive edge to a tax-avoiding firm by reducing its effective marginal cost $\frac{1-t^s\theta_i}{1-t^s}\varphi_i^{-1} < \varphi_i^{-1}$. Under standard conditions that guarantee a decreasing marginal revenue as a function of output, this competitive edge implies tax savings are optimally used to sustain more aggressive pricing strategies, which, in turn, boost firms' sales.[7] The increase in sales is achieved by lower costs and prices in our simple framework. Still, it could also be achieved by an increase in the perceived quality of the good in a model in which tax avoidance raises the return to marketing, advertising, or product innovation. This reduced-form modeling of tax planning is consistent with perfectly-legal loopholes. It is also consistent with various profit-shifting techniques as discussed in Appendix A.1.

Demand. We now parametrize firms' profits assuming consumers have CES preferences with an elasticity of substitution $\sigma > 1$. Total expenditure is denoted Y so that demand for the variety supplied by firm i at a price p_i is:

$$d(p_i; \mathcal{P}) = \frac{Y p_i^{-\sigma}}{\mathcal{P}^{1-\sigma}}\ , \tag{2}$$

. where $\mathcal{P}$ is the price-index: $\mathcal{P} = \left(\sum_{i=1}^{N} p_i^{1-\sigma}\right)^{\frac{1}{1-\sigma}}$. For most of what follows, we assume firms are price-index takers as in the seminal paper of Spence (1976). Monopolistically-competitive pricing leads to a constant and equal markup for all firms over their effective

[7]Such mechanism is in line with evidence in Baugh et al. (2018) that the avoidance of sales tax by online retailers allowed them to maintain a price advantage over brick-and-mortar competitors.

marginal cost; that is $p_i = \mathcal{M}\frac{1-t^s\theta_i}{1-t^s}{\varphi_i}^{-1}$, where $\mathcal{M}_i = \mathcal{M} = \frac{\sigma}{\sigma-1}$. We relax the assumption of monopolistic competition at the end of this section.

Tax avoidance: From theory to empirics. Our parameter θ_i has no direct counterpart in our dataset, but it can be mapped into observables such as the wedge between the statutory and the effective tax rates (ETRs). The ETR of firm i is given by the ratio of taxes over taxable income, that is, $t_i^e = t^s \frac{(p_i - \theta_i \varphi_i^{-1})q_i}{(p_i - \varphi_i^{-1})q_i}$, so that

$$t^s - t_i^e = t^s \left(\frac{\theta_i - 1}{\mathcal{M}_i \frac{1-t^s\theta_i}{1-t^s} - 1} \right). \tag{3}$$

The above equation shows a more favorable tax position for firm i corresponds to a higher θ_i. Using (2), we obtain sales $s_i(\boldsymbol{\varphi}, \boldsymbol{\theta}) = p(\varphi_i, \theta_i) \cdot d(p(\varphi_i, \theta_i), \mathcal{P}(\boldsymbol{\varphi}, \boldsymbol{\theta}))$ as a function of firm productivity φ_i, tax avoidance ability θ_i and the vectors of firm-level productivity and tax avoidance denoted $\boldsymbol{\varphi}$ and $\boldsymbol{\theta}$, respectively, encapsulated in the price-index. Firm sales s_i then depend on the level of tax avoidance as follows:

$$\begin{aligned} s_i(\boldsymbol{\varphi}, \boldsymbol{\theta}) &= c(\varphi_i)^{1-\sigma} \left(\frac{1 - t^s\theta_i}{1 - t^s} \right)^{1-\sigma} \mathcal{M}^{1-\sigma}\mathcal{P}^{\sigma-1}Y \\ &= z(\varphi_i) \times \tau(\theta_i) \times \kappa(\boldsymbol{\varphi}, \boldsymbol{\theta}). \end{aligned} \tag{4}$$

Firm-level sales thus consist of three (log) separable functions of, respectively, firm-level productivity ($z(\varphi_i)$), firm-level tax-avoidance intensity ($\tau(\theta_i)$), and industry-specific characteristics including the price index ($\kappa(\boldsymbol{\varphi}, \boldsymbol{\theta})$).

From tax avoidance to sales premium. It follows from equation (4) that, under monopolistically competitive pricing, the relative sales of any pair of firms i and j in the industry are given by

$$\frac{s_i\left(\boldsymbol{\varphi}, \boldsymbol{\theta}\right)}{s_j\left(\boldsymbol{\varphi}, \boldsymbol{\theta}\right)} = \frac{z(\varphi_i)\tau(\theta_i)}{z(\varphi_j)\tau(\theta_j)}. \tag{5}$$

This expression shows how differences in the intensity of tax avoidance distort relative sales in the economy. When the intensity of tax avoidance is the same across firms ($\tau(\theta_i) = \tau(\theta_j)$ for any pair (i, j)), sales dispersion in the economy is entirely driven by differences in productivity. Instead, an increase in tax aggressiveness by one firm to another leads to an increase in its relative sales. We now examine under which conditions such changes in tax aggressiveness increases concentration.

From micro to macro: concentration and aggregate efficiency. In our quantitative exercise, we use the CR4 – the combined market share of the four largest firms in an industry – as a measure of concentration. If a top 4 firm adopts a more aggressive tax-planning strategy, its market share increases relative to all firms, which leads to an increase in the CR4. If instead, a small firm adopts a more aggressive tax planning strategy, its market share increases relative to all firms, which reduces the CR4. Note that if tax avoidance is facilitated for all firms in a sector, firms' profitability may increase but their relative sales are unaffected.

We show in Appendix A.1 that the intuition is also valid if concentration is measured by the Herfindahl index (HHI) rather than the CR4. Formally, we prove the HHI increases *iff* the firm that increases tax avoidance has a sufficiently large market share ($\mathcal{S}_i > \frac{\mathcal{H}_{-i}}{1+\mathcal{H}_{-i}}$). Last, we investigate the impact of tax avoidance for efficiency, i.e. its impact on real output.

To do so, we show in Appendix A.1 that by linking firm production, firm productivity and CTA, output is given by:

$$\frac{Y}{\mathcal{P}} = \frac{K}{\sum_{i=1}^{i=N} \varphi_i^{-1} \mathcal{S}_i^{\frac{\sigma}{\sigma-1}}} \tag{6}$$

where K is the overall endowment in factor k used by firms. It is maximized when relative market shares reflect relative productivities. Instead, differences in the intensity to avoid taxes distort relative sales, as shown by equation (5). We discuss the implementation of equation (10) in Section 6.

Oligopoly pricing. The closed-form results derived above rest on CES demand and monopolistically competitive pricing assumptions. Accounting instead for oligopoly pricing implies markups are no longer constant across firms: the markup $\mathcal{M}_i$ then becomes an increasing function of a firm's market share, $\mathcal{M}_i = \mathcal{M}(\mathcal{S}_i)$.[8] Indeed, equation (4) still holds, whereas equation (5) becomes

$$\frac{\mathcal{S}_i}{\mathcal{S}_j}\left(\frac{\mathcal{M}(\mathcal{S}_i)}{\mathcal{M}(\mathcal{S}_j)}\right)^{\sigma-1} = \left(\frac{\varphi_i}{\varphi_j}\right)^{\sigma-1}\left(\frac{1-t_s\theta_j}{1-t_s\theta_i}\right)^{\sigma-1}.$$

The above equation implies that, under oligopoly again, a firm's sales increase with tax avoidance.[9] As shown in Appendix A.1, an increase in tax avoidance for the largest firm would increase its relative sales and market shares, increasing concentration as measured by the CR4 or HHI. If instead tax avoidance is facilitated for all firms, market concentration will not be affected. Furthermore, in practice, taking as a dependent variable $\mathcal{S}_i(\mathcal{M}(\mathcal{S}_i))^{\sigma-1}$ or its monopolistic approximation $\approx \mathcal{S}_i\left(\frac{\sigma}{\sigma-1}\right)^{\sigma-1}$ has virtually no impact on our estimates.

3 Data and facts

The theory predicts a positive impact of tax avoidance on firm-level sales and provides guidance on other factors that influence sales. We use detailed data from Compustat, a database of firm-level financial information from S&P Global Market Intelligence, to construct our variables of interest. The Compustat dataset contains data consolidated at the company level. Our analysis covers 1994-2017 when the U.S. had a worldwide taxation system.

Our empirical analysis focuses on firms headquartered in the U.S. and excludes subsidiaries. Consistent with prior research, we remove firms in the financial and utilities indus-

[8]For instance, under Cournot and Bertrand competition, and absent Ford effects, we get $\mathcal{M}(\mathcal{S}_i) = \frac{1}{\rho(1-\mathcal{S}_i)}$ and $\mathcal{M}(\mathcal{S}_i) = \frac{1-\rho\mathcal{S}_i}{\rho(1-\mathcal{S}_i)}$, respectively, where $\rho = \frac{\sigma-1}{\sigma}$.

[9]Note that under oligopoly, the wedge gap given by (3) holds with variable markups i.e. $t^s - t^e_i = t^s\left(\frac{\theta_i - 1}{\mathcal{M}(\mathcal{S}_)\,)_i\frac{1-t^s\theta_i}{1-t^s}-1}\right)$. A lower ETR requires a higher θ_i. An interesting implication of (3) is that an increase in tax avoidance implies the direct effect of tax savings on the ETR necessarily offsets the increase in markup.

tries because of their unique regulatory and institutional structures. The unbalanced dataset consists of 14,633 firms in 93 NAICS 3-digit industries. The dataset includes a wealth of financial information such as turnover, employment, domestic and foreign pre-tax income, as well as property, plant and equipment assets, and capital expenditures. The information on intangible assets includes acquired intangibles such as goodwill, blueprints, patents, and software. These variables are key in constructing the set of relevant controls at the firm level that is used in the empirical analysis below. Some of the observations in the dataset are missing, which decreases the size of our estimation sample to 10,914 firms. However, it covers more than 78% of total yearly sales, on average, from 1994 to 2017.

3.1 Tax avoidance

The literature in accounting and finance uses different measures to analyze tax avoidance. Our concept of tax avoidance is defined broadly as "tax planning activities that are legal, or that may fall into the gray area, as well as activities that are illegal. Thus, tax aggressive activities do not necessarily indicate that the firm has done anything improper" (Chen et al., 2010, pp. 41-42).

Four measures of tax avoidance. We use four measures of tax aggressiveness and show they yield similar results regarding the impact of tax avoidance on firm-level sales.

First, we follow the methodology proposed by Henry and Sansing (2018), which tracks the deviation between the actual amount of taxes paid ($TXPD_{is}$) and the amount that would have been paid if the pre-tax financial income (PI_i) were taxed at the statutory rate (τ):[10]

$$HS_{is} \equiv \frac{\sum_{t=1}^{S} \left(CTP_{it} - \tau \times PTI_{it}\right)}{\sum_{t=1}^{S} BVA_{it}} \ , \tag{7}$$

[10]Throughout our period of analysis, the profits of a U.S. company were taxed at 35%, regardless of the country where they were made.

where τ is the statutory tax rate and BVA_{is} is firm's i book value of assets over period s.[11] Firms that conduct aggressive tax-avoidance strategies have a value of cash tax paid that is smaller than the expected tax payment, so HS is negative.[12] When the firm has a disfavorable tax position, the value of the cash tax paid is larger than the expected tax payment, so HS is positive.[13] Following Dyreng et al. (2008) and Hanlon and Heitzman (2010), the measure is computed over several years (4, 6, or 8) to reduce the volatility found in measures that use annual data on cash tax paid.[14]

We further use alternative measures of tax aggressiveness based on Book Tax Differences (BTD) that is the total difference between a company's book income and taxable income.

Calculating BTD involves estimating taxable income, typically by adjusting the current tax expense using the statutory tax rate. Wilson (2009) found a positive link between BTD and tax sheltering. Several studies adopt variations of the book-tax difference concept as indicators of tax aggressiveness, such as Permanent Book-Tax Differences (PBTD). These studies mostly assume managers tend to favor tax strategies aimed at reducing income tax expenses. PBTD decrease taxable income and effectively reduce the firm's effective tax rate without affecting financial income reported to shareholders. PBTD are computed by subtracting deferred tax expenses from total book-tax differences and adjusting for the statutory rate (see e.g. Lisowsky et al., 2013, Frank et al., 2009). Another approach, as proposed by Frank et al. (2009), introduces the concept of Discretionary Permanent Differences (DTAX). This measure is derived by regressing total permanent book-tax differences

[11] Henry and Sansing (2018) also propose the market value of assets MVA_{is} as an alternative scalar. We show our results are robust to this alternative or when using a scale-free (binary) variable that takes the value of 1 when the HS tax gap is negative, and 0 otherwise.

[12] Several accounting papers show the HS-gap measure does capture aggressive tax-planning strategies of companies (see, e.g., Koester et al., 2017). Consistent with the view that the HS indicator captures agressive tax planning, figure A.1 shows multinationals with affiliates in tax havens have a consistently lower HS-gap than other multinationals.

[13] Firms are tax-disfavored due to unfavorable permanent and temporary book-tax differences, which may occur when expenses accrued for financial reporting purposes are deducted for tax purposes on a cash basis or when net operating loss can only be carried forward to offset future income.

[14] Because we use data that are consolidated worldwide and smoothed over the long run, we do not observe the tendency of firms to bunch around zero as observed when using unconsolidated data (see, e.g., Koethenbuerger et al., 2019).

on non-discretionary items unrelated to tax planning, using the resulting residuals as a proxy for tax aggressiveness.

Although BTD and PBTD measures are widely used in empirical studies, they are possibly affected by variations in earnings management. Desai and Dharmapala (2006) introduce the notion of tax sheltering residuals (TS_r), which seeks to differentiate between normal BTDs, unrelated to tax planning, and abnormal BTDs, attributable to tax planning efforts. The measure is obtained by isolating the residuals from estimating book-tax differences on total accruals, helping to identify the portion of BTDs that variations in earnings management cannot explain. We follow the literature and compute three alternative measures of tax aggressiveness based on book-tax differences: PBTD, DTAX, and TS_r.[15]

The correlations between the HS gap index and the measure of permanent book-tax differences (-0.739), the DTAX index (-0.736), and the TS_r measure (-0.419) are large and negative as expected. In line with Chen et al. (2010), the HS-gap and the BTD-based measures are proxies for the tax aggressiveness of the firm that reduces the tax burden. A legitimate concern is that this measure only captures differences in tax credits companies receive for activities such as R&D, and has little to do with tax avoidance. As we show later in section 4.3, our results are robust to controlling for R&D activities.

Tax avoidance in large and other firms. The conjecture regarding the correlation between tax avoidance and concentration rests on the premise that large firms have had more aggressive tax-planning strategies than smaller firms, which has strengthened concentration. Table 1 compares the evolution of tax avoidance of the largest firms and the other firms.

The group of large firms (*leaders*) gathers the top four largest firms within each 3-digit NAICS sector, whereas *other firms* comprises all other firms in Compustat. Tax avoidance is measured by the long-run HS-gap. It has increased from the mid-1990s to the most recent period, consistent with findings in Dyreng et al. (2017). The comparison of figures for *leaders*

[15] We consider the minimum values of PBTD, DTAX, and TS_r during a defined sample period s. We can therefore interpret their effects on sales as lower bounds.

Table 1: Evolution of tax avoidance by firms categories (1994-97, 2014-17)

	LR HS tax gap		
	1994-97	2014-17	Δ
Leaders	-.004	-.007	-.003
Other firms	- .0006	.0003	+.0009

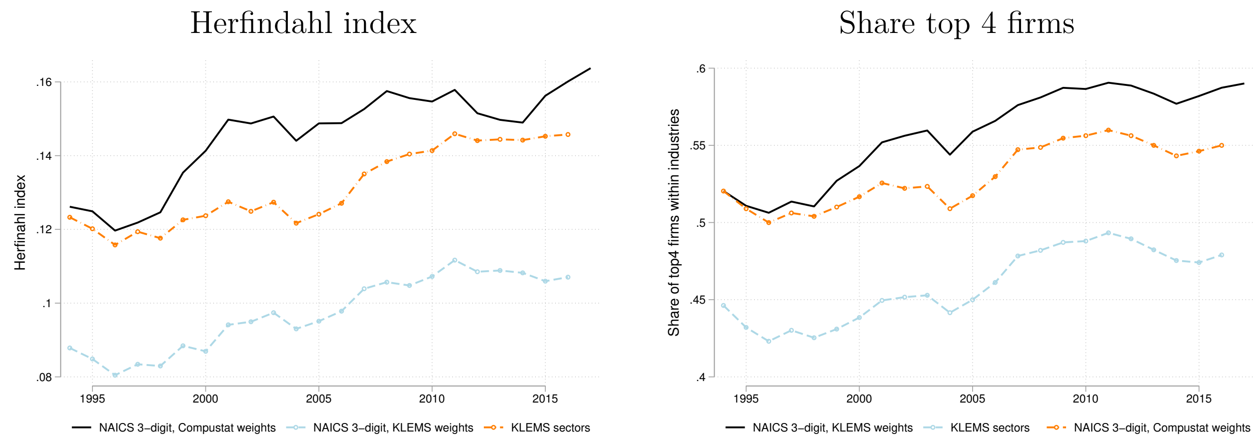
Note: *Leaders* is the group of firms that are among the four largest players (in terms of sales) in their 3-digit NAICS industry. *Other firms* are all the other firms in Compustat. The long-run HS tax gap is the ratio of the sum of cash tax paid minus .35 times the sum of profits over the sum of total assets computed within a group and a period.

and *other firms* further reveals the increase in tax avoidance has been more pronounced for the industry leaders. Tax avoidance as measured by the HS tax gap has dropped for large firms but not for smaller ones. The gap between industry leaders and other firms has also increased if one looks at the ETR, which dropped by six percentage points for the industry leaders and 4.6 percentage points for the other firms. As shown below, these average trends hide substantial heterogeneity in CTA dynamics across industries.

3.2 Industry Concentration

In Figure 1, we examine the evolution of industry concentration using either the Herfindahl index (HHI) or the share of the top four firms in sectoral sales (CR4). We define an industry as a NAICS 3-digit sector and show our descriptive results are robust using the E.U. KLEMS sector classification as an alternative definition of industry. To do so, we create a correspondence between KLEMS and NAICS 3-digit sectors. The aggregate level of concentration is a weighted average of our sector-level measures. The sector weights are computed using Compustat sectoral sales or KLEMS sectoral output data. The left panel of Figure 1 reports the evolution of the average HHI, and the right panel reports the evolution of the share of the top four firms in sectoral sales.

Figure 1: Evolution of concentration in the U.S. (1994-2017)



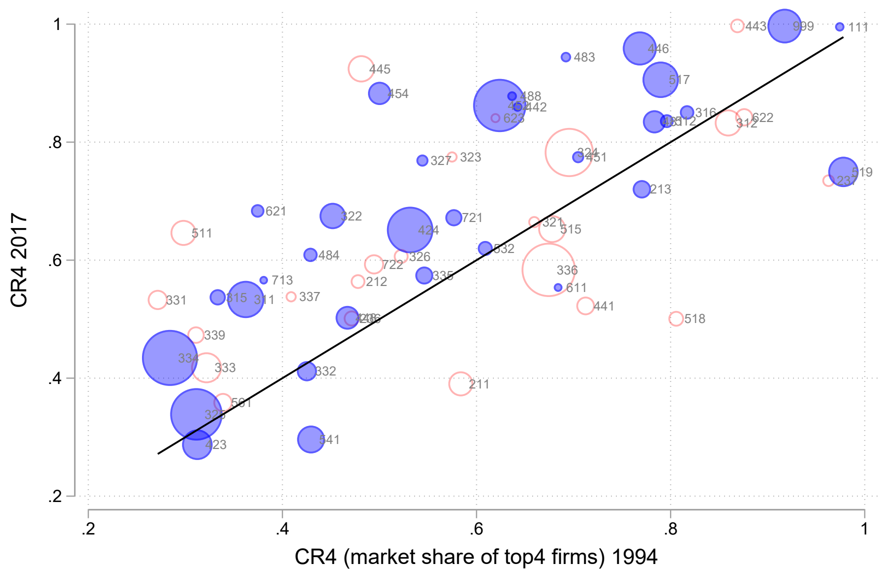
Notes: The HHI is computed at the sectoral level and then aggregated. Share top 4 is the share of industry sales made up by the four largest firms. "NAICS 3-digit KLEMS weights": sector-level concentration computed using Compustat data at the NAICS 3-digit industry level and then weighted using KLEMS data. "NAICS 3-digit, Compustat weights": concentration computed using Compustat data at the NAICS 3-digit industry level and then weighted using Compustat sectoral sales. "KLEMS sectors" is a weighted average of HHI computed from Compustat at the level of KLEMS sector.

The graphs display concentration measures computed with different aggregation schemes. Both panels confirm concentration has increased steadily in the U.S. since the mid-1990s. This observation is in line with evidence reported in Gutiérrez and Philippon (2018) and Grullon et al. (2019), among others. The documented trends are computed on a sample of U.S. publicly traded firms. Several papers show this subsample of the economy somewhat successfully tracks aggregate trends observed in comprehensive data such as the Census data (see, e.g., Covarrubias et al., 2020). As documented in Grullon et al. (2019) and discussed below, this aggregate trends mask differences in concentration dynamics across industries.

3.3 Concentration and tax avoidance at the industry level

Figure 2 presents a preliminary evidence of the link between the change in relative tax avoidance within industries and economic concentration. We compare the market share of the four largest firms (measured by the four-firm concentration ratio) across industries between 1994 and 2017, highlighting the industries in which the tax planning of the four largest firms has increased relative to smaller firms (blue circles). The level of tax avoidance is proxied by the HS measure described above. The figure shows that in many industries in the U.S., the CTA of the four largest firms has increased more than that of their competitors. This is especially valid in industries that have undergone a rise in concentration.

Figure 2: Concentration and tax avoidance in the U.S. (1994-2017)



Notes: Each circle is a NAICS 3-digit industry. Coordinates report the share of sales made up by the four largest firms in the industry (CR4). The size of the circle is proportional to the weight of the industry in the economy. Blue circles are industries in which the tax avoidance of the four largest firms has increased relative to small firms. Hollow circles are sectors in which the tax avoidance of the four largest firms has decreased relative to small firms. An increase (a decrease) in the relative tax avoidance of large firms is associated with an increase in (a reduction in) industry concentration in 31 out of 50 industries, accounting for 73% of aggregate sales.

4 Corporate tax avoidance and sales

This section focuses on the causal impact of CTA on sales at the firm-level. Implications for industry concentration dynamics are analyzed in the next sections.

4.1 Econometric analysis

Corporate tax avoidance and sales. We now turn to the empirical estimation of equation (4). Firm-level sales are determined by the intensity of CTA, factors related to the firm's productivity, and aggregate characteristics at the level of the sector, such as the price index. We estimate a log-linearized version of equation (4) throughout the empirical analysis:

$$\log s_{is} = \beta_0 + \beta_1 \log \tau_{is} + \mathbf{z}'_{is}\beta + \kappa_{k(i)s} + u_{is} \ . \tag{8}$$

The dependent variable refers to the log sales of firm i active in sector $k(i)$ in the last year of period s – or the (log) average sales across years within each period in a robustness check. Our preferred specification takes the sales at the end of the period to limit the simultaneity between sales and tax avoidance. This specification examines past tax avoidance's impact on current sales. We show the results are robust if one considers the average sales over the period rather than the end-of-period sales.

We consider several measures of tax avoidance τ_{is} as detailed in section 3.1. Because, reported tax and profit data may have significant year-to-year variations, we follow Dyreng et al. (2008) by smoothing data over several years. The measures thus capture firms' ability to pay a low amount of tax for a long period. We report the results using the four-year period and show they are qualitatively similar when using a longer six or eight-year period.

The vector z_{is} describes factors influencing firm-level productivity. In line with the theory and following the empirical literature, we use the richness of information in the Compustat dataset to construct important controls in our empirical analysis. Some research suggests productivity gains are important drivers of rising sales of large firms. Autor et al. (2020) argue the growth of *superstar* firms has been driven by productivity gains. We approximate the firm's productivity as the ratio of total sales to total employment. We also introduce a measure of intangibles because they are associated with productivity gains according to Crouzet and Eberly (2019). We calculate the intensity of firms in intangible assets as the ratio of intangible assets to total assets. We also include two indicators that have shown to be correlated with firm-level productivity: a dummy variable on acquisitions and payout and a dummy variable that accounts for the firm's multinational status.

The variables $\kappa_{k(i)s}$ are sector-time-specific factors common to all firms within each sector.

Our baseline model, therefore, identifies the impact of CTA across firms within a sector. We also use firm fixed effects to control for a broad set of unobserved firm attributes that explain the differences in the levels of sales: the firm's ability to manage tax avoidance, its corporate and managerial practices concerning tax avoidance, and its perception of the legal (tax) environment. u_{iks} is the error term.

Estimating equation (8) by least squares is unlikely to be consistent because large firms are more likely to follow more aggressive tax-planning strategies than smaller firms. Whereas potential reverse causality between size and tax avoidance justifies the use of an instrumentation strategy, the sign and magnitude of the bias of the OLS coefficient are unclear.

Instrumentation strategy. We present the instrumentation variable strategy (IV) used in our baseline analysis. In section 4.3, we present an alternative strategy exploiting the change in the reporting requirement of U.S. publicly listed firms that occurred in 1998. Both strategies lead to qualitatively similar results.

We use the audit probabilities disclosed by the Internal Revenue Service (IRS) to build the instrument for our measure of CTA. Hoopes et al. (2012) show stricter IRS monitoring implies a higher ETR. They further report that 72% of firms assess the probability of being audited when they make tax decisions. We collect data from the IRS annual Data Books, which disclose relevant information for eight asset classes across years. We construct the audit probability as the number of corporate tax return audits completed in the IRS's fiscal year t for an IRS asset-size group divided by the number of corporate tax returns received in the previous calendar year for the same IRS asset-size group upward.

These probabilities are correlated with firm size because the IRS has a higher audit rate for larger companies. We compute a measure of audit probability that is orthogonal to size- and year-specific patterns. The residual audit probabilities, $Audit_{g(i)t}^{adj}$, are constructed using the residuals of a regression of disclosed probabilities on asset class and year fixed-effects. Therefore, the mechanical correlation between the raw audit probability and firm size is broken.

The residual of the audit-probability regression captures the yearly fluctuations in audit probability specific to each asset class due to the inclusion of year-fixed effects in the regression. We visualize the audit probabilities by a class of assets over the sample period in Figure A.2 of Appendix A.2. The residual audit probabilities are not correlated with the asset class, preventing any mechanical correlation between our IV and firm size. A high residual audit probability for an asset class in a given year means that, within this year, firms in this asset class are relatively more likely to be audited.

Whereas adjusted audit probability is specific to a year and an asset class, our instrument is specific to a firm and a period. We take for each firm in each year its corresponding residual audit probability that we average over four years in our baseline estimation – or six or eight years in a robustness check. Firms may change asset classes within a period. We thus keep the asset class constant *within a period*.[16] The instrument is given by:

$$Adj.\ Audit_{is} = \sum_{t \in s} Audit^{adj}_{g(i)t} / N^S,$$

where $Adj.\ Audit_{g(i)t}$ is the adjusted audit probability of firms in asset-size class g and N^S is the number of years t in period s. The within-sector and period correlation between our instrument and the raw audit probabilities is low at about 19%.

Our identification strategy rests on the assumptions that (i) changes in tax avoidance at the level of firms are correlated to changes in the audit rates ($\mathbb{E}(\tau_{is}, Adj Audit_{is}) \neq 0$), ($ii$) fluctuations in audit probabilities within asset classes over time will solely affect total sales through the readjustment of companies' tax strategies ($\mathbb{E}(u_{is}, Adj Audit_{is}) = 0$).

Assumption (i) is likely to hold. The literature on tax enforcement finds that, all else equal, a decrease in tax enforcement is positively correlated with tax avoidance (Hoopes et al., 2012; Nessa et al., 2020). We verify this link in the first stage of our specification.

[16] More specifically, we set it to the asset class of the first year of the period. Results are very similar if the instrument includes asset class changes within a period.

There is no proper test of assumption (*ii*). It states that, *conditional* on our explanatory variables, audit probability shocks are not correlated with individual sales. Reassuringly, the literature suggests that audit probability shocks are associated to IRS funding and thus unrelated to the individual sales of US firms. Indeed, Nessa et al. (2020) show that IRS resources positively correlate with audit probability and the net revenue collected through tax enforcement. Fluctuations in the IRS enforcement budget resulted from Congress warfare between Democrats and Republicans. As Kiel and Eisinger (2018) report, the Republican-controlled Senate in 1997 and 1998 held a series of dramatic hearings on alleged abuses by the IRS.[17] Hoopes et al. (2012) argue this drop is tightly linked to cuts in the budget of the IRS. It is thus likely exogenous to individual companies' sales dynamics.

A related concern might be that growing firms have more resources to lobby on taxation, thus affecting audit probability. As we just discussed, there is no evidence in the literature that changes in IRS audit probability over the period have been driven by firms' lobbying. We show in section 4.3 that a firm's participation in lobbying on taxation or internal revenue code does not influence the effects of the residual audit probability on the level of CTA. Lobbyists have larger sales, but controlling the lobbying status of firms does not affect the impacts of tax avoidance on sales.

4.2 Baseline results

Table 2 displays the results of both the OLS and the second stage of the 2SLS regression analyses for the four tax aggressiveness measures. In Panel A, we present the outcomes when using the HS gap index. Panels B through D report the estimates for the permanent book-tax differences, discretionary permanent differences, and tax sheltering residuals indexes, respectively. Each panel includes the first stage coefficient of the adjusted audit probability, the sample size, and the Kleibergen-Paap F-statistic (KP F-stat) value. We

[17] A more recent reason for cutting the IRS funding is that the agency was chosen to monitor the Affordable Care Act. Instead, there were no cuts during the George W. Bush administration, and tax collection increased over this period (and the audit probability of large firms remained flat), which political commentators explain by the fact that the IRS was not an object of a dispute during this era.

include a set of sector- and period-specific effects to control for unobserved characteristics in each specification. We, therefore, identify the effect of each covariate using the variation in firm-level attributes across firms within sectors and periods. We also include firm fixed effects in some specifications. In this case, we use the variation of firm-level characteristics within the firm to identify the effect of tax avoidance and other covariates. We report robust standard errors in all specifications.[18] The full estimation results are displayed in Table A.1 and A.2.

We report the OLS results, including sector and period fixed effects in column (1) and firm fixed effects in column (2). In Panel A, we find a negative impact of the HS Gap measure on firm-level sales, which suggests a positive impact of tax avoidance on sales. The positive coefficients of the tax aggressiveness measures based on book-tax differences in panel B to D confirm this finding. The coefficients are significant at the 99% confidence level. In Table A.1 and A.2 which report the coefficients of the other covariates, we show that they have the expected signs and are highly significant at conventional levels. Firms with a larger share of intangibles and higher workforce productivity have larger sales. These findings support the results of Crouzet and Eberly (2019) and Autor et al. (2020) that industry leaders are often very good at producing intangible assets and are highly productive. In line with the literature that looks at the performance of multinational firms, we find these firms have larger sales than purely domestic firms (Antràs and Yeaple, 2014). The acquisition dummy variable is also positive and highly significant. These results are robust to including firm fixed effects in column (2).

We report the results of the 2SLS estimations using sector and period fixed effects in column (3) and firm fixed effects in column (4). In each panel, we report the coefficient of the adjusted audit probability from the first stage as well as the Kleibergen-Paap F-statistic (KP F-stat). We find a positive and highly significant impact of the residual audit

[18]The sampling and assignment mechanisms are not clustered in our setting. We thus follow Abadie et al. (2023), who argue that "if the sampling and assignment mechanisms are not clustered, one should not adjust the standard errors for clustering, irrespective of whether such an adjustment would change the standard errors." The results hold if standard errors are clustered at the firm's level and can be available upon request.

Table 2: Sales and tax avoidance – OLS and 2SLS estimates

Dep. Variable	Log Sales - End of Period (1)	(2)	(3)	(4)
	OLS		2SLS (2^{nd} Stage)	
A: HS tax gap	-2.648***	-1.038***	-5.408***	-4.085***
	(0.125)	(0.118)	(0.541)	(1.439)
1^{st} *Stage*: Adj. Audit Proba,			0.011***	0.002***
			(0.001)	(0.001)
Obs.	22,271	18,546	22,271	18,546
KP F-stat.			120.1	9.375
B: Permanent Book-Tax Differences	0.211***	0.067***	0.559***	0.333***
	(0.014)	(0.012)	(0.062)	(0.093)
1^{st} *Stage*: Adj. Audit Proba,			-0.165***	-0.039***
			(0.018)	(0.008)
Obs.	15,137	11,977	15,137	11,977
KP F-stat.			85.64	21.65
C: Discretionary Permanent Differences	0.207***	0.062***	0.554***	0.343***
	(0.015)	(0.012)	(0.062)	(0.097)
1^{st} *Stage*: Adj. Audit Proba,			-0.167***	-0.039***
			(0.018)	(0.009)
Obs.	14,918	11,784	14,918	11,784
KP F-stat.			83.38	19.86
D: Tax Sheltering Residuals	0.198***	0.065***	1.467***	0.388***
	(0.015)	(0.012)	(0.217)	(0.110)
1^{st} *Stage*: Adj. Audit Proba,			-0.061***	-0.034***
			(0.009)	(0.007)
Obs.	14,569	11,988	14,569	11,988
KP F-stat.			46.15	20.73
Other Controls	Yes	Yes	Yes	Yes
Sector × Period FE	Yes	Yes	Yes	Yes
Firm FE	No	Yes	No	Yes

Table A.1 and A.2 presents the full estimation results. Sample years: 1994-2017. The dependent variable is the firm's log sales at the end of the four-year window. OLS and 2SLS estimates with robust standard errors in parentheses. First-stage Kleibergen-Paap Wald F statistic reported. ***, **, and * significantly different from 0 at the 1%, 5%, and 10% confidence levels, respectively.

probabilities on the HS gap. This result is in line with Hoopes et al. (2012) and Nessa et al. (2020), which show U.S. firms undertake less aggressive tax positions when tax enforcement is stricter. This finding suggests our tax avoidance measure captures more than legal tax breaks.

As noted previously, the specification that includes sector and period fixed effects uses the variation in the IRS's audit probabilities across firms within sector and period. This finding suggests the heterogeneous cuts in the IRS's audit probabilities have contributed to tax avoidance across firms within sectors and periods.

The second-stage results confirm the causal impact of tax avoidance on firm-level sales. In Panel A, the coefficient of the HS_{is} measure is negative and highly significant while in Panel B to D, we find the positive and significant effects of the other tax aggressiveness measures.

We still find that reducing the IRS audit probability significantly increases firm-level tax avoidance in specifications that include firm fixed effects. The results confirm tax avoidance's causal and positive impact on firm-level sales.

In both specifications, the large value of the Kleibergen-Paap F-statistic (KP F-stat) confirms the strength of our instrument. The statistics yield values that are for most of them larger than 10 in the model that uses sector×period-specific effects and those that include firm-specific effects. These findings suggest the regression estimator is unlikely to suffer from weak-instruments bias.

4.3 Robustness of the results

The finding that CTA has a causal impact on sales is key to our analysis. In this section, we show this result is robust to the introduction of various controls, changes in the construction of our main variables of interest, and using an alternative identification strategy. The main findings of this robustness analysis are summarized below.

Lobbying activities and R&D. We first conduct a thorough investigation to ensure that our results are not subject to any potential bias from confounding factors such as tax lobbying activity of firms and R&D expenditures.

It is legitimate to ask whether heterogeneity in our tax avoidance measure reflects differences in R&D tax credits, which would lead to a different interpretation of our findings. An increase in R&D may spur sales and allow firms to save on taxes. Information on R&D expenditures is missing for about 40% of the firms and observations in our baseline sample. We show the positive impact of tax avoidance on sales is robust to the inclusion of R&D expenditures in this restricted sample. We lag the R&D variable by five periods because the effects of R&D on firm-level sales may not be contemporaneous.[19]

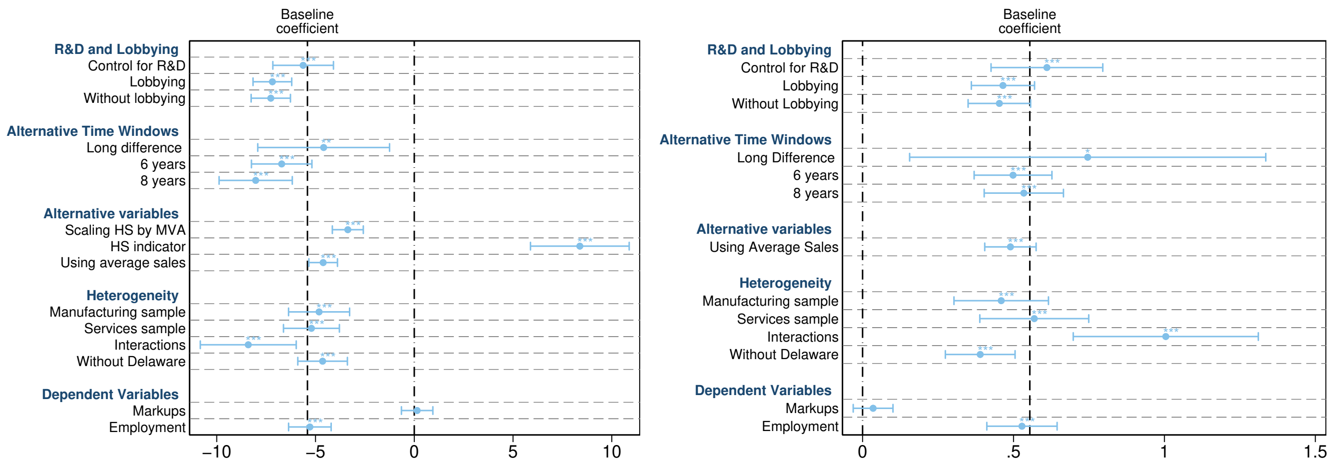
Arayavechkit et al. (2018) show the firm size and effective tax rate are positively correlated with their lobbying activities. We use data on lobbying activities of U.S. firms from Kim (2018) to create an indicator that provides information on the firm's participation in lobbying on taxation or the Internal Revenue Code. Firms that engage in lobbying on taxation issues are larger than non-lobbying firms, but this variable does not affect the causal effect of CTA on sales. Lobbying has no impact on the level of CTA, and including the lobbying dummy variable does not affect the effects of residual audit probability on aggressive tax-planning strategies. To mitigate the influence of lobbying activities, we estimate the effect of tax avoidance on sales in a sample of firms that do not engage in lobbying activities. The causal effect of tax avoidance on sales persists.

Other sensitivity analysis. We conduct additional tests to assess the causal relationship between tax avoidance and sales and summarize our key findings below. Figure 3 presents the second-stage coefficients of the HS tax gap (Figure 3a) and of the measure of discretionary permanent differences (Figure 3b) in various regression models used for our sensitivity analysis. The results are similar using the other measures of BTD-based tax aggressiveness.

First, we investigate whether our results are sensitive to the length of the time window used for identification. Our results hold when using a six- or eight-year window. However, the effects become statistically insignificant when we include firm-fixed effects, which is expected

[19]Using contemporaneous R&D expenses does not change the main results: an increase in tax avoidance positively affects sales. The coefficient of the contemporaneous R&D variable is, however, not significant.

given the smaller number of periods used for identification. Moreover, the results using a long difference in a sample of firms with positive sales in both the first and the last period of our sample confirm our findings on the causal impact of tax avoidance on sales. Additionally, using the average sales within a time window rather than the sales at the end of the period as an explanatory variable does not affect the results.



(a) HS Gap

(b) Discretionary Permanent Differences

Figure 3: Sensitivity analysis

Notes: The baseline coefficients are taken from Table 2 and are the semi-elasticity of the HS gap (coef=-5.408, s.e.=0.541) and measure of discretionary permanent differences (coef=-0.554, s.e.=0.062) in the regression that includes sector and period FE.

Second, we test our results' sensitivity to the alternative measure of the HS gap. Similar results are obtained when using the market value of assets instead of the book value to scale the measure. Using an indicator variable, we also create a scale-free measure of avoidance and find that tax-avoiding firms have larger sales. This indicator takes the value of 1 when the HS tax gap is negative, and 0 otherwise, so we expect a positive coefficient. Additionally, using the average sales within a time window rather than the sales at the end of the period as an explanatory variable does not affect the results in both panels.

Third, we investigate the heterogeneity in the impact of tax avoidance on sales across sectors and firms. We find that the positive impact is pervasive in manufacturing and services industries and is magnified for high-productivity firms and those that increased their share of intangible assets. Furthermore, we show that our results are not driven by firms incorporated in Delaware.

Fourth, we test the impact of tax avoidance on firm-level markups and find no significant effect, consistent with our baseline model under monopolistic competition and CES preferences. Finally, we show that tax avoidance positively and significantly impacts firm-level employment. Our results suggest that tax avoidance can increase employment by reducing effective marginal costs and increasing output.

Alternative identification strategies. We now show the robustness of the causal impact of tax avoidance on sales to an alternative identification strategy that exploits changes in how multinationals could report their earnings. We focus on two events that eased U.S. multinationals' avoidance: the check-the-box (CTB) regulations enacted in 1997 and the SFAS-131 enacted in December 1998. Under CTB regulations, U.S. firms can choose whether foreign subsidiaries should be treated as foreign corporations or disregarded as foreign entities of another corporation. In the latter case, payments between the disregarded foreign entities and other controlled foreign corporations of a U.S. firm are disregarded for U.S. tax purposes. SFAS 131 breaks the mandatory requirement to disclose geographic earnings by the jurisdiction in financial reporting. Both regulations have been shown to facilitate the avoidance of treated firms (Blouin and Krull, 2014; Congress, 2010; Sullivan, 2004; Hope et al., 2013; Herrmann and Thomas, 2000)

We compare the outcomes of multinationals relative to domestic firms during the 1993-1996 period before these two regulations were enacted with their outcomes during the 1999-2004 post-treatment period when both regulations were enacted.[20] More specifically, we use this experiment in two ways. First, we estimate a standard difference-in-differences equation in which sales are regressed on an indicator of the MNE status of the firm and its interaction with a dummy variable, which takes the value of 1 in the period post-treatment.

[20] The results hold if we use 1997-2004 as the post-treatment period.

We expect the interaction term to positively impact sales because the SFAS-131 and the CTB regulations help multinationals avoid taxes. The coefficient of the *Post* variable cannot be identified because it is perfectly collinear with the sector and year fixed effect. Second, we use the interaction term as an instrument for tax avoidance in the 2SLS specification.

Table 3 presents the results of our alternative strategy using both difference-in-differences and 2SLS estimations. The difference-in-differences strategy compares the sales of multinational firms (the treated group) with domestic firms (the control group), before and after the introduction of laxer regulations. The interaction between our treatment variable and the post-reform dummy is positive and significant, which means firms that benefited from the change in legislation experience increased sales. This finding holds when we use firm fixed effects.

Table 3: Sales and tax avoidance – quasi-experiment

Dep. Variable	Log Sales - End of Period					
	Diff-in-Diff	2SLS		Diff-in-Diff	2SLS	
		1^{st} Stage	2^{nd} Stage	1^{st} Stage	2^{nd} Stage	
HS tax gap			-4.910***			-9.251**
			(0.902)			(3.591)
Share of Intangible	1.568***	-0.030**	1.420***	1.610***	-0.034	1.300***
	(0.137)	(0.014)	(0.132)	(0.206)	(0.028)	(0.311)
Labor Prod.	0.963***	-0.065***	0.644***	0.668***	-0.037***	0.330**
	(0.027)	(0.004)	(0.063)	(0.059)	(0.008)	(0.160)
Acquisition	1.069***	-0.022***	0.963***	0.303***	-0.009***	0.220***
	(0.039)	(0.004)	(0.042)	(0.037)	(0.003)	(0.055)
MNE Status	1.313***	-0.008***	1.276***			
	(0.051)	(0.002)	(0.053)			
MNE $\times$ $Post_{1998}$	0.386***	-0.079***		0.134***	-0.014***	
	(0.073)	(0.007)		(0.046)	(0.004)	
Sector $\times$ Period FE	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	No	No	No	Yes	Yes	Yes
Obs.	11051	11051	11051	6178	6178	6178
Adj. R^2	0.468	0.184		0.889	0.358	
KP F-stat.			117			11.41

The dependent variable is the firm's log sales at the end of the six-year window. We have two periods of analysis starting in 1993. OLS and 2SLS estimates with robust standard errors in parentheses. First-stage Kleibergen-Paap Wald F statistic reported. ***, **, and * significantly different from 0 at the 1%, 5%, and 10% confidence levels, respectively.

The first stages of the 2SLS regressions use the HS-gap measure of tax avoidance as a dependent variable. Whatever the set of fixed effects included in the regressions is, we find negative and significant interaction coefficients, which confirms that multinational firms significantly increased their level of tax avoidance after 1998. The second-stage results show a negative and significant impact of the instrumented tax avoidance measure on sales. These results confirm the causal impact of tax avoidance on sales.

5 Impact of tax avoidance on concentration

Having established the causal effect of tax avoidance on sales, we explore its consequences for industry concentration. We build on the insights of our theoretical model and use the estimated semi-elasticity of CTA firms' sales (Table 2, column 6) to assess the quantitative impact of tax avoidance on the distribution of sales. To do so, we compute the counterfactual level of industry concentration in 1994 if firms resorted to their CTA strategy of 2017. This experiment is a way to single out the contribution of CTA to the 1994-2017 change in industry concentration.

Firm i's counterfactual level of sales can be written as $\hat{s}_i = z(\varphi_i) \times \tau(\hat{\theta}_i) \times \hat{\kappa}$. Using the expression of $z(\varphi_i)$ from equation (4), firm i's counterfactual sales and its market share are a function of observed sales and change in CTA:

$$\begin{aligned} \hat{s}_i &= \frac{\tau(\hat{\theta}_i)}{\tau(\theta_i)} \times \frac{\hat{\kappa}}{\kappa} \times s_i \\ \Leftrightarrow \frac{\hat{s}_i}{\sum_j \hat{s}_j} &= \frac{\tau(\hat{\theta}_i)/\tau(\theta_i) \times s_i}{\sum_j \tau(\hat{\theta}_j)/\tau(\theta_j) \times s_j}. \end{aligned} \tag{9}$$

Note the counterfactual market share of firm i in equation (9) - unlike the counterfactual level of sales - does not depend on the counterfactual price index. It only depends on (observed) individual sales s_i, and the impact of a change in tax avoidance of sales $\tau(\hat{\theta}_i)/\tau(\theta_i)$, which is back-out from the observed changed in tax avoidance and the sensitivity of tax avoidance to sales estimated in Table 2.

We do not observe the level of avoidance of each firm in 1994 and 2017 because of attrition.

To address this challenge, we consider the sales of the four largest firms in each industry in 1994 and the level of tax avoidance of the top 4 in 1994 and 2017. We use expression (9) to compute the counterfactual levels of CR4 (the combined market share of the top largest firms within an industry) across U.S. industries:

$$\widehat{CR4} = E(CR4_{1994}|Avoid = Avoid^{2017}, \varphi^{1994}) \ .$$

The deviation of the counterfactual from the observed CR4 in 1994: $\Delta_c = \widehat{CR4} - CR4_{1994}$ allows us to answer our question. Δ_c provides information on the change in concentration measured by the CR4 in 1994 had firms followed their CTA strategy of 2017. Besides Δ_c, we also report in Table 4 the observed concentration ratios and their evolutions between 1994 and 2017, $\Delta_o = CR4_{2017} - CR4_{1994}$.

Table 4: Evolution of observed and counterfactual industry concentration (in percent point)

NAICS code		Weight	CR4 (1994)	Δ_o	Δ_c
488	Support Activities for Transportation	0.1	74.5	16.93	9.56
454	Nonstore Retailers	1.9	49.8	40.13	4.14
721	Accommodation	0.7	62.9	3.83	3.31
621	Ambulatory Health Care Services	0.5	34.5	32.38	2.86
517	Telecommunications	3.2	77.6	12.10	2.33
334	Computer and Electronic Products	8.3	26.7	15.49	2.27
484	Truck Transportation	0.5	48.2	9.82	1.68
322	Paper Manufacturing	1.1	46.8	18.90	1.53
337	Furniture and Related Products	0.2	41.8	11.35	1.48
212	Mining (except Oil and Gas)	0.6	46.4	7.86	1.43
316	Leather and Allied Product Manufacturing	0.5	77.0	9.32	1.33
446	Health and Personal Care Stores	3.7	69.7	25.74	1.13
512	Motion Picture and Sound Recording	0.4	78.8	5.23	1.11
441	Motor Vehicle and Parts Dealers	1.1	73.6	-21.97	0.80
332	Fabricated Metal Products	0.8	39.0	2.39	0.67
541	Professionals Services	2.0	42.6	-11.33	0.56
424	Merchant Wholesalers	6.8	56.1	10.96	0.55
335	Electrical Equipment and components	0.6	57.5	-2.88	0.54
325	Chemical Manuf.	7.0	31.9	0.80	0.37
336	Transportation Equipment	6.7	67.2	-11.14	0.35
315	Apparel Manuf.	0.6	50.1	1.40	0.30
713	Recreation Indus.	0.1	52.4	4.66	0.25

Continued on next page

Table 4 – *Continued from previous page*

	NAICS code	Weight	CR4 (1994)	Δ_o	Δ_c
452	General Merchandise Stores	7.0	63.8	22.56	0.12
519	Other Information Services	2.8	99.0	-24.48	0.03
423	Merchant Wholesalers	2.4	36.6	-2.66	0.03
722	Food Services and Drinking Places	1.0	50.7	7.75	-0.01
333	Machinery Manuf.	2.2	33.4	7.35	-0.01
611	Educational Services	0.1	68.0	-3.29	-0.04
213	Support Activities for Mining	0.8	81.3	-9.99	-0.09
324	Petroleum and Coal Products	5.9	70.1	6.01	-0.11
236	Construction of Buildings	0.8	50.4	-1.64	-0.13
311	Food Manuf.	2.9	37.6	12.77	-0.13
448	Clothing Stores	1.3	49.5	1.75	-0.14
327	Nonmetallic Mineral Products	0.2	55.5	19.80	-0.16
481	Air Transportation	1.1	80.0	2.45	-0.16
515	Broadcasting (except Internet)	2.6	84.4	-18.57	-0.17
339	Miscellaneous Manuf.	0.8	28.6	21.23	-0.38
518	Data Processing and Related Services	0.8	87.8	-39.38	-0.48
326	Plastics and Rubber Products	0.4	56.9	10.18	-0.57
331	Primary Metal	0.8	30.5	25.80	-0.58
312	Beverage and Tobacco Products	1.5	86.0	-7.05	-0.82
532	Rental and Leasing	0.5	64.1	-2.15	-1.04
211	Oil and Gas Extraction	1.5	57.0	-19.93	-1.17
511	Publishing Industries	2.0	28.2	33.56	-1.34
622	Hospitals	0.9	80.7	0.90	-1.86
321	Wood Product Manufacturing	0.3	66.6	0.82	-1.98
561	Administrative Services	1.1	43.7	-7.02	-4.06

Our analysis of the U.S. economy shows that in around 40% of industries, responsible for half of total sales, increased tax avoidance by big companies has led to more industry concentration over the last two decades. This trend is seen in many key sectors including nonstore retail, telecommunication, computer and electronic product manufacturing or health and personal care stores.

This finding lines up with existing evidence. For example, Amazon's tax-avoidance practices have helped concentrate the nonstore retail market, giving it an edge over traditional retailers. Baugh et al. (2018) show the sales-tax avoidance of nonstore retailers gave them an advantage over brick-and-mortar retailers. A similar pattern is likely happening in the non-store retail sector for companies using their global online presence to cut down on corporate taxes.

Of course, tax avoidance is not the only ingredient that contributes to concentration in these industries. The aggressive tax planning of Apple or Amazon is likely to have been detrimental to their competitors, but these companies' growth is also the result of breaking innovation. Consistent with this view, the predicted changes in concentration are a fraction of the observed changes, which shows tax avoidance is an important but not the only determinant of industry concentration.

In 20% of industries, accounting for 10% of sales, we find that the reduced relative importance of CTA for large firms has led to a decrease in industry concentration, and we've observed a corresponding decline in concentration as well. This trend is evident in sectors like administrative and support services, with this category being the largest in that group.

In the third category of industries, accounting for 40% of all industries and 40% of total sales, CTA influences concentration in one direction, but other factors offset the effects of tax avoidance. For instance, in the petroleum and coal products manufacturing sector, which makes up about 6% of U.S. sales, we observe an increase in observed concentration, while the counterfactual CR4 show a slight decrease. The food and beverage stores sector is another example where tax avoidance appears to have had no significant role in the substantial increase in concentration.

Our analysis suggests that the choice of concentration measure does not significantly impact the observed trends. The sectors where large firms increase CTA and concentration remain consistent across CR3, CR4, and CR5, with high correlations in observed changes (98% and 96%) and in predicted counterfactuals (91% and 93%) among these measures.

A quantitative approach to assess the significance of tax avoidance in shaping concentration dynamics across US industries is to examine how much of the variation in concentration changes across these industries can be accounted for by changes predicted based on changes

Table 5: Explanatory power of CTA in changes in industry concentration

	Changes in:		
	CR3	CR4	CR5
$\Delta CR\ counterfactual$	4.76	4.66	4.92
	(1.71)	(1.60)	(1.27)
# obs.	47	47	47
R^2	12.2%	14.8%	19.6%

The dependent variable is the change in the observed concentration ratio (CR4, CR3 or CR5) between 2017 and 1994. The explanatory variable is the predicted change in the concentration ratio, based on the observed change in corporate tax avoidance of large vs smaller firms over the period. R^2 are obtained from a weighted regression where industries are weighted based on their aggregate sales. Sectors with less than 10 firms are excluded.

in the relative CTA between large and small firms within each industry. Specifically, we conduct regressions of the observed changes in concentration against the predicted changes, as presented in Table 4. These regressions are performed using three concentration measures (CR3, CR4, and CR5). The results of these regressions are reported in Table 5. The coefficient of determination (R^2) for these regressions ranges from 12.2% to 19.6% across the three measures, indicating that a significant portion of industry-specific concentration dynamics can be attributed to corporate tax avoidance.

6 Extension: the macroeconomic cost of tax avoidance

We show in this extension that CTA-induced distortions in firm-level sales are large enough to influence real production in many industries significantly. The model used through the analysis draws a link between production, firms' productivity, and CTA, leading to equation (10). Tax avoidance, however, is only one of the many sources of sales distortion in the economy. Thus, we adjust equation (10) as follows:

$$\frac{Y}{\mathcal{P}} = \frac{K}{\sum_{i=1}^{i=N} \widetilde{\varphi}_i^{-1} \mathcal{S}_i^{\frac{\sigma}{\sigma-1}}} \tag{10}$$

where the existence of other distortions imply that $\widetilde{\varphi}_i \neq \varphi_i$. The initial gap between $\widetilde{\varphi}_i$ and φ_i imply that tax avoidance reduction can decrease or increase real output. Concretely, large firms in 1994 might have been producing more than what would be predicted by their productivities φ_i. In this case, tax avoidance by large firms would increase distortions and thus decrease output by allocating more market shares to large firms. Alternatively, large firms in 1994 could produce relatively less than predicted by their relative productivity. In this case, tax avoidance by large firms would decrease distortions, and real output would increase.

In the quantitative exercise below, we remain agnostic about the type and magnitude of non-tax distortions at play. We derive the upper and lower bounds of real-output change implied by CTA changes between 1994 and 2017. Whether tax avoidance has the potential to affect real output can be directly observed by the wedge between the upper and lower bounds. If CTA does not play a role, the wedge should equal zero. In each sector, we consider two representative firms. The variable $\mathcal{S}$ denotes the market share of the four largest firms, and $1-\mathcal{S}$ aggregates the market shares of all other firms. By construction, the upper bound is larger than 1, and the lower bound is lower than 1. It reflects the ambiguous impact of CTA on real production depending on other distortions as described above.

We show in Appendix A.1 that the change in real production belongs to the following interval:

$$\widehat{\left(\frac{Y}{P}\right)} \in \left[\min\left\{\left(\frac{\mathcal{S}'}{\mathcal{S}}\right)^{\frac{\sigma}{\sigma-1}}, \left(\frac{1-\mathcal{S}'}{1-\mathcal{S}}\right)^{\frac{\sigma}{\sigma-1}}\right\}; \max\left\{\left(\frac{\mathcal{S}'}{\mathcal{S}}\right)^{\frac{\sigma}{\sigma-1}}, \left(\frac{1-\mathcal{S}'}{1-\mathcal{S}}\right)^{\frac{\sigma}{\sigma-1}}\right\}\right]. \quad (11)$$

$\mathcal{S}'$ denotes the counterfactual market shares of the four largest firms in 1994 as if they had a tax-avoidance level of 2017. Y' and $\mathcal{P}'$ are the counterfactual values of sector-level output and price index, respectively. These bounds are conservative intervals of the impact of tax avoidance as we abstract from inter-industry linkages.

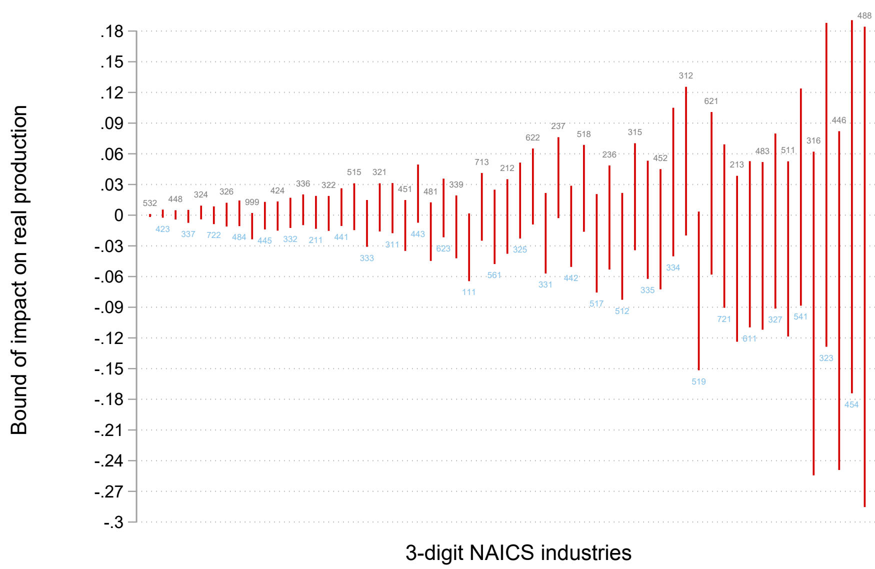
We use equation (11) to compute the upper and lower bounds for each industry by assuming a value of σ to be equal to 2. For the average industry, the CTA-induced change in the CR4 index is consistent with a movement in real production ranging from a drop of 6% to an increase of 5%. Note the results remain quantitatively important, assuming a larger elasticity σ equal to 4. In this event, the upper and lower bounds of production changes are

3% and -4%, respectively.

Determining the sign and exact magnitude of these changes would require an estimation of all other frictions in the economy, which is beyond the scope of this paper.

The upper and lower bounds for each sector are reported in Figure 4. We see that in many industries, the change in the CR4 induced by the change in tax avoidance of large versus small firms might be consistent with large swings in real production.

Figure 4: Bound of CTA impact on real production



Notes: each dash represents the upper and lower bound of the impact of CTA on real production. The 3-digit code of each sectors is reported at the end of the dash. Bounds are computed using formula in eq 11, calibrating $\sigma = 2$. Reading: for NAICS sector 212, the change in CTA could change real production from -3% to 3%.

The changes in market shares induced by CTA could have quantitatively important effects on production in many industries. Such industries include chemical manufacturing, electronic products, or non-store retailers, in which CTA has been shown to play a significant role in the increase in concentration. Real production in these industries could move between -2% and 5%, -4% and 10%, and -18% and 18%, respectively. In other industries, CTA has a small impact on real production, such as in the transportation equipment manufacturing industry, where the CTA effect on real production ranges from -0.5% to 0.7%. As shown earlier, this industry experienced a decrease in concentration, as did the petroleum and coal production manufacturing, in which concentration increased, but the CTA of large firms decreased relative to small firms. In this industry, the CTA effect on real production ranges from -1% and 2%.

To sum up, our results show that changes in CR4 induced by tax avoidance may negatively impact real production. Furthermore, note we have abstracted from sectoral linkages that would only magnify the impact on real output.[21] Policy-wise, we conclude that, beyond the increase in corporate tax revenue, curbing tax avoidance can induce a substantial change in allocative efficiency.

7 Conclusion

This study provides compelling evidence of the causal relationship between tax avoidance and firm-level sales. We demonstrate that tax avoidance gives a competitive edge to firms that engage in it. In several key industries, CTA has increased more for larger than smaller firms, reinforcing their dominant position and increasing concentration. Overall, changes in tax avoidance of large relative to small firms explains about 15% of the variation of concentration across US industries between 1994 and 2017. We also find that tax avoidance can distort market shares and influence real production in many industries beyond its impact on government revenues.

Our results have important implications for tax and competition policy, and our findings highlight the interdependence of these policies. We demonstrate that the enforcement of corporate tax policy can help curb industry concentration in industries where dominant firms have the most aggressive tax strategies, and that laxer tax enforcement has distorted sales in favor of avoiding firms. The European Commission's rulings against tax breaks given to Apple, Fiat, Amazon, and Starbucks further underscore the need for coordinated tax policies to ensure fair competition within the European Union.[22]

[21] For instance, if sectors are aggregated in a Cobb-Douglas fashion, the presence of distortions in a single sector affects the rest of the economy.

[22] According to Commissioner Margrethe Vestager, "We have to continue to use all tools at our disposal to ensure companies pay their fair share of tax $(\cdots)$ If Member States give certain multinational companies

Our study has broad implications for any reform of international taxation that changes the relative tax position of small and large firms, as such reforms could significantly impact firms' sales distribution and, thereby, real output. Our findings suggest that tax policy should be part of competition policy and that policymakers must work towards coordinated tax policies to ensure a level playing field for all firms.

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A Appendix

A.1 Theory

A.1.1 Micro-founding tax avoidance

Our reduced-form modeling of tax planning is consistent with perfectly-legal loopholes or even fiscal incentives designed to encourage investment. It also aligns with various profit-shifting techniques across affiliates in different jurisdictions. While the exact channel through which firms reduce their tax burden is not relevant per se, it should be noted that our instrument strategy captures practices whose legality is at least debatable, e.g., profit-shifting techniques through transfer pricing. Generally, a tax-avoiding firm may manipulate the value of intra-firm transactions (transfer pricing) to shift their tax base to low-tax jurisdictions.[23]

We detail three cases below. A profit-shifting firm may i) inflate costs by importing from an affiliate located in a low or zero tax jurisdiction a good or service beyond its "arm's length" ii) borrow from an affiliate in a tax haven, and deduct interest payments in the non-haven country while declaring them in a tax haven where they are not taxed iii) locate their intangible assets (e.g., intellectual property) in tax havens.

Transfer pricing. A common practice that firms adopt to shift profits to a low-tax jurisdiction is to inflate the costs of inputs (p^I) sourced from their affiliates in tax havens. Without loss of generality, we can assume taxes are almost nil in tax havens; that is, $t^H \approx 0$, and inputs are produced at almost no cost. Firm profits then read:

$$\pi_i \;=\; (1-t^s)\left(p_i - p^I - \varphi_i^{-1}\right) q_i(p_i) + p^I q_i(p_i)(1-t^H)\ . \tag{12}$$

Simplifying, we get

$$\pi_i \;=\; \left(p_i - \varphi_i^{-1}\right) q_i(p_i) - t^s\left(p_i - \varphi_i^{-1}\theta_i\right) q_i(p_i)\ , \tag{13}$$

where $\theta_i = 1 + \varphi_i p^I$.

If the profit-maximizing transfer price results from a trade-off between a lower effective tax rate and a concealment cost, θ can be endogeneized as in Davies et al. (2018). When large firms benefit from scale economies in their tax planning, $\theta(\varphi)$ increases with firm productivity, so large firms deviate more from the arm's-length price than small firms. A similar argument can be made if firms instead manipulate their export prices to foreign affiliates in tax havens downward.

Debt shifting In the spirit of Mintz and Smart (2004), capital is the sole factor of production. Technology is linear, and capital productivity is equal to φ. One affiliate in a tax haven may lend for free to the parent firm in a non-haven. We assume no outside debt can be issued so that the deduction in the non-haven corresponds exactly to the interest payments declared in the tax haven. We denote $b_i > 0$, the per-unit interest payment of capital in

[23] Note that low and high tax jurisdictions may coexist in one country: the discussion below does not imply that the bulk of tax avoidance is international rather than domestic.

the non-haven country. Overall, the firm uses q_i/φ_i units of capital so that total interest payments amount to $b_i\varphi_i^{-1}\varphi_i^{-1}q_i$. Firms' profits are then given by:

$$\pi_i = (1 - t^s)\left(p_i - \varphi_i^{-1}\right)q_i + t^s b_i\varphi_i^{-1}q_i - t^H b_i\varphi_i^{-1}q_i \ , \tag{14}$$

where t^H is the corporate tax rate in the tax haven. This leads back to our baseline equation with $\theta_i = \frac{t^s - t^H}{t^s}b_i$. In the absence of taxation in the tax haven, that is, $t^H = 0$, θ_i comes down to the per-unit interest payment b_i. As in Mintz and Smart (2004), b_i may be endogeneized, assuming borrowing is costly. In that case, firm size and profit shifting would be co-determined by φ_i.

Intangibles as an investment. Firms can invest in some intangible f to decrease their marginal cost of production, now written $\varphi^{-1}c(f)$, where $c_f < 0$, in the non-haven country. The tax-deducted share of this investment f is denoted γ. Absent profit-shifting motives, the investment is denoted f_0. The firm's profits are thus given by

$$\pi_i = \left(p_i - \varphi^{-1}c(f_0)\right)q_i(p_i) - f_0 - t^s\left(p_i - \varphi^{-1}c(f_0)\right)q_i(p_i) + t\gamma f_0 \ .$$

.

Now, assume the cost f is borne in a tax haven in the form of the production of an intangible that is not taxed. Assume, moreover, that this intangible may be imported at an inflated cost $\delta f > f$. The above equation becomes:

$$\pi_i = \left(p_i - \varphi^{-1}c(f_0)\right)q_i(p_i) - f - t^s\left(p_i - \varphi^{-1}c(f)\right)q_i(p_i) + t^s\delta\gamma f \ .$$

Taking the first-order conditions with respect to f_0 and f, respectively, the optimal investments f and f^* with and without tax avoidance are related by

$$\frac{c_f(f^*)}{c_f(f_0^*)} = \frac{1 - t^s\gamma\delta}{1 - t^s\gamma} < 1 \ .$$

It follows that firms engaged in profit shifting invest more $f^* > f_0^*$ with f^* so that their equilibrium productivity is higher, consistent with the competitive edge put forward in our baseline model.

A.1.2 Tax avoidance and concentration

The Herfindahl index defines concentration in the economy. Formally, if we denote the HHI by $\mathcal{H}$, we have

$$\mathcal{H} = \frac{\sum_{i\leq N} s_i^2}{\left(\sum_{i\leq N} s_i\right)^2}$$

where N is the overall number of firms that we omit in the expressions below for clarity.

Observing that $\mathcal{H} = 1 - \frac{2\sum_{j\neq k\leq N} s_k s_k}{\left(\sum_j s_j\right)^2}$, differentiating the above expression w.r.t. s_k, that $\mathcal{H}$ increases with the sales of firm i means

$$-\left(\sum_{j\neq i;j\leq N} s_j\right)\left(\sum_{j\leq N} s_j\right) + 2\left(\sum_{j\neq k\leq N} s_j s_k\right) > 0$$

which can be rearranged as

$$-\left(\sum_{j\neq i} s_j\right)^2 + 2\left(\sum_{j\neq k\neq i\leq N} s_j s_k\right) + s_i\left(\sum_{j\neq i\leq N} s_j\right) > 0$$

Introducing the Herfindahl index $\mathcal{H}_{-i} = \frac{\sum_{j\neq i\leq N} s_j^2}{\left(\sum_{j\neq i\leq N} s_j\right)^2}$ in the absence of firm i, we get:

$$-\mathcal{H}_{-i}\left(\sum_{j\neq i\leq N} s_j\right)^2 + s_i\left(\sum_{j\neq i\leq N} s_j\right) > 0$$

$$\frac{s_i}{\sum_{j\leq N} s_j} > \frac{\mathcal{H}_{-i}}{1+\mathcal{H}_{-i}}$$

where $\frac{s_i}{\sum_{j\leq N} s_j} = \mathcal{S}_i$ is the market share of firm i. The Herfindahl index increases when tax avoidance increases a large firm's market share, such that the above inequality is satisfied. It is straightforward to see that this condition is always verified for the largest firm and never for the smallest one.

A.1.3 The impact of tax avoidance on real production

Real production as a function of productivities and market shares. Real production is given by the CES aggregate

$$\frac{Y}{\mathcal{P}} = \left(\sum_{i=1}^{N} q_i^{\frac{\sigma-1}{\sigma}}\right)^{\frac{\sigma}{\sigma-1}} . \tag{15}$$

In the presence of additional non-tax distortions, market shares are given by $\mathcal{S}_i = \left(\frac{p_i}{\mathcal{P}}\right)^{1-\sigma}$ and $q_i = \widetilde{\varphi}_i k_i$, we obtain that the allocation of the factor of production between any pair of firms i and j is

$$\frac{k_i}{k_j} = \frac{\mathcal{S}_i^{\frac{\sigma}{\sigma-1}}\widetilde{\varphi}_i^{-1}}{\mathcal{S}_j^{\frac{\sigma}{\sigma-1}}\widetilde{\varphi}_j^{-1}}$$

and thus

$$q_i = \frac{\mathcal{S}_i^{\frac{\sigma}{\sigma-1}}}{\sum_{j\leq N}\mathcal{S}_j^{\frac{\sigma}{1-\sigma}}\widetilde{\varphi}_j^{-1}} K.$$

Plugging the above expression of output into (15) leads to the formula given in the main text:

$$\frac{Y}{\mathcal{P}}\Big/\frac{Y'}{\mathcal{P}'} = \frac{\sum_{i=1}^{i=n}\mathcal{S}'^{\frac{\sigma}{\sigma-1}}_i\widetilde{\varphi}_i^{-1}}{\sum_{i=1}^{i=n}\mathcal{S}_i^{\frac{\sigma}{\sigma-1}}\widetilde{\varphi}_i^{-1}}.$$

Bounds on real output Consider the special case of the real-production expression above for two firms with productivities φ_i and φ_j and market shares $\mathcal{S}$ and $1-\mathcal{S}$, respectively, we obtain that the change in real output from $\frac{Y}{\mathcal{P}}$ to $\frac{Y'}{\mathcal{P}'}$ induced by any shock that leaves unchanged factor productivities and non-tax distortions are given by

$$\widehat{\left(\frac{Y}{P}\right)} = \frac{Y}{\mathcal{P}} \Big/ \frac{Y'}{\mathcal{P}'} = \frac{\mathcal{S}'^{\frac{\sigma}{\sigma-1}} + \left(\widetilde{\varphi}_i/\widetilde{\varphi}_j\right)\left(1-\mathcal{S}'\right)^{\frac{\sigma}{\sigma-1}}}{\mathcal{S}^{\frac{\sigma}{\sigma-1}} + \left(\widetilde{\varphi}_i/\widetilde{\varphi}_j\right)\left(1-\mathcal{S}\right)^{\frac{\sigma}{\sigma-1}}}.$$

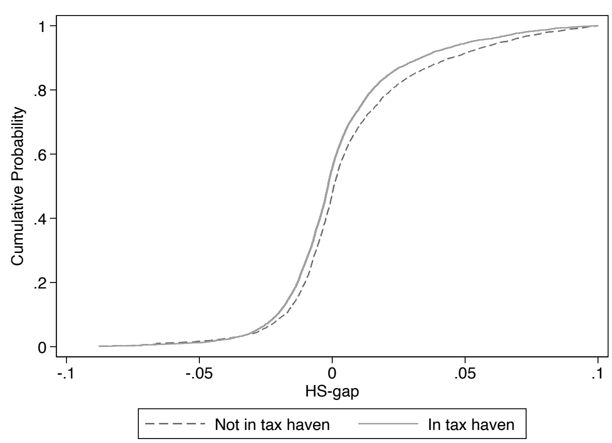
Evaluating this expression for the largest possible distortions, i.e., $\widetilde{\varphi}_i/\widetilde{\varphi}_j = 0$ and $\widetilde{\varphi}_i/\widetilde{\varphi}_j \to \infty$ pins down the bounds on real output changes.

A.2 Facts on presence in tax havens and audit probability

Presence in tax havens. We use the dataset provided by Dyreng and Lindsey (2009) to identify U.S. multinational presence in tax-haven countries. The information is available for a subsample of 7,148 firms spanning 1994-2014. For the tax-haven list, we follow the definition of Hines and Rice (1994) and add the Netherlands. See Souillard (2020a) for a discussion of the data.

Figure A.1 shows the cumulative distribution of the HS-gap measure for the group of MNEs present in tax havens and the group of MNEs that are not. The higher of the two lines in the plot is the cumulative distribution function (cdf) of the HS gap of multinationals that own affiliates in tax-haven countries. The lower line is the same for multinationals with no affiliates in tax havens. A higher cdf is consistent with lower $HS - gap$ for multinationals that have affiliates in tax havens. This evidence supports the idea that our tax avoidance measure captures the firm's aggressive tax planning strategies.[24]

Figure A.1: Cumulative distribution of the HS-gap measure across groups of MNEs

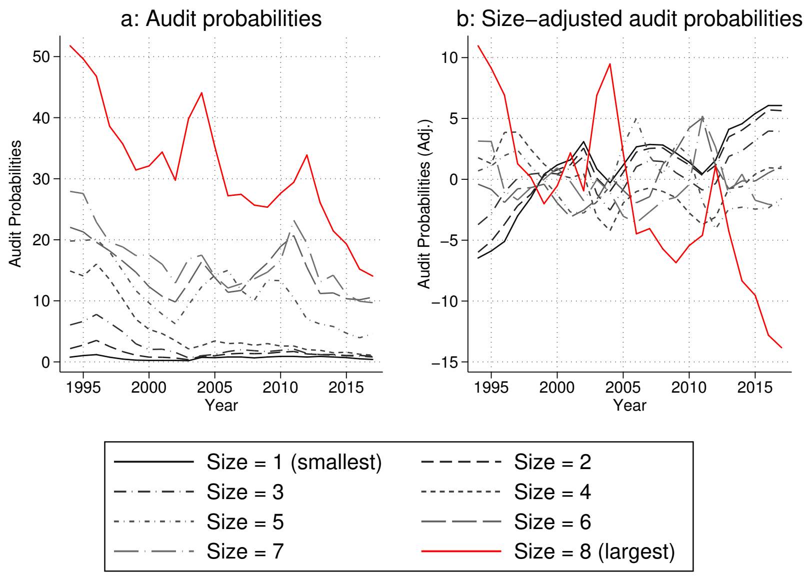


We use the Kolmogorov – Smirnov (KS) test to determine if any differences exist in the distribution of HS gaps for the group of MNEs present in tax havens and the group of MNEs that are not. The KS-test statistic is computed as the largest vertical distance (D) between the two cdfs. We find a maximal distance of 0.1017. This difference is computed to a null distribution to obtain the p-value for the test, which is 0.000. It indicates overwhelming evidence of a difference between the two distributions.

[24]We use the Kolmogorov – Smirnov (KS) test to determine any differences in the distribution of HS gaps for the two groups. The KS-test statistic is computed as the largest vertical distance (D) between the two cdfs. We find a maximal distance of 0.1017. This difference is computed to a null distribution to obtain the p-value for the test, which is 0.000. It indicates overwhelming evidence of a difference between the two distributions.

Audit probabilities. The IRS annual Data-Books disclose data to compute the average audit probability for each of the eight asset classes across our sample period. We compute the audit probability as the number of corporate tax return audits completed in the IRS's fiscal year t for an IRS asset-size group divided by the number of corporate tax returns received in the previous calendar year for the same IRS asset-size group. The size-adjusted audit probabilities are the residuals from a regression of raw probabilities on year and asset-size group dummy variables. The residuals provide information on the deviations from the predicted average audit probability within each year.

Figure A.2: Size-adjusted audit probabilities by asset class



The left panel of Figure A.2 shows the evolution of the audit probabilities as computed using the raw IRS disclosed data. It shows that the probability of an IRS audit has dropped for larger firms in the U.S. but remained relatively constant for the smallest firms. The drop in the audit probability for the smallest firms reporting in the first IRS asset class between 1994 and 2017 is 0.4%, whereas it is over 37% for the largest firms. In panel (b), we display the evolution of the adjusted audit probabilities. It shows a relative decline in the size-adjusted audit probabilities in the eighth class of assets from 1994 to 2017. The value of assets reported by firms in this class amount to $250,000,000 or more and corresponds to the largest firms. However, the decline in the adjusted probabilities is not linear across years, with large variations in the early 2000 and 2010.

Table A.1: Sales and tax avoidance – OLS and 2SLS estimates

Dep. Variable	Log Sales - End of Period					
	OLS		2SLS			
			1^{st} Stage	2^{nd} Stage	1^{st} Stage	2^{nd} Stage
HS tax gap	-2.648***	-1.038***		-5.408***		-4.085***
	(0.125)	(0.118)		(0.541)		(1.439)
Share of Intangible	1.167***	1.144***	-0.029***	1.041***	-0.071***	0.910***
	(0.082)	(0.076)	(0.011)	(0.090)	(0.019)	(0.151)
Labor Prod.	0.524***	0.484***	-0.054***	0.368***	-0.046***	0.345***
	(0.017)	(0.032)	(0.003)	(0.032)	(0.005)	(0.073)
Acquisition	1.243***	0.253***	-0.038***	1.130***	-0.006*	0.236***
	(0.029)	(0.017)	(0.004)	(0.035)	(0.003)	(0.020)
MNE Status	1.478***	0.317***	-0.071***	1.279***	-0.015***	0.274***
	(0.028)	(0.025)	(0.004)	(0.044)	(0.004)	(0.032)
Adj. Audit Proba,			0.011***		0.002***	
			(0.001)		(0.001)	
Sector × Period FE	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	No	Yes	No	No	Yes	Yes
Obs.	22,271	18,546	22,271	22,271	18,546	18,546
Adj. R^2	0.527	0.930	0.140		0.610	
KP F-stat.				120.1		9.375

Sample years: 1994-2017. The dependent variable is the firm's log sales at the end of the four-year window. OLS and 2SLS estimates with robust standard errors in parentheses. First-stage Kleibergen-Paap Wald F statistic reported. ***, **, and * significantly different from 0 at the 1%, 5%, and 10% confidence levels, respectively.

Table A.2: Sales and tax avoidance – BTD-based measures

Dep. Variable	Log Sales - End of Period											
	Permanent book-tax difference (PBTD)				Discretionary Permanent Differences (DTAX)				Tax sheltering residuals (TS_r)			
	1^{st} St.	2^{nd} St.	1^{st} St.	2^{nd} St.	1^{st} St.	2^{nd} St.	1^{st} St.	2^{nd} St.	1^{st} St.	2^{nd} St.	1^{st} St.	2^{nd} St.
Tax Aggressiveness		0.559***		0.333***		0.554***		0.343***		1.467***		0.388**
		(0.062)		(0.093)		(0.062)		(0.097)		(0.217)		(0.110)
Share of Intangible	0.027	1.030***	0.623**	0.983***	0.020	1.046***	0.660**	0.972***	0.094	0.912***	0.444*	1.019**
	(0.168)	(0.121)	(0.264)	(0.142)	(0.171)	(0.123)	(0.269)	(0.147)	(0.097)	(0.173)	(0.235)	(0.141)
Labor Prod.	0.488***	0.344***	0.221***	0.481***	0.492***	0.344***	0.216**	0.457***	0.173***	0.339***	0.220***	0.469**
	(0.036)	(0.032)	(0.084)	(0.050)	(0.037)	(0.032)	(0.087)	(0.052)	(0.019)	(0.041)	(0.066)	(0.049)
Acquisition	0.384***	1.094***	0.044	0.242***	0.394***	1.090***	0.039	0.234***	0.131***	1.110***	0.049	0.237**
	(0.046)	(0.043)	(0.041)	(0.023)	(0.046)	(0.043)	(0.040)	(0.023)	(0.028)	(0.054)	(0.037)	(0.024)
MNE Status	0.693***	1.178***	0.189***	0.225***	0.700***	1.174***	0.189***	0.223***	0.204***	1.243***	0.141***	0.233**
	(0.050)	(0.049)	(0.056)	(0.036)	(0.051)	(0.050)	(0.057)	(0.037)	(0.029)	(0.057)	(0.052)	(0.036)
Adj. Audit Proba,	-0.165***		-0.039***		-0.167***		-0.039***		-0.061***		-0.034***	
	(0.018)		(0.008)		(0.018)		(0.009)		(0.009)		(0.007)	
Sector × Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	No	No	Yes	Yes	No	No	Yes	Yes	No	No	Yes	Yes
Obs.	15,137	15,137	11,977	11,977	14,918	14,918	11,784	11,784	14,569	14,569	11,988	11,988
Adj. R^2	0.139		0.601		0.137		0.601		0.290		0.442	
KP F-stat.		85.64		21.65		83.38		19.86		46.15		20.73

Sample years: 1994-2017 (24 years). The dependent variable is the firm's log sales at the end of the four-year windows. 2SLS estimates with robust standard errors in parentheses. First stage Kleibergen-Paap Wald F statistic reported. ***, **, and * significantly different from 0 at the 1%, 5%, and 10% confidence levels respectively.