

Supplemental Appendix

Market Externalities of Tax Evasion

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A. Data Appendix

A1. Data Construction

Our data cover the universe of Italian firms operating over the 2005–2019 period. The dataset is the outcome of a collaboration between the Bank of Italy and the Italian National Statistical Agency and combines statistical, administrative, and fiscal sources at the firm level (see Abbate, Ladu and Linarello 2017).

The construction of the dataset starts from the definition of the universe of active firms using the Business Register (ASIA), which includes all enterprises active for at least six months. ASIA provides firm identifiers, legal form, location, industry classification (NACE), employment, and revenues.

Economic variables are recovered by integrating multiple sources according to a hierarchical order. Annual balance sheet data for incorporated firms, collected by the Chambers of Commerce and processed by CERVED, are used whenever available. The second main data source is the Sectoral Studies (Fiscal Authority Survey). All firms are required to report taxable income through tax returns, and, depending on their legal form and accounting regime, to complete specific survey questionnaires. These surveys aim to assess firms' income-generating capacity and tax compliance. In practice, almost all firms must submit them alongside their tax returns, providing detailed information on revenues and costs. For manufacturing and services, the common questionnaire (Sectoral Study F) closely resembles a balance sheet and allows us to recover turnover, value added, and labor costs. Information is more limited for professional activities (Sectoral Study G), but these data remain a valuable source for a large number of firms. Finally, starting from 2011, income tax returns (Modello Unico) are used to obtain economic information for firms not covered by the previous sources, which is often the case for micro firms operating under special tax regimes such as the *Regime dei Minimi*.

Although the dataset combines information drawn from different sources, Abbate, Ladu and Linarello (2017) construct variables using harmonized definitions and consistency checks to ensure comparability across firms; differences across sources reflect differences in data availability rather than differences in data quality.

For the period 2005–2011, value added is not directly observed for a subset of firms due to incomplete fiscal information. Following Abbate, Ladu and Linarello (2017), missing values are imputed using the median value added per worker computed within cells defined by industry, size class, location, and legal form. The quantitative importance of the imputation is limited: the share of imputed value added reaches at most approxi-

mately 11 percent of firms in the late 2000s and is lower in other years. Abbate, Ladu and Linarello (2017) further show that aggregate value added dynamics computed with and without imputation track each other closely, both at the economy-wide and sectoral level, suggesting that the imputation procedure does not materially distort measured outcomes.

A2. Characteristics of Eligible Firms

Eligible firms are sole proprietorships that satisfy the revenue cutoff in force in each year and, when applicable, the age restriction specified by the preferential tax regime. Table A1 reports summary statistics for eligible and non-eligible individual firms.

Table A1: Characteristics of Individual Firms by Eligibility Status

	Non-Eligible		Eligible	
	Mean (1)	SD (2)	Mean (3)	SD (4)
Revenue	114,457.2	318,100.2	55,866.96	130,278.4
Value-added	43,279.25	89,649.58	25,609.3	46,576.4
Value-added per-worker	23,003.21	41,059.74	23,554.41	44,130.21
N of workers	1.786	1.835	1.121	0.497
Age	16.119	10.988	13.515	10.945
Entry rate	0.038	0.192	0.018	0.133
Exit rate	0.086	0.280	0.075	0.264
Number of observations	24,030,451		17,181,093	

Note: This table reports summary statistics for individual firms, distinguishing between those eligible and not eligible for the preferential tax regime. The sample includes sole proprietorships over the 2005–2019 period.

Eligible firms operate at a smaller scale in terms of revenue and value added and employ fewer workers on average, consistent with the low revenue thresholds defining eligibility. However, value added per worker is very similar across eligible and non-eligible firms, indicating that eligible firms are smaller in scale but not characterized by systematically lower productivity at the worker level. In addition, eligible firms are not disproportionately short-lived: their average age and exit rates are comparable to those of non-eligible individual firms, suggesting that they do not primarily consist of temporary or transitory activities. Over the period considered, eligible firms account for approximately 42 percent of all individual firms, a substantial segment of the firm population.

Table A2 reports the sectoral distribution of firms by legal form and eligibility status. Eligible firms are predominantly active in wholesale and retail trade (27.9 percent), professional, scientific and technical activities (24.5 percent), construction (11.9 percent), and human health and social work (9.1 percent). These sectors account for the majority of eligible firms and also represent substantial shares of non-eligible and non-individual firms. The presence of eligible firms in the same broad sectors as other firms indicates that they operate within economically relevant markets characterized by potential com-

Table A2: Sector Composition by Firm Type and Eligibility Status

	Non-Individual (%)	Non-Eligible Indiv. (%)	Eligible Indiv. (%)
Mining and quarrying	0.12	0.02	0.01
Manufacturing	13.90	7.19	5.32
Electricity, gas, steam and air conditioning	0.50	0.03	0.05
Water supply and waste management	0.45	0.08	0.04
Construction	12.62	12.51	11.94
Wholesale and retail trade	23.89	26.28	27.88
Transport and storage	3.13	2.83	2.74
Accommodation and food services	9.74	6.74	3.84
Information and communication	3.48	1.25	1.88
Financial and insurance activities	1.61	3.53	0.95
Real estate activities	12.37	1.16	1.38
Professional, scientific and technical	7.01	19.65	24.45
Administrative and support services	4.16	2.61	2.89
Education	0.77	0.50	0.67
Human health and social work	1.88	8.38	9.12
Arts, entertainment and recreation	1.52	1.33	1.61
Other services	2.86	5.93	5.25

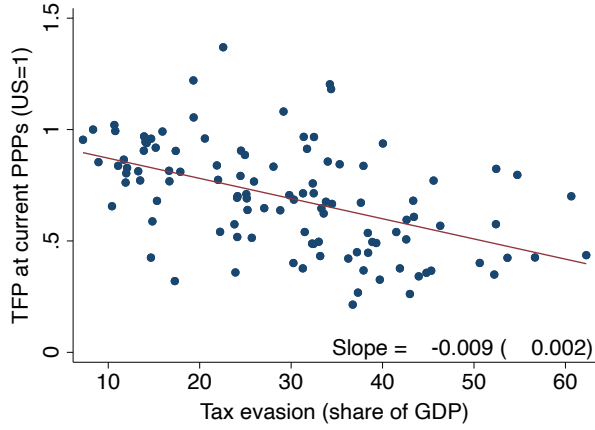
Note: This table reports the distribution of firms across broad industry groups, distinguishing between non-individual firms, non-eligible individual firms, and eligible individual firms. Shares are expressed in percentages.

petitive interactions.

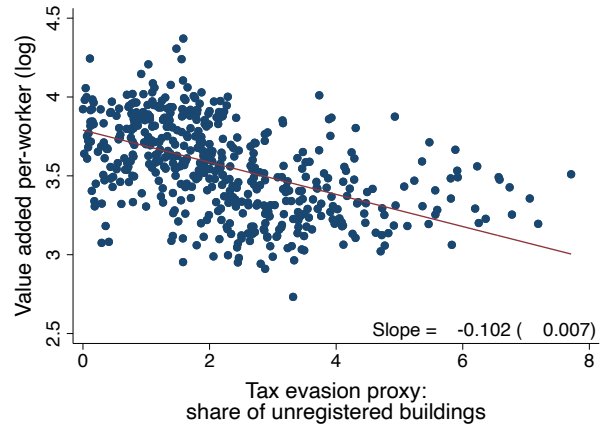
B. Additional Tables and Figures

Figure B1: Self-Employment Share, Tax Evasion, and Productivity

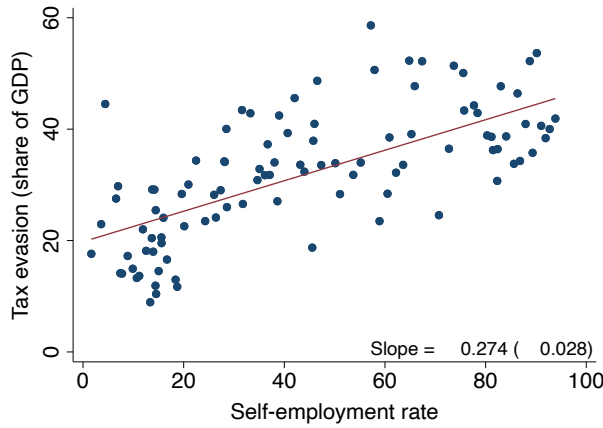
A. Productivity and tax evasion:
cross-country correlation



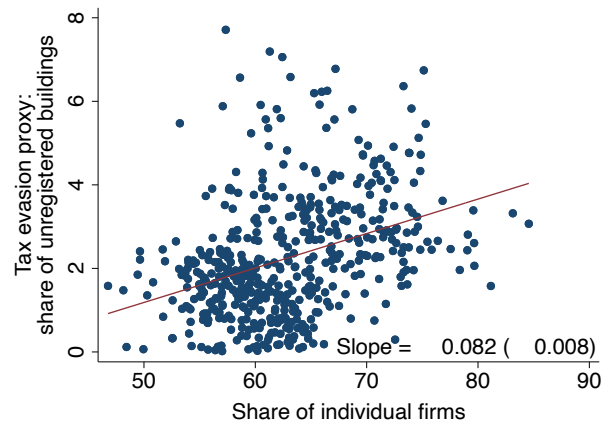
B. Productivity and tax evasion:
cross-Italian LLM correlation



C. Tax evasion and self-employment:
cross-country correlation



D. Tax evasion and self-employment:
cross-Italian LLM correlation



Notes: The top panel presents the relationship between productivity and tax evasion. The left-hand side graph provides a cross-country comparison between total factor productivity at current power purchasing prices (from the Penn World Table) and a proxy for the size of the shadow economy (from Medina and Schneider 2018). The right-hand side graph plots the relationship between the log of the value added per worker (from the Italian Institute of Statistics) and the share of unregistered buildings (from the Italian Internal Revenue Agency) across Italian local labor markets. The bottom panel depicts the association between tax evasion and the share of self-employed taxpayers. The left-hand side graph provides a cross-country comparison between a proxy for the size of the shadow economy (from Medina and Schneider 2018) and the self-employment rate (World Bank data). The right-hand side graph plots the relationship between the share of unregistered buildings (from the Italian Internal Revenue Agency) and the self-employment rate across Italian local labor markets.

Table B1: Preferential Tax Regimes for Individual Firms

Regime	Period	Revenue cutoff	Age cutoff	Tax rate (%)	Tax rate by age?	Tax base by industry?	Cutoff by industry?	VAT duties?
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
<i>Agevolato:</i> Law 388/2000	2000-2011	30,000	3 years	10%	No	Yes	No	No
<i>Minimi:</i> Law 244/2007	2008-2011	30,000	No	20%	No	Yes	No	No
<i>Minimi agevolato:</i> Law 98/2011	2012-2014	30,000	5 years	5%	No	Yes	No	No
<i>Forfettario I:</i> Law 140/2014	2015-2018	15,000 -40,000	No	5%-15%	Yes	Yes	Yes	No
<i>Forfettario II:</i> Law 145/2018	2019-	65,000	No	5%-15%	Yes	Yes	No	No

Note: This table presents the main features of the preferential tax scheme for individual firms in Italy. For each preferential tax regime for individual firms, the table provides the following information: *column 1:* the regime's denomination and the relevant law; *column 2:* the period when the regime applies; *column 3:* the revenue cutoff (in €) determining eligibility (with minimum and maximum values in case of different cutoffs by industries); *column 4:* the firm's age cutoff determining a change in the preferential tax rate; *column 5:* the preferential tax rate (or the tax rate below and above the age cutoff in case the tax rate varies by age); *column 6:* whether the tax rate varies by firm's age; *column 7:* whether the tax base varies by industry; *column 8:* whether the revenue cutoff determining eligibility varies by industry; *column 9:* whether the regime grants exemption from value-added tax fulfillment.

Table B2: Industry-Level Information on Preferential Tax Regimes for Individual Firms

Industry	NACE-2007	Revenue cutoff (€)				Profitability coefficient (%)
		2005-2014 (1)	2015 (2)	2016-2018 (3)	2019 (4)	
Food & beverage	10-11	30,000	35,000	45,000	65,000	0.4
Wholesale and retail trade	45; 46.2-46.9; 47.1-47.7; 47.9	30,000	40,000	50,000	65,000	0.4
Street vendors of food and drinks	47.81	30,000	30,000	40,000	65,000	0.4
Street vendors of other products	47.82-47.89	30,000	20,000	30,000	65,000	0.54
Construction and real estates	41-43; 68	30,000	15,000	25,000	65,000	0.86
Trade intermediaries	46.1	30,000	15,000	25,000	65,000	0.62
Accommodation and catering	55-56	30,000	40,000	50,000	65,000	0.40
Professional activities	64-66; 69-74; 85; 86-88	30,000	15,000	30,000	65,000	0.78
Other industries		30,000	20,000	30,000	65,000	0.67

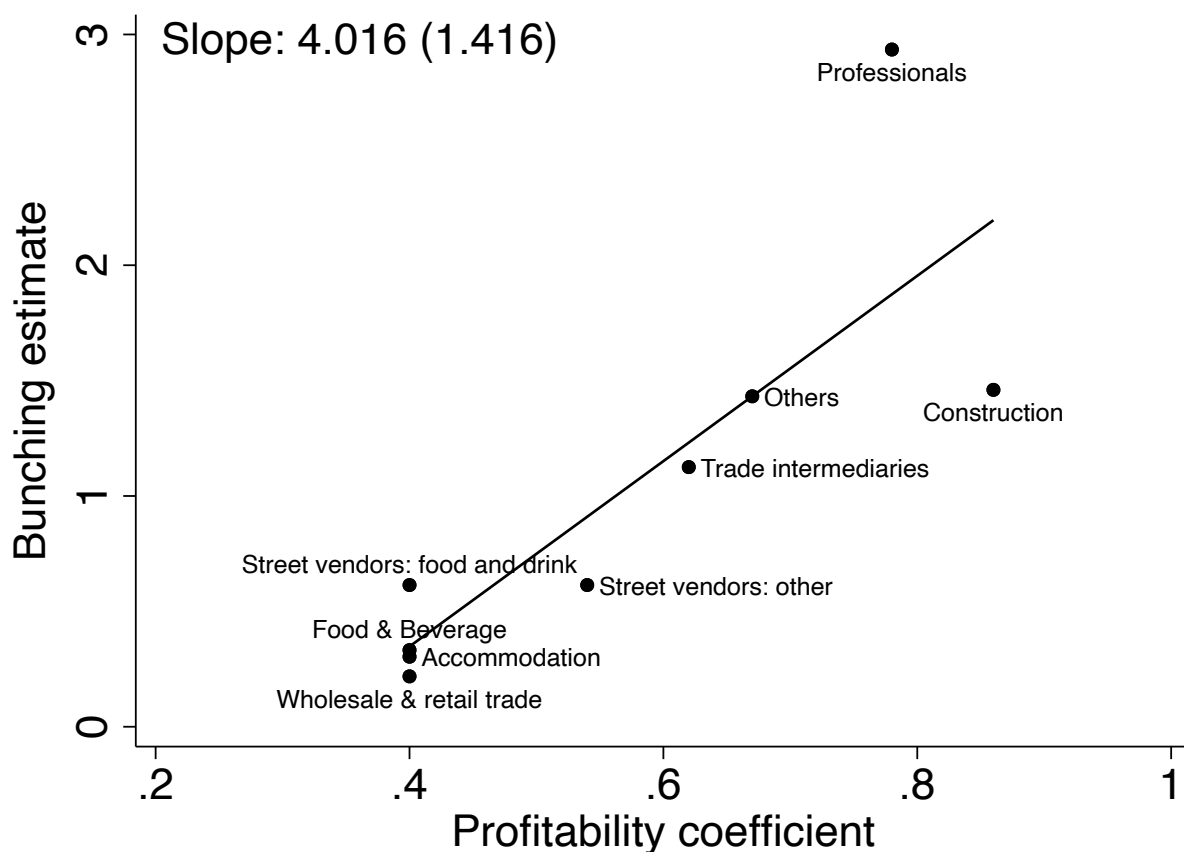
Note: For each industry and regime, the table presents information on the revenue cutoff determining eligibility for the preferential tax regime (columns 1-4), and the profitability coefficient, which is the portion of the tax base subject to the preferential tax rate (column 5).

Table B3: Personal Income Tax Schedule

Taxable income (€ per-year)	Marginal tax rate (%)
A. Period 2005-2006	
<4,500	0
4,501-26,000	23
26,001-33,500	33
33,501-100,000	39
> 100,000	45
B. Period 2007-2019	
< 4,500	0
4,501-15,000	23
15,001-28,000	27
28,001-55,000	38
55,001-75,000	41
> 75,000	43

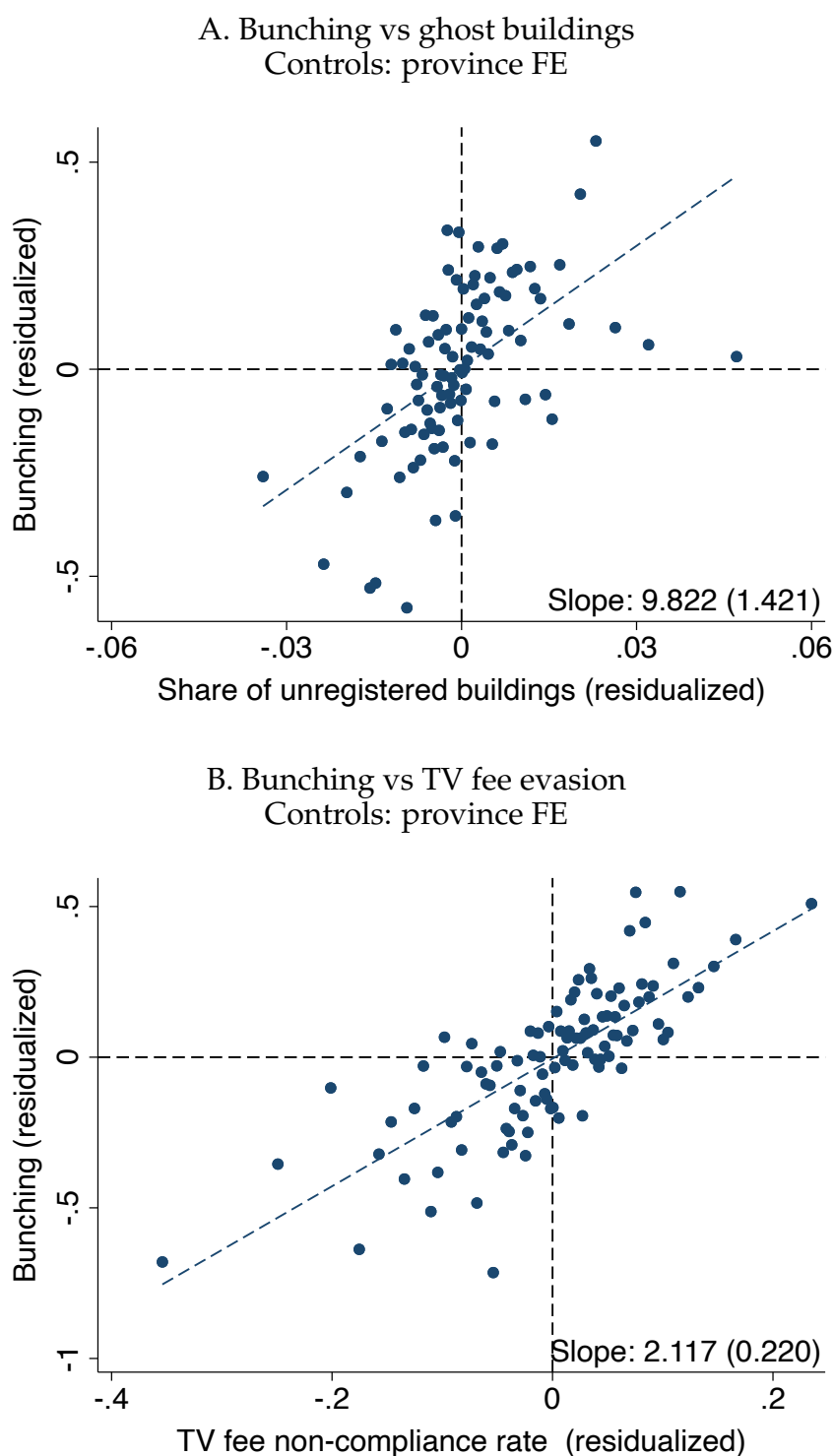
Note: This table displays information on the Italian personal income tax (IRPEF) over the 2005-2006 period (top panel) and 2007-2019 period (bottom panel). The tax base is defined as net of deductible expenses, such as social security and welfare contributions or donations to non-profit organizations.

Figure B2: Bunching Response and Profitability Coefficient



Notes: This figure relates the bunching estimate to the industry-specific profitability coefficient. Bunching estimates are calculated as described in Section 3.1. The profitability coefficient follows the classification of the Italian government (see Section 2.2 for details): it varies from a minimum of 40 percent for street vendors, the food and beverage industry, and the accommodation and catering industry, to a maximum of 86 percent for the construction and real estate industry. The graph also reports the estimated slope and the associated standard error.

Figure B3: Correlating the Bunching Rate with Other Tax Evasion Indicators - Further Specifications



Notes: Each panel compares the bunching rate (vertical axes) with the share of unregistered buildings (horizontal axis; right-hand side graph) and the national TV fee non-compliance rate (horizontal axis; left-hand side graph), controlling for province fixed effects. These variables are depicted in 100 equal-sized bins and show the line of best fit. Each graph also reports the estimated slope and the associated standard error.

C. Bunching Responses to Tax Notches

C1. Bunching Theory

To provide an example of how these notches affect firm behavior, let us consider the case of self-employed individuals' responses to the tax rate notch. Assume that each self-employed individual, indexed by i , sells her output at a price p , and makes two choices: tax evasion (e_i) and labor supply (l_i). Each self-employed individual reports taxable income (i.e., the individual firm's revenue) $\bar{y}_i = y_i - e_i$, where y_i is her true revenue ($y_i = pl_i$). The preferential tax regime provides a (flat) tax rate, τ^p , that is lower than the basic (regular) tax rate, τ^b , when a self-employed individual reports revenue below a certain cutoff K . Therefore, self-employees reporting $\bar{y}_i > K$ face the tax rate τ^b , while those reporting $\bar{y}_i \leq K$ are eligible for the preferential tax scheme and pay τ^p . The preferential tax scheme creates a notch in the budget constraint of the self-employed: a discontinuity in the choice set of taxable versus net income. This notch induces self-employed individuals, who would otherwise report more revenue, to instead bunch right at the tax notch.

Figure C1 provides a simple illustration of how self-employed individuals would respond to the preferential tax scheme notch. Panel A presents a budget set diagram; panel B the density distributions. Before the introduction of the preferential tax scheme, self-employees report taxable income, y , that maximizes their own utility subject to their budget constraint. Income is distributed according to a smooth density distribution $h(y)$ and any heterogeneity is due to preferences or idiosyncratic shocks. With the introduction of the preferential tax scheme, self-employees will face a tax notch at income level y^* . The notch generates a region of strictly dominated choice in the income interval $(y^*, y^* + \Delta y^D]$, where it is possible to increase both leisure and consumption (net income) by moving to the notch point y^* . At this income level, a self-employee can maximize net income by paying the preferential tax rate τ^p instead of the basic tax rate τ^b . All self-employees located in the income interval $(y^*, y^* + \Delta y^*]$, where the bunching region is larger than the area of strictly dominated choice, $\Delta y^* > \Delta y^D$, will respond to the preferential tax scheme notch by bunching.

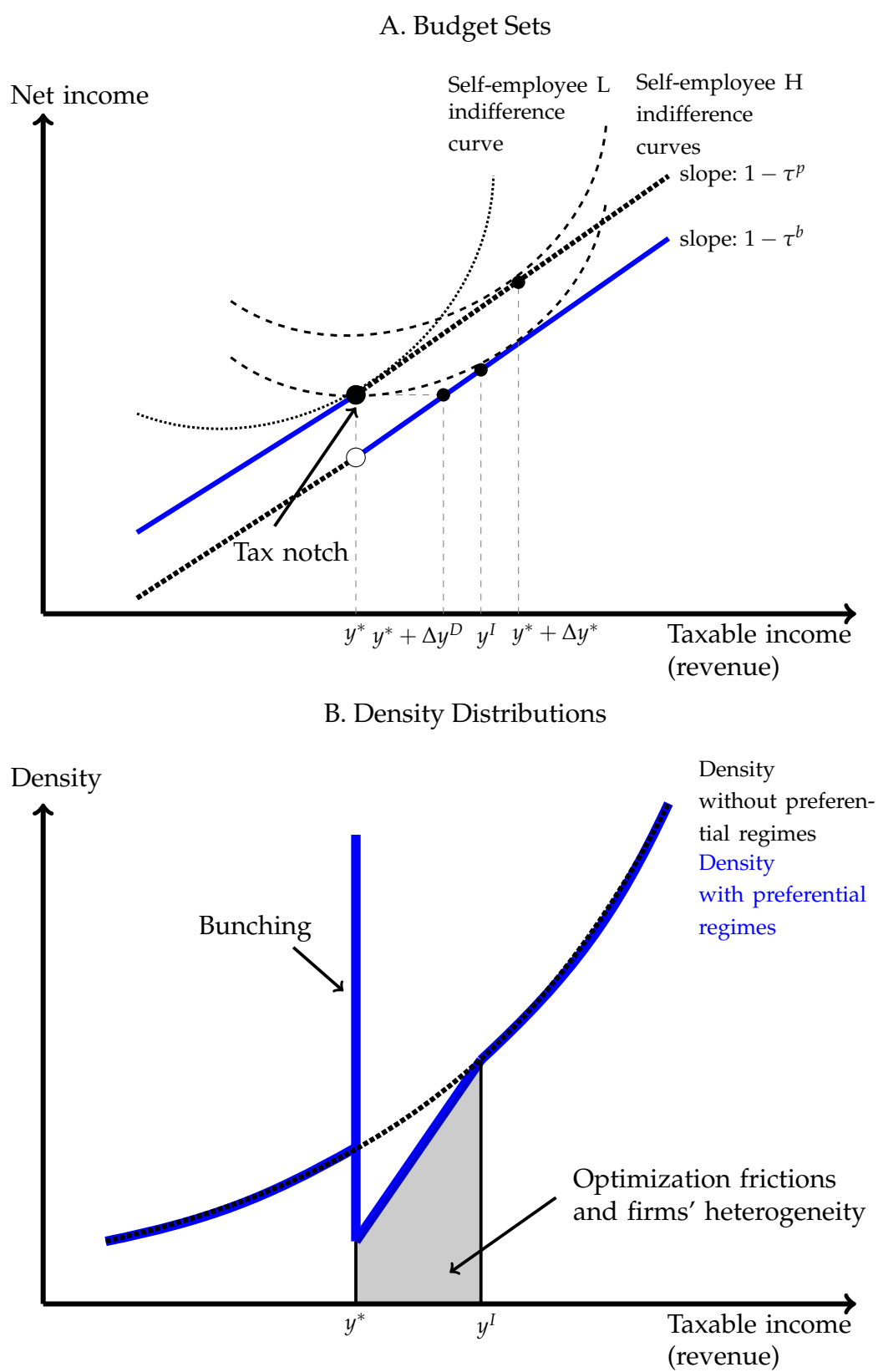
In this example, we illustrate responses from two types of "bunchers". We define self-employed L as the one with the lowest income before the introduction of the preferential tax scheme, y^* ; self-employed H as the one with the highest income, $y^* + \Delta y^*$. The preferential tax scheme leads self-employed L to continue to choose income y^* and benefit from an increase in net income. Self-employee H will also bunch at the tax notch because is exactly indifferent between the notch point y^* and the interior point y^I . Self-employed L and H represent the two extreme cases: each self-employed between L and H will bunch at the preferential tax scheme notch. Therefore, because no one is willing to locate between the tax notch y^* and the interior point y^I , this model would predict a density hole in the segment $(y^*, y^I]$ and excess bunching at the notch y^* .

In practice, the predictions of this benchmark model can be questioned due to optimization frictions, such as adjustment costs or inattention, and heterogeneity in the advantages of the preferential regime across firms. For instance, since revenue-based taxation is advantageous to taxpayers with high-profit margins, the standard regime might be convenient for self-employees with large (deductible) costs, such as those with many employees, significant investments, and high operating expenses. Changes in the location of the eligibility cutoff, which varies across industries and over time, and in the size of the preferential tax rate, also introduce heterogeneities in the size of the tax notch across industries, by firm's age, and over time. These heterogeneity and optimization frictions might prevent some self-employed individuals from bunching, creating a significant density mass in the (otherwise empty) strictly dominated region.

Assuming that the counterfactual density $h_0(y)$ is roughly constant on the bunching segment $(y^* + \Delta y^*)$, we can express excess bunching at the tax notch as a function of the counterfactual density and the marginal buncher:

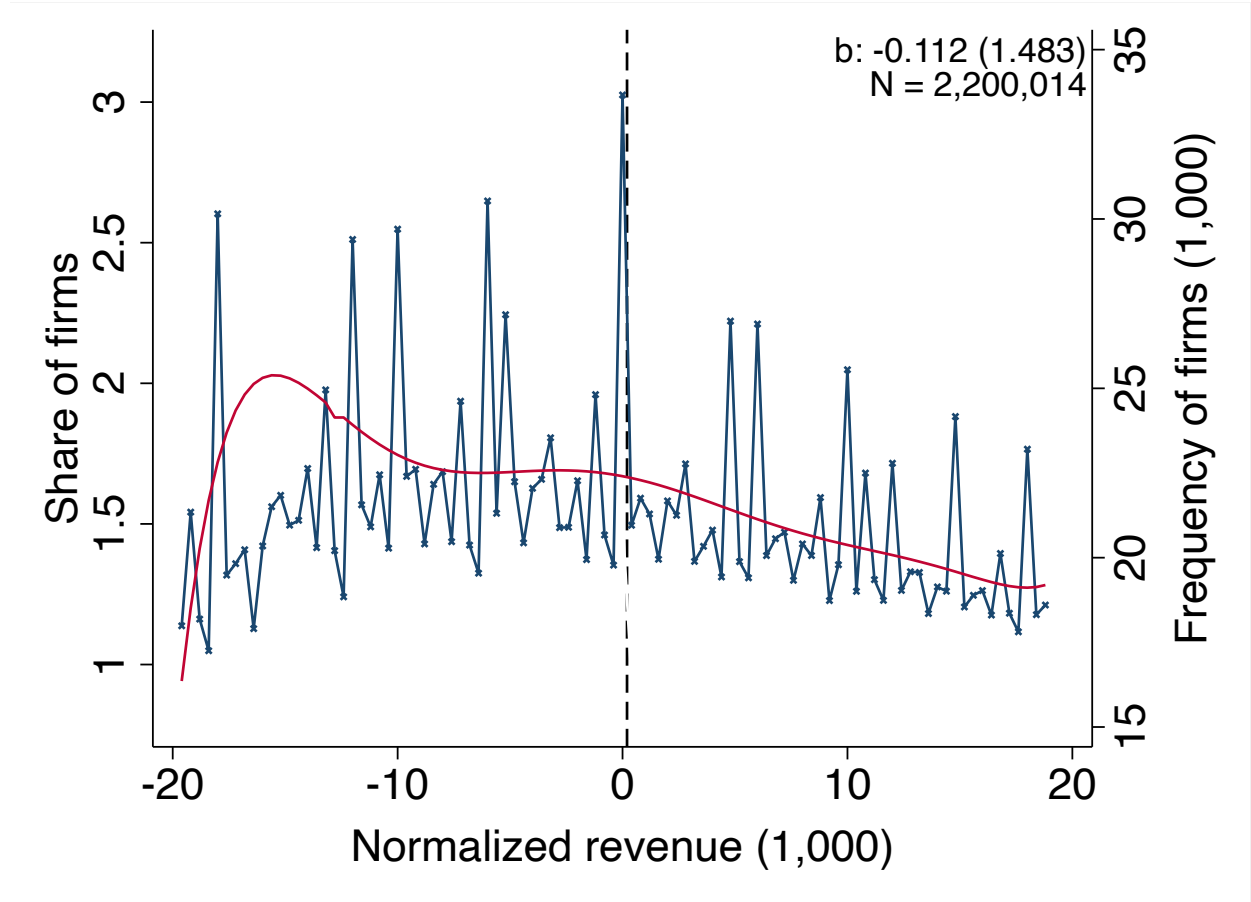
$$(C1) \quad B = \int_{y^*}^{y^* + \Delta y^*} h_0(y) dy \approx h_0(y^*) \Delta y^*.$$

Figure C1: Bunching Responses to the Tax Notch



C2. Revenue Distribution of Non-Individual Firms

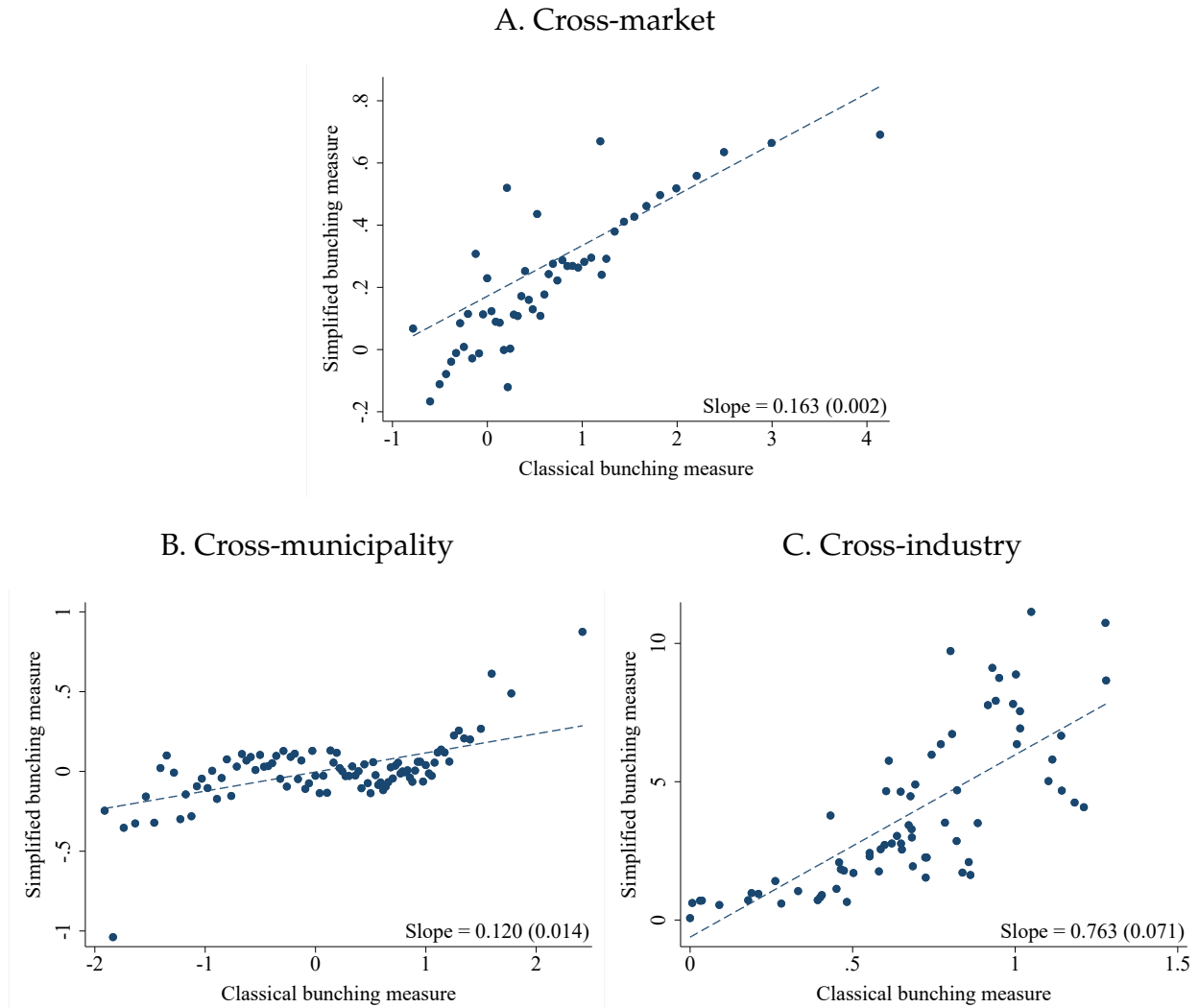
Figure C2: Revenue Distribution of Non-Individual Firms



Notes: The figure shows the empirical distribution of revenue (blue dots) and the counterfactual distribution (red curve). The sample is based on the pooled data of *non-individual firms* (i.e., potential competitors) for the period 2005-2019. The horizontal axes report normalized revenues: the distance from the revenue cutoff determining eligibility for preferential tax schemes. The bins are €400 wide. We also report estimates of the excess mass of firms as a share of average density in the bunching region and its associated bootstrapped standard error (b), and the number of observations.

C3. Correlating “Simplified” and “Classical” Bunching Measures

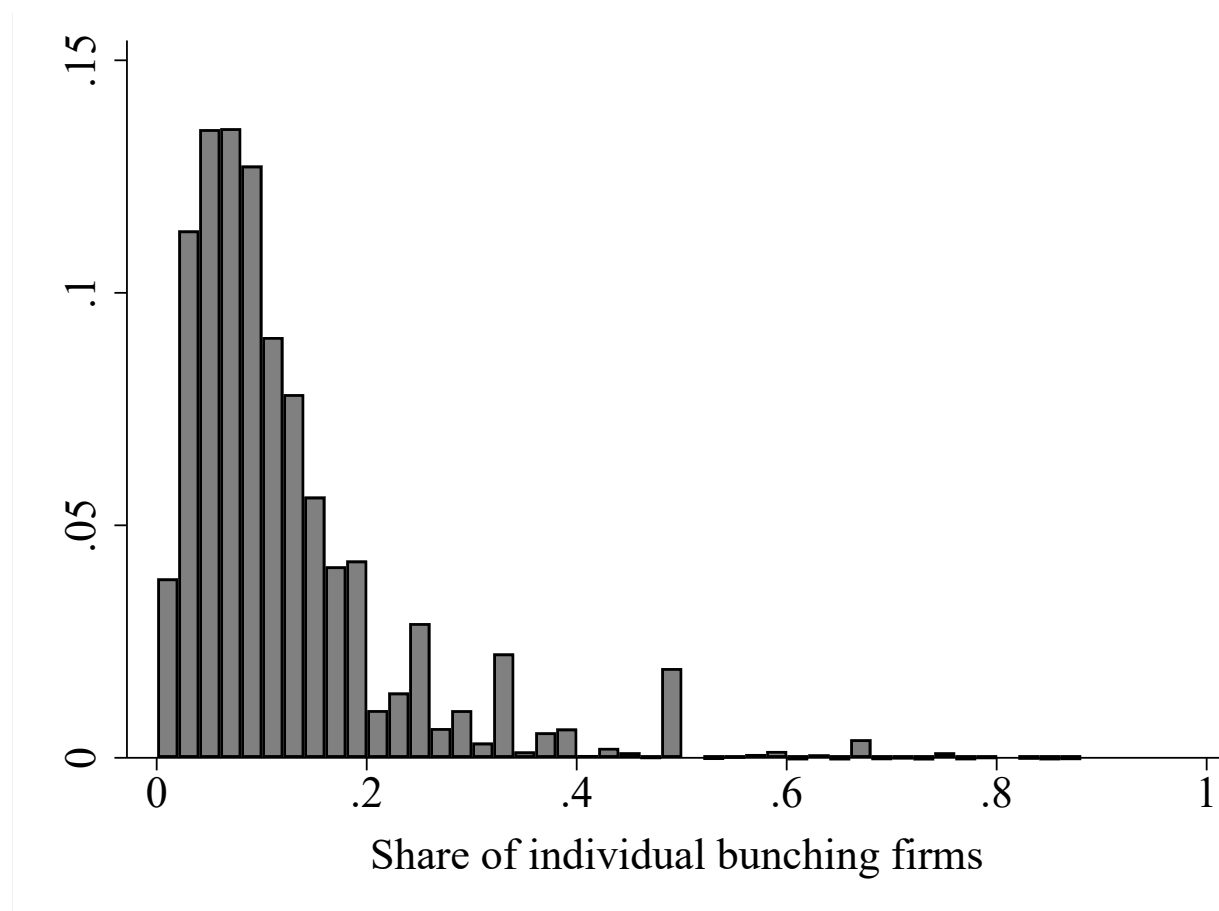
Figure C3: Correlating “Simplified” and “Classical” Bunching Measures



Notes: This figure compares the “simplified” bunching measure (vertical axes) with the “traditional” bunching measure (horizontal axes). We depict the cross-market correlation in panel A; the cross-municipality correlation in panel B; and the cross-industry correlation in panel C. Each graph depicts the two variables in 100 equal-sized bins and shows the line of best fit. The two measures are standardized to have mean 0 and standard deviation equal to 1. Each graph also reports the estimated slope and the associated standard error.

C4. The Distribution of the Bunching Rate

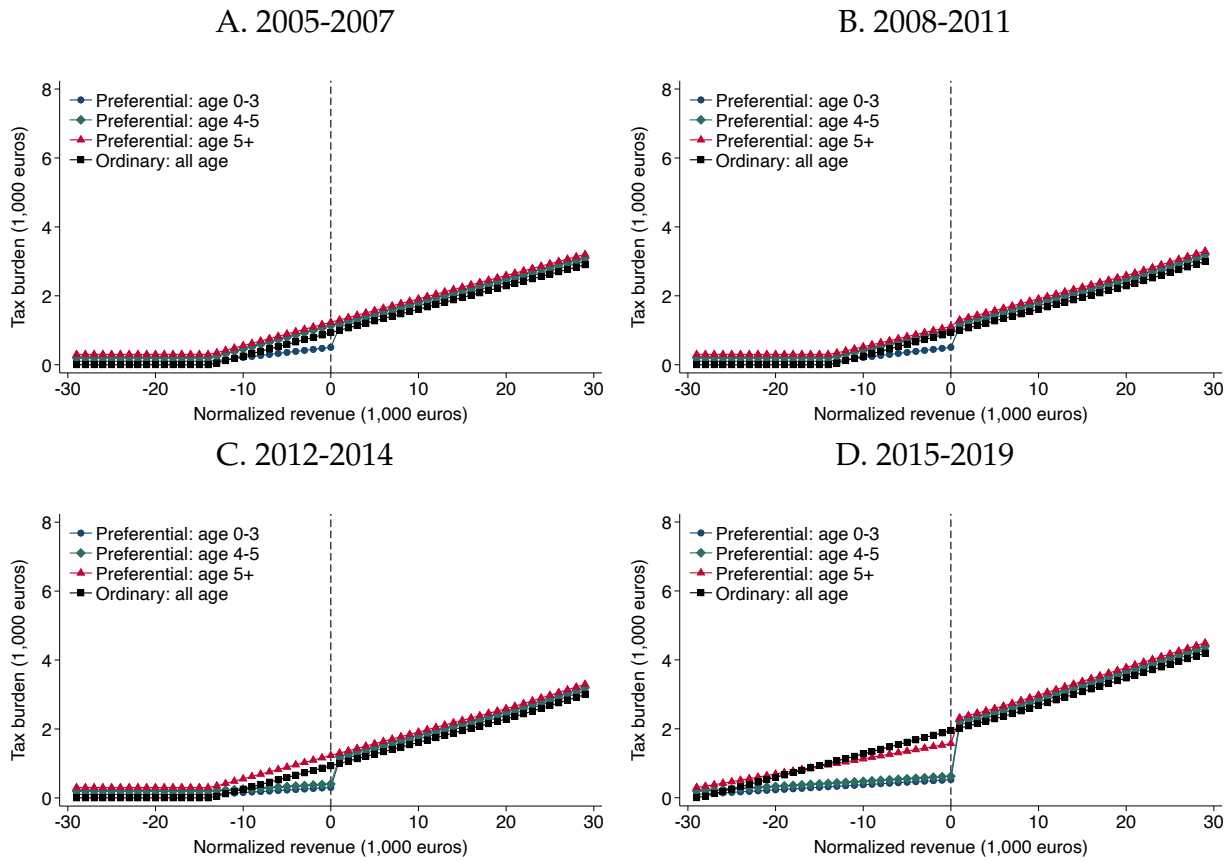
Figure C4: The Distribution of the Bunching Rate



Note: The histograms show the distribution of the bunching rate, the share of individual bunching firms reporting revenue of at least €5,000 below the eligibility cutoff for preferential tax regimes. The bars' width is 2 percentage points.

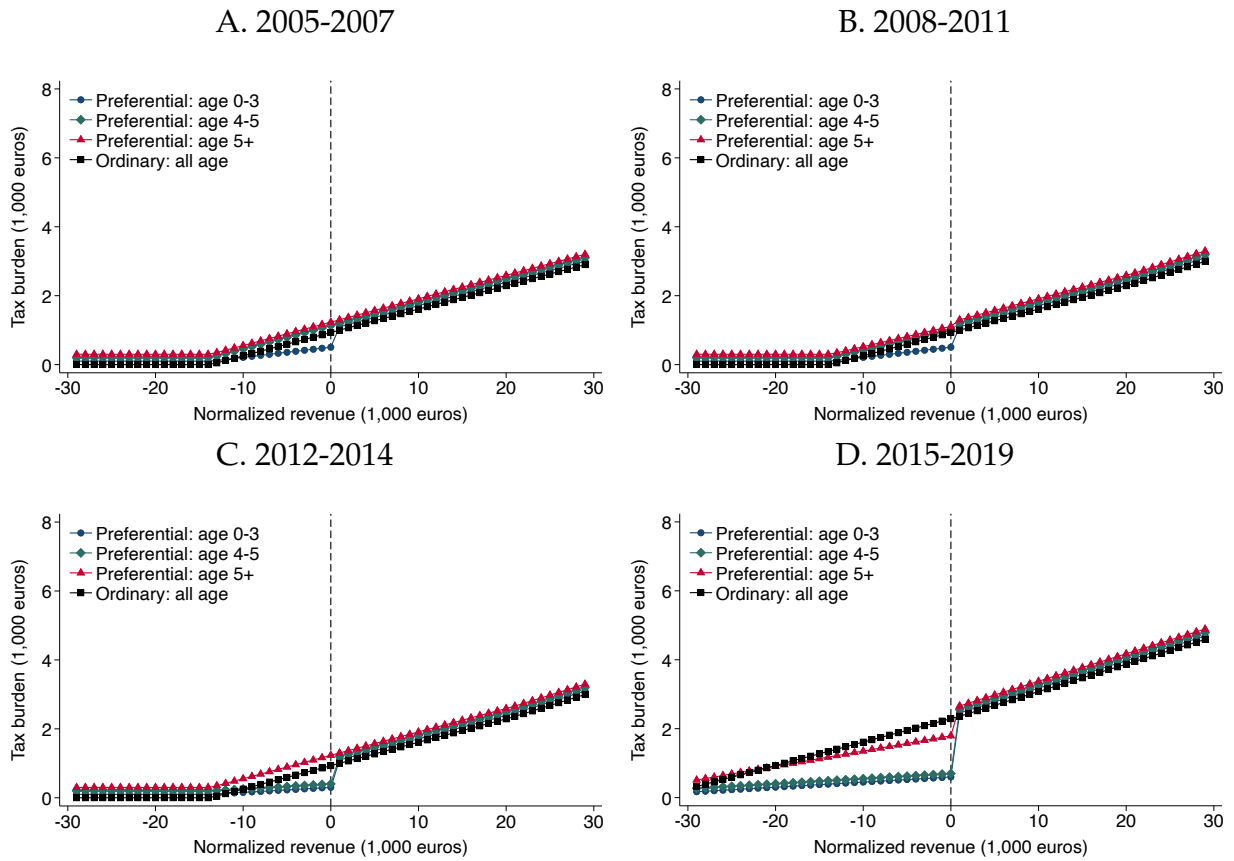
C5. Heterogeneity in the Size of the Tax Notch by Industry

Figure C5: Food and Beverage Industry



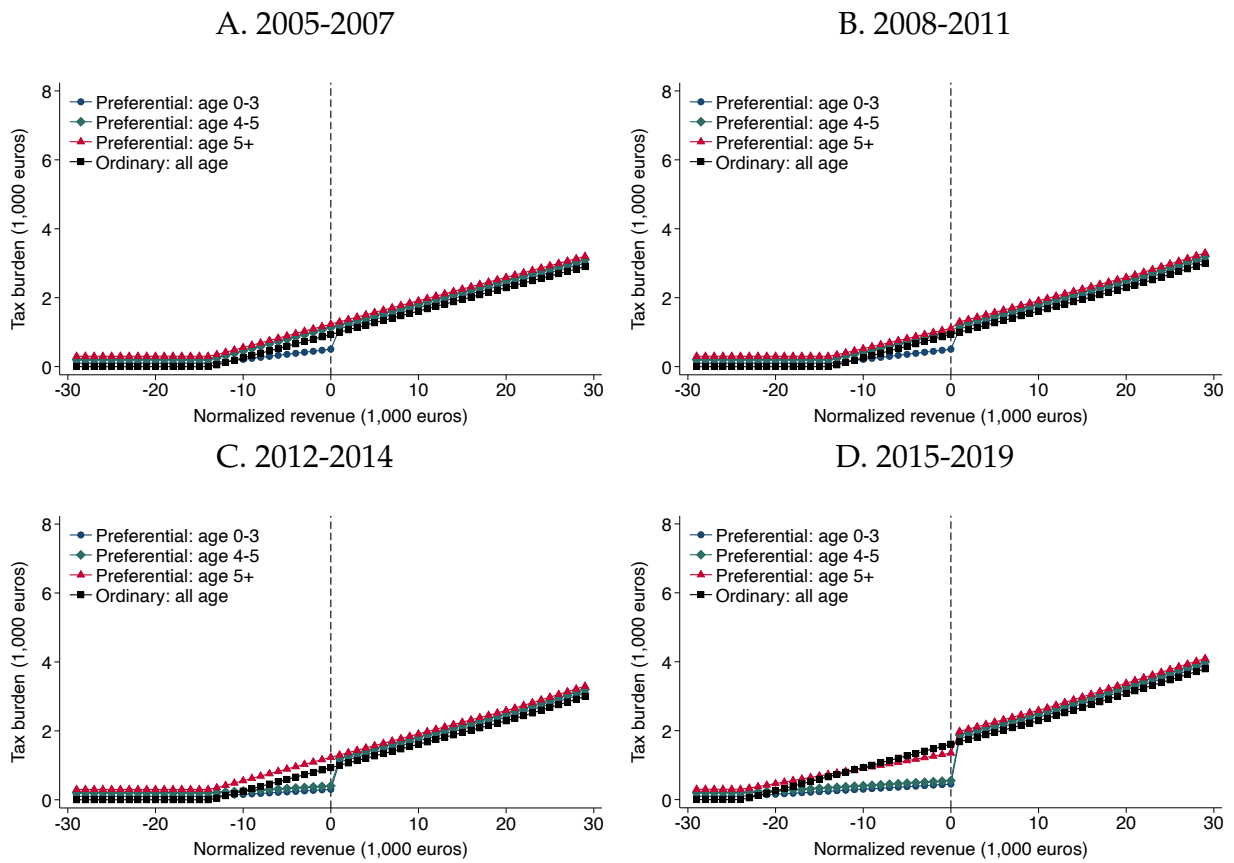
Note: This figure summarizes the main sources of policy changes that create variation in the size of the tax notch for a firm operating in the food and beverage industry. The figure plots the tax burden (vertical axes) by revenue bins, normalized to the eligibility cutoff for preferential regimes (horizontal axes). We plot the tax burden for a representative firm under the preferential and ordinary tax regimes for firms in the age cohorts 0-3, 4-5, and 5+.

Figure C6: Retail Trade Industry



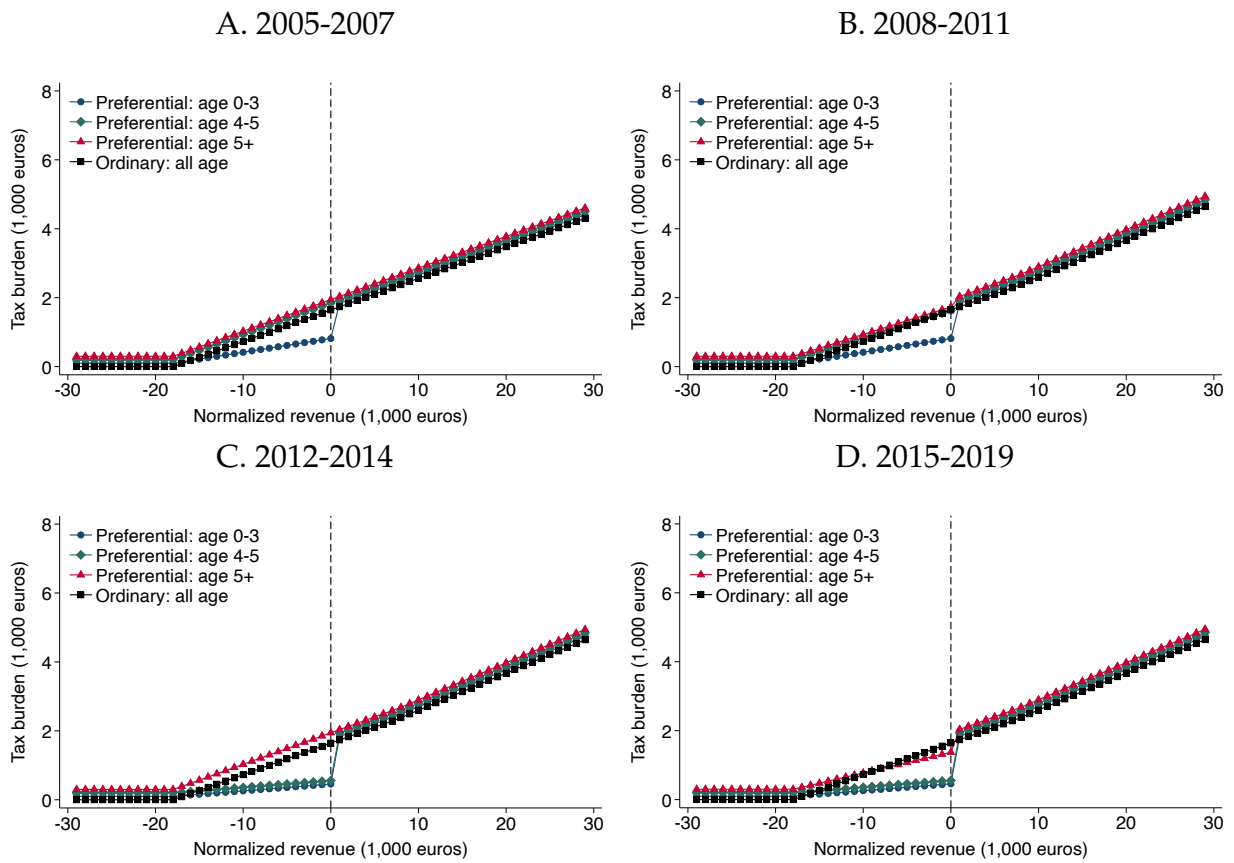
Note: This figure summarizes the main sources of policy changes that create variation in the size of the tax notch for a firm operating in the retail trade industry. The figure plots the tax burden (vertical axes) by revenue bins, normalized to the eligibility cutoff for preferential regimes (horizontal axes). We plot the tax burden for a representative firm under the preferential and ordinary tax regimes for firms in the age cohorts 0-3, 4-5, and 5+.

Figure C7: Street Vendors in Food and Beverage Industry



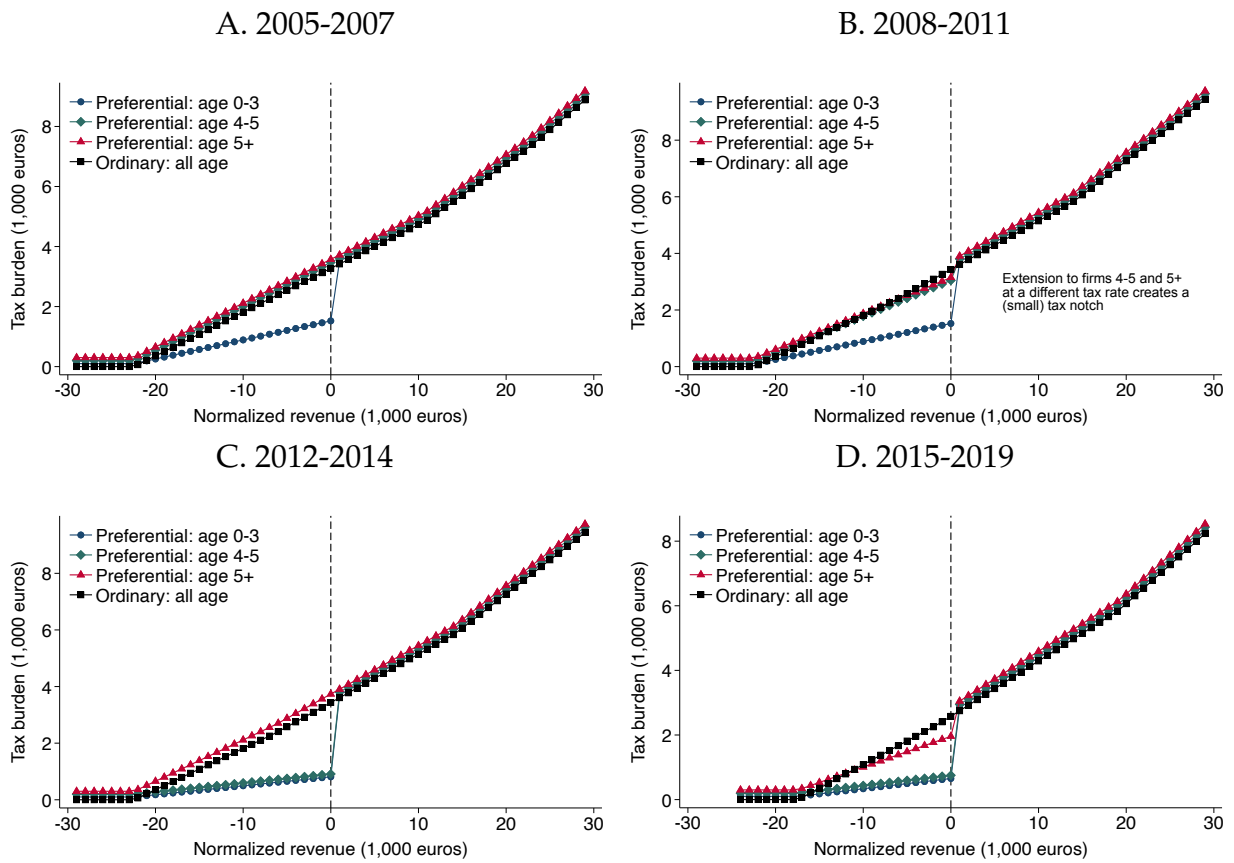
Note: This figure summarizes the main sources of policy changes that create variation in the size of the tax notch for a firm operating in the street vendors of the food and beverage industry. The figure plots the tax burden (vertical axes) by revenue bins, normalized to the eligibility cutoff for preferential regimes (horizontal axes). We plot the tax burden for a representative firm under the preferential and ordinary tax regimes for firms in the age cohorts 0-3, 4-5, and 5+.

Figure C8: Street Vendors in Other Industries



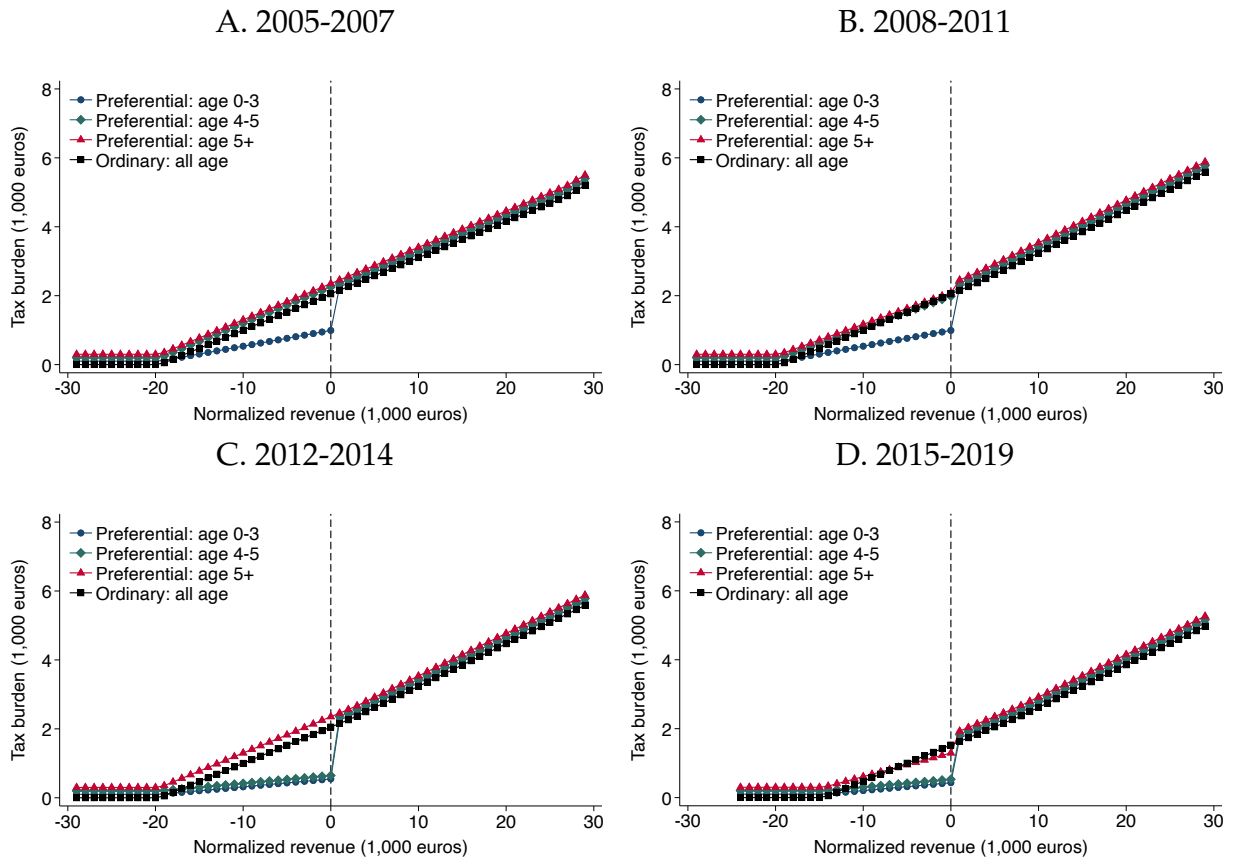
Note: This figure summarizes the main sources of policy changes that create variation in the size of the tax notch for a firm operating in the street vendors of other industries. The figure plots the tax burden (vertical axes) by revenue bins, normalized to the eligibility cutoff for preferential regimes (horizontal axes). We plot the tax burden for a representative firm under the preferential and ordinary tax regimes for firms in the age cohorts 0-3, 4-5, and 5+.

Figure C9: Construction Industry



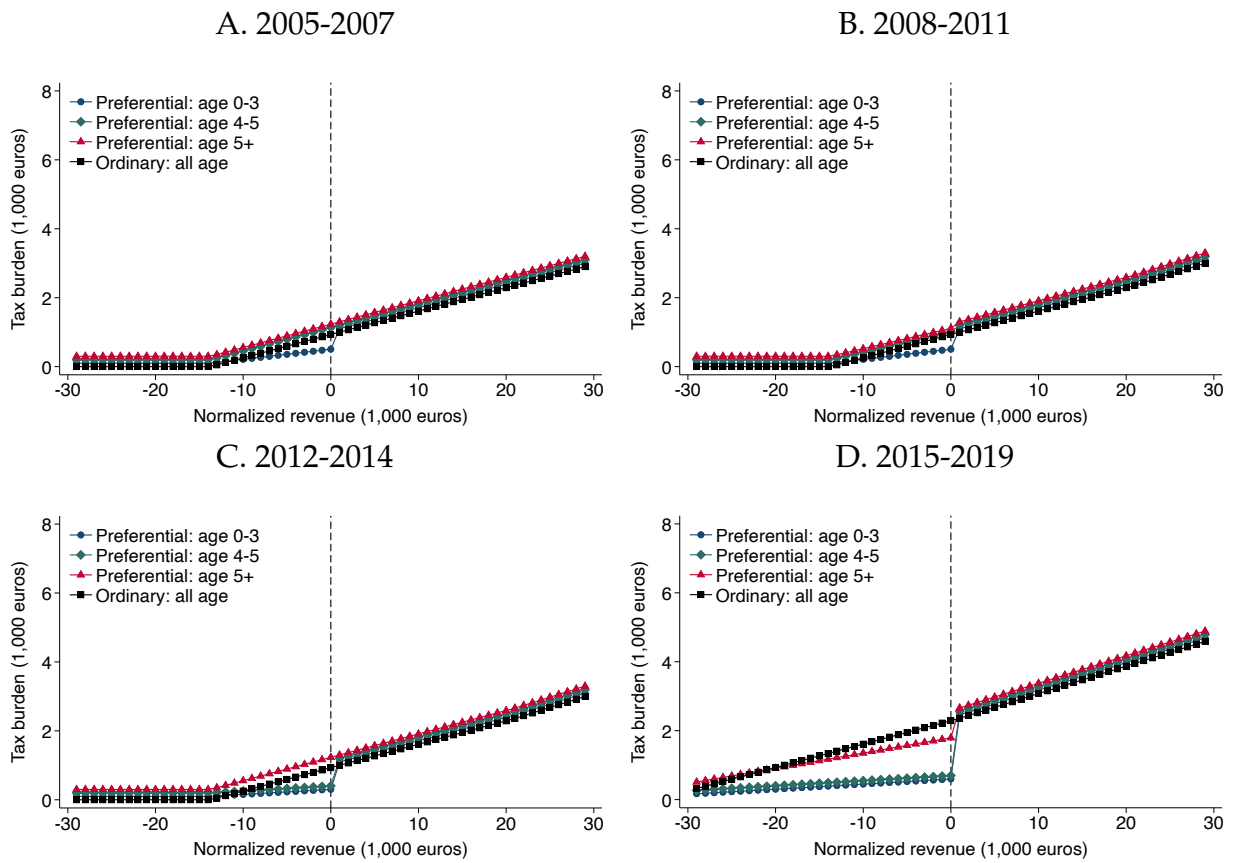
Note: This figure summarizes the main sources of policy changes that create variation in the size of the tax notch for a firm operating in the construction industry. The figure plots the tax burden (vertical axes) by revenue bins, normalized to the eligibility cutoff for preferential regimes (horizontal axes). We plot the tax burden for a representative firm under the preferential and ordinary tax regimes for firms in the age cohorts 0-3, 4-5, and 5+.

Figure C10: Trade Intermediaries Industry



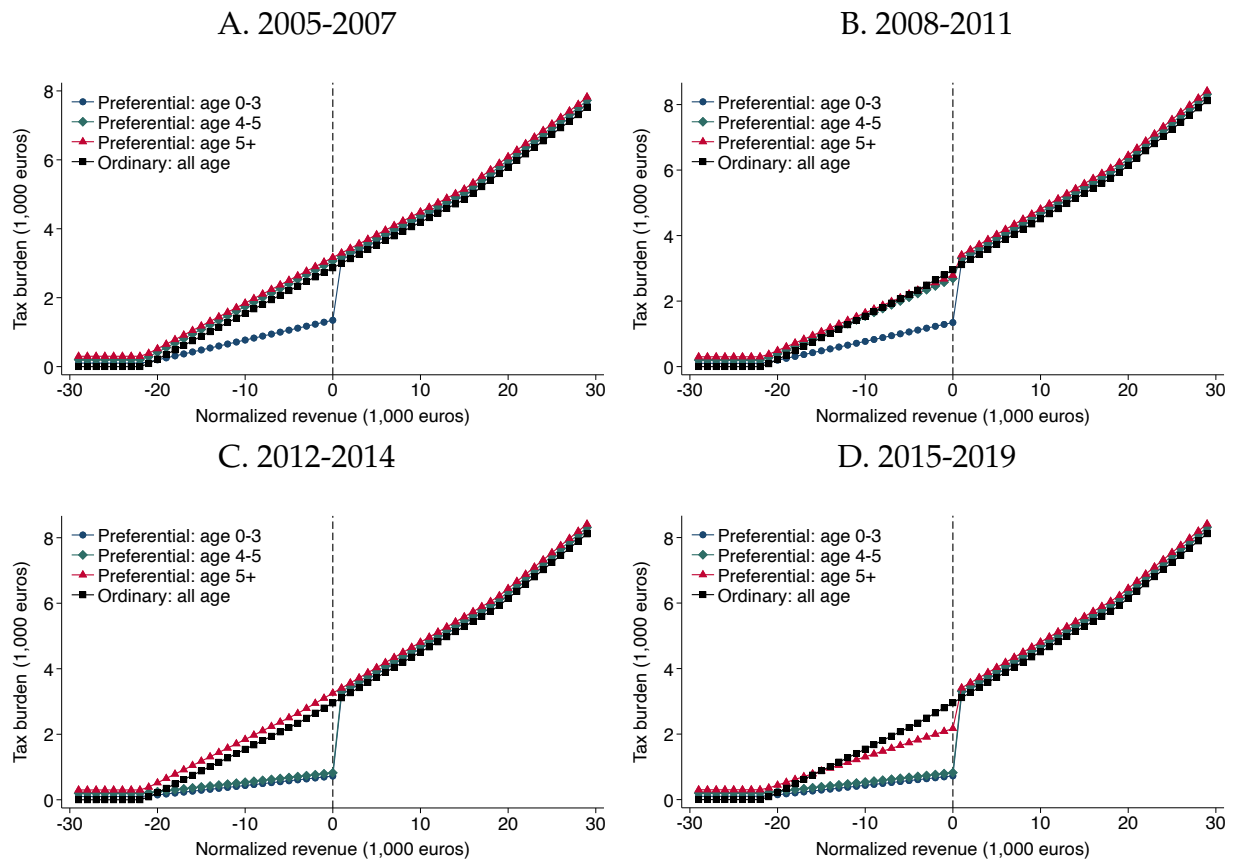
Note: This figure summarizes the main sources of policy changes that create variation in the size of the tax notch for a firm operating in the trade intermediaries industry. The figure plots the tax burden (vertical axes) by revenue bins, normalized to the eligibility cutoff for preferential regimes (horizontal axes). We plot the tax burden for a representative firm under the preferential and ordinary tax regimes for firms in the age cohorts 0-3, 4-5, and 5+.

Figure C11: Accommodation and Catering Industry



Note: This figure summarizes the main sources of policy changes that create variation in the size of the tax notch for a firm operating in the accommodation and catering industry. The figure plots the tax burden (vertical axes) by revenue bins, normalized to the eligibility cutoff for preferential regimes (horizontal axes). We plot the tax burden for a representative firm under the preferential and ordinary tax regimes for firms in the age cohorts 0-3, 4-5, and 5+.

Figure C12: Professional Activities Industry



Note: This figure summarizes the main sources of policy changes that create variation in the size of the tax notch for a firm operating in the professional activities industry. The figure plots the tax burden (vertical axes) by revenue bins, normalized to the eligibility cutoff for preferential regimes (horizontal axes). We plot the tax burden for a representative firm under the preferential and ordinary tax regimes for firms in the age cohorts 0-3, 4-5, and 5+.

C6. Last-Digit Heaping

The evidence in Figure 7 shows that firms just below the eligibility threshold are disproportionately likely to report revenues ending in zero. While informative, that evidence is descriptive and based on the normalized revenue distribution pooled across firms. In this section, we develop a complementary indicator of strategic reporting that can be constructed at the same level of aggregation as our bunching measure, namely the municipality $\times$ industry $\times$ year market.

The idea is to exploit irregularities in the distribution of the *last digit* of reported revenue. In naturally generated accounting data, the last digit should be approximately uniformly distributed, reflecting idiosyncratic variation in transactions, aggregation, and reporting. By contrast, when values are strategically rounded or deliberately chosen, reported numbers tend to heap at salient digits, especially zero. The identifying assumption behind this test is that real adjustments in activity — such as changes in hours, quantities, or scale — may shift the level of revenue, but should not by themselves create a discrete spike in the propensity to report “clean” last digits.

Formally, let $L \in \{0, 1, \dots, 9\}$ denote the last digit of reported revenue. Under the null of no systematic rounding or manipulation, a natural benchmark is

$$\Pr(L = \ell) \approx 0.1 \quad \forall \ell \in \{0, \dots, 9\}.$$

Strategic reporting, instead, can generate excess mass at salient digits, especially $\ell = 0$, because firms round to focal values or mechanically report in coarser units.

Guided by this literature, we define the following market-year indicator:

$$\text{Heaping}_{mt} \equiv \Pr(L = 0 \mid m, t),$$

that is, the share of firms in market m and year t whose reported revenue ends in zero. This measure has two advantages for our purposes. First, unlike the graphical evidence in the main text, it is *time-varying*. Second, it is constructed at the same market level as our bunching measure and can therefore be directly related to bunching and to policy-induced variation in incentives to remain below the threshold.

We use this measure to provide three additional pieces of evidence. First, Table C1 shows that markets with more bunching also display more last-digit heaping. This relationship is estimated using panel regressions at the market level with market and year fixed effects, and remains statistically significant when we include municipality-year and industry-year fixed effects. Relative to the cross-sectional patterns shown in Figure 7, this evidence is more informative because it exploits within-market variation over time and is therefore not driven by time-invariant differences in local accounting norms, sectoral cash intensity, or persistent enforcement environments. It is also substantially more precise, as the panel structure allows us to absorb high-dimensional confounders and pool repeated observations over time.

Second, panel A establishes that bunching and heaping co-move, but both variables are potentially endogenous. We therefore examine whether our instrument $Z_{m,t}$ — which captures policy-induced variation in the incentive to locate below the threshold — also predicts heaping. If bunching primarily reflects strategic reporting, then an increase in $Z_{m,t}$ should raise manipulation incentives and therefore increase heaping. Panel B of Table C1 confirms this prediction: markets more exposed to exogenous increases in the tax incentive to bunch also exhibit significantly more last-digit heaping.

Third, we expect the effect of $Z_{m,t}$ on heaping should be localized just below the eligibility threshold for preferential tax regimes. If, instead, the instrument were capturing broader forces inducing diffuse misreporting, one would expect nonzero effects across a wider range of the revenue distribution. Figure C13 tests this implication by estimating the reduced-form effect of $Z_{m,t}$ on heaping separately by distance-to-threshold bins. The effect is concentrated entirely in the bin just below the threshold and is statistically indistinguishable from zero elsewhere. This pattern closely mirrors the localization of bunching responses and strengthens both the strategic reporting interpretation and the plausibility of the instrument's exclusion restriction.

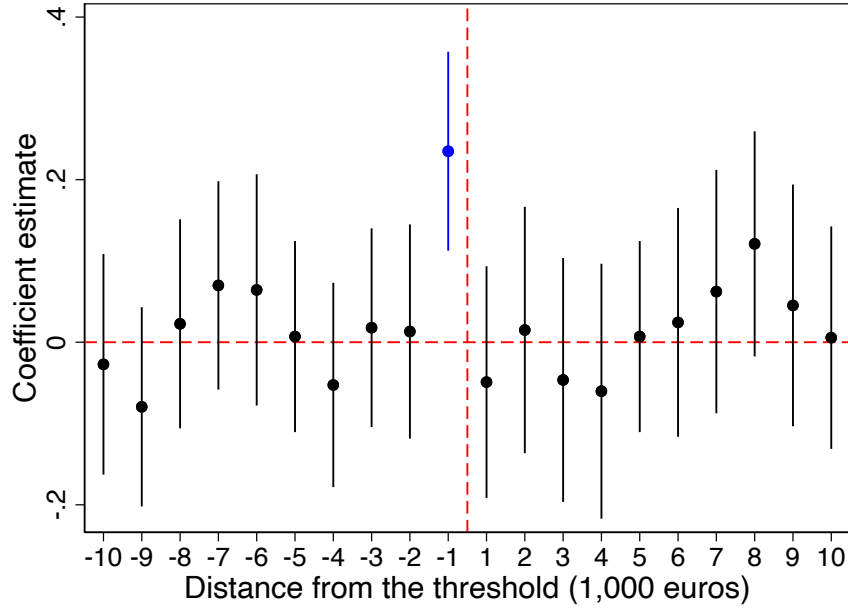
Taken together, these results show that bunching is closely associated with a time-varying, market-level indicator of suspicious reporting behavior. The fact that heaping co-moves with bunching, responds to exogenous variation in tax incentives, and emerges only just below the eligibility threshold provides additional support for interpreting bunching as a manipulation response rather than as a purely real adjustment in economic activity.

Table C1: Heaping Predicts Bunching and Correlates with the Instrument

	(1)	(2)
A. Outcome: $Heaping_{m,t}$		
$B_{m,t}$	0.066 (0.007)	0.015 (0.006)
B. Outcome: $Heaping_{m,t}$		
$Z_{m,t}$	0.235 (0.062)	0.146 (0.056)
Observations	430,722	430,722
Market FE	Yes	Yes
Year FE	Yes	Yes
Year-industry FE	No	Yes
Year-municipality FE	No	Yes

Notes: The unit of observation is a market–year, defined as municipality $\times$ industry $\times$ year. The dependent variable is $Heaping_{m,t}$, the share of firms in market m and year t whose reported revenue has last digit equal to zero (i.e., $\Pr(L = 0 \mid m, t)$). Panel A regresses $Heaping_{m,t}$ on the market-level bunching measure $B_{m,t}$ (share of firms reporting revenues just below the preferential-regime eligibility threshold). Panel B regresses $Heaping_{m,t}$ on the instrument $Z_{m,t}$, which captures exogenous variation in incentives to locate below the threshold. All specifications include market fixed effects and year fixed effects. In column (2), we additionally include municipality $\times$ year and industry $\times$ year fixed effects. Standard errors are clustered at the market level.

Figure C13: The Instrument Predicts Heaping Just in the Bin Below the Threshold



Notes: This figure reports coefficient estimates from bin-specific reduced-form regressions of last-digit heaping on the instrument. The unit of observation is a market-year, defined as municipality $\times$ industry $\times$ year. For each revenue-distance bin b (in €1,000 increments) around the eligibility threshold, we estimate

$$\text{Heaping}_{mt}^{(b)} = \alpha_b + \pi_b Z_{mt} + \mu_m + \lambda_t + \varepsilon_{mt}^{(b)},$$

where $\text{Heaping}_{mt}^{(b)}$ is the share of firms in market m and year t whose reported revenue falls in bin b and has last digit equal to zero, Z_{mt} is the instrument, μ_m are market fixed effects, and λ_t are year fixed effects. Points plot $\hat{\pi}_b$ and whiskers denote 95% confidence intervals; standard errors are clustered at the market level. The estimates are concentrated in the bin immediately below the threshold, with coefficients statistically indistinguishable from zero in bins farther below or above.

C7. Where Does the Cross-Market Variation Come From?

The event-study design exploits cross-market differences in exposure to the 2015 reform, measured by the pre-reform share of individual firms older than five years in each municipality $\times$ industry market. This appendix provides additional intuition for the substantial dispersion in this measure and clarifies whether it is driven primarily by sectoral or geographical heterogeneity. The key point is that, once markets are defined at the intersection of municipality and industry, the exposure measure reflects both persistent differences in firm survival across sectors and differences in local business dynamism within sector across places.

To quantify the relative importance of sectoral and geographical sources of variation, we estimate a sequence of fixed effects-only regressions at the market level, always including year fixed effects: (i) the share of firms older than five years on municipality and year fixed effects; (ii) the same variable on industry and year fixed effects; and (iii) on municipality, industry, and year fixed effects jointly. We use the adjusted R^2 as a summary measure of explanatory power, since these specifications differ substantially in the number of fixed effects. The results show that sectoral heterogeneity is quantitatively more important than geographical heterogeneity: municipality and year fixed effects explain about 8.6 percent of the variation, whereas industry and year fixed effects explain 27.4 percent. Including both sets of fixed effects raises explanatory power to 32.6 percent. This indicates that firm age structure is driven largely by persistent sectoral characteristics, while still leaving a non-trivial role for place-specific factors.

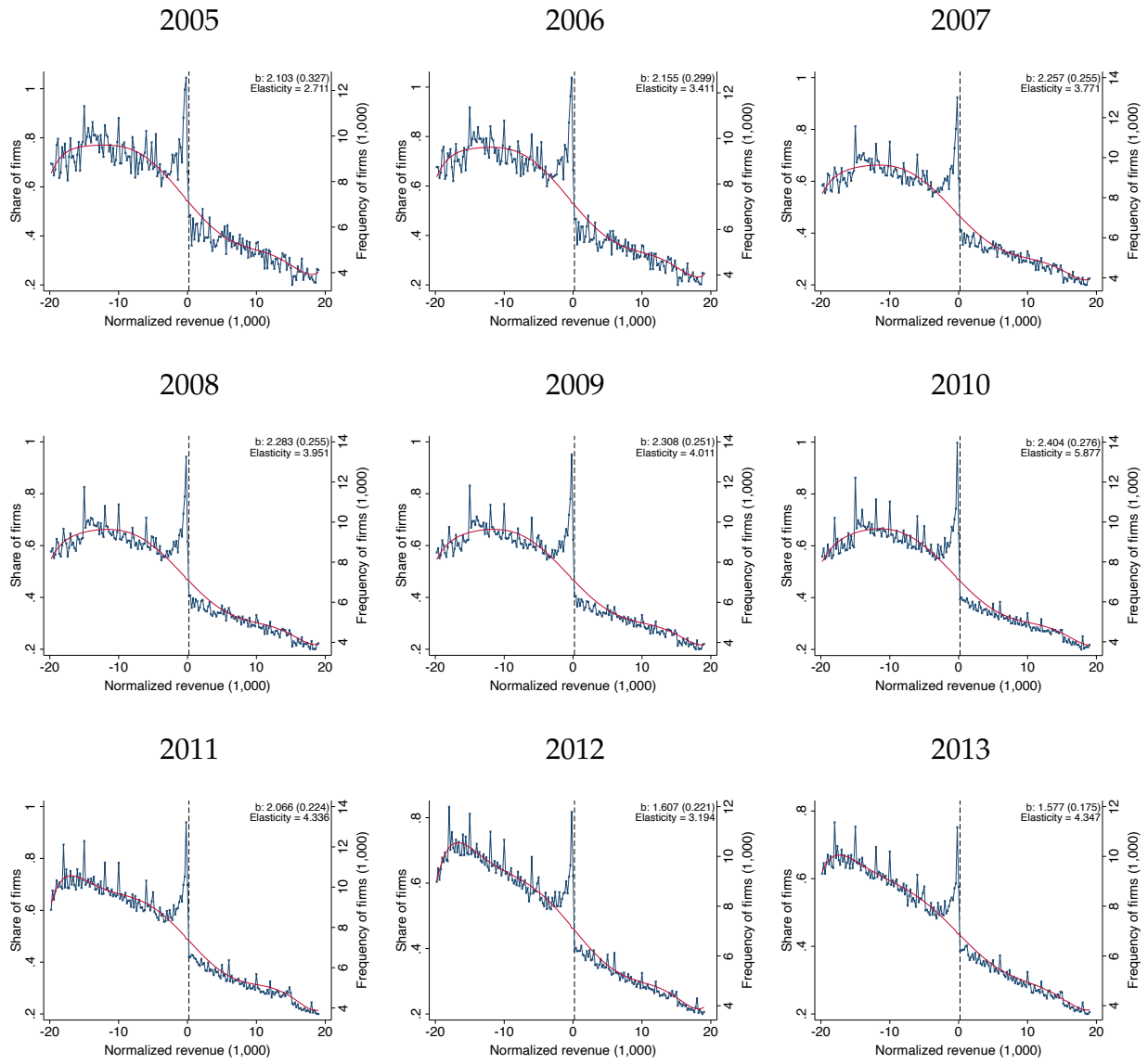
The sectoral component is economically intuitive. Industries with the highest shares of firms older than five years are those in which survival is supported by regulation, relationship capital, and stable local demand. In our data, the share of older individual firms reaches 72.2 percent in architecture and engineering/technical testing, 71.2 percent in human health activities, and 68.2 percent in legal and accounting activities. These sectors typically involve licensing requirements, reputational capital, client-specific relationships, and stronger switching costs, all of which increase the value of accumulated intangible capital and make established firms more likely to survive. High values also appear in personal services (66.7 percent), retail trade (60.8 percent), and auxiliary financial services (60.0 percent), where repeat clientele and local networks generate persistence.

Geography also matters. To illustrate the spatial component of the design, we compute the share of firms older than five years for the ten largest Italian cities. The resulting values range from 22.8 percent in Milan to 33.6 percent in Palermo, with southern metropolitan areas generally displaying higher shares of older individual firms than the main northern growth poles. This pattern is consistent with well-known differences in business dynamism across Italian macro-areas: lower shares of old firms reflect greater entry and turnover, whereas higher shares are consistent with lower churn and stronger persistence of incumbent micro-businesses.

These descriptive patterns help interpret the DiD design but do not drive identification

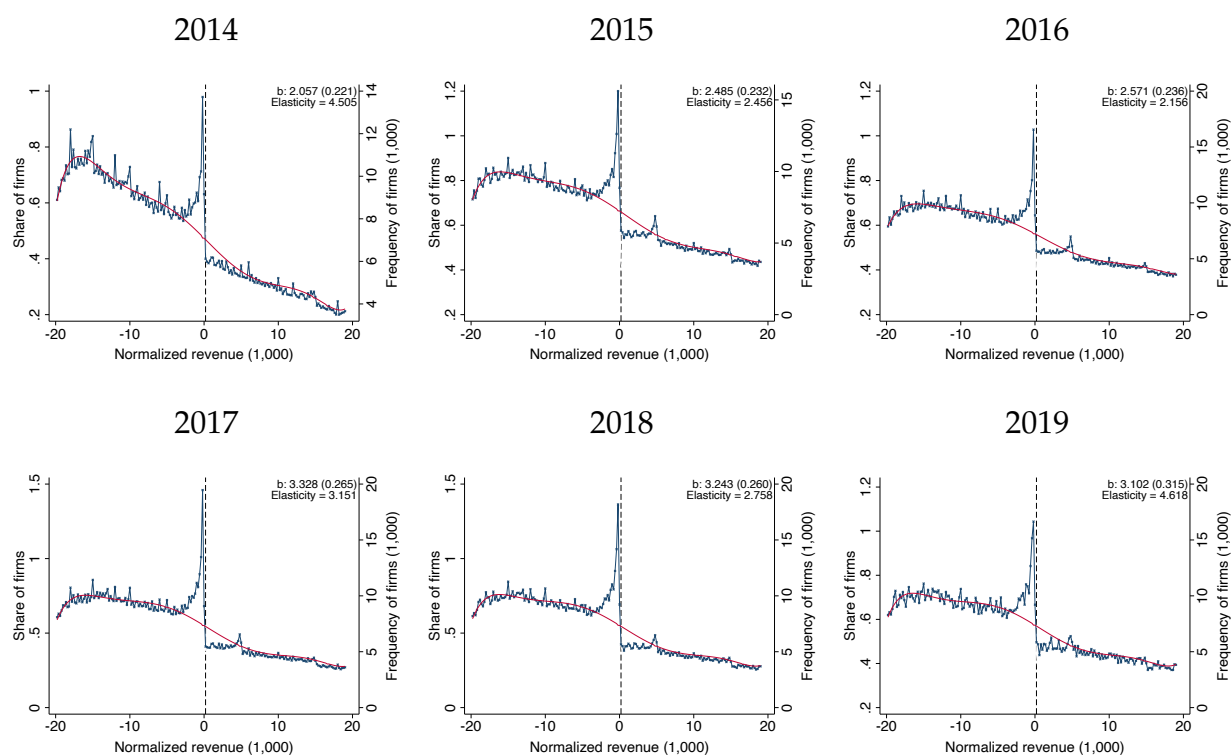
mechanically. In the baseline specifications, firm fixed effects absorb all time-invariant heterogeneity at the firm and market level, while municipality-year and industry-year fixed effects absorb any shocks common to all firms in the same place-year or sector-year environment. The event-study is therefore identified from within-firm changes over time after netting out common local and sectoral shocks, not from static differences in the average age structure of markets.

Figure C14: Bunching Estimates by Year (I/II)



Notes: Each graph shows the empirical distribution of revenue (blue dots) and the counterfactual distribution (red curve) in a specific year. The horizontal axes report normalized revenues. The bins are €200 wide.

Figure C15: Bunching Estimates by Year (II/II)



Notes: Each graph shows the empirical distribution of revenue (blue dots) and the counterfactual distribution (red curve) in a specific year. The horizontal axes report normalized revenues. The bins are €200 wide.

D. Theoretical Framework

This section develops a simple monopolistic competition model with CES preferences. Firms eligible for the preferential regime face a lower effective wedge, which lowers their optimal price and mechanically raises their market share. Eligibility for the preferential regime is governed by a revenue cutoff, which endogenizes regime choice and generates bunching: firms whose unconstrained revenue would lie above the threshold may still access the low-tax regime either by under-reporting revenue or by real scale-down. These two margins have different implications for competition: under-reporting preserves low prices and shifts demand away from compliant competitors, whereas real scale-down requires contracting activity and thus attenuates competitive pressure. The model is used to deliver empirical predictions for bunching responses and competitor spillovers.

D1. Setting

Consider a market with a unit mass of consumers and a continuum of horizontally differentiated varieties indexed by $i \in \mathcal{I}$. Preferences are CES:

$$(D1) \quad U = \left(\int_{i \in \mathcal{I}} q_i^{\frac{\sigma-1}{\sigma}} di \right)^{\frac{\sigma}{\sigma-1}},$$

where q_i denotes consumption of variety i and $\sigma > 1$ is the elasticity of substitution across varieties.

Let total expenditure on the differentiated sector be E . The associated CES price index is

$$(D2) \quad P \equiv \left(\int_{i \in \mathcal{I}} p_i^{1-\sigma} di \right)^{\frac{1}{1-\sigma}},$$

where p_i is the consumer price of variety i .

Firms are monopolistic competitors: each variety is supplied by a single firm that takes aggregate expenditure E as given.¹

Utility maximization subject to the budget constraint yields the standard Marshallian demand system:

$$(D3) \quad q_i = \left(\frac{p_i}{P} \right)^{-\sigma} \frac{E}{P},$$

¹For simplicity, we hold fixed both aggregate expenditure and the set of active varieties, so there is no endogenous entry or exit. Allowing for endogenous entry would add the standard love-of-variety channel as well as additional extensive margin responses to preferential taxation. We abstract from these margins for three reasons. First, our empirical patterns concern reallocation among incumbents operating within the same market. Second, strategic entry into the regime is not allowed *ex lege*. Third, our empirical design exploits variation *at the notch* and finds little evidence of entry or exit responses.

so that firm i 's gross revenue is

$$(D4) \quad r_i \equiv p_i q_i = E \left(\frac{p_i}{P} \right)^{1-\sigma}.$$

Hence, revenue depends only on the firm's price relative to the aggregate price index. Lower relative prices translate mechanically into higher demand and higher sales. The revenue market share of firm i is

$$(D5) \quad s_i \equiv \frac{r_i}{E} = \left(\frac{p_i}{P} \right)^{1-\sigma},$$

which shows that market shares are a power function of relative prices, with elasticity governed by $\sigma - 1$. This feature is central for the analysis below: any tax cut that lowers a firm's effective marginal cost also lowers its equilibrium price and reallocates expenditure toward that firm.

For simplicity, we consider the case where all firms share the same technology and use a single variable input with unit cost w , so marginal costs are common across firms and equal to $mc = w$.² Firms face an ad-valorem wedge on sales revenue, denoted by τ_i , interpretable as an effective tax rate on the revenue tax base. Let $r_i \equiv p_i q_i$ denote gross (pre-tax) revenue, net-of-tax revenue is then $(1 - \tau_i)r_i$.

We distinguish two tax regimes. Under the ordinary regime, the effective wedge is τ^H ; under the preferential regime, the effective wedge is τ^L , with $\tau^L < \tau^H$. Eligibility for the preferential regime requires reported revenue to lie weakly below a cutoff $\bar{r} > 0$. In the next subsection, we first analyze the benchmark in which τ_i is taken as given (i.e., firms cannot manipulate or adjust their revenue to sort into the preferential regime).

D2. A Benchmark with Heterogeneous Tax Wedges

In this benchmark, firms exogenously face different wedges τ_i and cannot adjust behavior to satisfy the eligibility constraint. The purpose of this benchmark is to isolate the pure intensive margin competitive effect of heterogeneous taxation: how a lower ad-valorem wedge translates into lower prices and larger market shares under CES demand.

Because firms are otherwise identical in technology, the only source of within-market asymmetry is the tax wedge.

Proposition 1 (Pricing with an ad-valorem wedge). *For a firm facing a tax wedge τ_i , the profit-maximizing price is a constant markup over marginal cost, scaled by the net-of-tax rate:*

$$(D6) \quad p_i = \frac{\sigma}{\sigma - 1} \frac{w}{1 - \tau_i}.$$

²Productivity differences can be incorporated in the standard way to study additional questions, such as distortions in allocative efficiency or implications for market concentration, as, e.g., in Martin, Parenti and Toubal (2025). Since our main results hold even in the absence of productivity heterogeneity, we rely on representative firms in this analysis.

Proof. Under CES demand, write firm i 's demand as $q_i(p_i) = Ap_i^{-\sigma}$, where $A \equiv EP^{\sigma-1}$. Profits are

$$(D7) \quad \pi_i(p_i) = (1 - \tau_i)p_i q_i - wq_i = [(1 - \tau_i)p_i - w] Ap_i^{-\sigma}.$$

The first-order condition is

$$(1 - \tau_i) \left(1 - \frac{1}{\sigma}\right) p_i = w,$$

which yields

$$p_i = \frac{\sigma}{\sigma - 1} \frac{w}{1 - \tau_i}.$$

Proposition 1 shows that the ad-valorem wedge acts like a proportional increase in effective marginal cost. A higher wedge raises the price required for the firm to obtain a given net return per unit sold. Tax differences thus map directly into price differences.

Proposition 2 (Tax wedges and reallocation of demand). *Fix the aggregate price index P . If two firms a and b face wedges $\tau_a < \tau_b$, then $p_a < p_b$ and $s_a > s_b$. More generally, the elasticities of market share and revenue with respect to the net-of-tax rate are:*

$$(D8) \quad \frac{\partial \log s_i}{\partial \log(1 - \tau_i)} = \sigma - 1 > 0, \quad \frac{\partial \log r_i}{\partial \log(1 - \tau_i)} = \sigma - 1 > 0.$$

Proof. By Proposition 1, price satisfies $p_i \propto (1 - \tau_i)^{-1}$, so $\tau_a < \tau_b$ implies $p_a < p_b$. From CES demand,

$$s_i = \frac{r_i}{E} = \left(\frac{p_i}{P}\right)^{1-\sigma},$$

so a lower price implies a larger market share. Log-differentiating,

$$d \log s_i = (1 - \sigma) d \log p_i.$$

Since $d \log p_i = -d \log(1 - \tau_i)$, it follows that

$$d \log s_i = (\sigma - 1) d \log(1 - \tau_i).$$

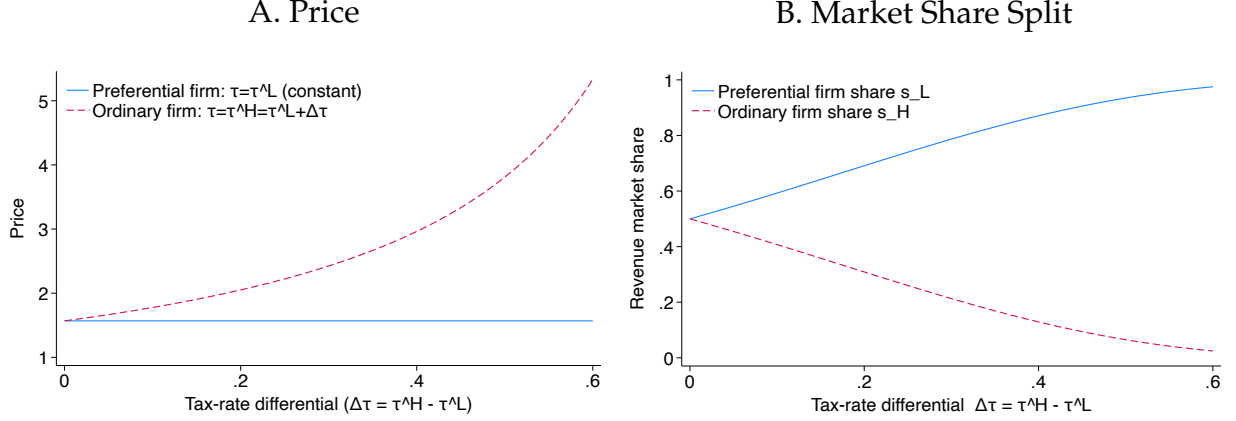
Because $r_i = Es_i$ and aggregate expenditure E is fixed, the same elasticity applies to revenue.

This proposition delivers a simple competitive mechanism. Preferential taxation creates a *tax-induced cost advantage*: firms facing a lower wedge charge lower prices and attract a larger share of expenditure.

Figure D1 illustrates this mechanism in the case of two firms that differ only in the wedge they face. Panel A plots equilibrium prices as the tax differential $\Delta\tau \equiv \tau^H - \tau^L$ varies. Consistent with Proposition 1, the firm subject to the lower wedge charges a

lower price, and the price gap widens monotonically with the tax differential. Panel B reports the corresponding split in revenue market shares. As $\Delta\tau$ increases, expenditure is progressively reallocated toward the low-wedge firm.

Figure D1: Competitive Distortions from Preferential Taxation



Notes: The figure illustrates the benchmark CES monopolistic competition model with heterogeneous revenue wedges and no bunching. Panel A plots equilibrium prices implied by Proposition 1, $p(\tau) = \frac{\sigma}{\sigma-1} \cdot \frac{w}{(1-\tau)}$, for two otherwise identical firms that differ only in the wedge they face: a preferential regime firm with $\tau = \tau^L$ (solid line) and an ordinary regime firm with $\tau = \tau^H = \tau^L + \Delta\tau$ (dashed line). We normalize the common cost shifter to $w = 1$ and set the elasticity of substitution to $\sigma = 4$. The latter value lies within the range commonly used in empirical and quantitative applications of CES demand. For example, Broda and Weinstein (2006) estimate an average elasticity of substitution around 4 across three-digit goods in the United States over 1990–2001. We fix $\tau^L = 0.15$ and vary the tax rate differential over $\Delta\tau \in [0, 0.6]$ (so $\tau^H \in [0.15, 0.75]$). Panel B reports the implied revenue market shares under CES demand, $s_i = p_i^{1-\sigma} / (p_L^{1-\sigma} + p_H^{1-\sigma})$.

D3. Eligibility Cutoff and Bunching

A distinctive feature of preferential tax regimes is that firms qualify only if reported revenue lies below a cutoff. This creates a notch in the effective tax schedule. Relative to the benchmark with heterogeneous but exogenous wedges, the cutoff introduces an additional distortion: firms whose unconstrained optimum lies above the threshold may *bunch* just below it in order to obtain the preferential tax rate (Saez 2010; Chetty, Friedman, Olsen and Pistaferri 2011; Kleven and Waseem 2013).

Let *true* revenue be $r_i \equiv p_i q_i$ as in (D3) and *reported* revenue be $\hat{r}_i \leq r_i$. Eligibility for the preferential tax rate τ^L requires $\hat{r}_i \leq \bar{r}$, where $\bar{r}$ denotes the statutory cutoff. In addition to a lower tax rate, the preferential regime also provides *non-tax advantages*, such as simplified bookkeeping and lower compliance costs (Aghion, Gravoueille, Lequien and Stantcheva 2024). We summarize these advantages through a firm-specific lump-sum benefit $B_i \geq 0$.³

³Because B_i is a lump-sum gain from eligibility, it enters the firm's payoff as an additive constant conditional on qualifying for the regime. Hence, when eligibility is already secured and the cutoff is slack, B_i does not vary with p_i , q_i , or r_i , and therefore drops out of the firm's pricing first-order condition.

Two types of bunching responses. Consider a firm whose *unconstrained* optimum under the preferential wedge τ^L would imply true revenue above the cutoff, $r_i^{L*} \equiv r_i(p_i^{L*}) > \bar{r}$, where p_i^{L*} is the constant markup price under τ^L from Proposition 1. Such a firm can qualify for τ^L in two conceptually distinct ways:⁴

- i. **Under-reporting (evasion-induced bunching).** The firm keeps its *real* choices at the unconstrained optimum under τ^L , i.e. it charges p_i^{L*} and sells $q_i^{L*} \equiv q_i(p_i^{L*})$, but understates revenue to satisfy the eligibility rule. Let reported revenue be $\hat{r}_i \leq r_i^{L*}$, so that evaded revenue are $e_i \equiv r_i^{L*} - \hat{r}_i \geq 0$. Under-reporting entails a real resource cost $\kappa_i(e_i)$ with $\kappa_i(0) = 0$, $\kappa_i'(e) > 0$, and $\kappa_i''(e) > 0$, capturing concealment effort, expected penalties, and other evasion-related expenditures.⁵ The payoff with under-reporting is

$$(D9) \quad \pi_i^{\text{ev}} = \pi_i(p_i^{L*}; \tau^L) + B_i - \kappa_i(r_i^{L*} - \hat{r}_i),$$

where $\pi_i(p; \tau)$ denotes operating profits given price p and wedge τ . Relative to remaining in the ordinary regime (optimal price p_i^{H*} under τ^H), the net gain from qualifying via under-reporting is

$$(D10) \quad \Delta V_i^E \equiv [\pi_i(p_i^{L*}; \tau^L) - \pi_i(p_i^{H*}; \tau^H)] + B_i - \kappa_i(r_i^{L*} - \hat{r}_i).$$

- ii. **Scale-down (real bunching).** The firm qualifies by adjusting *real* activity if true revenue lies above the cutoff. A parsimonious representation is a constraint on true revenue: the firm chooses its price under τ^L subject to $r_i(p_i) \leq \bar{r}$,

$$(D11) \quad \max_{p_i} \pi_i(p_i; \tau^L) + B_i \quad \text{s.t. } r_i(p_i) \leq \bar{r}.$$

Under CES demand, $r_i(p) = E \left(\frac{p}{\bar{p}} \right)^{1-\sigma}$, and $\frac{dr_i(p)}{dp} = E(1-\sigma) \left(\frac{1}{\bar{p}} \right)^{1-\sigma} p^{-\sigma} < 0$, so $r_i(\cdot)$ is strictly decreasing in p . Therefore, if $r_i^{L*} = r_i(p_i^{L*}) > \bar{r}$, feasibility requires a price $\bar{p}_i$ satisfying $r_i(\bar{p}_i) = \bar{r}$, which exists uniquely and obeys $\bar{p}_i > p_i^{L*}$. Real scale-down thus implies a strictly *higher* price and lower quantity than the unconstrained optimum under τ^L . Let $\pi_i^{L, \text{cap}}(\bar{r})$ denote the maximized operating profits in (D11).

⁴These two cases should be interpreted as polar mechanisms. In practice, firms may combine revenue underreporting with some real adjustment. The empirical predictions are monotone in the relative importance of the evasion component: the more bunching reflects underreporting rather than true scale reduction, the more likely we are to observe lower prices and negative spillovers on compliant competitors.

⁵We let the evasion cost function be firm-specific because even under identical regulation, firms may differ in their ability to under-report. For instance, some firms may have better managerial ability to conceal transactions, while others operate in industries where a smaller share of revenue is third-party reported and thus easier to misstate. Following Chetty (2009), the cost $\kappa_i(e)$ should be interpreted as the real resource cost of sheltering revenue. It includes, for example, time and resources devoted to concealment, or the loss from operating in cash rather than through traceable electronic payments.

The net gain from qualifying via real scale-down is then

$$(D12) \quad \Delta V_i^R \equiv \left[\pi_i^{L, \text{cap}}(\bar{r}) - \pi_i(p_i^{H*}; \tau^H) \right] + B_i.$$

Because the cap forces the firm away from its unconstrained optimum under τ^L , we have $\pi_i^{L, \text{cap}}(\bar{r}) \leq \pi_i(p_i^{L*}; \tau^L)$, with strict inequality when $r_i^{L*} > \bar{r}$.⁶ The benefit B_i rationalizes why such distortion may still be privately optimal: even if operating profits fall, qualifying for the preferential regime can remain attractive because it delivers B_i , which can be particularly valued by firms facing large hassle costs.

Firms need not react uniformly to the eligibility cutoff. Heterogeneity in non-tax benefits B_i and in misreporting costs $\kappa_i(\cdot)$ implies that some firms may not seek eligibility at all, while others qualify either by underreporting or by real scale-down. To make this distinction transparent, it is useful to distinguish between the *extensive margin* (whether the firm bunches at all) and the *modes of adjustment* conditional on bunching (how the firm gets below the cutoff).

Extensive margin. Let p_i^{H*} and p_i^{L*} denote the unconstrained optimal prices under wedges τ^H and τ^L , respectively, and let $r_i^{L*} \equiv r_i(p_i^{L*})$. The value of staying in the ordinary regime is defined as $V_i^H \equiv \pi_i(p_i^{H*}; \tau^H)$. If the firm enters the preferential regime, it obtains the lower wedge τ^L and the lump-sum benefit B_i , but may have to incur a distortion cost to satisfy the cutoff. Let $C_i^E \equiv \kappa_i(r_i^{L*} - \hat{r}_i)$ denote the cost of misreporting, and let $C_i^R \equiv \pi_i(p_i^{L*}; \tau^L) - \pi_i^{L, \text{cap}}(\bar{r}) \geq 0$ denote the profit loss from real scale-down when the cap binds. The best attainable value from accessing the preferential regime is then:

$$V_i^L \equiv \pi_i(p_i^{L*}; \tau^L) + B_i - \min\{C_i^E, C_i^R\},$$

and the firm enters the preferential regime if

$$(D13) \quad V_i^L \geq V_i^H.$$

This extensive margin condition is driven by the wedge differential $(\tau^H - \tau^L)$, which raises operating profits under τ^L via lower prices and higher demand, and by the non-tax benefit B_i , net of the cheapest way of meeting the cutoff.

Prediction 1 (Bunching increases with the tax rate differential). *Ceteris paribus, bunching at the eligibility cutoff is stronger when the tax rate differential between the ordinary and preferential regimes, $\Delta\tau \equiv \tau^H - \tau^L$, is larger. This is because an increase in $\Delta\tau$ raises the net gain from entering the preferential regime and relaxes the extensive margin condition in (D13), expanding*

⁶Real scale-down may similarly occur through quantity restrictions rather than price changes (e.g., limiting capacity, shortening opening hours, or turning away customers). Holding the price fixed at p_i^{L*} , satisfying the cutoff requires choosing a quantity $\bar{q}_i$ such that $p_i^{L*}\bar{q}_i = \bar{r}$, i.e. $\bar{q}_i = \bar{r}/p_i^{L*} < q_i^{L*}$ whenever $r_i^{L*} > \bar{r}$. Our main prediction would still hold.

the set of firms for which entering is optimal. Therefore, holding constant $\kappa_i(\cdot)$ and B_i , the mass of firms choosing to enter the preferential tax regime is weakly increasing in $\Delta\tau$, implying a larger bunching mass below $\bar{r}$.

Modes of adjustment. Conditional on entering the preferential regime, firms choose the bunching response that minimizes the cost of satisfying the cutoff. Under-reporting is preferred to real scale-down if the evasion cost is below the profit loss implied by real scale-down:

$$(D14) \quad \underbrace{\kappa_i(r_i^{L*} - \bar{r})}_{\text{misreporting cost}} < \underbrace{\pi_i(p_i^{L*}; \tau^L) - \pi_i^{L, \text{cap}}(\bar{r})}_{\text{profit loss from real constraint}}.$$

Therefore, B_i and $\Delta\tau$ primarily shift the extensive margin, changing the attractiveness of entering the preferential regime. In contrast, heterogeneity in $\kappa_i(\cdot)$ and in the profit loss C_i^R governs the intensive margin. Firms with low $\kappa_i(\cdot)$ disproportionately sort into evasion-driven bunching, while firms with high misreporting costs but sufficiently large gains from entry choose to qualify via real scale-down. This distinction is especially relevant in our institutional setting. Since eligibility for the preferential regime is associated with a lower probability of audit, then the expected penalty component of $\kappa_i(\cdot)$ is reduced for firms that succeed in qualifying. This implies a low $\kappa_i(\cdot)$ around the cutoff, enlarging the set of firms satisfying (D14), and implying that observed bunching predominantly reflected under-reporting.

Prediction 2 (Bunching is stronger where evasion costs are lower). *If bunching primarily reflects under-reporting, then bunching should be more prevalent in industries or municipalities with lower misreporting costs $\kappa_i(\cdot)$, since a downward shift in $\kappa_i(\cdot)$ expands the set of firms satisfying the condition in (D14).*

Figure D2 summarizes these predictions in the (κ_i, L_i) space, where $L_i \equiv \pi_i(p_i^{L*}; \tau^L) - \pi_i^{L, \text{cap}}(\bar{r})$ is the profit loss from qualifying via real scale-down. The downward sloping boundary $\kappa_i = L_i$ shows that, conditional on entering the preferential regime, firms under-report when evasion is cheaper than real adjustment ($\kappa_i \leq L_i$) and scale down when the real profit loss is lower ($L_i < \kappa_i$). The dashed L-shaped threshold captures the extensive margin: firms enter the preferential regime only if the net gain from entry, $T \equiv \Delta\tau + B_i$, exceeds the least-cost method of qualifying, i.e. $\min\{\kappa_i, L_i\} \leq T$; otherwise they remain in the ordinary regime. A higher tax rate differential $\Delta\tau$ (or, similarly, a larger non-tax benefits B_i) shrinks the ordinary region, increasing bunching.

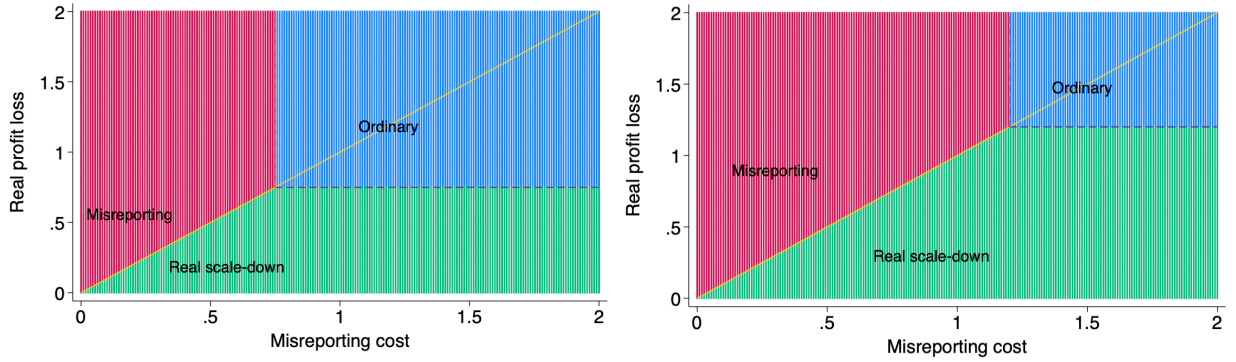
D4. Implications for Market Competition

We now study how bunching responses influence market competition. Consider a set of potential bunchers $\mathcal{B}$ and a set of compliant competitors $\mathcal{C}$ that cannot access τ^L and

Figure D2: Firm Choice at the Cutoff

A. Low incentive

B. High incentive



Notes: The figure depicts firms' optimal response to a revenue cutoff determining access to a preferential tax regime. The horizontal axis reports the evasion (misreporting) cost evaluated at the amount needed to qualify, $\kappa_i \equiv \kappa_i(r_i^{L^*} - \bar{r})$; the vertical axis reports the operating profit loss from qualifying via real scale down, $L_i \equiv \pi_i(p_i^{L^*}; \tau^L) - \pi_i^{L, \text{cap}}(\bar{r})$. The 45-degree line $\kappa_i = L_i$ is the modes of adjustment boundary: conditional on entering the preferential regime, firms under-report when $\kappa_i \leq L_i$ (red region) and scale down when $L_i < \kappa_i$ (green region). The "ordinary" region (in blue) captures the extensive margin: firms remain in the ordinary regime when both qualification methods are too costly relative to the net gain from entering, i.e. when $\min\{\kappa_i, L_i\} > T$, where $T \equiv \Delta\tau + B_i$ is the sum of the operating profit gain from accessing the preferential tax regime. Panel A uses a low-entry-incentive calibration $T = 0.75$, while Panel B uses a high-entry-incentive calibration $T = 1.20$, so the dashed L-shaped boundary shifts outward.

therefore always face τ^H (hence choose $p_j^{H^*}$). Bunching responses generate spillovers from $\mathcal{B}$ to $\mathcal{C}$ depending on *how* firms in $\mathcal{B}$ satisfy the eligibility constraint.

Proposition 3 (Evasion-driven bunching reduces competitors' revenue). *Suppose a positive mass of firms in $\mathcal{B}$ switches from τ^H to τ^L via evasion-driven bunching. Then, those firms set lower prices and gain revenue market share relative to the no-bunching counterfactual, and compliant competitors lose market share and revenue. In particular, for any $j \in \mathcal{C}$,*

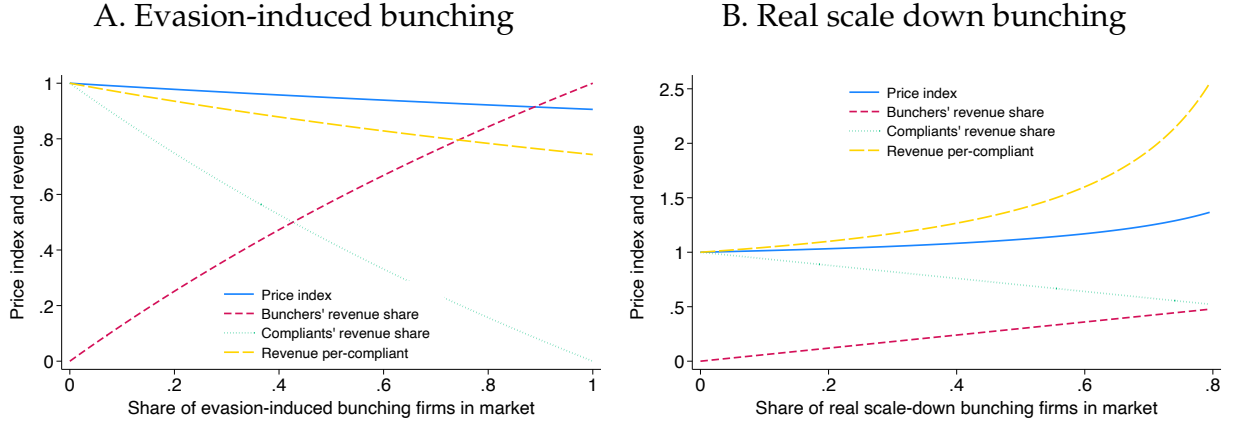
$$(D15) \quad \Delta \log s_j < 0 \quad \text{and} \quad \Delta \log r_j < 0.$$

Proof. By Proposition 1, a firm facing wedge τ sets $p_i^*(\tau) = \frac{\sigma}{\sigma-1} \frac{w}{1-\tau}$. Hence, when a bunching firm switches from τ^H to $\tau^L < \tau^H$, its price falls: $p_i^{L^*} < p_i^{H^*}$, so $p_i^{1-\sigma}$ rises. Because $P^{1-\sigma} = \int_{k \in \mathcal{I}} p_k^{1-\sigma} dk$, an increase in $p_i^{1-\sigma}$ for a positive measure of firms increases $P^{1-\sigma}$ and therefore lowers P . For any compliant competitor $j \in \mathcal{C}$, $p_j = p_j^{H^*}$ is unchanged, so $s_j = (p_j/P)^{1-\sigma}$ falls when P falls. With E fixed, $r_j = Es_j$ falls as well.

Panel A of Figure D3 illustrates the reallocation implied by Proposition 3. The figure is constructed by varying the mass of evasion-driven bunchers (b) in the market. Bunching firms price at the unconstrained preferential regime optimum $p^{L^*} = \frac{\sigma}{\sigma-1} \frac{w}{1-\tau^L}$, while compliant firms price at the ordinary regime optimum $p^{H^*} = \frac{\sigma}{\sigma-1} \frac{w}{1-\tau^H}$. For each b , the CES price index is computed as $P(b) = [b(p^{L^*})^{1-\sigma} + (1-b)(p^{H^*})^{1-\sigma}]^{1/(1-\sigma)}$ and plotted relative to the no-bunching benchmark $P(0)$. Group revenues are obtained from CES demand, $r_k(\alpha) = E(p_k/P(\alpha))^{1-\sigma}$, aggregated by group as $R_B(b) = br_B(b)$ and

$R_C(\alpha) = (1 - b)r_C(b)$, and displayed as shares of aggregate expenditure, $R_B(b)/E$ and $R_C(b)/E$. The main patterns follow directly from the CES aggregator: as b increases, a larger measure of varieties is priced at the lower p^{L*} , which raises $\int p_i^{1-\sigma} di$ and therefore lowers the price index $P(\alpha)$. The fall in $P(\alpha)$ implies that compliant firms become relatively more expensive, so their revenue share $R_C(b)/E$ declines, while the bunching group's revenue share $R_B(b)/E$ rises.

Figure D3: Spillover Effects from Bunching Responses



Notes: The figure illustrates spillover effects from bunching responses to get eligibility from a preferential tax regime. The horizontal axis reports the share of bunching firms in the market (b). In both panels, the solid line plots the CES price index $P(b)$, with $P(0) = 1$. The dashed lines report group revenue shares (revenues divided by E) for bunching firms and compliant firms, computed from CES demand and the implied price index. The long dashed line reports revenue per compliant firm, $r_C(b)$, with $r_C(0) = 1$. Panel A (evasion-induced bunching) assumes bunching firms obtain the preferential wedge $\tau^L = 0.15$ and set the unconstrained optimal price $p^{L*} = \frac{\sigma}{\sigma-1} \frac{w}{1-\tau^L}$, while compliant firms face $\tau^H = 0.23$ and price at p^{H*} . As b increases, a larger mass of varieties is priced at $p^{L*} < p^{H*}$, which lowers $P(b)$ and reallocates expenditure toward bunchers: the bunchers' revenue share rises and the compliant group's revenue share (and per-firm revenue) falls. Panel B (real scale-down bunching) assumes bunching firms qualify for τ^L by satisfying a binding true revenue cap $r_B = \bar{r}$ ($\bar{r}/E$ set at 0.60 for graphical purposes), implying a constrained price $p_B(b)$ above the unconstrained preferential optimum. As b increases, the constrained bunchers' contribution to $\int p^{1-\sigma} di$ falls, so $P(b)$ rises, attenuating competitive pressure on compliant firms: revenue per compliant firm increases even though the compliant group's revenue share declines mechanically as the mass of compliant firms $(1 - b)$ shrinks. In this panel, we restrict the support of b to $b \leq 0.8$ to avoid the mechanical divergence of $P(b)$ as $b \rightarrow 1$ under a binding cap with $\bar{r} < E$. In both graphs, we set $\sigma = 4$ and normalize w and E equal to 1.

Proposition 4 (Real scale-down bunching does not reduce competitors' revenue). *Suppose that firms in $\mathcal{B}$ qualify via real scale-down, i.e. they obtain τ^L but choose prices subject to the binding true-revenue constraint $r_i(p_i) = \bar{r}$. Then, their contribution to $P^{1-\sigma}$ is weakly lower than under the unconstrained τ^L optimum, and compliant competitors' market shares and revenues weakly increase. In particular, for any $j \in \mathcal{C}$,*

$$(D16) \quad \Delta \log s_j \geq 0 \quad \text{and} \quad \Delta \log r_j \geq 0.$$

Proof. Under CES demand, $r_i(p) = E(p/P)^{1-\sigma}$ is strictly decreasing in p . If the constraint binds for $i \in \mathcal{B}$, then relative to the unconstrained preferential optimum (which by definition yields $r_i^{L*} > \bar{r}$), feasibility requires a higher price $\bar{p}_i > p_i^{L*}$ satisfying $r_i(\bar{p}_i) = \bar{r}$. Since

$1 - \sigma < 0$, raising p_i lowers $p_i^{1-\sigma}$, reducing the firm's contribution to $P^{1-\sigma} = \int p_k^{1-\sigma} dk$. A reduction in $P^{1-\sigma}$ (for a positive measure of firms) raises P . With competitors' prices fixed at $p_j^{H^*}$, $s_j = (p_j/P)^{1-\sigma}$ and $r_j = Es_j$ weakly increase.

Panel B of Figure D3 visualizes the core implication of Proposition 4. As the share of real scale down bunching firms in the market increases, each bunching firm must satisfy the binding true revenue cap $r_i(p_i) = \bar{r}$, which under CES demand requires setting a higher price (or equivalently restricting quantities) relative to the unconstrained preferential optimum. This real contraction lowers bunchers' contribution to the CES aggregator $\int p_i^{1-\sigma} di$ (i.e., to $P^{1-\sigma}$), so the equilibrium price index $P(b)$ rises. With compliant firms' prices fixed at the ordinary regime level, a higher $P(b)$ makes compliant varieties relatively cheaper, shifting expenditure toward them. The graph shows that revenue per compliant firm increases with b , while the compliant group's aggregate revenue share declines mechanically because the mass of compliant firms shrinks.

Propositions 3 and 4 imply that observing revenue losses for compliant competitors supports evasion-driven bunching as the operative mechanism, because real scale-down does not predict competitor losses.

Prediction 3 (Compliant firms lose revenue under evasion-induced bunching). *If bunching reflects evasion-induced access to τ^L , then compliant competitors experience market share and revenue losses as the mass of bunchers rises, as implied by Proposition 3.*

Prediction 4 (The price index falls under evasion-induced bunching). *If bunching reflects evasion-induced access to τ^L , then the market-level CES price index decreases with the share of bunchers, because a larger mass of firms price at $p^{L^*} < p^{H^*}$ (Proposition 1).*

D5. Welfare Effects

This section discusses the welfare implications through the channel highlighted by our empirical evidence: evasion-induced bunching at the eligibility cutoff. We study welfare in terms of how this form of tax evasion affects (i) consumers via the CES price index, (ii) incumbent firms via reallocation of revenues and resource costs of misreporting, and (iii) the government via foregone tax receipts. Importantly, this is not a full welfare evaluation of preferential tax regimes: we abstract from other potential margins such as entry and exit, variety creation, employment and job creation, occupational choice, and long-run productivity responses. The welfare discussion should thus be interpreted as a partial equilibrium analysis of the bunching-induced externalities of preferential taxation.

We express social welfare as

$$(D17) \quad W = \underbrace{U(P, E)}_{W_C} + \underbrace{\int \pi_i di}_{W_F} + \underbrace{\lambda \int \tau_i \hat{r}_i di}_{W_G},$$

where W_C captures consumer welfare as a function of the price index P and aggregate expenditure E , W_F denotes aggregate firm profits, and W_G is the government component valued at the marginal cost of public funds, $\lambda \geq 0$.⁷

Consumers. Using Proposition 1 and the CES price index, the equilibrium price index can be written as

$$(D18) \quad P = \left(\int p_i^{1-\sigma} di \right)^{\frac{1}{1-\sigma}} = \frac{\sigma}{\sigma-1} w \left(\int (1-\tau_i)^{\sigma-1} di \right)^{-\frac{1}{\sigma-1}}.$$

To see this, note that Proposition 1 implies

$$p_i^{1-\sigma} = \left(\frac{\sigma}{\sigma-1} w \right)^{1-\sigma} (1-\tau_i)^{\sigma-1}.$$

When firms switch from τ^H to τ^L without any change in technology (the defining feature of evasion-induced bunching), their optimal price falls mechanically by Proposition 1, so $p_i^{L*} < p_i^{H*}$. Since $1-\sigma < 0$, this raises $P^{1-\sigma} = \int p_i^{1-\sigma} di$ and therefore lowers P . Holding nominal expenditure E fixed, the fall in the CES price index raises real expenditure, E/P , and therefore increases the consumer component of welfare. Panel A of Figure D3 illustrates this mechanism: as the share of evasion-driven bunchers rises, $P(b)$ falls.

Firms. Under CES demand, evaluating (D7) at the optimal constant markup price yields operating profits

$$(D19) \quad \pi_i = \frac{1}{\sigma} (1-\tau_i) r_i,$$

To account for non-tax benefits and evasion costs, define the firm's private payoff $v_i \equiv \pi_i + B_i - \kappa_i(e_i)$.

For a firm that switches into the preferential regime via misreporting, the payoff change relative to remaining in the ordinary regime is

$$(D20) \quad \Delta v_i = \frac{1}{\sigma} \left[(1-\tau^L) r_i^L - (1-\tau^H) r_i^H \right] + B_i - \kappa_i(e_i).$$

Since firms select into evasion-induced bunching only when it weakly dominates both staying in the ordinary regime and qualifying via real scale-down (condition $\Delta V_i^E \geq \max\{0, \Delta V_i^R\}$), switchers satisfy $\Delta v_i \geq 0$ by revealed preference.

For a compliant competitor j that remains under τ^H , the wedge is unchanged and the effect operates through reallocation of demand via the price index. Under evasion-

⁷When $\lambda = 1$, this is equivalent to assuming that each euro of tax revenue is returned one-for-one to households (or firms) as a lump-sum rebate, so that public funds enter welfare as private income.

induced bunching, Proposition 3 implies $\Delta r_j < 0$, hence

$$(D21) \quad \Delta \pi_j = \frac{1}{\sigma} (1 - \tau^H) \Delta r_j < 0.$$

With E fixed (so $\int r_i di = E$), an increase in evasion-induced bunching that moves a set $\mathcal{B}$ from τ^H to τ^L changes aggregate operating profits only because some revenue is now earned under the lower wedge:

$$(D22) \quad \int \Delta \pi_i di = \frac{1}{\sigma} (\tau^H - \tau^L) \int_{\mathcal{B}} r_i^L di.$$

Adding non-tax benefits and evasion costs yields the change in aggregate private firm payoffs,

$$(D23) \quad \int \Delta v_i di = \frac{1}{\sigma} (\tau^H - \tau^L) \int_{\mathcal{B}} r_i^L di + \int_{\mathcal{B}} B_i di - \int_{\mathcal{B}} \kappa_i(e_i) di.$$

Thus, the firm-side private surplus rises if and only if the wedge-gap gain plus non-tax benefits dominate real evasion resource costs. We treat $\kappa_i(e_i)$ as a real resource cost that lowers social welfare,⁸ and B_i as a real resource savings. If, instead, they were a pure transfer, they would affect the distribution of payoffs across agents but not total surplus.

Government. Government receipts are taxes on *reported* revenue $T \equiv \int \tau_i \hat{r}_i di$. An increase in evasion-induced bunching lowers T through two channels.

First, there is a direct revenue loss from switching to the preferential tax regime and misreporting. Each bunching firm $i \in \mathcal{S}$ moves from paying τ^H on truthful revenue r_i^H to paying τ^L on the cutoff $\hat{r}_i = \bar{r}$. The per-firm change in tax payments is

$$(D24) \quad \Delta T_i^{\text{direct}} = \tau^L \bar{r} - \tau^H r_i^H = - \underbrace{(\tau^H - \tau^L) \bar{r}}_{\text{tax rate difference at the cutoff}} - \underbrace{\tau^H (r_i^H - \bar{r})}_{\text{tax base removed from the ordinary regime}} < 0.$$

Second, there is the revenue loss from competitors. By Proposition 3, compliant competitors experience $\Delta r_j < 0$ and their tax payments fall by:

$$(D25) \quad \Delta T^{\text{spill}} = \int_{j \in \mathcal{C}} \tau^H \Delta r_j dj < 0.$$

⁸Following Chetty (2009), firms typically incur private costs when reducing reported revenue (e.g., expected fines or penalties). However, these private costs need not be social costs: fines are transfers to the government, and some other penalties are transfers to third parties. In such cases, sheltering changes the distribution of surplus but does not reduce total output, and hence does not mechanically generate dead-weight loss. By contrast, the component of sheltering that uses *real resources* (e.g., time and inputs spent on concealment, operating in cash rather than electronic payments, or other distortions that lower productivity or output) is an efficiency cost and should enter social welfare with a negative sign. Therefore, we interpret $\kappa_i(e_i)$ as capturing the part of evasion that shrinks the economic pie.

This is the fiscal analogue of the competitor profit loss in (D21).

Overall, evasion-induced bunching delivers an unambiguously negative effect on the public sector. In the welfare aggregator (D17), these effects are amplified with a high marginal value of public funds.

Welfare as a function of the bunching share. We now propose a simple exercise that translates the welfare decomposition in (D17) into a quantitative mapping from the share of evasion-induced bunchers in a market, $b \in [0, 1]$, to consumer, firm, and government components of welfare.

Under Proposition 1 and normalizing $w = 1$, $p_L = \frac{\sigma}{\sigma-1} \frac{1}{1-\tau^L}$ and $p_H = \frac{\sigma}{\sigma-1} \frac{1}{1-\tau^H}$. With a share b of varieties priced at p_L and $1 - b$ priced at p_H , the CES price index is

$$(D26) \quad P(b) = \left[b p_L^{1-\sigma} + (1-b) p_H^{1-\sigma} \right]^{\frac{1}{1-\sigma}}.$$

Holding nominal expenditure E fixed, the consumer component is proportional to real expenditure. In log terms, we use the money-metric index $W_C(b) \equiv \log\left(\frac{E}{P(b)}\right)$, and report changes relative to $b = 0$.

Under CES demand, per-variety revenue is $r_k(b) = E (p_k/P(b))^{1-\sigma}$ for $k \in \{L, H\}$. Aggregate revenues by group are $R_B(b) = b r_L(b)$, and $R_C(b) = (1-b) r_H(b)$, with $R_B(b) + R_C(b) = E$ by construction. Aggregate operating profits are

$$(D27) \quad W_F(b) \equiv \int \pi_i di = \frac{1}{\sigma} \left[(1-\tau^L) R_B(b) + (1-\tau^H) R_C(b) \right].$$

To incorporate non-tax benefits and evasion costs in the firm component, we assume average per-buncher mass real benefit $\bar{B}$ and real evasion cost $\bar{\kappa}$, so the firm component becomes

$$\tilde{W}_F(b) = W_F(b) + b \bar{B} - b \bar{\kappa}.$$

We assume that compliant firms report truthfully, $\hat{r}_H = r_H$, while bunchers report a reporting rate $\mu \in (0, 1]$ such that $\hat{r}_L = \mu r_L$. Tax receipts are then

$$(D28) \quad T(b) = \tau^L \mu R_B(b) + \tau^H R_C(b),$$

and the government component of welfare is $W_G(b) \equiv \lambda T(b)$.

We calculate the welfare change relative to a counterfactual with no bunching ($b = 0$):

$$\Delta W(b) = [W_C(b) - W_C(0)] + [\tilde{W}_F(b) - \tilde{W}_F(0)] + [W_G(b) - W_G(0)].$$

A minimal calibration requires $(\sigma, \tau^L, \tau^H, \lambda, \mu, \bar{B}, \bar{\kappa})$ and a normalization for E . Table D1 reports the benchmark calibration.

Figure D4 plots the welfare impact as a function of the market share of bunching firms b . The horizontal axis reports b , while the vertical axis reports the change in welfare rela-

Table D1: Benchmark Calibration

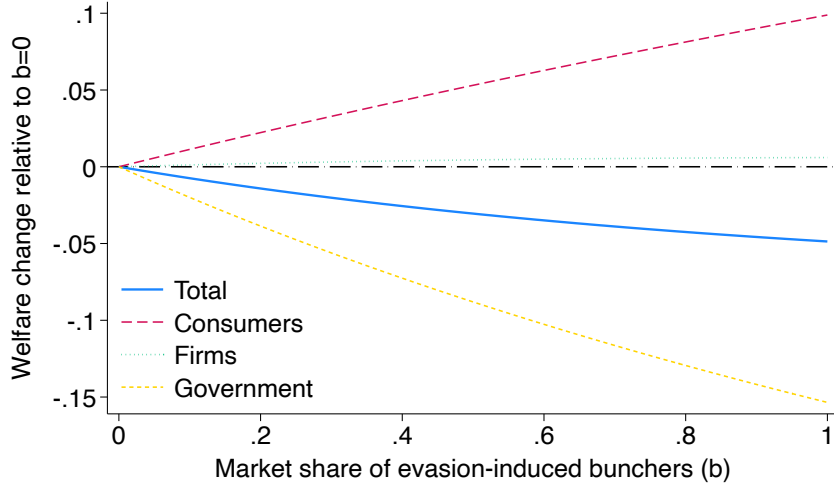
Parameter	Symbol	Baseline	Justification or source
Elasticity of substitution	σ	4	Within the range typically estimated in empirical work on product substitution and trade elasticities. For instance, Broda and Weinstein (2006) find that the average σ was around 4 among three-digit goods in the US over the 1990-2001 period. We perturb σ over a wide range [2–8] as a robustness check.
Preferential tax rate	τ^L	0.15	Institutional setting.
Ordinary tax rate	τ^H	0.23	Institutional setting.
Aggregate expenditure	E	1	Normalization: welfare is expressed in units of aggregate expenditure.
Reporting rate for bunchers	μ	0.51	Our z_u estimate suggests that, in the absence of the notch, the marginal buncher would have reported approximately €15,000 more in revenue (i.e., around 51 percent of their reported revenue); see Figure 3 in the paper.
Marginal cost of public funds	λ	1	We perturb λ in the range [0–3] to assess sensitivity to alternative values.
Lower tax compliance costs	$\bar{B}$	0.026	Tax compliance costs account for 2.6% of turnover for micro-sized enterprises according to a survey from European Commission (2022); we perturb $\bar{B}$ in the [0–0.1] range as a robustness.
Real resource costs of evasion	$\bar{\kappa}$	0.04	Estimated scaling factor of the evasion cost function from Aghion, Gravouille, Lequien and Stantcheva (2024) is 0.11: we transformed in share of expenditure; we perturb $\bar{\kappa}$ in the [0–0.1] range as a robustness.

Notes: Baseline values used to construct $\Delta W(b)$ and its decomposition into consumers, firms, and government as the share of evasion-induced bunchers b varies.

tive to the no-bunching benchmark $b = 0$. The solid blue line shows total welfare, $\Delta W(b)$. The dashed red line reports the consumer component, measured by the change in real expenditure $\Delta \log(E/P(b))$ holding E fixed, where the price index $P(b)$ is constructed from Proposition 1 by assuming that a mass b of varieties is priced at the preferential regime optimum p^{L^*} and a mass $1 - b$ at the ordinary regime optimum p^{H^*} . The dotted green line reports the firm component, which combines changes in aggregate operating profits and include non-tax benefit $\bar{B}$ and the real evasion resource cost $\bar{\kappa}$. The dashed yellow line reports the government component, $\lambda \Delta T(b)$, where tax receipts depend on reported revenue.

Two main patterns emerge. First, the consumer component is positive and increasing in b : as more varieties are priced at $p^{L^*} < p^{H^*}$, the CES price index falls and real

Figure D4: Welfare Effects of Evasion-Induced Bunching



Notes: The figure reports model-implied welfare changes from evasion-induced bunching as a function of the market share of bunching firms $b \in [0,1]$. The vertical axis plots changes relative to the no-bunching benchmark $b = 0$. The *Consumers* line is $\Delta W_C(b) = \log(E/P(b)) - \log(E/P(0))$, where $P(b) = [b p_L^{1-\sigma} + (1-b) p_H^{1-\sigma}]^{1/(1-\sigma)}$ is the CES price index implied by Proposition 1 with $p_L = \frac{\sigma}{\sigma-1} \frac{1}{1-\tau^L}$ and $p_H = \frac{\sigma}{\sigma-1} \frac{1}{1-\tau^H}$, normalizing $w = 1$ and $E = 1$. The *Firms* line is the change in aggregate private firm payoffs, combining operating profits and reduced-form non-tax benefits and evasion costs: $W_F(b) = \frac{1}{\sigma} [(1-\tau^L)R_B(b) + (1-\tau^H)R_C(b)] + b\bar{B} - b\bar{\kappa}$, with $R_B(b) = b E (p_L/P(b))^{1-\sigma}$ and $R_C(b) = (1-b) E (p_H/P(b))^{1-\sigma}$. The *Government* line is $\Delta W_G(b) = \lambda [T(b) - T(0)]$ with tax receipts $T(b) = \tau^L \mu R_B(b) + \tau^H \mu R_C(b)$, assuming compliant firms report truthfully and bunchers report a fraction μ of true revenue. The *Total* line is $\Delta W(b) = \Delta W_C(b) + [W_F(b) - W_F(0)] + \Delta W_G(b)$. Baseline calibration: $\sigma = 4$, $\tau^L = 0.15$, $\tau^H = 0.23$, $\lambda = 1$, $\mu = 0.51$, $\bar{B} = 0.026$, and $\bar{\kappa} = 0.04$.

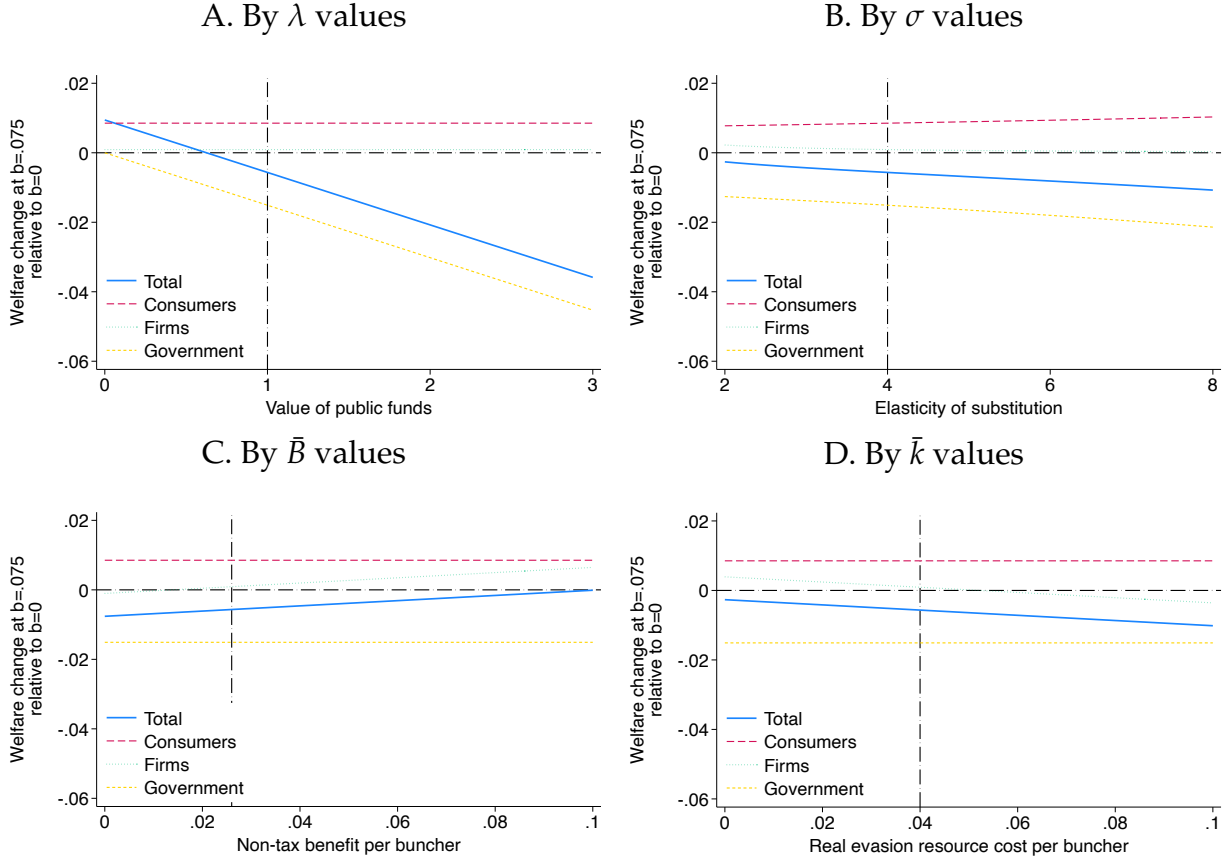
consumption E/P rises. Second, the firm component is small and mildly positive in the benchmark calibration: reallocation of demand toward low-wedge firms mechanically raises aggregate operating profits. Third, the government component is strongly negative and decreasing in b : bunchers both face a lower wedge and reduce the reported tax base through underreporting.

The total welfare effect is negative. Evaluated at the bunching rate observed in our data ($b = 7.5\%$), the calibration underlying the figure implies $\Delta W_C(0.075) \approx +0.9\%$, $\Delta W_F(0.075) \approx +0.1\%$, and $\Delta W_G(0.075) \approx -1.5\%$, for a total effect of $\Delta W(0.075) \approx -0.6\%$ relative to $b = 0$ (in units consistent with the normalization $E = 1$). This partial equilibrium analysis thus suggests that the observed bunching level reduces total social welfare by about 0.6%.

Robustness in welfare effects of evasion-induced bunching. Figure D5 assesses the sensitivity of the welfare decomposition to key calibration parameters. In each panel, we hold the bunching share fixed at $b = 0.075$, corresponding to the baseline bunching rate in our data, and plot the change in welfare relative to the no-bunching benchmark $b = 0$. Each panel varies one of the calibrated parameters at a time, holding the remaining

parameters at their benchmark values. In each graph, the vertical dashed line refers to the calibrated baseline value.

Figure D5: Robustness in Welfare Effects of Evasion-Induced Bunching



Notes: The figure reports robustness of the calibrated welfare effects of evasion-induced bunching. In each panel, we fix the market share of bunchers at $b = 0.075$ (the baseline bunching rate in the data) and plot welfare changes relative to the no-bunching benchmark $b = 0$. The solid line reports total welfare, $\Delta W = \Delta W_C + \Delta W_F + \Delta W_G$, while the dashed/dotted lines report the consumer, firm, and government components, respectively, computed from the quantitative exercise in Section D5. The consumer component is $\Delta W_C = \log(E/P(b)) - \log(E/P(0))$, where $P(b) = [b p_L^{1-\sigma} + (1-b) p_H^{1-\sigma}]^{1/(1-\sigma)}$ and $p_L = \frac{\sigma}{\sigma-1} \frac{1}{1-\tau^L}$, $p_H = \frac{\sigma}{\sigma-1} \frac{1}{1-\tau^H}$, normalizing $w/\varphi = 1$ and $E = 1$. The firm component is the change in aggregate private firm payoffs, $W_F(b) = \frac{1}{\sigma} [(1-\tau^L)R_B(b) + (1-\tau^H)R_C(b)] + b\bar{B} - b\bar{k}$, with $R_B(b) = bE(p_L/P(b))^{1-\sigma}$ and $R_C(b) = (1-b)E(p_H/P(b))^{1-\sigma}$. The government component is $\Delta W_G = \lambda [T(b) - T(0)]$, where tax receipts are $T(b) = \tau^L \mu R_B(b) + \tau^H R_C(b)$ assuming compliant firms report truthfully and bunchers report a fraction μ of true revenue. Each panel varies one parameter at a time (A: $\lambda \in [0, 3]$; B: $\sigma \in [2, 8]$; C: $\bar{B} \in [0, 0.10]$; D: $\bar{k} \in [0, 0.10]$), holding all others at the benchmark calibration $\tau^L = 0.15$, $\tau^H = 0.23$, $\mu = 0.51$, $E = 1$, $\lambda = 1$, $\sigma = 4$, $\bar{B} = 0.026$, and $\bar{k} = 0.04$ (Table D1).

Panel A varies λ . Since λ only scales the government term in (D17), the consumer and firm components are invariant to λ , while the government component becomes mechanically more negative as λ increases.

Panel B varies σ . As σ increases, substitution across varieties becomes more elastic, strengthening the competitive channel through which low-wedge firms gain market share and the price index responds. In the figure, this manifests as a modestly larger consumer gain (more responsive $P(b)$) alongside a more negative government component (because the tax base shifts more strongly toward underreporting bunchers and away

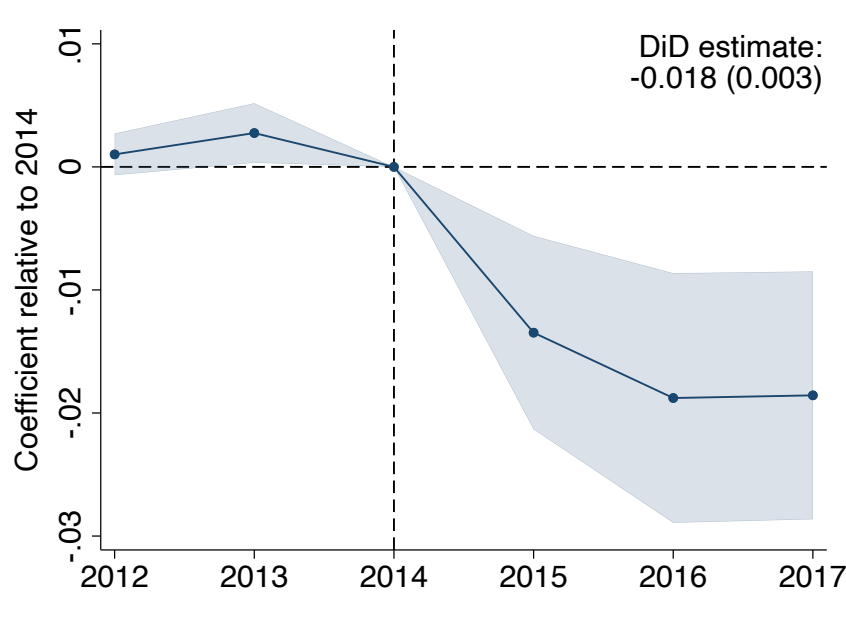
from compliant firms). The net effect is that total welfare becomes (weakly) more negative with higher σ over the considered range.

Panel C and D vary $\bar{B}$ and $\bar{\kappa}$. By construction, these parameters affect only the firm component. As $\bar{B}$ rises, the firm line shifts upward, and total welfare increases. Nevertheless, for the plotted range, the total welfare effect remains negative because the fiscal loss continues to dominate. Increasing evasion resource costs makes the welfare impact of bunching more negative. For sufficiently large values of $\bar{\kappa}$, even the firm component turns negative: the real resources dissipated in concealment more than offset the mechanical operating profit gain from earning revenue under the lower tax wedge.

E. Robustness Checks and Alternative Specifications

E1. Effect of the 2015 Reform on the Average Effective Tax Rate

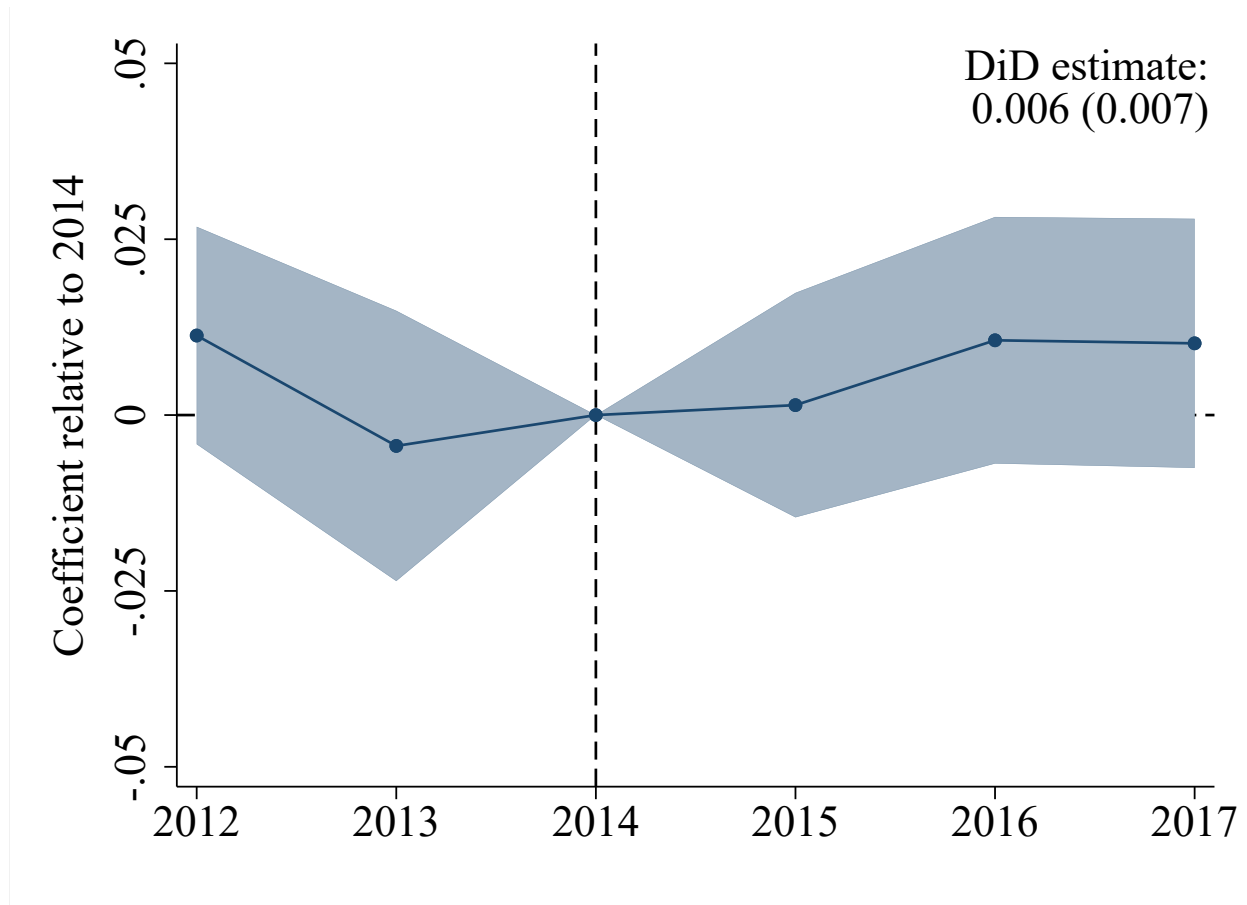
Figure E1: Effect of the 2015 Reform on the Average Effective Tax Rate



Notes: The figure depicts the event-study coefficient estimates and 95 percent confidence intervals from market-level standard errors obtained. The outcome variable is the market-level average effective tax rate. The event-study is based on the 2015 reform, which created a tax notch for firms older than 5 years, but kept the notch for other firms unaffected. Therefore, it creates a differential exposure across markets depending on the fraction of individual firms older than 5 years. We classify markets as more (less) exposed if in the top (bottom) tercile of the market share of individual firms older than 5 years. The specification includes market and year fixed effects. The figure also reports the coefficient estimate and its associated standard error from a specification of the average effective tax rate on the interaction between a dummy for more exposed markets and the post-2015 dummy. Estimates are weighted by the average number of firms in each market.

E2. Event-Study Evidence on Larger-Size Competitor Firms

Figure E2: Event-Study Evidence on Larger-Size Competitor Firms



Note: The figure depicts the event-study coefficient estimates and 95 percent confidence intervals from market-level standard errors obtained from equation (5). The outcome variable is the log of the revenue of competitor firms that are larger in size compared to our baseline group. The event-study is based on the 2015 reform, which created a tax notch for firms older than 5 years, but kept the notch for other firms unaffected. Therefore, it creates a differential exposure across markets depending on the fraction of individual firms older than 5 years. We classify markets as more (less) exposed if in the top (bottom) tercile of the market share of individual firms older than 5 years. Each specification includes firm and year fixed effects. The figure also reports the coefficient estimate and its associated standard error from a specification of the outcome of interest on the interaction between a dummy for more exposed markets and the post-2015 dummy.

E3. OLS estimate

Table E1: OLS Effects of Bunching Firms' Market Share on Competitors' Revenue by Size

	Outcome: log(revenue)				
	Sample: firms with revenue				
	<100k (1)	100-250k (2)	250-500k (3)	500-1,000k (4)	>1,000k (5)
$B_{m,t}$	0.00034 (0.000)	0.00081 (0.000)	0.00132 (0.000)	0.00077 (0.000)	0.00030 (0.000)
Observations	8,751,859	5,485,722	3,273,322	2,320,568	3,880,411
Firm FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes

Note: This table reports OLS estimates of the effect of the market share of individual bunching firms on the log revenue of their competitors, disaggregated by competitor firm size. Column (1) includes competitor firms with average revenue below €100,000; column (2) includes those with revenue between €100,000 and €250,000; column (3) covers firms reporting between €250,000 and €500,000; column (4) includes those between €500,000 and €1,000,000; and column (5) focuses on firms with revenue exceeding €1,000,000. All specifications control for firm and year fixed effects. Markets are defined at the municipality–industry–year level. Individual bunching firms are defined as those reporting revenue within €5,000 below the eligibility threshold for preferential tax regimes. Competitors are non-individual firms operating in the same market and reporting less than €500,000 revenue. Standard errors, clustered at the market level, are reported in parentheses.

E4. The Instrument Does Not Predict Firm Exit

Table E2: The Instrument Does Not Predict Firm Exit

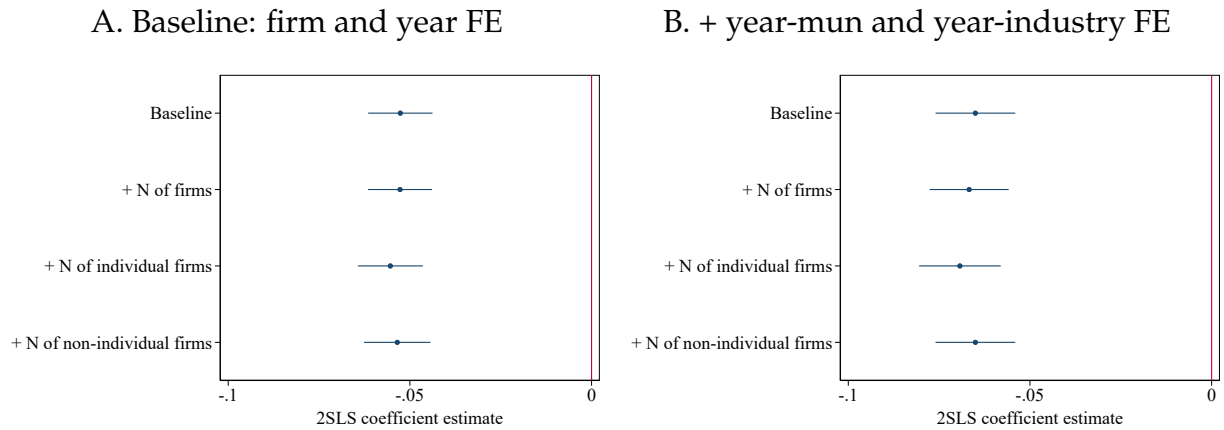
	Outcome: log(N of exiting firms)
$Z_{m,t}$	-0.020 (0.076)
Observations	430,722
Market FE	Yes
Year FE	Yes

Notes: The unit of observation is a market–year, defined as municipality $\times$ industry $\times$ year. The dependent variable is the log of the number of individual firms eligible for preferential tax regimes exiting the market. The specification includes market fixed effects and year fixed effects. Standard errors are clustered at the market level.

E5. Sensitivity to Additional Controls

In this section, we test the sensitivity of our estimates to additional time-varying factors. Although our baseline model relies on a very demanding set of fixed effects - firm, industry-year, and municipality-year fixed effects - the exclusion restriction of our 2SLS model might be threatened by omitted changes in competitors' firm outcomes due to market-specific time-varying channels unrelated to individual firms' revenue manipulation. For instance, changes in the generosity of the preferential tax schemes might stimulate business creation. If this is the case, then competitor firms' outcomes would be influenced by tougher (fair) competition, rather than through (unfair) competition driven by tax evasion. To show that this is not the source of our findings, we check for significant changes in our coefficient estimate when conditioning on the number of firms operating in a market in any given year. As suggested in Altonji, Elder and Taber (2005) and Oster (2019), significant changes in coefficient estimates imply the potential importance of unobserved confounders. In Figure E3, we provide reassuring evidence that our coefficient estimate remains remarkably stable when controlling for time-varying market-specific changes in i) the number of firms; ii) the number of individual firms; iii) the number of non-individual firms. Similar results are obtained when we replicate the analysis on other firm outcomes (results available upon request).

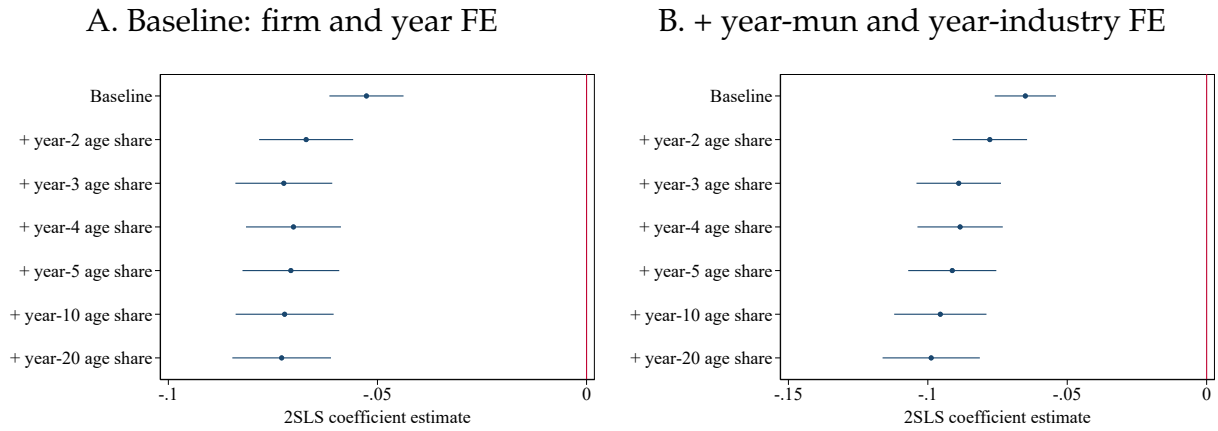
Figure E3: Sensitivity to Additional Controls



Notes: This figure tests the robustness of our baseline 2SLS estimates on the log of firm revenue, computed from a model with firm fixed effects and year fixed effects (“Baseline” in the left-hand side graph); and with firm, year-municipality, and year-industry fixed effects (“Baseline” in the right-hand side graph). Each figure reports the 2SLS coefficient estimate and 95 percent confidence intervals from market-level clustered standard errors. On top of the baseline controls, we add the following controls: i. number of firms in the market (second row); ii. number of individual firms in the market (third row); number of non-individual firms in the market (fourth row). Market-level bunching rate is calculated as the fraction of individual firms that report revenue just below the cutoff determining eligibility for preferential tax schemes. Individual bunching firms are those reporting revenue between -€5,000 and 0 relative to the eligibility cutoff for preferential regimes, while competitors are non-individual firms operating in the same market and reporting less than and reporting less than €500,000 revenue. The instrument, described in equation (4), is based on policy changes and a market’s (fixed) demographic composition. Each market is defined at the municipality-industry-year level.

E6. Robustness for “Incomplete” Shift-Share Design

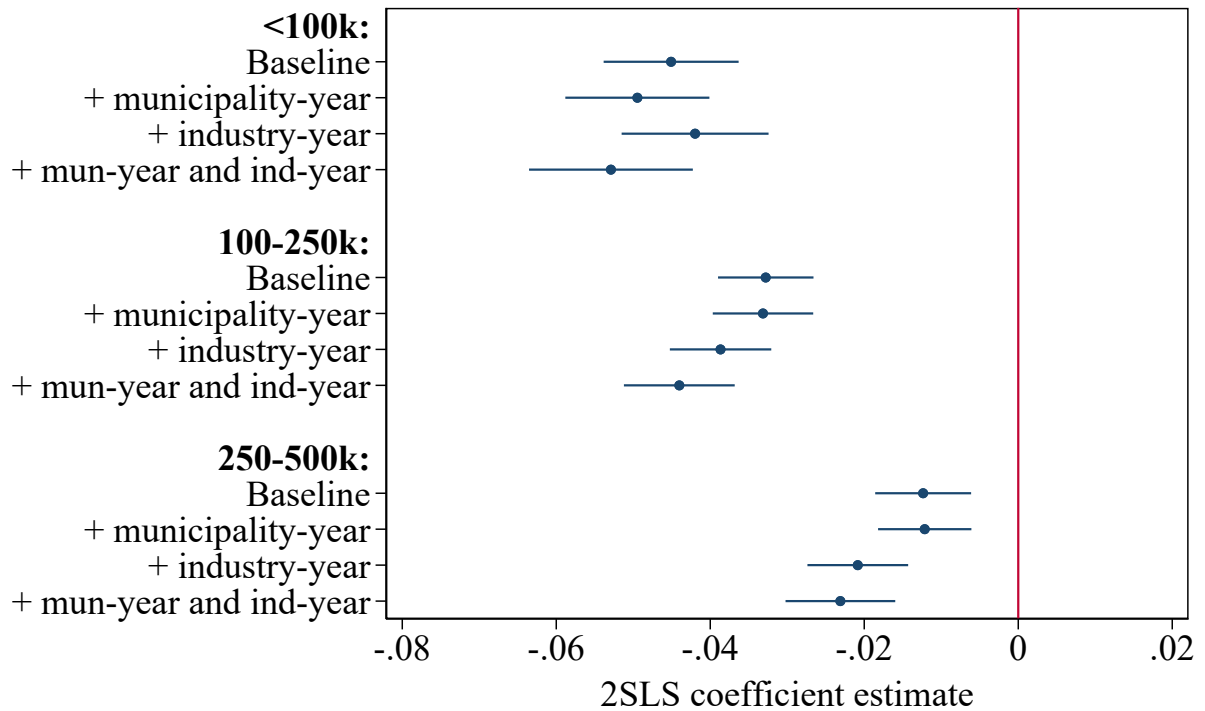
Figure E4: Robustness for “Incomplete” Shift-Share Designs



Notes: This figure tests the robustness of our baseline 2SLS estimates on the log of firm revenue, computed from a model with firm fixed effects and year fixed effects (“Baseline” in the left-hand side graph); and with firm, year-municipality, and year-industry fixed effects (“Baseline” in the right-hand side graph). Each figure reports the 2SLS coefficient estimate and 95 percent confidence intervals from market-level clustered standard errors. On top of the baseline controls, we control for groups (2, 3, 4, 5, 10, or 20) of market shares interacted with year fixed effects. The market-level bunching rate is calculated as the fraction of individual firms that report revenue just below the cutoff determining eligibility for preferential tax schemes. Individual bunching firms are those reporting revenue between -€5,000 and 0 relative to the eligibility cutoff for preferential regimes, while competitors are non-individual firms operating in the same market and reporting less than €500,000 revenue. The instrument, described in equation (4), is based on policy changes and a market’s (fixed) demographic composition. Each market is defined at the municipality-industry-year level.

E7. Alternative Specifications

Figure E5: Alternative Specifications



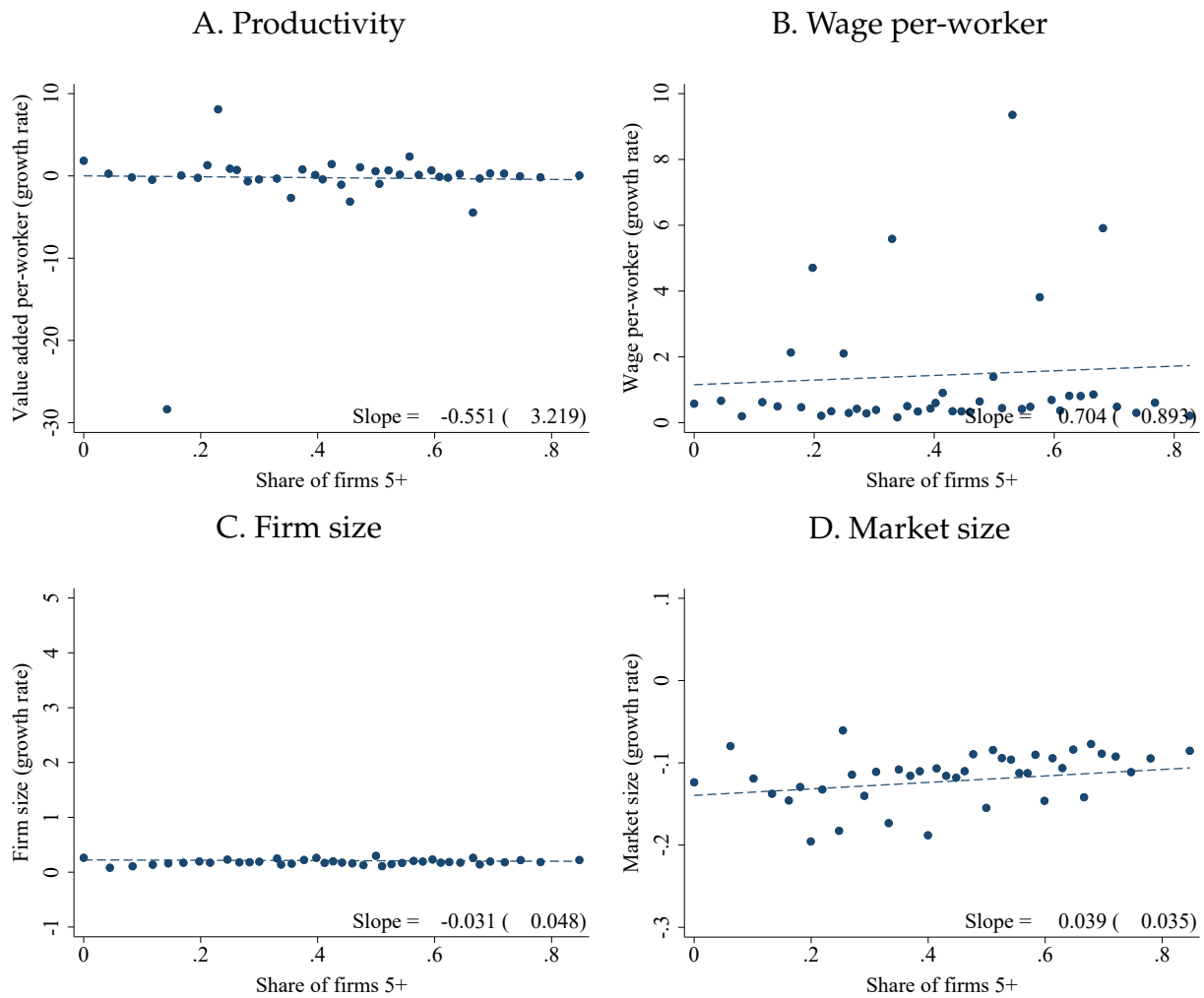
Notes: This figure presents 2SLS estimates and 95 percent confidence intervals of how the market share of individual bunching firms affects their competitors' log revenue by firm size group. In each group, the first row reports estimates from the baseline model with firm and year fixed effects; the second row includes firm and year-industry fixed effects; the third row incorporates firm and year-municipality fixed effects; and the fourth row accounts for firm, industry-year, and municipality-year fixed effects. The instrument, detailed in equation (4), is constructed using policy changes and a market's fixed demographic composition. Markets are defined at the municipality-industry-year level. Individual bunching firms are those reporting revenue between -€5,000 and 0 relative to the eligibility cutoff for preferential regimes, while competitors are non-individual firms operating in the same market.

E8. The Correlates of Market Shares

Our research design reflects differential exogenous exposure, due to the (fixed) market demographic composition, to common policy shocks. It rests on the assumption that *changes* (rather than *levels*) of firms' revenue are similar across markets with a different share of young (or old) firms.⁹ The central identification concern is thus that the market shares predict firm outcomes through channels other than those due to policy changes in the incentive to bunch. For example, markets with high versus low exposure may have features that predict change in the outcome through channels other than the bunching rate, violating the exclusion restriction. As recommended by Goldsmith-Pinkham, Sorkin and Swift (2020), one way to assess this possibility is to look at the correlates of the shares. If these correlates suggest other channels through which the shares affect outcomes, then we might be skeptical of the identifying assumption. Guided by both theoretical and empirical evidence on firm growth predictors (see, e.g., Kumar, Rajan and Zingales 1999; Luttmer 2011), Figure E6 plots the relationship between the share of firms older than 5 in a market and four predictors of firm growth: firm productivity (value added per worker); wage per worker; firm size; market size (total number of firms). We find no significant association, validating our identifying assumption.

⁹As emphasized in Goldsmith-Pinkham, Sorkin and Swift (2020), "In particular, the key question researchers should have in mind is whether the correlates of the levels of the shares predict *changes* in the outcome. For the empirical strategy to be valid, it is fine if the level of the correlates is related to the level of the outcome." (p. 2605).

Figure E6: The Correlates of Market Age Shares



Notes: This figure shows scatter-plots comparing various predictors of firm growth (vertical axes) on the market share of firms created from more than 5 years (horizontal axes). Panel A focuses on firm productivity (value added per worker); panel B on the wage per worker; panel C on firm size; panel D on market size (total number of firms). Each graph compares the growth rate in each predictor (over the 2005-2019 period) versus the market share observed in 2005. We plot 50 equal-sized bins and show the line of best fit. Each graph also reports the estimated slope and the associated market-level clustered standard error from regressions on the first-difference of each predictor variable on the market share of firms created from more than 5 years.

E9. Sensitivity to Alternative Market Definitions

Our empirical strategy depends on the market definition. In this section, we verify the robustness of our results after perturbation of the boundaries of the market. We modify our baseline market definition along two dimensions. First, we use a finer industry classification: the NACE 3-digit level. Second, we move to a wider geographical unit of analysis: from municipality to the local labor market (LLM). Finally, we focus on a sample of firms operating in markets with a high level of localism. In this case, market boundaries are more naturally delineated, and our estimates should be less sensitive to attenuation bias due to measurement errors in market definition. Highly localized markets are defined as those where the share of firms identifying competitors within the same municipality is above the median, based on data from ISTAT, Censimento Industrie e Servizi. Table E3 shows that our baseline estimates remain substantially similar.

Table E3: Sensitivity to Alternative Market Definitions

	<i>Outcome: log(revenue)</i>			
	(1)	(2)	(3)	(4)
<i>A. Market definition: municipality-NACE 3-digit</i>				
$B_{m,t}$	-0.019 (0.003)	-0.016 (0.003)	-0.020 (0.003)	-0.019 (0.004)
<i>B. Market definition: LLM-NACE 2-digit</i>				
$B_{m,t}$	-0.059 (0.004)	-0.065 (0.005)	-0.065 (0.005)	-0.076 (0.005)
<i>C. Market definition: LLM-NACE 3-digit</i>				
$B_{m,t}$	-0.027 (0.003)	-0.027 (0.003)	-0.028 (0.003)	-0.030 (0.003)
<i>D. Market definition: Highly local markets</i>				
$B_{m,t}$	-0.067 (0.003)	-0.081 (0.003)	-0.069 (0.003)	-0.096 (0.003)
Observations (panels A-C)	17,510,903			
Observations (panel D)	12,462,841			
Firm FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Year-industry FE	No	Yes	No	Yes
Year-municipality FE	No	No	Yes	Yes

Note: This table presents 2SLS estimates of how the market share of individual bunching firms affects their competitors' log revenue. We test the sensitivity to different market definitions: i. municipality and 3-digit industry code (panel A); ii. local labor market and 2-digit industry code (panel B); iii. local labor market and 3-digit industry code (panel C); iv. markets with a high level of localism (panel D), defined as those where the share of firms identifying competitors within the same municipality is above the median. Column 1 reports estimates from a model with firm and year fixed effects; column 2 includes firm and year-industry fixed effects; column 3 incorporates firm and year-municipality fixed effects; and column 4 accounts for firm, industry-year, and municipality-year fixed effects. The instrument, detailed in equation (4), is constructed using policy changes and a market's fixed demographic composition. Individual bunching firms are those reporting revenue between -€5,000 and 0 relative to the eligibility cutoff for preferential regimes, while competitors are non-individual firms operating in the same market and reporting less than €500,000 revenue. Standard errors, clustered at the market level, are shown in parentheses.

E10. Alternative Measures of the Bunching Rate

Table E4 shows that our estimates remain consistent when we use alternative measures of the bunching rate. We employ four alternative definitions: i. the log of the share of bunching firms +1 (panel A); ii. the number of bunching firms (panel B); iii. the log of the number of bunching firms +1 (panel C); iv. the ratio between the number of firms below the cutoff [-€5,000;0] and those above [0; €10,000] (panel D). The latter is a variable more closely related to the “classical” measure of bunching. Although the interpretation of the first-stage effect varies depending on the measure used to quantify the baseline bunching rate, we observe that regardless of the method used to measure bunching, it has a negative impact on the outcome variables of interest.

Table E4: Sensitivity to Alternative Bunching Definitions

	<i>Outcome: log(revenue)</i>			
	(1)	(2)	(3)	(4)
<i>A. Bunching measure: log of share of bunching firms</i>				
$B_{m,t}$	-0.681 (0.090)	-0.526 (0.055)	-0.680 (0.080)	-0.610 (0.058)
<i>B. Bunching measure: N of bunching firms</i>				
$B_{m,t}$	-0.003 (0.001)	-0.002 (0.001)	-0.002 (0.001)	-0.002 (0.001)
<i>C. Bunching measure: log of N of bunching firms</i>				
$B_{m,t}$	-0.859 (0.205)	-0.455 (0.049)	-0.542 (0.071)	-0.383 (0.035)
<i>D. Bunching measure: classical bunching estimate</i>				
$B_{m,t}$	-0.646 (0.149)	-1.243 (0.339)	-0.709 (0.122)	-1.699 (0.450)
Observations	17,510,903			
Firm FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Year-industry FE	No	Yes	No	Yes
Year-municipality FE	No	No	Yes	Yes

Note: This table tests the robustness of our estimates to different bunching measures. We employ four alternative definitions: i. the log of the share of bunching firms +1 (panel A); ii. the number of bunching firms (panel B); iii. the log of the number of bunching firms +1 (panel C); iv. the ratio between the number of firms below the cutoff [-€5,000;0] and those just above [0; €5,000] (panel D). Column 1 reports estimates from a model with firm and year fixed effects; column 2 includes firm and year-industry fixed effects; column 3 incorporates firm and year-municipality fixed effects; and column 4 accounts for firm, industry-year, and municipality-year fixed effects. The instrument, detailed in equation (4), is constructed using policy changes and a market's fixed demographic composition. Competitors are non-individual firms operating in the same market and reporting less than €500,000 revenue. Standard errors, clustered at the market level, are shown in parentheses.

E11. Effect by Distance to Treated Firms

In this robustness check, we replicate our baseline externality specification, but redefine the “competitor set” to include all firms operating in the same province (or, alternatively, the same region) as the treated municipality, excluding firms located in the treated municipality itself. Intuitively, if the relevant product market is very local – as the mechanism in the paper suggests – spillovers should attenuate quickly with distance. Under this interpretation, effects on competitors in the broader province/region outside the municipality should be smaller.

Consistent with this prediction, Table E5 below shows no statistically significant effect of bunching on competitors’ revenues when the outcome is aggregated over competitors located elsewhere in the same province or region (i.e., after excluding the treated municipality). This result supports the view that the competitive externality we document is local.

Table E5: Effect by Distance to Treated Firms

	<i>Outcome: log(aggregate revenue)</i>		
	(1)	(2)	(3)
<i>Competitor: same province excluding treated muni</i>			
$B_{m,t}$	0.006 (0.004)	0.001 (0.033)	0.004 (0.010)
<i>Competitor: same region excluding treated muni</i>			
$B_{m,t}$	0.005 (0.004)	0.007 (0.034)	0.004 (0.009)
Observations	2,277,450	2,277,450	2,277,450
Municipality FE	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Mun-ind FE	No	Yes	Yes
Year-ind FE	No	Yes	Yes
Year-mun FE	No	No	Yes

Notes: This table assesses the geographic scope of competitive spillovers using placebo competitor sets defined outside the treated municipality. We redefine the outcome as the log of aggregate revenue of competitor firms located in the same province (top panel) or the same region (bottom panel) as municipality m , excluding all competitors operating in municipality m . Estimates are from a 2SLS model. If markets are highly local, spillovers should attenuate with distance, implying small and statistically insignificant effects outside the treated municipality. All specifications include municipality, industry, and year fixed effects; columns 2 and 3 add progressively richer interactions (municipality $\times$ industry, year $\times$ industry, and year $\times$ municipality) as indicated. Observations are collapsed at the market (municipality-industry-year) level.

E12. Heterogeneity by Import Exposure

We examine heterogeneity across industries with differing scope for market power and price-setting. We measure import exposure at the industry level using customs-based trade flows, constructing an import-penetration index that scales the value of imports in a given sector by the size of the sector's domestic market. Industries are then classified as "high import exposure" if their import-penetration index lies above the median in the sample. This proxy captures the extent to which domestic producers face competition from foreign suppliers and, correspondingly, the degree to which product market conditions are driven by international (instead of local) forces.

We find that spillover effects are significantly attenuated in industries with above-median exposure to imports (see Table E6 below). The attenuation in high-import sectors is economically intuitive for several reasons.

First, in tradable industries with substantial import penetration, prices are to a large extent disciplined by international competition and (directly or indirectly) by world-price benchmarks, leaving less scope for local firms to strategically adjust prices in response to cost advantages enjoyed by bunchers.

Second, sectors with high import (and typically export) exposure tend to be populated by larger firms, for which competition from small individual firms operating near the preferential regime threshold is less consequential, both because their customer base is broader and because they are less likely to operate in the same narrow set of local segments.

Third, in these tradable sectors, the relevant market is inherently wider than the municipality: firms compete in regional, national, or international product markets. As a result, any municipality-level shock to the prevalence of bunching firms represents a smaller component of the competitive environment, mechanically diluting the spillover effect relative to non-tradable sectors where competition and pricing are predominantly local.

Table E6: Heterogeneity by Import Exposure

	Outcome: log(revenue)
$B_{m,t}$	-0.064 (0.006)
$B_{m,t} \times \mathbb{1}(\text{Import exposure} > \text{median})$	0.020 (0.004)
Observations	17,510,903
Firm FE	Yes
Year FE	Yes

Notes: This table examines whether the 2SLS estimate of the effect of individual bunching firms' market share on their competitors' revenue varies by industry-level (2-digit NACE) import exposure. Each specification includes firm and year fixed effects. Markets are defined at the municipality-industry-year level. Individual bunching firms are those reporting revenue between -5,000 and 0 relative to the eligibility threshold for preferential regimes; competitors are non-individual firms operating in the same market and reporting less than €500,000 revenue. Standard errors clustered at the market level are reported in parentheses.

E13. Markup Compression for Competitors

This section provides additional evidence on the mechanism through which bunching firms affect their competitors. In our baseline approach (Table 4), we use the ratio of revenue to operating costs as an accounting-based proxy for markups. Under the assumption that operating costs mainly capture variable inputs and approximate marginal costs, a decline in this ratio is consistent with stronger competitive pressure and lower price–cost margins (De Loecker and Warzynski 2012). Yet, our data do not contain firm-level output prices or physical quantities, so we cannot cleanly decompose the response into price and quantity margins. Moreover, if operating costs include quasi-fixed components that adjust only slowly in the short run, the revenue-to-cost ratio could fall even in the absence of immediate price changes.

To address this concern, we examine the composition of operating costs more closely. In our data, operating costs include purchases of intermediate inputs and services, labor costs, expenses for the use of third-party assets, and other operating expenses, while excluding financial costs, depreciation, and extraordinary items. A potential concern is that the item *use of third-party assets* may contain relatively rigid expenditures, such as rents under medium- to long-term leases, alongside more flexible items such as short-term leasing of machinery or vehicles. If this component were slow to adjust, it could mechanically amplify declines in the revenue-to-cost ratio.

Table E7 addresses this issue. Column (1) reports the baseline markup specification, where markup is defined as revenue over total operating costs. Column (2) reports an adjusted markup measure that excludes *use of third-party assets* from operating costs. The estimated effect of bunching firms on competitors' markup is very similar across the two definitions: a one percentage point increase in the market share of bunching firms reduces the baseline markup measure by 0.035 points and the adjusted markup measure by 0.038 points. The close correspondence between the two estimates suggests that the baseline result is not mechanically driven by rigid rental expenditures or other quasi-fixed components embedded in this cost category.

Table E7: Markup Compression for Competitors

	Baseline markup (1)	Adjusted markup (2)
$B_{m,t}$	-0.035 (0.008)	-0.038 (0.008)
Observations	17,510,903	17,510,903
Firm FE	Yes	Yes
Year FE	Yes	Yes

Note: This table reports 2SLS estimates of how the market share of individual bunching firms affects competitors' markup. In column (1), markup is defined as the ratio of revenue to operating costs. In column (2), we use an adjusted markup measure that excludes the item *use of third-party assets* from operating costs. This component may include quasi-fixed items such as rents under medium- to long-term contracts, alongside more flexible short-term leasing expenditures tied to production activity. All specifications include firm and year fixed effects. The instrumental variable, described in equation (4), leverages policy changes and the fixed demographic composition of each market. Markets are defined at the municipality-industry-year level. Individual bunching firms are those with reported revenue between –€5,000 and 0 relative to the eligibility threshold for preferential regimes; competitors are non-individual firms operating in the same market and reporting less than €500,000 revenue. Standard errors, clustered at the market level, are reported in parentheses.

E14. Heterogeneity by Multi-Firm Exposure

A potential concern is that part of the negative effect of bunching on non-eligible firms may reflect income shifting across commonly owned firms rather than genuine competitive spillovers. For example, an owner operating both a non-eligible company and a sole proprietorship under the preferential regime could reassign contracts or invoice sales through the lower-tax entity, mechanically lowering reported revenue in the former without any real loss of customers. If this margin were quantitatively important, our estimates would partly capture reallocation within ownership networks rather than competition across distinct firms.

Our main administrative data do not contain firm-to-owner identifiers, so we cannot directly observe ownership links and test this mechanism at the firm level. To nevertheless assess its relevance, we merge our baseline dataset with external information from the Italian Chamber of Commerce, available to the Bank of Italy, reporting at the individual level the number of firms owned by each physical person. Using these data, we construct, for each province $\times$ NACE 2-digit cell, the average number of firms per owner. We then define an indicator $1\{MultiFirm > \text{median}\}$, equal to one for cells above the median of this ownership-prevalence measure. We work at the province level, rather than the municipality level, because the latter provides too little variation in the ownership measure.

The intuition is straightforward. If within-owner shifting is an important driver of our main results, the effect of bunching on non-eligible firms should be more negative in places where multi-firm ownership is more prevalent, since these are the cells in which owners have greater scope to reallocate revenue, contracts, or invoicing across legal entities within the same ownership network. We test this by augmenting our baseline specification with an interaction between the bunching rate and the high-multi-firm indicator, in the same spirit as the heterogeneity analysis in Table 6:

$$y_{i,t} = \beta_1 B_{m(i),t} + \beta_2 \left(B_{m(i),t} \times 1\{MultiFirm > \text{median}\} \right) + \gamma_i + \delta_t + u_{i,t}.$$

Table E8 reports the results. The interaction term is small and statistically indistinguishable from zero. Hence, markets with a greater prevalence of multi-firm owners do not display systematically stronger negative effects of bunching on non-eligible firms. This pattern is difficult to reconcile with a quantitatively important role for within-owner income shifting.

Table E8: Heterogeneity by Multi-Firm Exposure

	Outcome: log(revenue)
$B_{m,t}$	-0.059 (0.006)
$B_{m,t} \times \mathbb{1}(\text{MultiFirm exposure} > \text{median})$	0.006 (0.003)
Observations	17,510,903
Firm FE	Yes
Year FE	Yes

Notes: This table tests whether our 2SLS estimate varies with the scope for within-owner revenue shifting. We proxy the likelihood of within-owner shifting using an external dataset on individual (physical-person) ownership: for each province×NACE 2-digit cell, we compute the prevalence of multi-firm ownership (average number of firms per owner) and define an indicator equal to one if this measure is above the median. We then interact the market-level bunching share with this indicator. Each specification includes firm and year fixed effects. Markets are defined at the municipality-industry-year level. Individual bunching firms are those reporting revenue between -5,000 and 0 relative to the eligibility threshold for preferential regimes; competitors are non-individual firms operating in the same market and reporting less than €500,000 revenue. Standard errors clustered at the market level are reported in parentheses.

F. Dynamic Olley-Pakes Productivity Decomposition

We examine the effects of a variation of the bunching rate on variation of the productivity at the market level employing the decomposition approach proposed by Melitz and Polanec (2015). This approach allows us to assess, for any period, the relative contribution of three groups of firms: the ones that survive (i.e., incumbents), entrants, and exiting firms. For incumbents, it is possible to further distinguish the contribution of two more components: (i) the variation in the efficiency of individual firms (i.e., within margin); and (ii) the reallocation of resources to firms characterized by different productivity levels (i.e., between margin). Therefore, for each market and any year, we decompose the productivity growth into four main components: the productivity growth of incumbent firms, the covariance between employment shares and productivity (which measures the extent of reallocation), the contribution of entering firms, and the contribution of exiting firms.

Formally, we split firms into entrants (E), exiters (X), and incumbents (S), and we define Φ_{gt} and w_{gt} as the aggregate productivity and the share of employment in the group $g \in \{E, X, S\}$ at time t . Then:

$$(F1) \quad \Phi_1 = \Phi_{S1}\omega_{S1} + \Phi_{X1}\omega_{X1};$$

$$(F2) \quad \Phi_2 = \Phi_{S2}\omega_{S2} + \Phi_{E2}\omega_{E2},$$

and the difference between Φ_1 and Φ_2 is:

$$(F3) \quad \Phi_2 - \Phi_1 = (\Phi_{S2} - \Phi_{S1}) + \omega_{E2}(\Phi_{E2} - \Phi_{S2}) + \omega_{X1}(\Phi_{S1} - \Phi_{X1})$$

We can rewrite the former equation as:

$$(F4) \quad \underbrace{\Delta\Phi}_{\text{productivity change}} = \underbrace{\Delta\varphi_S}_{\text{prod incumbents}} + \underbrace{\Delta Cov_S}_{\text{cov prod and size}} + \underbrace{\omega_{E2}(\Phi_{E2} - \Phi_{S2})}_{\text{entry gain/loss}} + \underbrace{\omega_{X1}(\Phi_{S1} - \Phi_{X1})}_{\text{exit gain/loss}}$$

where $\Delta\varphi_S$ measures the gain deriving from average productivity changes, ΔCov_S the increase due to reallocation of workers toward more productive firms, $\omega_{E2}(\Phi_{E2} - \Phi_{S2})$ the gain from new firm entering the market and $\omega_{X1}(\Phi_{S1} - \Phi_{X1})$ the contribution of firm exiting the market.

Note that the contribution of the selection margin at the productivity decomposition depends on the reference productivity level for entrants and exiters. Namely, entrants

generate positive productivity growth if (and only if) they have higher productivity than the remaining (surviving) firms in the same time period when entry occurs. Exiters, in turn, generate positive productivity growth if (and only if) they have lower productivity than the remaining (surviving) firms in the same time period when the exit occurs.

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