Who Bears the Burden of Local Taxes?

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Abstract

We study the incidence of local taxes on the welfare of heterogeneous residents. A structural model of imperfectly mobile households who differ in terms of income and family status allows us to back out type-specific preferences for local public goods. We calibrate the model with plausibly causal tax-base and housing-price elasticity estimates. Based on municipality-level data for Switzerland, we find that households with children perceive locally provided public services as a normal good, whereas households without children perceive them as an inferior good. This in turn implies that the burden of local taxes is mainly borne – linearity of taxes and capitalization into lower housing prices notwithstanding – by top-25% income households without children.

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Introduction

The distributional effects of taxation are among the most studied topics in public finance. At the national level, the tax system most evidently redistributes through the progressivity of rate schedules and because of differential avoidance opportunities.\(^1\) In this paper, we focus on an alternative setting, designed for the study of heterogeneous tax incidence at the local level: changes in personal tax rates are taken to be linear, and location choices are considered as the only avoidance option. Here, distributional effects arise for two other reasons: capitalization of tax rates into housing prices combined with non-homothetic housing demand, and heterogeneous preferences for publicly provided goods.

Consider a linear increase in a jurisdiction’s tax rate, associated with a corresponding increase in public expenditure. This policy change will to some extent get capitalized into housing prices. Suppose that capitalization is such that tax increases lead to lower equilibrium housing prices. If higher-income households spend a smaller share of their budget on housing than lower-income households, then capitalization will reduce higher-income households’ direct loss from the higher tax rate relatively less. This will leave them with a bigger tax-induced welfare loss compared to their lower-income fellow residents. However, to the extent that higher-income households attach relatively more weight to publicly provided goods, they will benefit more from the expenditure increase. Even a linear change in taxation will therefore not be distributionally neutral, with uncertainty about both the ordering and the sign of welfare effects across household types. The key determinants of these welfare effects are capitalization rates and household-type specific preferences for publicly provided goods – quantities that we estimate.

Local taxation, at the level of cities, counties, school districts or municipalities, accounts for important shares of public revenue in many countries: 16% in Switzerland, 15% in the United States, 10% in Canada, 9% in Spain and 8% in Germany.\(^2\) Most of these taxes correspond to the setting that we analyze: they are levied on the property or income of local residents and are used to finance locally provided public goods.\(^3\) Local government spending typically includes schooling, which in turn shapes residential sorting by income and household demographic structure.\(^4\) Therefore, residents’ heterogeneous incomes and preferences imply that local taxes have distributional implications even if statutory tax schedules are characterized by no or weak progressivity.

Based on a new panel dataset for Swiss municipalities, we indeed find substantial variation in incidence across household types. For households without dependent children, the welfare effect of a local tax increase ranges from strongly positive for below-median income

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\(^1\)On the distributional effects of heterogeneous avoidance opportunities, see e.g. Johns and Slemrod (2010) and Kopczuk, Marion, Muehlegger and Slemrod (2016).
\(^2\)Averages for the period 2000-2016. Data from the OECD Fiscal Decentralization Database. This list includes only countries with a three-tier jurisdictional architecture. In some two-tier federations, the local share is even higher (e.g. 34% in Sweden, 28% in Denmark).
\(^3\)In the United States, some 72% of local tax revenues are from property taxation, and some 5% are from income taxation (Annual Census of State and Local Government Finance, 2016). In Switzerland, the corresponding shares are 7% and 63% (see Section 2.1).
\(^4\)Primary and secondary education accounts for 40% of local government spending in the U.S. (Annual Census of State and Local Government Finance, 2016). In Switzerland, 27% of local expenditure goes to schooling (see Section 2.1).
households to strongly negative for top-quartile income households. In contrast, the welfare effect of a local tax increase is approximately zero across income classes for households with dependent children. Underlying these welfare effects are (a) a strongly positive relationship between income and preferences for local public goods among households with children, and (b) the reverse relationship for households without children. The implication is that local public services enter the preferences of households with children as a normal good, whereas for households without children they are an inferior good. While our estimations exploit variations in local taxes on income and in rental prices, our analytical framework as well as our qualitative findings apply also to other residence-based local taxes and to homeowners.

Moreover, we investigate how the tax base responds to tax rates depending on household type. We find that structural tax base elasticities with respect to tax rates are positive for below-median income households (+0.29 and +0.01 for households without and with children, respectively), strongly negative for top-quartile income households without children (-0.69) and insignificant for the other household types. The housing price elasticity with respect to local income taxes is -0.15. We also use cross-sectional variation in topographical constraints and local administrative environment as supply shifters. We weakly reject the assumption of perfectly price inelastic housing supply, finding an average long-run supply elasticity of 0.73. The price elasticity of housing demand is estimated to be around −1.1.

As we focus on changes in local taxes within a given functional labor market or commuting area, we can treat wages as exogenous with respect to location choices. This allows us to take account of residential mobility while assuming a constant labor income. When residents are mobile, higher personal tax rates will to some extent be capitalized into lower housing prices. In much of the literature on sub-national public finance, following Tiebout (1956) and Oates (1969), residential mobility is costless. In that case, the incidence of local taxes is fully borne by landowners, the immobile factor. In reality, moving costs exist even at the local level, and hence the welfare of renter households will also be affected by changes in local taxation. We therefore depart from the standard assumption by allowing households to have idiosyncratic prior preferences over locations, and thus non-zero moving costs, even within a given labor market.

Specifically, we first develop a spatial equilibrium model featuring imperfectly mobile residents renting housing from absentee landlords, with upward-sloping local housing supply (Section 1). Households choose where to reside among jurisdictions that offer different public expenditure levels, financed by a proportional income tax on residents. We allow residents’ valuation of the locally provided public good to vary across household types, without imposing any prior restrictions on the estimated relationship. Household types are defined in terms of income (to allow for preference non-homotheticity) and the presence or absence of dependent children (to account for the fact that schooling is one of the main locally financed and provided public goods). The incidence of changes in local tax rates on households depends on their their type-specific ‘bid-rent’ price, i.e. their marginal willingness to accept higher taxes and spending through lower housing costs. We use equilibrium conditions for location choices and for local housing markets to derive theoretical reduced-form effects of a tax increase on the number of households per type and on housing prices. The theoretical
reduced-form elasticities are determined by three key parameters: type-specific preferences for the local public good, the price elasticity of housing supply, and the (type-invariant) dispersion of idiosyncratic locational preferences that drives the price elasticity of housing demand.

Our estimation of reduced-form parameters is based on 1.6 million transaction-level rental price postings covering all of Switzerland over the 2004-2014 period, and on local tax rates and income-class-specific taxpayer counts for the 2,352 Swiss municipalities over that same period. The Swiss institutional setting, featuring three jurisdictional layers, each with large autonomy to tax and spend, allows us to instrument changes in local tax rates – a unique advantage of our setting that permits us to estimate causal effects of local taxes. Specifically, we exploit cross-section and time variation in municipal tax rates at state (‘canton’) borders that we instrument with neighboring state-level tax rates.

A key feature of this paper is that we allow households in different income classes to value the local public good differently. Figure 1 provides illustrative survey evidence for Switzerland. Stated support for government spending increases with income until the third income quintile and then decreases for the top two quintiles. Such preferences will obviously differ further across household types and public goods.\(^5\) To the best of our knowledge we are among the first to structurally estimate the relationship between revealed public-good preferences and household types.\(^6\)

Our study contributes to four main strands of the literature. First, we build on and contribute to an active literature on the incidence of subfederal taxation that carefully takes account of capitalization effects. In a seminal study, Suárez Serrato and Zidar (2016) use structural estimation to apportion the incidence of U.S. state corporate tax rates to workers, landowners and firm owners. They estimate that some 40 percent of the gain from state-level corporate tax cuts accrue to firm owners and 30-35 percent accrue to workers.\(^7\) Löffler and Siegloch (2018) focus on local property taxation in Germany and explicitly consider locally provided public goods. They find that renter households bear one fifth of the incidence of property taxes. Schönholzer and Zhang (2017), study municipal annexations in the United States and similarly find most of the incidence of local public spending to accrue to property owners. Importantly, they estimate substantial valuations by residents of locally provided public goods.

Our paper differs from this work along the following main dimensions. Most importantly, we estimate distributional effects by disaggregating residents by family type, income and age. We study the incidence of local income taxes that influence residential location decisions

\(^5\)See Appendix Figure A.4.1 for stated preferences on education spending and unemployment benefits.
\(^6\)Epple, Romer and Sieg (2001) allow preference parameters for the local public good to follow a log-normal distribution. Using a structural model calibrated for the Boston Metropolitan Area, they find evidence for a within-community dispersion of the preference parameter. Suárez Serrato and Wingender (2016) study the incidence of federal government spending at the local level and structurally estimate separate preference parameters for “skilled” and “unskilled” workers. We also complement Eugster and Parchet (2019), who use the Swiss language border to show the effect of culture on preferred tax levels, without, however, considering heterogeneity across household types.

\(^7\)The share of corporate-tax incidence falling on workers has been found to be even higher in smaller jurisdictions. Based on reduced-form empirical moments, Fuest, Peichl and Siegloch (2018) estimate that half of the gains from cuts to municipal business tax rates in Germany accrue to workers. This effect is mainly driven by small, single-plant (and thus immobile) firms.
within a given labor market, i.e. in a jurisdiction that is connected to surrounding jurisdictions through commuting. The wage (distribution) is therefore exogenous in our case, which allows us to study differential welfare effects across household types. Methodologically, we address a key identification issue by instrumenting local tax rates. We moreover use housing demand and supply shifters to estimate housing demand and supply elasticities that are particularly important as they govern the effect of local policies on welfare (Kline and Moretti, 2014).

Second, we contribute to a well developed empirical literature on the capitalization of taxes in housing prices. Basten, Ehrlich and Lassmann (2017) also draw on Swiss micro-geographic data to estimate the capitalization of local taxes into housing prices. Similar to most of the empirical literature on the capitalization of local policies or amenities, they use a (border) regression discontinuity framework and assume that households are perfectly mobile and housing demand is perfectly elastic. Reduced-form estimates of house price responses then serve directly as a measure of willingness to pay (through housing prices), with the incidence of the tax assumed to be fully borne by the immobile factor. We take a structural approach to estimate the elasticities that need to be calibrated for an analysis of incidence on different types of imperfectly mobile households, taking account not only of non-homothetic preferences for housing but also of heterogenous preferences for local public goods across

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8The assumption of locally exogenous wages has empirical support: Löffler and Siegloch (2018) find no effect of local property taxes on local wages, which is all the more remarkable considering that their German sample municipalities are on average almost 20 times larger than our Swiss sample municipalities. This is of course not to deny that labor supply and wages are affected by subfederal income taxation at larger spatial scales, such as that of U.S. states (see, e.g., Zidar, 2019).


10See, e.g., Black (1999); Reback (2005); Bayer, Ferreira and McMillan (2007); Fack and Grenet (2010); Cellini, Ferreira and Rothstein (2010); Black and Machin (2011); Boustan (2013); Gibbons, Machin and Silva (2013).
household types.

Third, we complement the empirical literature on the mobility response of households to tax changes.\(^\text{11}\) This literature is largely focused on top-income taxpayers and provides limited guidance about the mobility response of middle-income and lower-income households. Tax-induced mobility has previously been found to be significant in the case of Switzerland, probably due to the combination of high degree of fiscal decentralization and a small spatial scale.\(^\text{12}\) We link type-specific tax base elasticities to taxpayers’ marginal willingness to pay and study the distributional effects of local tax changes.

Fourth, our results shed light on the empirical relationship between local spending and the demographic composition of households.\(^\text{13}\) We show that a key dimension behind different preferences for locally provided public goods is the interaction between household income and the presence of children, and not the two single dimensions alone.\(^\text{14}\)

The paper proceeds as follows.\(^\text{15}\) In Sections 1 and 2, we present a model of local labor and housing markets and the data that will inform our empirical estimations. In Section 3 we estimate reduced-form elasticities of tax bases and housing prices with respect to local tax rates. Section 4 reports our baseline structural type-specific incidence estimates. In Section 5, we present some extensions of the baseline estimations, and Section 6 concludes.

1 Model

In this Section, we develop a model of residential location choice, housing markets and local public good provision. First, we assume a public sector that uses a proportional income tax to provide a potentially rival publicly provided good, and we characterize location choices and housing demand by households in different income classes.\(^\text{16}\) Second, we model housing supply in an absentee landlord setting. Third, we use the model to investigate the effect of tax rate changes on the number of residents in different income classes, on housing prices, and on the incidence of local taxes across household types.

1.1 Housing demand

We assume a functional labor market that consists of \(J\) municipalities. This labor market is populated by a unit continuum of \(I\) households that rent dwelling space from atomistic absentee landlords and take housing prices as given. Households are heterogeneous in their income and preferences for public goods, and they are budget constrained. Within an income class, households have identical preferences for housing and public goods, but they derive

\(^{11}\)See, e.g., Kleven, Landais and Saez (2013); Moretti and Wilson (2017); Agrawal and Foremny (2019).

\(^{12}\)See, e.g., Schmidheiny and Slotwinski (2018); Brühlhart, Gruber, Krapf and Schmidheiny (2016); Martínez (2017); Widmann (2018).

\(^{13}\)See, e.g., Harris, Evans and Schwab (2001); Hilber and Mayer (2009); Aaberge, Bhuller, Langørgen and Mogstad (2010); Figlio and Fletcher (2012); Bertocchi, Dimico, Lancia and Russo (2017); Aaberge, Eika, Langørgen and Mogstad (2019).

\(^{14}\)On residential income segregation by households with and without children, see, e.g., Owens (2016). For evidence on residential sorting by household type according to differences in exogenous local amenities (rather than local public goods), see, e.g., Chen and Rosenthal (2008) and Albouy and Faberman (2018).

\(^{15}\)Appendix A.1 offers a schematic overview of the different building blocks of the paper.

\(^{16}\)For simplicity, we use the term “public goods” as equivalent to “publicly provided goods”. Our setting can be easily extended to other resident-based taxes such as a property tax and to homeowners as in, e.g., Epple and Romer (1991).
idiosyncratic utility from exogenously given local amenities.\(^{17}\)

Specifically, each of the \(i \in \mathcal{I}\) renter households belongs to a discrete income class \(m \in \mathcal{M}\), within which everybody’s income equals \(w_m\). Households maximize the log Stone-Geary utility of residing in municipality \(j \in \mathcal{J}\) by choosing consumption levels of a freely tradable numeraire composite good \(z_{mj}\) and dwelling size \(h_{mj}\) at a price \(p_j\), subject to their after-tax income \((1 - \tau_j)w_m.\(^{18}\)

The indirect utility of household \(i\) with income \(m\), based on its choice of location \(j\), is

\[
V_{imj} = \kappa + \ln[(1 - \tau_j)w_m - p_j \nu_h] - \alpha \ln(p_j) + \delta_m \ln(g_j) + \ln(A_{ij}),
\]

where \(\kappa\) is a constant, \(\alpha \in (0,1)\) is a taste parameter for housing, \(\nu_h\) is the Stone-Geary parameter capturing the minimum amount of housing required, and \(\delta_m\) is a taste parameter for the local public good.\(^{19}\) We assume a balanced budget for the public sector with \(\tau_j \sum_m w_m N_{mj} = N_j^\theta g_j\), where \(\theta \in [0,1]\) indicates the degree of rivalness in the consumption of the public good.\(^{20}\) The number of residents, \(N_{mj}\), is defined below. We assume local amenities \(A_{ij}\) to be fixed.

At this stage, it is useful to define the change in the housing price (‘bid-rent’ price change) a household with in income class \(m\) would require to be indifferent toward a given change in the local tax rate:

\[
\frac{dp_j \tau_j}{d\tau_j p_j} \bigg|_{dV_{imj}=0} = -\left[\frac{\tau_j}{(1 - \tau_j)S_{mj}} - \frac{\delta_m}{\alpha} \left(1 - \frac{\nu_h}{h_{mj}^*}\right) \left(\frac{dg_j \tau_j}{dg_j g_j}\right)\right],
\]

where \(S_{mj} \equiv p_j h_{mj}^*/(1 - \tau_j)w_m\) represents the housing expenditure share and \(h_{mj}^*\) is the households’ Marshallian demand for housing space.\(^{21}\) \(\frac{dg_j \tau_j}{dg_j g_j}\) is the elasticity of public good provision with respect to the local tax rate. Using the balanced budget constraint, we have

\[
\frac{dg_j \tau_j}{g_j d\tau_j} = 1 + \sum_m (\gamma_{mj} - \theta s_{mj}) \frac{dN_{mj}}{N_{mj}} \frac{\tau_j}{d\tau_j},
\]

where \(\gamma_{mj} \equiv w_m N_{mj} / \sum_m w_m N_{mj}\) represents income class \(m\)’s share of municipality \(j\)’s tax base, \(s_{mj}\) is the proportion of households in income class \(m\), and \(\frac{dN_{mj}}{N_{mj}} \frac{\tau_j}{d\tau_j}\) is the elasticity of the number of residents in income class \(m\) with respect to the local tax rate.

Expression (2) determines the income-specific marginal willingness to pay rent (MWPR) for a (small) tax rate change. We impose no prior on \(\delta_m\) such that the MWPR could be positive.

\(^{17}\)When we take the model to the data, we shall further distinguish household types by family status (children/no children) and age (pensioners/non-pensioners). We do not incorporate these dimensions in the model, as they would complicate it without yielding additional insight.

\(^{18}\)We employ a Stone-Geary utility function for two reasons. First, it allows for a full range of housing demand elasticities with respect to rental prices of housing, i.e. \(|\eta h p| \in (0, +\infty)\). Second, in a setting with income heterogeneity, households belonging to different income classes are likely to have different expenditure shares on housing (see Appendix Figure A4.2). The Stone-Geary utility function incorporates this feature.

\(^{19}\)See the Online Appendix W.1 for detailed derivations.

\(^{20}\)If \(\theta = 0\), \(g_j\) is a pure public good. \(\theta = 1\) in turn represents the fully rival case, where \(g_j\) is a publicly provided private good.

\(^{21}\)With Cobb-Douglas preferences, we would have \(\nu_h = 0\) and \(S_{mj} = \alpha\).
or negative, and can increase or decrease non-monotonically with income.

We incorporate imperfect residential mobility by modeling location-specific amenities $A_{ij}$, consisting of a common location-specific component $\bar{A}_j$ and a location-specific idiosyncratic preference component $\xi_{ij}$. The household’s objective is therefore to maximize

$$\max_j V_{imj} = \kappa + \ln[(1 - \tau_j)w_m - p_j\nu_h] - \alpha \ln(p_j) + \delta_m \ln(g_j) + \bar{A}_j + \xi_{ij},$$  

where household $i$ will choose municipality $j$ if their indirect utility is higher there than in any other municipality $j' \neq j$. The variable $u_{mj}$ defines the systematic valuation of municipality $j$, common to all households in income class $m$.

We make the standard assumption that the idiosyncratic component $\xi_{ij}$ follows an i.i.d. Gumbel distribution with mean zero, variance $\sigma^2$ and scale parameter $\lambda = \frac{\pi \sigma}{\sqrt{6}}$. This scale parameter serves to model residential mobility. At one extreme, as $\lambda \to \infty$ ($\sigma \to 0$), the idiosyncratic attachment to location disappears and all households with income $m$ choose identically. At the other extreme, as $\lambda \to 0$ ($\sigma \to \infty$), idiosyncrasies dominate the systematic valuation of locations $u_{mj}$, and the population in each jurisdiction is fixed. We model $\lambda$ as a constant parameter across income groups, such that the differential incidence of local taxes is explained by non-homothetic preferences rather than differential mobility.

The share of households in income class $m$ choosing to reside in municipality $j$ is then given by

$$N_{mj} \equiv Pr\left(V_{imj} > V_{imj'} \forall j' \neq j\right) = \frac{\exp(\lambda u_{mj})}{\sum_{j'} \exp(\lambda u_{mj'})}, \quad \text{with } \sum_m \sum_j N_{mj} = 1. \quad (5)$$

Aggregate demand for housing in municipality $j$ is

$$H_j^d = \sum_m N_{mj} \cdot h_{mj}^*, \quad \forall j \in J,$$

which is the sum of all households across income classes $m \in M$ choosing to live in municipality $j$, multiplied by their corresponding Marshallian demands for housing.

**1.2 Housing supply**

We model housing as a homogeneous good produced with constant returns to scale using non-land capital and land. Housing is supplied by developers at increasing marginal cost and sold to atomistic absentee landlords who then rent it out to residents.

The total dwelling stock in municipality $j$ is equal to

$$H_j^s = B_j p_j^{\nu_h}, \quad \forall j \in J,$$

22This appears to be a reasonable assumption in the Swiss case. Basten et al. (2017) have observed the marginal willingness to migrate to be “remarkably homogeneous” (p. 677) across income quartiles. Evidence for the United States also points toward relatively minor heterogeneity in worker mobility, conditional on the intensity of relevant localized demand shocks (e.g. Notowidigdo, forthcoming; Suárez Serrato and Wingender, 2016; Bayer, McMillan, Murphy and Timmins, 2016). Nonetheless, we show in Section 5.2 that our main results are robust to allowing for differential mobility costs by allowing $\lambda$ to vary across household types.
where $B_j$ is a constant and $\eta_j^{s,p}$ represents the housing supply elasticity with respect to housing prices. Housing supply is allowed to vary across locations according to the tightness of topographical and administrative constraints on construction (Saiz, 2010; Hilber and Vermueen, 2016).

In this simple framework, housing supply does not depend on local income tax rates. This may not be an accurate representation of many empirical settings (ours included) in which, for example, rental income is taxed in the jurisdiction where the dwelling is located. In Appendix Section A.2.1, we carefully address the implications of a dependence of housing supply on local income tax rates, used as demand shifters, for the empirical identification of $\eta^{s,p}$.

### 1.3 Equilibrium

The model’s equilibrium is characterized by three main equations:

$$N_j = \sum_m N_{mj} \text{ with } N_{mj} = \frac{\exp(\lambda u_{mj})}{\sum_j \exp(\lambda u_{mj})} \forall j \in J, \quad (8a)$$

$$H_j^d = H_j^s \forall j \in J, \quad (8b)$$

$$g_j = \tau_j N_j^{-\theta} \sum_m w_m N_{mj} \forall j \in J, \quad (8c)$$

where (8a) describes the population, (8b) governs the housing market, and (8c) is the government budget constraint for each jurisdiction $j$.\(^{23}\) In what follows, we concentrate on the first-order effects of a tax change in a jurisdiction $j$ on its tax base and housing price. We therefore abstract from the effects of $j$’s tax policy on housing prices and public good provision in other jurisdictions.\(^{24}\) Total log-differentiating these equations and stacking them into a system of equations yields

$$A_j \dot{y}_j = B_j \dot{z}_j, \quad (9)$$

where $\dot{y}_j = [\dot{N}_{1j}, \ldots, \dot{N}_{Mj}, \dot{p}_j]'$ is the vector of endogenous variables and $\dot{z}_j$ is the exogenous variable.\(^{25}\) More explicitly, the elements of matrices $A_j$ and $B_j$ are given by

$$A_j = \begin{bmatrix}
\frac{1-\delta_1(\gamma_1-\theta_1)}{\alpha} & \left(1 - \frac{\nu_h}{\pi_1j}\right) & \cdots & \left(1 - \frac{\nu_h}{\pi_{Mj}}\right) & \frac{1-\delta_1(\gamma_M-\theta_M)}{\alpha} \\
-\frac{\delta_2(\gamma_1-\theta_1)}{\alpha} & \left(1 - \frac{\nu_h}{\pi_{2j}}\right) & \cdots & \left(1 - \frac{\nu_h}{\pi_{Mj}}\right) & \frac{1-\delta_2(\gamma_M-\theta_M)}{\alpha} \\
\cdots & \cdots & \cdots & \cdots & \cdots \\
-\frac{\delta_M(\gamma_1-\theta_1)}{\alpha} & \left(1 - \frac{\nu_h}{\pi_{Mj}}\right) & \cdots & \cdots & \frac{1-\delta_M(\gamma_M-\theta_M)}{\alpha} \\
\pi_1j & \cdots & \cdots & \cdots & \pi_{Mj} - \rho_j
\end{bmatrix}$$

\(^{23}\)We provide evidence in Section 5.3 that the balanced-budget assumption largely holds in Swiss municipalities.\(^{24}\)Like in Suárez Serrato and Zidar (2016), this is consistent with households being ‘myopic’: they do not anticipate the effect of their own and other households’ location decision on public good provision and housing prices in other jurisdictions. Alternatively, one might assume an economy composed by an infinite number of small jurisdictions.\(^{25}\)In this paper, we use the notation $\dot{x} \equiv d\dot{x}/x$ for any variable $x$. 


and

\[
B_j = \begin{bmatrix}
\delta_1 \left(1 - \frac{\nu_m}{H_{mj}}\right) - \frac{\tau_j}{(1 - \tau_j)S_{mj}} \\
\vdots \\
\delta_M \left(1 - \frac{\nu_m}{H_{mj}}\right) - \frac{\tau_j}{(1 - \tau_j)S_{mj}} \\
\alpha \sum_m \pi_{mj} S_{mj}
\end{bmatrix},
\]

where \(\pi_{mj} \equiv \frac{H_{mj}}{H_j}\) is income class \(m\)’s share of aggregate housing demand, \(\gamma_{mj} \equiv w_m N_{mj} / \sum_m w_m N_{mj}\) represents income class \(m\)’s share of the municipality \(j\)’s tax base, and \(s_{mj}\) is the proportion of households in income class \(m\). The term \(\rho_j \equiv \sum_m \pi_{mj} (1 - (1 - \alpha) \frac{\nu_m}{H_{mj}}) + \eta^{sp}_j\) collects other parameters, notably \(j\)’s housing supply elasticity.

The diagonal elements of the upper block in matrix \(A_j\) represent how a given income class reacts to a tax rate shock, and off-diagonal elements in a given row represent how that same income class reacts to other income classes’ location decision, i.e. they represent feedback effects between heterogeneous households through public good provision. The matrix \(B_j\) captures direct effects of tax rate changes on local tax bases and housing prices, holding fixed the between-equation interdependencies collected in matrix \(A_j\).

Pre-multiplying equation (9) by \(A_j^{-1}\) yields the reduced-form version of the system of equations, which is given by

\[
\dot{\mathbf{y}}_j = A_j^{-1} B_j \dot{\tau}_j,
\]

where \(A_j^{-1} B_j\) represents the reduced-form theoretical moments that will be used in the structural estimation of the preference for public goods parameters (\(\delta_m\), in equation 15 below).

### 1.4 Incidence

We now have the elements in hand for analyzing welfare effects of local taxes on different household types.

We follow Kline and Moretti (2014) by defining aggregate renter household welfare as

\[
W^R = \sum_m s_m \cdot E[\max_j \{u_{mj} + \xi_{ij}\}].
\]

Assuming location-specific idiosyncratic preferences to be Gumbel distributed, aggregate household welfare is given by

\[
W^R = \sum_m s_m \cdot \frac{1}{\lambda} \log \left( \sum_j \exp(\lambda u_{mj}) \right),
\]

where \(s_m\) is the population share of income class \(m\).

The effect of a small change in the income tax rate of municipality \(j\) on household welfare in income class \(m\) is given by

\[
\frac{dW^R_m}{d\ln \tau_j} = \alpha N_{mj} \left(1 - \frac{\nu_m}{H_{mj}}\right)^{-1} \left\{ \frac{-\tau_j}{(1 - \tau_j)S_{mj}} - \frac{\delta_m}{\alpha} \left(1 - \frac{\nu_m}{H_{mj}}\right) \left(1 + \sum_m (\gamma_{mj} - \theta s_{mj}) \frac{dN_{mj}}{N_{mj}} \frac{\tau_j}{d\tau_j} \right) \right\},
\]

for\(m\) with the condition

\[
\frac{dp^*_j \tau_j}{d\tau_j} \eta^{sp}_j.
\]

(11)
where \( \eta^{p^{*r}} \) is the change in the equilibrium housing price. The aggregate change in household welfare is then \( \frac{dW_R}{d\ln\tau_j} = \sum_m s_m \cdot \frac{dW_m}{d\ln\tau_j} \).

Inspection of equation (11) highlights that the sign of household incidence in a given income class \( m \) is determined by the differential between households’ marginal willingness to pay rent and the change in equilibrium rental prices. Household welfare increases if the tax-induced change in the equilibrium housing price (i.e. capitalization) is larger in absolute value than the household’s bid-rent price, and vice-versa.

Landlords’ utility is defined as rental revenue less the cost of supplying location-\( j \) housing. The inverse supply curve is \( p_j = \left( \frac{H^*}{\pi_j} \right)^{1/\eta^{s_P}} \). Producer surplus is therefore given by

\[
W^L = \int_0^{H^*} \left( p^*_j - \left( \frac{x}{B_j} \right)^{1/\eta^{s_P}} \right) dx = \frac{p^*H^*}{(1 + \eta^{s_P})}.
\]

The change in the landlord’s welfare after a change in the local tax rate is then

\[
\frac{dW^L}{d\ln\tau_j} = p^*H^* \left( \frac{dp^*_j}{d\tau_j} \frac{\tau_j}{p^*_j} \right). \tag{12}\]

Landlords’ welfare is driven by changes in equilibrium housing prices: to the extent that changes in taxation capitalize into housing prices, their incidence is borne by the absentee owners.

### 1.5 From theory to empirics

The empirical analogue of equation (9) is

\[
A\dot{y}_j = B\dot{\tau}_j + e_j, \tag{13}\]

where \( e_j \) represents structural error terms. The reduced-form version of the system of equations is given by

\[
\dot{y}_j = A^{-1}B\dot{\tau}_j + A^{-1}e_j, \tag{14}\]

where \( \eta = [\eta^{N_1}, \ldots, \eta^{N_M}, \eta^{P}]' \) is the vector of reduced-form moments.\(^{26}\)

Our aim is to find the parameter vector \( \vartheta = [\delta_1, \ldots, \delta_M] \) that best matches the moments \( m(\vartheta) = \eta \) to their reduced-form empirical counterparts \( \hat{\eta} \). For a given set of calibrated parameters, we use classical minimum distance (CMD) structural estimation (Chamberlain, 1984) to find

\(^{26}\)Hereinafter, reduced-form elasticities of a variable \( x \) with respect to \( \tau \) are denoted \( \eta^x \) instead of \( \eta^{x,\tau} \) to save on notation, unless explicitly stated otherwise. Note also that the empirical estimates of reduced-form moments are \( j \)-invariant. We therefore drop subscript \( j \) on matrices \( A \) and \( B \).
\[
\hat{\vartheta} = \arg \min_{\vartheta \in \Theta} [\hat{\eta} - m(\vartheta)]' \hat{V}^{-1} [\hat{\eta} - m(\vartheta)],
\]

where \( \hat{V}^{-1} \) is the inverse of the variance-covariance matrix from the reduced-form empirical estimation of the vector \( \hat{\eta} \).

This structural estimation relies on two building blocks:

1. joint estimation of two responses to changes in taxation, contained in the vector \( \hat{\eta} \):
   - the elasticity of the tax base with respect to the local tax rate (the “tax base elasticity”), and
   - the elasticity of the housing price with respect to the local tax rate (the “capitalization elasticity”),

and

2. the calibration of two parameters:
   - the elasticity of housing supply with respect to the housing price (the “housing supply elasticity”, \( \eta^{s,p} \)), and
   - the idiosyncratic location preference dispersion parameter \( \lambda \), derived from an estimate of the elasticity of housing demand with respect to the housing price (the “housing demand elasticity”, \( \eta^{d,p} \)).

We take advantage of the Swiss setting (Section 2) to identify and jointly estimate tax base and capitalization elasticities while instrumenting local income tax rates (Section 3). We also exploit (instrumented) local income tax variation as a demand shifter to estimate the housing supply elasticity. In a similar fashion, we use cross-sectional variation in topographical constraints and local administrative environment as supply shifters to estimate the housing demand elasticity that will inform our calibration of \( \lambda \) (Appendix Section A.2). The other parameters of matrices \( \mathbf{A} \) and \( \mathbf{B} \) (\( \gamma_{mj}, s_{mj}, h^*_{mj}, \pi_{mj}, \rho_j \) and \( S_{mj} \)) as well as income tax rates \( \tau_j \) will be calibrated with observed values (Section 4). Appendix A.1 offers a schematic overview of the different building blocks of the paper.

## 2 Empirical setting

### 2.1 Institutional background

Switzerland is a highly decentralized country composed of 26 cantons and 2,352 municipalities.\textsuperscript{27} The three layers of government enjoy significant autonomy in taxation and public spending. According to the OECD Fiscal Decentralization Database, Switzerland has OECD’s highest local revenue share, followed by the United States and Canada. Gauged by the share of autonomously raised municipal taxes, Switzerland is the third-most decentralized OECD

\textsuperscript{27}The municipality count refers to 2014, our final sample year. Due to municipal mergers, this number has been gradually decreasing. In 2004, our first sample year, the municipality count stood at 2,780.
country, after Finland and Iceland, but with a somewhat higher local tax share than the United States, Canada, Spain and Germany.\footnote{See Brülhart, Bucovetsky and Schmidheiny (2015).}

Our focus in this paper is on the municipal (“local”) level. Most municipalities are small. In 2014, the average municipal population was 3,256, for a maximum of 382,000 (Zurich). Nonetheless, municipalities are important in fiscal terms. In 2014, municipal spending accounted for 23% of consolidated public expenditure and 34% of consolidated personal income tax revenue.\footnote{The summary statistics cited in this and the following paragraphs are taken from SFSO (2017).} Municipalities are largely autonomous over most of their budget, including schooling (27% of average municipal expenditure), transport and environmental services (19%), general administration (11%) and recreation and culture (7%). In contrast, for some categories, the level of spending is mainly driven by canton-level or federal-level mandates. This primarily concerns social transfers (19% of municipal expenditure) and policing (6%).\footnote{The precise allocation of responsibilities between cantons and municipalities is complex and varied. The most comprehensive available account is given by Rühli (2012). All municipal tasks are to some extent affected by canton-level regulations and co-financing, but in only 2 of the 13 tasks identified in that study (policing and business development) does the average financial and executive weight of the canton dominate that of the municipalities. In 21 of 26 cantons, school districts perfectly overlap with municipalities, and in the remaining five cantons this is also the case for the majority of school districts, with a recent trend towards further integration of schooling into the general-purpose municipal administrations. Rühli (2012) also documents a trend towards increasing formal inter-municipal cooperation, with close to 40% of municipal tasks being shared through formal agreements with neighbor municipalities. In terms of our study this implies spatially correlated municipal policies.}

On the revenue side, municipalities have considerable decision-making powers as well. In 2014, some 64% of municipal revenue were raised through their own taxes, of which 63% were by personal income taxes. Property-related taxes, however, are relatively unimportant in international comparison, accounting for less than 7% of revenue.\footnote{We can only state an upper bound for the share of property-related taxes, as the corresponding category in the financial statistics also includes tax revenue that is not related to property taxes.}

Municipal tax policy in most cases consists of setting a single number: a multiplier on the canton-level tax schedule that determines the municipal share of the sub-federal tax take. Municipal tax multipliers can be adapted annually by municipal parliaments or citizen assemblies. Hence, within-canton variation in local income tax rates is almost perfectly captured by municipal tax multipliers.\footnote{We also take account of the fact that parishes levy their own (small) tax multipliers.}

Cantonal laws define statutory tax schedules and, combined with federal-level legislation, determine deductions and exemptions for the definition of the tax base. Municipalities, however, have no say over tax schedules, deductions and exemptions. Canton multipliers applied to the basic statutory tax schedule are determined annually by cantonal parliaments. Changes to the definition of the tax base or tax schedule are more infrequent, as they imply changes in cantonal tax laws and are thus typically subject to referenda.

Unlike income taxes, housing-related tax rates are mostly set at the canton level, with revenue sharing between cantons and municipalities.\footnote{Thus, housing tax rates largely cancel out in estimations featuring canton fixed effects. We will however have to take account of the minority of municipalities that set their own property tax rate.} Three such taxes are applied: First, 19 of the 26 cantons levy an annual property tax, computed as a fraction of the assessed value of the property.\footnote{The highest tax rate amounts to 0.3% of the assessed value (canton of Fribourg).} Second, when property ownership is transferred, sellers pay a real estate-
specific capital gains tax at a rate that is decreasing in ownership tenure. The real estate capital gains tax is levied in all cantons. Third, 18 out of the 26 cantons apply a property transaction tax.\textsuperscript{35}

An important aspect of real estate taxation in all of Switzerland is that owner-occupiers pay income taxes on imputed rents. Imputed rents are generally set somewhat below estimated market values, with federal guidelines stipulating at least 70\% of estimated market rent. Mortgage interest and renovation costs are tax deductible. Hence, the implied tax subsidy for owning relative to renting is significantly smaller in Switzerland than in countries that do not tax imputed rents. Indeed, at a first approximation, the Swiss tax system can be considered roughly neutral between renting and owning.\textsuperscript{36} Hence, our qualitative results should be valid not only for the considered population of renters but also for owner-occupiers, conditional on equal incomes and family status.

\subsection{Data}

We have assembled a unique municipality-level dataset covering the period 2004-2014. Our most important observed variables are personal income tax rates, housing prices, housing stocks, taxpayer counts by income bracket and local public expenditure. Table 1 provides summary statistics for all municipality-level variables. Information is presented separately for the full sample of 1,815 municipalities for which we have housing price data in 2004/2005 and 2013/2014, in columns (1)-(3), and for the subsample of 524 municipalities located close to canton borders, in columns (4)-(6). In columns (7)-(8), we report differences between the sample means of border and non-border municipalities.

We first need a measure of household \textit{income} to attribute taxpayers to income classes. We use taxable income according to the definition used for federal income taxation, which offers us a measure that is consistent across years and cantons. Our main focus is on three income classes: below-median income, the third quartile, and the top quartile. Quartile boundaries are calculated annually using the universe of federal income tax records.\textsuperscript{37} Importantly, we distinguish between households with and without dependent children. Among households without dependent children, we moreover distinguish between pensioner and non-pensioner households as a proxy for age. This last distinction is prone to some reporting errors (see Section 5.1) and available only for a subset of municipalities and years. We will therefore not use it for our baseline estimates.

For each of the nine household types (by income class, family type and age), we compute a representative average \textit{tax rate} using the consolidated cantonal, municipal and church tax liability as a percentage of gross wage income for representative households.\textsuperscript{38}

\textsuperscript{35}The mean tax rate is 0.5\% of the transaction price, with an upper bound of 3.3\% (canton of Neuchâtel).
\textsuperscript{36}The relative effect of the taxation of imputed rents on owners and renter households depends on the mortgage interest rate. As valuations on average remain fixed over 15 years but the mortgage interest deduction changes annually along with actual payments, the system favors homeowners in periods of high interest rates but disadvantages them in periods of low interest rates. According to estimations by the Swiss Federal Tax Administration, the system is approximately neutral for interest rates in the range 2.5-3.5\%, which comprises Swiss mortgage rates over our sample period.
\textsuperscript{37}The 75th (50th) percentile incomes for married households were CHF 101,000 (CHF 68,000) in 2014 (USD 111,000 and 75,000, respectively, using the 2014 exchange rate of CHF 1.10 per USD, which we consider throughout this paper).
\textsuperscript{38}Representative tax rates for the different household types are based on tax rates computed by the Swiss
Federal Tax Administration for discrete taxable income levels that range from CHF 10,000 to CHF 1,000,000 (USD 11,000 to USD 1,100,000 in 2014). These data are published for a sample of the largest municipalities. We can draw on earlier work, where we have extended this dataset to all municipalities (Parchet, 2019). Tax rates for specific income values (quartile boundaries) are obtained through linear interpolation between the nearest income levels reported in the official statistics. Published tax rates are reported relative to gross income. Based on the average difference between taxable and gross incomes documented by Peters (2005), we multiply our basic taxable-income measure by 1.25 for correspondence.
We focus on the following three main representative tax rates:

- **households with children (non-pensioners):** consolidated tax rates on income of married couples with two dependent children and a taxable income at, respectively, the median, the 75th and the 95th percentile of the nationwide distribution,

- **households without children (non-pensioners):** corresponding tax rates for unmarried taxpayers without dependent children,

- **pensioner households:** corresponding tax rates for married pensioners without dependent children.

In our baseline estimates, where we do not distinguish between pensioner and non-pensioner households, we use non-pensioner tax rates for households with dependent children. For

<table>
<thead>
<tr>
<th>Panel D: Public expenditure (in CHF million)</th>
<th>Mean (S.D.)</th>
<th>Min</th>
<th>Max</th>
<th>Mean (S.D.)</th>
<th>Min</th>
<th>Max</th>
<th>Difference (S.E.)</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Main sample</strong> (border &amp; non-border municipalities)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Total</td>
<td>27.91 (20.49)</td>
<td>0.13</td>
<td>8541.32</td>
<td>19.72 (4.96)</td>
<td>0.41</td>
<td>654.78</td>
<td>-12.09 (6.99)</td>
<td>0.181</td>
</tr>
<tr>
<td>Education</td>
<td>5.60 (2.83)</td>
<td>0.00</td>
<td>1020.63</td>
<td>3.39 (1.65)</td>
<td>0.00</td>
<td>145.96</td>
<td>-0.05 (0.784)</td>
<td></td>
</tr>
<tr>
<td>Social</td>
<td>5.32 (3.13)</td>
<td>0.02</td>
<td>1497.00</td>
<td>3.81 (2.27)</td>
<td>0.04</td>
<td>152.93</td>
<td>-2.13 (1.07)</td>
<td>0.164</td>
</tr>
<tr>
<td>Administration</td>
<td>2.77 (9.77)</td>
<td>0.03</td>
<td>832.37</td>
<td>2.97 (5.05)</td>
<td>0.04</td>
<td>188.54</td>
<td>-8.16 (0.773)</td>
<td></td>
</tr>
<tr>
<td>Roads</td>
<td>0.20 (1.34)</td>
<td>0.01</td>
<td>998.72</td>
<td>1.33 (4.49)</td>
<td>0.01</td>
<td>81.49</td>
<td>-1.34 (1.24)</td>
<td>0.277</td>
</tr>
<tr>
<td>Police</td>
<td>1.53 (2.48)</td>
<td>0.00</td>
<td>584.54</td>
<td>0.87 (6.55)</td>
<td>0.01</td>
<td>51.29</td>
<td>-0.89 (0.653)</td>
<td>0.141</td>
</tr>
<tr>
<td>Health</td>
<td>1.87 (2.81)</td>
<td>0.00</td>
<td>1089.62</td>
<td>0.98 (5.6)</td>
<td>0.00</td>
<td>127.24</td>
<td>-1.52 (1.37)</td>
<td>0.246</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Panel E: Time-invariant control variables (municipality-level)</th>
<th>Mean (S.D.)</th>
<th>Min</th>
<th>Max</th>
<th>Mean (S.D.)</th>
<th>Min</th>
<th>Max</th>
<th>Difference (S.E.)</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Share of developed land (1979-1995)</td>
<td>0.23 (0.18)</td>
<td>0.03</td>
<td>1.00</td>
<td>0.20 (0.15)</td>
<td>0.04</td>
<td>0.99</td>
<td>-0.14 (0.08)</td>
<td>0.000</td>
</tr>
<tr>
<td>Time-to-permit fixed effect coefficients (1997-2003)</td>
<td>0.80 (0.73)</td>
<td>-3.60</td>
<td>5.09</td>
<td>0.79 (0.66)</td>
<td>-2.29</td>
<td>1.78</td>
<td>-0.02 (0.33)</td>
<td>0.539</td>
</tr>
<tr>
<td>Index of accessibility</td>
<td>4.83 (2.13)</td>
<td>1.00</td>
<td>10.00</td>
<td>4.35 (1.86)</td>
<td>1.00</td>
<td>8.00</td>
<td>-0.68 (0.100)</td>
<td>0.000</td>
</tr>
<tr>
<td>Index of exposure to natural risks</td>
<td>0.27 (4.41)</td>
<td>1.00</td>
<td>10.00</td>
<td>0.72 (2.34)</td>
<td>1.00</td>
<td>10.00</td>
<td>-0.64 (0.122)</td>
<td>0.000</td>
</tr>
<tr>
<td>Index of architectural heritage</td>
<td>0.70 (6.60)</td>
<td>1.00</td>
<td>30.00</td>
<td>0.71 (6.81)</td>
<td>1.00</td>
<td>30.00</td>
<td>0.58 (0.348)</td>
<td>0.095</td>
</tr>
<tr>
<td>Hours of sunlight</td>
<td>0.74 (1.46)</td>
<td>0.00</td>
<td>8.10</td>
<td>0.91 (1.13)</td>
<td>0.00</td>
<td>8.10</td>
<td>0.24 (0.066)</td>
<td>0.000</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Panel F: Local autonomy in property taxation (canton-level)</th>
<th>Mean (S.D.)</th>
<th>Min</th>
<th>Max</th>
<th>Mean (S.D.)</th>
<th>Min</th>
<th>Max</th>
<th>Difference (S.E.)</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>No common multiplier</td>
<td>0.82 (0.38)</td>
<td>0.00</td>
<td>1.00</td>
<td>0.88 (0.33)</td>
<td>0.00</td>
<td>1.00</td>
<td>0.07 (0.018)</td>
<td>0.000</td>
</tr>
<tr>
<td>Property tax</td>
<td>0.68 (0.46)</td>
<td>0.00</td>
<td>1.00</td>
<td>0.55 (0.50)</td>
<td>0.00</td>
<td>1.00</td>
<td>-0.13 (0.025)</td>
<td>0.000</td>
</tr>
<tr>
<td>Transaction tax</td>
<td>0.36 (0.48)</td>
<td>0.00</td>
<td>1.00</td>
<td>0.33 (0.47)</td>
<td>0.00</td>
<td>1.00</td>
<td>-0.09 (0.025)</td>
<td>0.117</td>
</tr>
</tbody>
</table>

Notes: The main sample consists of all border and non-border municipalities for which rental prices are available in both 2004/2005 and 2013/2014. It includes 1,815 municipalities (1,800 for the number of pensioner and non-pensioner taxpayers, hence means do not always add up, 1,437 for public expenditure data). The border sub-sample contains 244 municipalities (240 for the number of pensioner and non-pensioner taxpayers, 196 for public expenditure data). The share of developed land is the ratio of developed land to developable land (total surface minus unproductive areas, taking into account topography). Time-to-permit fixed effects are municipality fixed effect coefficients from a regression of building permit approval time on observable characteristics of the project, municipality and year fixed effects. No common multiplier indicates municipalities that are allowed to set a different multiplier for their income tax and real estate capital gains taxes. Property tax and Transaction tax are dummy variables for municipalities that are allowed to levy a property tax or a real-estate transaction tax, respectively. (S.D.) means standard deviation and (S.E.) means standard error. Standard errors in column (7) are clustered at the municipality level.
childless households, we use a weighted average of tax rates for unmarried taxpayers without children and tax rates for pensioner couples without children, where the weights are based on the nation-wide tax base shares in 2004. Finally, as a measure of the overall representative tax rate in a municipality and year (needed, e.g., for estimating the elasticity of housing prices), we compute weighted averages of the 90th-percentile tax rates for married taxpayers with two children, unmarried taxpayers without children and pensioner couples without children.

Panel B of Table 1 shows that there exists considerable variation in local income tax rates within Switzerland, with the highest sample rate exceeding the lowest rate by a factor of around five for most of our representative tax rates. Figure 2 illustrates this variation in the cross-section and over time, mapping the local tax rates for unmarried taxpayers without children at the 95th income percentile (approximately CHF 120,000 in 2004). Figure 2 (a) shows that tax rates can vary within geographically small regions, thus allowing residents to change their tax bill by relocating within commuting zones. In our empirical analysis, we exploit time variation, illustrated in Figure 2 (b). This variation is substantial as well: the scale attached to the map shows that tax rate changes ranged from -6.8 to +3.4 percentage points, for a sample average tax rate of 18.7 percent (Table 1, Panel B).

Information on housing prices is taken from rental postings. The basic dataset available to us covers very close to the universe of Swiss online and print offers – some 1.6 million rental postings in total. The mean monthly rent for a 100 m² apartment is CHF 1,655 (USD 1,820), but price variations are large (see the summary statistics in Panel A of Table 1). In addition to rental prices, postings report object-level characteristics including floor space, the number of rooms and information on recent renovations. Rental prices provide an accurate measure of market prices, because posted rents are very close to transaction rental prices in Switzerland, where negotiation over posted rents is rare. In order to control for heterogeneous housing characteristics, we use residuals from an object-level regression of log rental prices on floor size (cubic polynomial), the number of rooms, the interaction between size and number of rooms, a dummy for recent renovations, municipality and year fixed effects.

For a proxy measure of local administrative efficiency, used as a supply shifter, we draw on the universe of individual-level building permits issued in Switzerland over the 1997-2003 period (i.e. prior to our main data period of 2004-2014). Our permits data include the projected costs, building type (e.g. a garage), type of project (e.g. renovation), and the number of structures (e.g. two garages). We compute, for all successful applications, the duration from the initial request to the award of the building permit, measured in months. We then perform a hedonic regression of time-to-permit on the observable characteristics of the project and municipality and year fixed effects. The estimated coefficients on the municipality fixed effects then serve as our proxy measure for local administrative efficiency (TTP).

Maps of average housing prices per municipality and of changes in these prices over time are presented in Appendix Figure A4.3. These are raw prices per square meter, without conditioning on dwelling characteristics. Data on rental postings, building permit requests and amenities are confidential and were kindly provided by Wüest Partner AG. This consultancy firm collect property advertisement information daily from all relevant websites and newspapers. Our dataset therefore covers essentially all arm’s-length rental offers. Exceptions not covered by our data include some postings in case of simultaneous new rentals in multi-unit buildings, and offers publicized only via informal local notice boards or word-of-mouth.

Negotiation over purchase prices, however, is as common in Switzerland as it is elsewhere. Hence, posted prices are a more reliable measure in the rental market than in the owner market.
Figure 2: The geography of local taxes in Switzerland

Notes: Figure (a) shows the consolidated cantonal, municipal and church income tax rates (in %) for unmarried taxpayers without children at the 95th income percentile. Figure (b) shows the difference in the consolidated income tax rates between 2014 and 2004. Gray lines represent municipality borders. Thick black lines represent canton borders. White areas are lakes, and light gray shaded areas are uninhabited mountains.

As a second shifter of housing supply, we consider topographic constraints. We draw on a cross section of data indicating the most relevant form of land use within 100×100m grid cells across Switzerland for the period 1979 to 1985. We combine this information with digital height model data that report the gradient of the surface.\textsuperscript{41} We define ‘developable’ land as

\textsuperscript{41}Both data sets are produced by the Swiss Federal Statistics Office. The land use data are publicly available here. They distinguish 17 land-use types, which we aggregate into four broader categories. The first category is ‘developed land’, consisting of (i) industrial and commercial areas, (ii) residential and public buildings, (iii) transport areas, (iv) special infrastructure and (v) recreational areas. The second category is ‘agricultural land’ and consists of (i) horticultural and viticultural areas, (ii) arable land, (iii) meadows and (iv) pastures. The third
the total surface area minus unproductive areas, forests and remaining unbuilt land with a slope greater than 20 percent (gradient of 11.3 degrees). The ratio of developed land to developable land yields the share of developed land (SDL).

We also collected time-invariant municipality-level amenity measures including indices for accessibility, exposure to natural risks (e.g. landslides), architectural heritage and winter sunlight hours (Panel E of Table 1).

For the estimation of housing demand and supply elasticities, reported in Appendix A.2, we compute the municipal housing stock as habitable residential floor space net of demolitions (dwelling space) at annual intervals for 2004-2014. In those estimations, we need to take account of the fact that cantons differ in the autonomy they grant to their municipalities with respect to property taxation. Where municipalities are allowed to set specific taxes on property values or transactions, these taxes will likely affect supply as well as demand, and local tax multipliers can no longer be interpreted as pure demand shifters (see Appendix Section A.2.1). We capture the degree of local autonomy through three binary variables. The no common multiplier variable is set to one for cantons that allow municipalities to apply a different multiplier for the income tax and for real estate capital gains taxes, and to zero where municipalities do not have that option. The property tax variable is set to one where municipalities are allowed to levy an annual tax on property values, and to zero otherwise. Finally, the transaction tax variable is set to one where municipalities are allowed to levy a real-estate transaction tax or such a tax exists at the cantonal level, and to zero otherwise.

Finally, we collected data on municipal public expenditure. Except for some 170 large municipalities, municipal public accounts are reported only to the cantonal authorities but not to the federal level. This forced us to gather these data from cantonal and, in some cases, municipal archives. We succeeded in obtaining broadly comparable expenditure data for 1,437 municipalities. The summary statistics in Panel D of Table 1 confirm that schooling (which includes pre-school services) is the largest municipal expenditure category, followed by social spending (which is largely non-discretionary) and administration.

Columns 7 and 8 of Table 1 show differences in means of our municipality-level variables between the border and non-border sub-samples. Municipalities in the border sample have somewhat lower housing prices and higher tax rates than those in the non-border sample. There are no statistically significant differences in households size nor composition, but the border sample does not include the most populous municipalities, as evident in columns (3) and (6). This also explains the lower share of developed land in the border sample. As a consequence, housing supply elasticities might differ between the two samples. We test the implications of different housing supply elasticities in Section 4.3.

category contains forests. Finally, we define ‘unproductive areas’ as including (i) lakes, (ii) rivers, (iii) unproductive vegetation, (iv) barren land and (v) glaciers and perpetual snow. The Digital Height Model (DHM25) data have been developed by the Geographic Information System group at the University of Lausanne.

42Forest areas in Switzerland are protected by federal law and can only be cleared in case of an evident public interest, in which case an identical surface has to be reforested within the same region.

43We thank the Federal Statistical Office for granting us access to confidential data from the Swiss Federal Registry for Buildings and Housing.

44The lower share of expenditure for schooling in our data (20%) compared to the aggregated statistics reported by SFSO (2017) (27%) is in part due to the existence in five cantons of single-purpose school districts, for which we do not have data.
3 Reduced-form responses to tax changes

Based on the data described in Section 2, we can estimate the vector of reduced-form moments \( \hat{\eta} \) of equation (13): elasticities with respect to local income tax rates (a) of municipality-level counts of taxpayers for each of our six household types (tax-base elasticities) and (b) of municipality-level average housing prices. We start by estimating a panel model featuring municipality and canton-year fixed effects. We then turn to an instrumental variable strategy to address the endogeneity of local tax rates.

3.1 Empirical models

3.1.1 OLS estimation

Even though our main estimates will be those estimated from long first-differences in a cross-border IV design, we begin the analysis by showing panel OLS estimations and then gradually building up to our preferred empirical model.

Our baseline panel-data estimating equation is

\[
\ln y_{jt} = \eta^y \ln \tau^y_{jt} + \phi_j + \phi_{ct} + \varepsilon^y_{jt}, \tag{16}
\]

where \( y_{jt} \) is either the count of taxpayers belonging to a specific household type, or housing prices in municipality \( j \) and canton \( c \) at time \( t \in [2004, \ldots, 2014] \), and \( \ln \tau^y_{jt} \) is the log consolidated (canton + municipal + church) tax rate as relevant to the associated regressand \( y \). Municipality fixed effects, \( \phi_j \), absorb time-invariant factors, and \( \phi_{ct} \) is a canton-year fixed effect such that our identification comes from municipalities in the same canton changing their tax multipliers at different points in time. Standard errors are clustered at the municipality level. Since housing price data are more reliable in larger municipalities, we weight our main regressions by the log of population in 2000.

Identifying causal effects of local tax rates is challenging for two reasons. First, local tax rates are decided by residents and could therefore respond directly to changes in the tax base. For example, an inflow of high-income taxpayers could strengthen the position of residents favoring lower tax rates; or municipalities could decide to lower their tax burden to mitigate the outflow of such taxpayers. Second, changes in local tax rates could be correlated with several unobserved time-varying factors that also influence location decisions, giving rise to omitted variable bias in the estimation of equation (16). We therefore move to an instrumental variable strategy to address the potential endogeneity of local tax rates.

3.1.2 Instrumenting local tax rates

Following the approach developed in Parchet (2019), we take advantage of the fact that, in Switzerland, three layers of government tax the same tax base. Cantonal borders create spatial discontinuities in fiscal policies across areas that are otherwise highly integrated. We implement a cross-border pairwise-comparison strategy and exploit changes in neighbor-canton tax rates as a source of exogenous variation. This variation is used to instrument differential changes in tax rates between neighboring municipalities located on opposite sides of canton borders.
We therefore restrict the sample to municipalities adjacent to a canton border, and we pair each of these municipalities with the nearest municipality in the neighboring canton provided their population centroids are separated by a road distance of no more than 10 kilometers.\footnote{For a map of the border-municipality sample, see Appendix Figure A4.4. Summary statistics are given in Table 1.} We then apply a cross-canton spatial difference estimation strategy, instrumenting the differential in the consolidated tax rate with the differential in the cantonal tax rate.

Our baseline panel-data estimating equation thus becomes

\[ \nabla \ln y_{jkt} = \eta \nabla \ln \tau_{jkt} + \phi_{jk} + \phi_{ct} + \epsilon_{jkt}, \]

(17)

where \( \nabla \) indicates the cross-canton spatial difference within pairs of municipalities \( jk \) in two neighboring cantons, \( c \) and \( d \), with \((j \in c) \neq (k \in d \neq c)\). Municipality-pair directional fixed effects, \( \phi_{jk} \), absorb time-invariant factors, and \( \phi_{ct} \) is an origin canton-year fixed effect such that our identification comes from municipalities in the same canton but bordering different neighboring cantons. Differentials in local tax rates, \( \nabla \ln \tau_{jkt} \), are instrumented with the corresponding differential in canton-level tax rates \( \nabla \ln \tau_{cdt} \). Standard errors are clustered two-ways, at the level of origin and destination municipalities. Regressions are weighted by the log of population in 2000 of the smaller municipality in the pair.

To be valid, this estimation strategy has to satisfy several conditions. First, tax base changes in border municipalities should not systematically affect canton-level fiscal policy. The implied assumption is that border municipalities are small compared to the overall (population) size of the canton.\footnote{Note that, due to spatial differencing, the identifying assumption requires the neighboring cantonal policy to be independent from the tax base in municipalities \( j \) and \( k \), and not only from municipality \( j \) as in Parchet (2019). Alternatively, one could estimate equation (16) without spatial differencing and instrument the tax rate of municipality \( j \) with the neighbor-canton tax rate. This, however, generally results in weak first-stage regressions (where the first stage corresponds to the reduced-form estimates in Parchet, 2019).} Second, canton-level tax changes should not be driven by unobserved factors that also change the attractiveness of border municipalities. In that respect, spatial differencing controls for common shocks at the local level (in, e.g., the local labor market) and at the cantonal level (due to tax competition for example), and canton-year fixed effects control for changes in canton-wide policies.

For the exclusion restriction to be valid, taxpayers should react to changes in cantonal tax differentials only because of the changed consolidated tax burdens. A concern would arise if municipal and cantonal tax rates were used to provide different types of public goods that are valued unequally by taxpayers. We can assume here that taxpayers care only about their total tax bill (and a “consolidated” public good), irrespective of whether the public services they consume are financed at the municipal or the cantonal level. A less stringent version of this assumption is that taxpayers do not distinguish the levels of government involved in the financing of specific public services. This is a reasonable assumption given the complexity of the financing of sub-federal public expenditure. With this identification strategy, we depart from our modeling assumption of a public good provided by one level of government. In our empirical setting, households consume locally (i.e. through their residence) a bundle of public services potentially provided by different levels of governments, and we structurally estimate their valuation of this bundle of public goods.
We first estimate tax-base and housing-price elasticities with respect to tax rates by running separate panel regressions for each dependent variable. Then, we perform a joint estimation on long first-differences for the period 2004/2005 to 2013/2014. In this alternative approach, we estimate the reduced-form moments \( \eta = [\eta^{N_1}, ..., \eta^{N_6}, \eta^{P}] \) jointly, using three-stage least squares estimation (3SLS), and again instrumenting municipality-pair-level differentials in consolidated tax rates with the corresponding differential in canton-level tax rates.

Specifically, the seven estimating equations are

\[
\nabla \Delta \ln N_{jk}^{1} = \eta^{N_1} \nabla \Delta \ln \tau_{jk}^{1} + \mu^{N_1} \nabla X_{jk} + \phi_{c}^{N_1} + \epsilon_{jk}^{N_1}, \quad (18a)
\]

\[
\vdots
\]

\[
\nabla \Delta \ln N_{jk}^{6} = \eta^{N_6} \nabla \Delta \ln \tau_{jk}^{6} + \mu^{N_6} \nabla X_{jk} + \phi_{c}^{N_6} + \epsilon_{jk}^{N_6}, \quad (18f)
\]

\[
\nabla \Delta \ln P_{jk} = \eta^{P} \nabla \Delta \ln \tau_{jk}^{P} + \beta_1 \nabla SDL_{jk} + \beta_2 \nabla TTP_{jk} + \mu^{P} \nabla X_{jk} + \phi_{c}^{P} + \epsilon_{jk}^{P}, \quad (18g)
\]

where \( \Delta \) represents the long difference between the averages for 2013-2014 and 2004-2005. We also control for the vector \( X \) of time-invariant municipal characteristics. In the housing-price elasticity equation, we in addition control for topographical constraints and local administrative efficiency.\(^{47}\) The first-difference strategy has the advantage of removing municipality-pair fixed effects for the joint estimation of seven equations. Moreover, it parallels our identification of the housing demand elasticity (for which we use cross-sectional variation in supply shifters) in Appendix A.2.

### 3.2 Results

Table 2 presents a range of estimation results, beginning with OLS estimations on the full data sample and then gradually building up towards our preferred empirical model.

First, we report estimates from simple panel models featuring municipality and a canton-year fixed effects. For the results shown in Panel A of Table 2, we use all municipalities for which housing prices are available. In Panel B, we restrict the sample to the border municipalities later retained in IV estimations. The two samples yield very similar results: a mostly negative correlation between changes in local tax rates and changes in taxpayer counts, with the magnitude of the correlation increasing with income. Similarly, local tax increases are associated with lower housing prices.

Panel C of Table 2 presents results for the cross-border spatial difference specification of equation (17). Most of the estimated coefficients are smaller in absolute value than in Panel B, suggesting that spatial differencing controls for time-varying confounding factors that are common among proximate jurisdictions.

Instrumenting local tax differentials with canton-level tax differentials in Panel D of Table 2 further shrinks the estimated coefficients. However, we still observe a negative and statistically significant tax base elasticity for households without children and above-median income. The estimated effect on housing prices, however, is no longer statistically significant.

\(^{47}\)See Appendix A.2 for details.
Table 2: Tax base and rental price elasticities: OLS and 2SLS results

<table>
<thead>
<tr>
<th>Panel</th>
<th>OLS estimation on all municipalities</th>
<th>OLS estimation on origin municipality</th>
<th>OLS estimation on destination municipality</th>
<th>OLS estimation on origin and destination municipality</th>
<th>Income tax rate</th>
<th># of observations</th>
<th># of municipalities</th>
<th>Municipality fixed effect</th>
<th>Canton-year fixed effect</th>
</tr>
</thead>
<tbody>
<tr>
<td>Panel A</td>
<td>OLS estimation on all municipalities</td>
<td>Income tax rate</td>
<td>$-0.086^{**}$</td>
<td>$-0.334^{***}$</td>
<td>$-0.649^{***}$</td>
<td>$0.136^{*}$</td>
<td>$0.099$</td>
<td>$-0.201^{***}$</td>
<td>$-0.172^{***}$</td>
</tr>
<tr>
<td>Panel B</td>
<td>OLS estimation on origin municipalities</td>
<td>Income tax rate</td>
<td>$-0.149^{***}$</td>
<td>$-0.308^{***}$</td>
<td>$-0.621^{***}$</td>
<td>$0.376^{**}$</td>
<td>$0.080$</td>
<td>$-0.189^{*}$</td>
<td>$-0.217^{***}$</td>
</tr>
<tr>
<td>Panel C</td>
<td>OLS pairwise difference estimation on origin municipalities</td>
<td>Income tax rate</td>
<td>$-0.075$</td>
<td>$-0.264^{***}$</td>
<td>$-0.540^{***}$</td>
<td>$0.005$</td>
<td>$0.085^{*}$</td>
<td>$-0.082$</td>
<td>$-0.171^{***}$</td>
</tr>
<tr>
<td>Panel D</td>
<td>IV pairwise difference estimation on origin municipalities</td>
<td>Income tax rate</td>
<td>$-0.001$</td>
<td>$-0.168^{*}$</td>
<td>$-0.345^{***}$</td>
<td>$0.008$</td>
<td>$0.079$</td>
<td>$-0.036$</td>
<td>$-0.095$</td>
</tr>
<tr>
<td>Panel E</td>
<td>IV pairwise difference estimation on origin municipalities: distributed lag model</td>
<td>Income tax rate (cumulative effect)</td>
<td>$0.048$</td>
<td>$-0.126^{***}$</td>
<td>$-0.526^{***}$</td>
<td>$0.011$</td>
<td>$0.034$</td>
<td>$-0.071$</td>
<td>$-0.157^{*}$</td>
</tr>
<tr>
<td>Panel F</td>
<td>IV pairwise difference estimation on origin municipalities</td>
<td>Income tax rate</td>
<td>$0.128$</td>
<td>$-0.342^{*}$</td>
<td>$-0.814^{***}$</td>
<td>$0.007$</td>
<td>$0.014$</td>
<td>$-0.053$</td>
<td>$-0.283$</td>
</tr>
<tr>
<td>Panel G</td>
<td>IV pairwise difference estimation on origin municipalities</td>
<td>Income tax rate</td>
<td>$0.158$</td>
<td>$-0.356^{*}$</td>
<td>$-0.750^{***}$</td>
<td>$0.012$</td>
<td>$0.048$</td>
<td>$-0.102$</td>
<td>$-0.288$</td>
</tr>
</tbody>
</table>

Notes: Cluster robust standard errors reported in parentheses. In panels A and B, standard errors are clustered at the municipality level. In the remaining panels, standard errors are two-way clustered at origin and destination municipality level. In municipalities with zero taxpayer in a given category, ln(0) has been replaced by 0 (0 never occurs). Regressions in panel E employ a standard distributed lag approach estimating $\ln p_{j,t} = \rho \ln p_{j,t-1} + \sum_{k=1}^{K} \beta_k (\ln p_{j,t-k} + \ln p_{j,t-k-1}) + \phi_{j,t} + \psi_{j,t}$, so that we may interpret $\rho$ directly as the long-term effect. Controls in Panel G include (time-invariant) indices of accessibility, exposure to natural risks, architectural heritage, and hours of sunlight. In columns (7) we in addition control for topographical constraints and local administrative efficiency. **$p<0.01$, ***$p<0.05$, $p<0.1$. 

Long difference model between the averages 2013-2014 and 2004-2005

Panel A: OLS estimation on all municipalities

Panel B: OLS estimation on origin municipalities

Panel C: OLS pairwise difference estimation on origin municipalities

Panel D: IV pairwise difference estimation on origin municipalities

Panel E: IV pairwise difference estimation on origin municipalities: distributed lag model

Panel F: IV pairwise difference estimation on origin municipalities

Panel G: IV pairwise difference estimation on origin municipalities

Panel H: Housing Prices

<table>
<thead>
<tr>
<th>Panel</th>
<th>Controls</th>
<th>Canton-year fixed effect</th>
<th>Canton-year fixed effect</th>
<th>Canton-year fixed effect</th>
<th>Canton-year fixed effect</th>
<th>Canton-year fixed effect</th>
<th>Canton-year fixed effect</th>
<th>Canton-year fixed effect</th>
</tr>
</thead>
<tbody>
<tr>
<td>Panel A</td>
<td>Canton-year fixed effect</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
</tr>
<tr>
<td>Panel B</td>
<td>Canton-year fixed effect</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
</tr>
<tr>
<td>Panel C</td>
<td>Canton-year fixed effect</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
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<td>YES</td>
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</tr>
<tr>
<td>Panel D</td>
<td>Canton-year fixed effect</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
</tr>
<tr>
<td>Panel E</td>
<td>Canton-year fixed effect</td>
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<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
</tr>
<tr>
<td>Panel F</td>
<td>Canton-year fixed effect</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
</tr>
<tr>
<td>Panel G</td>
<td>Canton-year fixed effect</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
</tr>
</tbody>
</table>

Notes: Cluster robust standard errors reported in parentheses. In panels A and B, standard errors are clustered at the municipality level. In the remaining panels, standard errors are two-way clustered at origin and destination municipality level. In municipalities with zero taxpayer in a given category, ln(0) has been replaced by 0 (0 never occurs). Regressions in panel E employ a standard distributed lag approach estimating $\ln p_{j,t} = \rho \ln p_{j,t-1} + \sum_{k=1}^{K} \beta_k (\ln p_{j,t-k} + \ln p_{j,t-k-1}) + \phi_{j,t} + \psi_{j,t}$, so that we may interpret $\rho$ directly as the long-term effect. Controls in Panel G include (time-invariant) indices of accessibility, exposure to natural risks, architectural heritage, and hours of sunlight. In columns (7) we in addition control for topographical constraints and local administrative efficiency. **$p<0.01$, ***$p<0.05$, $p<0.1$. 

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Our IV-based elasticity estimates being smaller in absolute value than the non-instrumented estimates is consistent with endogenous local tax setting due to reverse causation. The magnitude of the “shrinkage” of estimated elasticities through instrumenting the tax rates suggests elasticities reported in the overwhelming majority of related studies that do not instrument tax rates may be significantly overestimated.

Conversely, estimated tax-base and capitalization elasticities are biased towards zero to the extent that it takes time for households to move and for rents to adjust. In Panel E of Table 2, we therefore augment equation (17) with two lags, themselves instrumented with the corresponding lags of the cross-border canton-level tax differentials. We report implied long-term effects and their standard errors, based on the sum of the contemporaneous and the lagged coefficients. As expected, estimated three-year tax-base and house-price capitalization elasticities are larger in absolute value than their one-year counterparts.\(^{48}\) The long-term capitalization elasticity is statistically significant at 10%.

We further exploit the panel structure of our data to test for the validity of our instrumental variable strategy. Figure 3 shows the cumulative effect of canton-level tax differentials (the instrument) on (a) the number of top-quartile households without children and (b) housing prices. Interpreting our panel estimates as a combination of individual event studies, and as a check of the assumption of common pre-trends, we plot the sum of the coefficients and their corresponding standard errors from estimating equation (17) augmented with two lags (in black) and two lags and two leads (in gray) using $\nabla \ln \tau_{ct}$ instead of $\nabla \ln \tau_{jt}$ as regressors. We find no evidence of changes in the municipality-level number of high-income and potentially mobile households in advance of canton-level tax changes (upper panel in Figure 3). Long-term (reduced-form) tax base elasticities are also very similar in the model with and without leads. Results for housing prices (lower panel in Figure 3) similarly lead us to rule out changes in the dependent variable pre-dating changes in cantonal tax rates.

Next, we turn to the long first-differences model. Panel F of Table 2 presents estimates based on differences between the averages for 2013-2014 and 2004-2005. Results are qualitatively similar to the distributed lag model presented in Panel E, with estimated tax base elasticities of households without children and the housing price elasticity larger in the 10-year first-difference model. In Panel G, we in addition control for differences in amenities across municipalities, and, in the rental-price regressions, for differences in topographical constraints and local administrative efficiency. Estimated coefficients are not sensitive to the inclusion of these variables as controls.

Overall, the results in Table 2 show that it is mainly above-median income households without children that move in response to local tax differentials, whereas households with children seem largely unresponsive. We also find evidence of capitalization of local tax rates into local housing prices. The estimated elasticities grow in absolute value as the time window is lengthened, suggesting delayed mobility and capitalization responses.

\(^{48}\)For below-median-income households without children, we moreover observe that instrumenting and lengthening the time window turns the tax elasticity from negative to positive (compare Panels E and A of Table 2). This is consistent with two-way causation, whereby the arrival of such households allows municipalities to lower their tax rates as these households’ (current) consumption of local public goods is below-average, but such households nonetheless prefer to move to municipalities with higher tax rates and thus more generous provision of local public goods.
Figure 3: Cumulative tax base and housing price elasticities

Notes: The figures show the cumulative effect of our instrument on the number of households without children and top-quartile income (upper panel) and on housing prices (lower panel). It plots the sum of the coefficients and their corresponding standard errors from estimating equation (17) augmented with two lags (in black) and two lags and two leads (in gray) using $\nabla \ln \tau_{ct}$ instead of $\nabla \ln \tau_{jt}$ as regressor.

Finally, Table 3, presents the three-stage least squares joint estimates of equations 18a-18g. Panel A presents the results with standard errors bootstrapped at the municipality-pair level. This comes at the cost of not weighting regressions by municipality size. Panel B presents the results for weighted regression but standard errors are assumed to be homoskedastic. Both specifications lead to similar results. Our structural estimation in Section 4 will be based on the weighted regression estimates of Panel B.

Again, we find that reduced-form tax base elasticities decrease strongly and monotonically
with income for households without children but do not vary much with income for households with children. Moreover, we find estimated elasticities to be positive for below-median income households. These results strongly suggest (a) that preferences are non-homothetic, (b) that households perceive taxes not as a net income losses but consider them jointly with the public goods supplied in return, and (c) that they hold heterogenous preferences over those public goods.

Our baseline estimated reduced-form housing price elasticity of -0.31 is close to the long-difference estimates of Table 2 and well within the range of estimates reported by Basten et al. (2017), based on a border discontinuity framework.

4 Estimation of structural parameters and incidence: baseline

With the estimated reduced-form elasticities in hand, we can progress towards estimating the structural model given by equation (15). We begin by making the simplifying assumption that renter households are myopic, in the sense that they do not anticipate the effect of their own location decision, nor the effect of the location decision of other income groups, on the provision of the public good. That is, for renter households, \(\hat{y}_j = \hat{\tau}_j\) instead of \(\hat{y}_j = \hat{\tau}_j + \sum_{m} (\theta s_{m} - \theta s_{m}) \hat{N}_{mj}\). Under this assumption, matrix \(A\) becomes considerably simpler: off-diagonal elements of the upper block are zero, on-diagonal elements do not depend on \(\delta m\). This proves useful for two reasons. First, the housing demand elasticity varies with income only though different housing expenditure shares. We can therefore calibrate the idiosyncratic location dispersion parameter \(\lambda\) directly using our empirical estimate of the

\[\gamma_{mj} - \theta s_{mj} = 0 \ \forall m \in M\] in matrix \(A_j\) from equation (9). Note also that \(\theta\) drops out.

---

Table 3: Tax base and rental price elasticities: 3SLS estimation

<table>
<thead>
<tr>
<th></th>
<th>Households without children</th>
<th></th>
<th>Households with children</th>
<th></th>
<th>Housing prices</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Bottom 50 (1)</td>
<td></td>
<td>Next 25 (2)</td>
<td></td>
<td>Top 25 (3)</td>
</tr>
<tr>
<td>Income tax rate</td>
<td>0.282***</td>
<td></td>
<td>-0.064</td>
<td></td>
<td>-0.633***</td>
</tr>
<tr>
<td></td>
<td>(0.063)</td>
<td></td>
<td>(0.077)</td>
<td></td>
<td>(0.118)</td>
</tr>
<tr>
<td>Panel A: unweighted regression, bootstrapped standard errors</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Income tax rate</td>
<td>0.010**</td>
<td></td>
<td>0.064</td>
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<tr>
<td></td>
<td>(0.005)</td>
<td></td>
<td>(0.049)</td>
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<td>(0.075)</td>
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<td></td>
<td>-0.330***</td>
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<td>(0.121)</td>
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<td>Controls</td>
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<td>Origin canton FE</td>
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<td># of observations</td>
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<td># of municipalities</td>
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<td>Instrument</td>
<td>Cantonal income tax rate differential</td>
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<tr>
<td>Estimator</td>
<td>3SLS</td>
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<tr>
<td>Panel B: weighted regression, homoskedastic disturbances</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: Standard errors reported in parentheses. Each column refers to an equation from 18a-18g. The equations are estimated jointly using three stage least squares. The sample consists of cross-canton pairs of municipalities with a pairing road distance of 10 km. Panel A bootstraps the standard errors with 250 iterations of the unweighted 3SLS estimations. Panel B regressions are weighted by the log population in 2000 of the smallest municipality in the pair. The consolidated personal income tax rate differentials are instrumented by the cantonal personal income tax rate differentials. Controls include (time-invariant) indices of accessibility, exposure to natural risks, architectural heritage, and hours of sunlight. In column (7) we in addition control for topographical constraints and local administrative efficiency. ***p<0.01, **p<0.05, *p<0.1.
housing demand elasticity.\textsuperscript{50} Second, we can formally derive the first-order conditions of the minimization problem and therefore compute standard errors. Estimates of a ‘fully fledged’ model that includes indirect effects will be shown in Section 4.3.

### 4.1 Calibration

In order to implement our structural estimation, we need to calibrate a number of additional parameters. Panel A of Table 4 presents these calibrated values.

First, we draw on data from the Swiss Household Panel (SHP) to calibrate taste and expenditure parameters related to housing. The housing taste parameter $\alpha$ follows from households’ Marshallian housing demand equation $h_{mj}^* = \nu_h + \frac{\alpha(1-\tau_j)w_m - \nu_h}{p_j}$, which can be rewritten as $S_{mj} = S_{mj}^{\min} + \alpha \left[1 - S_{mj}^{\min}\right]$, where $S_{mj}^{\min} \equiv \frac{p_j\nu_h}{(1 - \tau_j)w_m}$ is the expenditure share on minimum housing consumption. We compute $\alpha$ as $\frac{S - S_{mj}^{\min}}{S^{\min}}$, where $S$ is the average expenditure share on housing (defined as annual rent over disposable income) calculated using SHP data for the years 2000 to 2004 ($\bar{S} = 0.24$). We proxy the expenditure share on minimum housing needs, $S_{mj}^{\min}$, using the average rent paid by bottom-5% income renter households in the SHP data, computed separately for different household types. Similarly, the type-specific expenditure share on minimum housing needs ($\frac{\pi_j^{\min}}{\pi_j}$) is obtained by the average rent paid by bottom-5% income renters (differentiating households with and without children) over the average rent paid in each income class. Aggregate housing shares ($\pi_m$) are likewise calculated directly from the SHP data.

We calibrate proportional income tax rates $\tau_j$ in matrix $B$ (Section 1.3) by the group-averaged consolidated income tax rates for 2000-2004.\textsuperscript{51} Table 4 shows that these representative tax rates range from 5% (bottom-50 households with children) to 24% (top-25 households without children).

The scale parameter $\lambda$ is calibrated using our empirical estimate of the housing demand elasticity (Appendix Section A.2.2). It’s calibrated value of 1.9 suggests that residents are far from perfectly mobile even across small Swiss municipalities.

The housing supply elasticity comes from estimates presented in Appendix Section A.2.2. The estimated value of 0.73 implies that to assume perfectly inelastic housing supply would not be appropriate in our setting.

Finally, population shares $s_m$ are computed from federal income tax statistics.\textsuperscript{52} Of our six baseline household types, fully 44% belong to the category bottom-50 without children.

\textsuperscript{50}From equation (6), the housing demand elasticity is $\eta_{dp} \equiv \frac{H_{dp}}{p} = \sum_m \pi_m \left( \frac{S_{mj}}{p_j} + \frac{h_{mj}^*}{p_j} \right)$. In the myopic model, in which renters assume $\dot{y}_j = \dot{\tau}_j$, $\frac{S_{mj}}{p_j} = -\alpha \lambda \left(1 - \frac{\nu_h}{\pi_m}\right)^{-1}$ such that $\eta_{dp} = -\alpha \lambda \sum_m \pi_m \left(1 - \frac{\nu_h}{\pi_m}\right)^{-1} - 1 - \left(1 - \frac{\nu_h}{\pi_m}\right) \sum_m \pi_m \frac{h_{mj}^*}{p_j}$. The scale parameter $\lambda$ can therefore be calculated from our estimate $\hat{\eta}_{dp}^m$ presented in Appendix Section A.2.2, given the calibration of $\alpha$, $\pi_m$, and $\frac{\nu_h}{\pi_m}$.

\textsuperscript{51}Consolidation is across the federal, cantonal, municipal and parish levels. In the calibrations, we include the federal income tax rate that in Section 3 and Appendix Section A.2 is absorbed by fixed effects. Moreover, for households with children, we proxy the tax rate depending on the average number of children per household in each quartile (1.5, 1.7 and 1.8 children for the first two quartiles, the third quartile and the top quartile, respectively) by using a linear interpolation between the tax rates for married couples without children and with two children.

\textsuperscript{52}In the ‘myopic’ model, these population shares do not enter the structural estimation but are used as weights to estimate average incidence across resident groups.
The least frequent household type are bottom-50 taxpayers with children, accounting for only 6% of the total. This difference reflects the fact that households with children on average have higher incomes than households without children.

4.2 Estimates of structural parameters

Armed with the reduced-form parameters of Section 3 and the calibrated values of Section 4.1, we can estimate the structural public-goods preferance parameters through iterative minimization of equation (15). We take as starting values for the baseline minimization procedure the average estimated parameters resulting from repeating the procedure with 1,000 different starting values drawn randomly from a uniform distribution (see Section 4.3 and Appendix Figure A4.7).

Panels B and C of Table 4 show point estimates and standard errors of our baseline structural estimation, and Figure 4 provides a corresponding illustration. In Panel B of Table 4, we present our estimates of the preference parameter for the public good, $\delta_{m}$, for each household type. We find this preference parameter to increase with income among households with children, while decreasing with income among childless households. This is consistent with the local public good being a normal good for households with children but an inferior good for households without children. Importantly, these results suggest that researchers should not focus on heterogeneity across income or across demographic groups separately, but on the interaction between the two.

Implied structural elasticities are shown in Panel C of Table 4. The structural tax base...
elasticiies are reassuringly close to the reduced-form elasticities presented in Table 3, with high-income households without children the most strongly deterred by higher local taxes and bottom-50 households without children responding positively to higher taxes.

Panel C of Table 4 also reports the estimated marginal willingness to pay rent (MWPR) to compensate for higher taxes per group, as defined by equation 2. Our estimates are negative for five of the six household types; the exception being below-median income households without children, whose preferences for local public goods outweigh their disutility from higher tax burdens. Conversely, we obtain large negative estimates of the MWPR for top-25% households without children. Hence, at the margin, these households derive greater disutility from taxation and its effect on the cost of housing than the utility provided by local public goods. For households with children, however, we observe no relationship between income and the sign of the MWPR: all three income classes have a negative marginal willingness to pay rent, and the differences between them are not statistically significant.

Applying equation (11) allows us to compute household type-specific welfare effects of an increase in local taxes. These effects are reported as "resident incidence" in Table 4 and illustrated by the bars in the left-hand panel of Figure 4. Among residents, the negative incidence of local taxes is borne entirely by above-median income households without children. All other household types are either indifferent or would gain from a marginal increase in local taxes.
taxation and the associated local public goods. Averaged across households, the estimated incidence is slightly negative, suggesting local taxes to be above their optimum level. This average, however, is not statistically significantly different from zero, consistent with local tax rates being at their optimal level in utilitarian terms.

This result contrasts with the structural estimate of the housing price elasticity of \(-0.146\). In our model, this is entirely borne by absentee landlords. In reality, we estimate that around a third of landlords are resident in the same municipality.\(^5\) Private landlords very likely belong to the top-25% income class. Considering the effect on resident landlords would therefore exacerbate the negative incidence we estimate for the top-25% income class, but it would not qualitatively affect the different welfare effects we estimate across the six household types.

Appendix Figure A.4.5 shows corresponding results for a decomposition based on four income classes, where we divide the top income quartile into households with income between the 75th and the 90th percentile and those in the top income decile. These results corroborate the strongly progressive effects of local taxes, as the incidence among renters is significantly more negative again for top-10% income households without children. We find, however, the four-class specification to be quite sensitive to starting values used for the minimization process (see below). We thus focus on the three-class specification.

### 4.3 Robustness

Our baseline incidence estimates might be sensitive to three choices in particular: starting values employed for the iterative minimization of equation (15), calibrations of housing supply and idiosyncratic location preferences, and our assumption that households are ‘myopic’ with respect to the effect of location choices on the local supply of public goods. We now address these three issues in turn.

To explore the sensitivity of our results with respect to starting values of \(\delta_m\), we re-estimated the model for 1,000 different starting values drawn randomly from a uniform distribution \(\in [0, 0.5]\). Appendix Figure A.4.6 plots the average incidence and residents’ preference for the public good estimates plus/minus one standard deviation for our specifications with three classes (top panel) and with four classes (bottom panel).\(^5\) It turns out that starting values have next to no impact on our estimates for the three-class specification (top panel), as the structural estimates are distributed tightly around the values of our baseline estimates. The four-income-class specification is however quite sensitive to starting values (bottom panel).

In Figure 5, we investigate the sensitivity of our incidence estimates to different calibrated values for the housing supply elasticity \(\eta_{h,p}\) and the idiosyncratic location preference dispersion parameter \(\lambda\). Each dot in those graphs corresponds to a specific combination of \(\eta_{h,p}\) and \(\lambda\) chosen using an interval of plus/minus one standard error of their respective estimated values presented in Appendix Section A.2.2. Unsurprisingly, the incidence on households depends chiefly on the location preference dispersion parameter that governs their mobility, while the incidence on absentee landlords depends on the housing supply elasticity. More mobile households (higher \(\lambda\)) bear a lower burden (in absolute value) of local taxes. This

\(^{5}\)See https://www.bfs.admin.ch/bfsstatic/dam/assets/4262589/master.

\(^{54}\)The full distribution of the estimated residents’ preference parameters and of the chi-square statistic are reported in Appendix Figure A4.7 and A4.8.
is especially true for households without children. Note, however, that our results are not qualitatively affected by the calibration of $\eta^{sp}$ and $\lambda$: households with children are roughly indifferent to marginal changes in local income taxes, bottom-50% households without children benefit from local tax increases, whereas above-median households without children are hurt by higher taxes.

So far, our structural estimates were derived from the ‘myopic’ version of the model presented in Section 1. A fully fledged model that incorporates feedback effects between income classes and the production of the public good yields qualitatively unchanged estimates (see Appendix Table A3.4 and Appendix Figure A4.9). Quantitatively, the fully fledged model implies somewhat higher valuations of the public good for most household types. Notably, the incidence on top-quartile households with children turns positive. This is explained by the fact that this model takes into account the positive effect of lower housing prices on public-good provision through retaining residents (the off-diagonal elements in upper-block of matrix $A$). The fully fledged model, however, has the drawback of being less analytically tractable than our ‘myopic’ version. We cannot derive the first-order conditions of the minimization problem and therefore cannot analytically compute standard errors.
Figure 5: Households’ and landlords’ incidence for different values of $\eta^{\pi-p}$ and $\lambda$

<table>
<thead>
<tr>
<th>Households without children</th>
<th>Households with children</th>
<th>Landlords</th>
</tr>
</thead>
<tbody>
<tr>
<td><img src="image1.png" alt="Graph" /></td>
<td><img src="image2.png" alt="Graph" /></td>
<td><img src="image3.png" alt="Graph" /></td>
</tr>
<tr>
<td><img src="image4.png" alt="Graph" /></td>
<td><img src="image5.png" alt="Graph" /></td>
<td><img src="image6.png" alt="Graph" /></td>
</tr>
<tr>
<td><img src="image7.png" alt="Graph" /></td>
<td><img src="image8.png" alt="Graph" /></td>
<td><img src="image9.png" alt="Graph" /></td>
</tr>
</tbody>
</table>

Notes: Each graph is based on 10,000 different combinations of calibrated values for $\eta^{\pi-p}$ and $\lambda$ in a range given by an interval of plus/minus one standard error of their respective estimated values presented in Appendix Section A.2.2. The interval for $\lambda$ is truncated at 0.5 to avoid model instability. The black dots indicate baseline estimates. The z-axis scale is common to all panels.
Table 5: Tax base elasticities for pensioners and non-pensioners: 3SLS estimation

<table>
<thead>
<tr>
<th></th>
<th>Household types</th>
<th>3SLS estimation</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Non-pensioners</td>
<td></td>
</tr>
<tr>
<td></td>
<td>Pensioners</td>
<td></td>
</tr>
<tr>
<td>Income tax rate</td>
<td>Bottom 25</td>
<td>-0.114***</td>
</tr>
<tr>
<td></td>
<td>Next 25</td>
<td>-0.396***</td>
</tr>
<tr>
<td></td>
<td>Top 25</td>
<td>-0.928***</td>
</tr>
<tr>
<td></td>
<td>Bottom 25</td>
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</tr>
<tr>
<td></td>
<td>Next 25</td>
<td>-0.029</td>
</tr>
<tr>
<td></td>
<td>Top 25</td>
<td>-0.543***</td>
</tr>
<tr>
<td></td>
<td>Bottom 25</td>
<td>0.014***</td>
</tr>
<tr>
<td></td>
<td>Next 25</td>
<td>0.071***</td>
</tr>
<tr>
<td></td>
<td>Top 25</td>
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</tr>
<tr>
<td></td>
<td>Housing prices</td>
<td>-0.399***</td>
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<td>Controls</td>
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<td></td>
</tr>
<tr>
<td>Origin canton FE</td>
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<td></td>
</tr>
<tr>
<td># of observations</td>
<td>218</td>
<td></td>
</tr>
<tr>
<td># of municipalities</td>
<td>48</td>
<td></td>
</tr>
<tr>
<td>Instrument</td>
<td>Cantonal income tax rate differential</td>
<td></td>
</tr>
<tr>
<td>Estimator</td>
<td>3SLS</td>
<td></td>
</tr>
</tbody>
</table>

Notes: Heteroskedastic standard errors reported in parentheses. The equations are estimated jointly using three stage least squares. The sample consists of cross-canton pairs of municipalities with a pairing road distance of 10 km. Regressions are weighted by the log population in 2000 of the smallest municipality in the pair. The consolidated personal income tax rate differentials were instrumented by the cantonal personal income tax rate differentials. Controls include (time-invariant) indices of accessibility, exposure to natural risks, architectural heritage, and hours of sunlight. In column (1) we in addition control for topographical constraints and local administrative efficiency. Significance levels: ***p<0.01, **p<0.05, *p<0.1.

5 Extensions

5.1 Decomposition by age

So far, we have posited that income and family status were the key dimensions driving the heterogeneity of households’ valuation of local public goods. Another dimension likely to be important is age. In this subsection, we therefore divide childless households into pensioner (“old”) and non-pensioner (“young”) categories, based on a variable in the federal income tax statistics indicating whether households receive a pension. This variable is recorded with some inconsistencies, forcing us to clean the dataset and to exclude several years from the sample.55

Table 5 presents the three-stage least squares joint estimates for the nine household types and housing prices.56 Results for housing prices, for households with children and for pensioner households without children are similar to our baseline estimates reported Table 3, while estimates for non-pensioner households without children are all negative and lower in absolute value. While the estimates in Tables 5 and 3 are not exactly comparable, as the samples are different, they suggest that younger households are more strongly deterred by taxes than older households.

Results of the structural estimation are shown in Figure 6.57 Our estimated public-good preferences, shown in the right-hand panel, are similar for pensioner and non-pensioner

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55 One source of measurement error is that the pension variable includes invalidity benefits. On average, 9% of “pensioner” households are below the pension age (64 for women and 65 for men). The median age of invalidity benefit recipients is around 53. Another source of imprecision is that cantonal tax authorities have different reporting practices (especially for married couples) that in some cases change over time. The calculated share of pensioner households at the canton level can therefore jump between years by several percentage points (up to 13 percentage points, for an average pensioner share of 23%). We dropped observations for cantons where such jumps occurred at the beginning (Thurgau and Ticino) or end (St. Gallen) of our sample period, and where they affected a single year (Basel-Stadt). We also dropped observations for the canton of Vaud between 2005 and 2008 because of evident reporting errors. For the cantons in which discrete jumps happened in the middle of our observation period, we inferred for each municipality the number of pensioner households with the canton-level increase netted out (Geneva, Glarus, Fribourg, Solothurn, Valais). We moreover dropped observations within the top and bottom percentiles of the residuals from regressing the total number of pensioner households on municipality fixed effects and canton-year fixed effects.

56 Analogous to Table 2, Appendix Table A3.5 shows the corresponding OLS and 2SLS specifications for non-pensioners and pensioner households without children.

57 Details of the calibrated values are reported in Appendix Table A3.6.
Figure 6: Structural estimates by income class and age

Notes: In the left-hand-side panel, light gray bars represent the tax incidence experienced by non-pensioner households without children, gray bars represent the tax incidence experienced by pensioner households without children and the dark gray bars the tax incidence experienced by households with children. The black bar represents the group-size-weighted average of resident incidence. Incidence is the percent-change in resident welfare with respect to one percent increase in income taxes as defined by equation (11). Bar widths represent relative population sizes. Whiskers represent 95% confidence intervals.

households although there is some indication that high-income pensioners might have stronger tastes for local public goods than younger high-income households. The main difference with respect to our baseline results is the estimated tax incidence for the bottom-50% non-pensioner households (left-hand panel of Figure 6), which is now only slightly positive and similar to households with children. This suggests that the significantly positive tax incidence estimated for bottom-50 childless households in the baseline specification (Figure 6) may have been mainly driven by older households who are less mobile and more reliant on local public goods. These results by age are however only indicative, as, unlike in our baseline estimates, they are sensitive to starting values (see Appendix Figures A.4.10 and A.4.11).

5.2 Mobility

Up to now, we have tied the differential incidence of local tax changes to heterogeneous preferences for the local public good. We shall now investigate an alternative mechanism: heterogeneous mobility costs.

We structurally estimate the scale parameter $\lambda$ that governs residents’ idiosyncratic attachment to locations. We allow this parameter to vary freely across household types. To do so, we change two assumptions. First, we work with the fully fledged model that incorporates
feedback effects between income classes. When the focus is on heterogeneous mobility costs, this version of the model should be considered, because the mobility of one household type can affect location decisions of other types through the provision of the local public good. Second, we now force the public-goods preference parameter to be constant across types. We set its value to the average of the estimated parameters reported in Appendix Table A3.4, weighted by taxpayer population shares.

Appendix Figure A.12 reports the average incidence and scale parameter plus/minus one standard deviation resulting from estimating the model with 1,000 different starting values. These starting values are drawn from a uniform distribution \( \in [0.5, 3.5] \), analogously to Figure 5.58 These results suggest that mobility costs vary across family status rather than across income classes: not surprisingly, households with children are less mobile (lower \( \lambda \)) than households without children. However, especially the estimated \( \lambda \) parameter for the top-25\% households with children is implausibly small, as it implies next to no mobility for these households, even though these households bear higher local taxes.

Finally, we combine heterogeneous public-goods preferences with heterogeneous mobility costs. Similar to Figure 5, Appendix Figure A.14 reports the estimated incidence for different calibrated values of two idiosyncratic location preference dispersion parameters \( \lambda \), one for households without children and one for households with children. As the results show, the implied incidence depends on own mobility (the higher is \( \lambda \), the higher is mobility and the lower is the incidence), but not on the mobility of the other household type. Therefore, assuming a constant \( \lambda \) across household types does not appear to be a strong constraint. Especially, assuming lower mobility for households with children than for childless households, would not qualitatively change our findings.59

### 5.3 The local budget constraint

In this paper, identification of the distributional effects of local tax changes relies on a key simplification: the municipal budget is balanced, such that there exists a one-to-one relationship between increased tax revenue following a local income tax hike and increased availability of the local public good. In reality, changes in tax revenue might not always map one-for-one into changes in public good provision, e.g. in the presence of public-sector rent extraction or corruption (Diamond, 2017), or in the case of net public (dis-)saving (Pettersson-Lidbom, 2010). Here, we therefore employ our IV strategy to test for the existence of a balanced budget constraint at the municipality level.

Our identification strategy exploits upper-level tax changes in neighboring cantons as a source of exogenous variation for consolidated tax differentials. Hence, an exogenous increase in the consolidated tax differential between two municipalities located in adjacent cantons is driven by a decrease in the neighboring cantonal tax rate. Consistent with our negative tax base elasticities for top-income households, we expect that higher tax differentials lead to a worsening in municipal tax revenue (differentials). This is what we find in the first column.

58The full distributions of the estimated scale parameters and of the chi-square statistic are reported in Appendix Figure A.13.

59This can be seen by moving to the south-east in the first six panels of Appendix Figure A.14. Incidence estimates remain remarkably stable in five out of the seven panels. Only in the first and fourth panels, the incidence would be different than our baseline (and, as expected, would be higher the lower is mobility).
of Table 6, which reports the results of our baseline distributed-lag (Panel A) and long first difference (Panel B) models for total municipal revenue.\textsuperscript{60}

In column (2), we test the effect on total expenditure. Importantly, we find an effect that is quantitatively similar to column (1), consistent with a binding local budget constraint.\textsuperscript{61}

In columns (3) to (8) we test for (endogenous) changes in the composition of expenditure. Results are unfortunately not informative, as standard errors are large and often exceed the estimated coefficients. Note that the effect on educational expenditure (column 3) is among the less imprecisely estimated coefficients and is consistent across both panels. These results therefore suggest that lower tax revenue is associated with a decrease in educational spending but to a lesser extent than other categories.

6 Conclusion

This paper studies the differential welfare effects of local income taxes on heterogeneous households and absentee landlords. This question is important for three reasons. First, according to the standard assumption of perfectly mobile local-level residents, land, the immobile factor, bears the full incidence of local policies. However, residential mobility is costly, even at the local level, and hence welfare effects on resident non-owners need to be considered too. Second, preferences for local public goods have hitherto been studied without considering heterogeneity. In this paper, we show that these preferences differ substantially across income classes and family types. Third, studies of taxpayer mobility typically focus on

\textsuperscript{60}In all specifications of Table 6, we use as dependent variable the residuals from a regression on canton-year fixed effects in order to take into account canton-level changes in public accounting standards for municipality finances as well as changes of task allocations between different levels of governments.

\textsuperscript{61}Note that total expenditure is also directly affected by the change in the composition of the tax base. Hence, we cannot directly test the local budget constraint as expressed in equation (8c). Note also that we cannot use total expenditure directly in our structural model, as we do not know how much cantons spend in a given municipality.

Table 6: Public good elasticities

<table>
<thead>
<tr>
<th>Panel A: Distributed lag model</th>
<th>Total Revenue (1)</th>
<th>Total Expenditure (2)</th>
<th>Education (3)</th>
<th>Social (4)</th>
<th>Admin. (5)</th>
<th>Roads (6)</th>
<th>Police (7)</th>
<th>Health (8)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Income tax rate (cumulative effect)</td>
<td>-0.407**</td>
<td>-0.342**</td>
<td>-0.257</td>
<td>-0.314</td>
<td>-0.190</td>
<td>-0.002</td>
<td>-0.669</td>
<td>-0.898</td>
</tr>
<tr>
<td># of observations</td>
<td>13,702</td>
<td>13,702</td>
<td>13,702</td>
<td>13,702</td>
<td>11,522</td>
<td>13,690</td>
<td>11,464</td>
<td></td>
</tr>
<tr>
<td>Kleibergen-Paap F Stat</td>
<td>398</td>
<td>398</td>
<td>398</td>
<td>398</td>
<td>325</td>
<td>398</td>
<td>325</td>
<td></td>
</tr>
<tr>
<td>Panel B: Long difference model between the averages 2013-2014 and 2004-2005</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Income tax rate</td>
<td>-0.367**</td>
<td>-0.556***</td>
<td>-0.256</td>
<td>-0.965</td>
<td>-0.132</td>
<td>-0.084</td>
<td>-0.284</td>
<td>-0.622</td>
</tr>
<tr>
<td># of observations</td>
<td>1,398</td>
<td>1,398</td>
<td>1,398</td>
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</tr>
<tr>
<td>Kleibergen-Paap F Stat</td>
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<td>351</td>
<td>296</td>
<td>351</td>
<td>296</td>
<td></td>
</tr>
</tbody>
</table>

Notes: This table reports the results of the IV pairwise difference estimation model for border municipalities. Specifications of Panels A and B correspond to Panel E and G of Table 2, respectively. Cluster robust standard errors at origin and destination municipality level are reported in parentheses. Regressions in Panel A employ a standard distributed lag approach by estimating $\ln y_{ijkt} = \beta_0 + \sum\limits_{s=1}^{4} \beta_s \ln \tau_j + \phi_{jk} + \phi_{ct} + \phi_{jkt}$. Hence, one can interpret $\hat{\beta}$ directly as the long-term effect. Controls in Panel B include (time-invariant) indices of accessibility, exposure to natural risks, architectural heritage, and hours of sunlight. **p < 0.01, ***p < 0.001.

3 Column 5 reports the result of using canton fixed effects instead of canton-year fixed effects. In columns 3 and 4, the estimated coefficients are quantitatively similar. The estimated effect on total expenditure is consistent across both panels. However, the estimated effect on education is less precise in column 3. In columns 5 to 8, the estimated coefficients are consistent across both panels. These results therefore suggest that lower tax revenue is associated with a decrease in educational spending but to a lesser extent than other categories.
top-income households. We also consider taxpayers in lower brackets of the income distribution, and we link type-specific tax base elasticities to measures of households’ willingness to trade off housing costs against local tax burdens.

We estimate the incidence of local tax changes in three main steps. First, we allow for heterogeneous households to have different valuations for locally provided public goods, funded through proportional local income taxes. Second, we exploit cross-section and time variation in Swiss municipal tax rates at canton borders that we instrument with neighboring canton-level tax rates. This enables us to obtain plausibly causal reduced-form elasticities of tax bases and housing prices with respect to tax rates, and of housing demand and supply elasticities with respect to housing prices. Third, we search for the preference parameters that best match our theoretical moments with the reduced-form elasticities.

We find large variation in the incidence of local income tax increases: for childless households, the incidence ranges from strongly positive for the bottom-50% income households – driven, it turns out, largely by pensioner households – to strongly negative for the top-25% and top-10% income households. For households with children, the estimated incidence is approximately zero across all income classes. Our structural estimates imply a positive relationship between preferences for local public goods and income among households with children. This relationship is reversed among households without children.

Our results show that local taxation has distributional effects even in the absence of a progressive rate schedule, for two reasons. First, to the extent that households exhibit non-homothetic housing demand, the capitalization of tax rates into housing prices will affect them differently; and, second, heterogeneous preferences for publicly provided goods imply that different households perceive local tax changes differently. This might help explain the absence of empirical evidence for perfect sorting of households across income lines: households at different income levels will differ significantly in their valuation of local bundles of tax rates and public goods depending on their demographic type.

Our analysis is predicated on the implicit assumption that residents continuously update their optimal location choice. In reality, residential moves are infrequent. The average tenancy of renter households in Switzerland is approximately 6 years, and that of owner occupiers is 16 years. Together with the fact that our elasticity estimates are identified from ten-year changes, this implies that our estimates ought to be interpreted as long-run welfare effects.

One striking result of our analysis is the low estimated preference weight of locally provided public goods among relatively poor households with children. Given that these households probably stand to gain most from municipal services – especially primary schools – this result is counterintuitive. One conceivable cause are information frictions: households in that class might face higher information costs about local public goods than other households.

Another area where more detailed data could offer deeper insights is the mapping between household income and housing properties. In this paper we consider housing as a homogeneous good, implying that all household types operate on a single integrated housing market. As a further step toward capturing heterogeneity in all relevant dimensions, it

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62 Note that, unlike e.g. in the United States, much of expenditure relevant to seniors, such as care homes and home help, is decided on and funded at the local level.

63 Own calculations based on Wüest Partner data.
would be useful to allow for different housing-market segments to have unequal relevance across household types. We leave such a possible extension for future work as well.

References


A Appendix

A.1 Schematic overview

Section 3
Tax base and rental price elasticities with respect to tax rates

Endogenous variables (Equation 18):
- Tax payer counts (N)
- Rental price (P)

Exogenous variables:
- Consolidated income tax rates (r),
- Amenities (X),
- Share of developed land (SDL),
- Administrative efficiency (TTP)

Results (joint estimation in Table 3):
- Tax base elasticities with respect to tax rates (\( \eta # \)),
- Rental price elasticity (with respect to tax rates \( \eta $ \)).

Section 2
Theoretical model

Housing Demand (Equation 1):
- Imperfectly mobile renters
- Heterogeneous in their income and preference for local public good

Housing Supply (Equation 7):
- Absentee landlords
- Constant upward-sloping housing supply

Local government (Equation 8c):
- Homogeneous public good financed through a proportional income tax (balanced budget)

Equilibrium conditions

Log-differentiation

Express changes in population and housing prices as function of changes in tax rates (system of equations)

Results (joint estimation in Table 3):
- Tax base elasticities with respect to tax rates (\( \eta ^\text{P} \)),
- Rental price elasticity (with respect to tax rates \( \eta ^\text{F} \)).

Appendix Section A.2
Housing demand and housing supply price elasticities

Endogenous variables (Equation A.1):
- Rental prices (F)
- Dwelling space (H)

Exogenous variables:
- Consolidated income tax rates (r),
- Amenities (X),
- Share of developed land (SDL),
- Administrative efficiency (TTP)

Simultaneous Equation IV

Identifying variation: cross-canton border spatial difference

Instrument: cantonal tax differentials

Parameters:
- Lambda (\( \lambda \)) using (\( \eta ^\text{P}, \eta ^\text{F} \)),
- Housing supply elasticity (\( \eta ^\text{F,H} \))

Calibration (Table 4):
Swiss Household Panel:
- Housing tenures (t),
- Minimal housing expenditures (\( t_{\text{min}} \)),
- Expenditure shares on housing (\( s_{\text{min}} \)),
- Aggregate housing shares (\( s_{\text{min}} \)),

Tax rate database:
- Average income tax rates (\( t_{\text{avg}} \)),

Tax base database:
- Taxpayer population shares (\( s_{\text{avg}} \))

3SLS

Identifying variation: cross-canton border spatial difference

Instrument: cantonal tax differentials

Empirical moments:
- Reduced form elasticities (\( \beta \)),
- Variance-covariance matrix (\( \Omega \))

Theoretical moments (\( \mu (\theta) \))

Section 5
Structural estimation

Section 3
Tax base and rental price elasticities with respect to tax rates

Endogenous variables (Equation 18):
- Tax payer counts (N)
- Rental price (P)

Exogenous variables:
- Consolidated income tax rates (r),
- Amenities (X),
- Share of developed land (SDL),
- Administrative efficiency (TTP)

Results (joint estimation in Table 3):
- Tax base elasticities with respect to tax rates (\( \eta # \)),
- Rental price elasticity (with respect to tax rates \( \eta $ \)).

Resident welfare

Public good preferences

Preference parameter for the local public good (\( \delta_m \))

\( Q1 \& Q2 \)
\( Q3 \)
\( Q4 \)
\( \text{w/o children} \)
\( \text{with children} \)
\( \text{wgt sum} \)

Income distribution percentile

41
A.2 Housing supply and demand

Here, we describe our estimation of the price elasticity of housing supply and housing demand, two parameters required for our structural estimation. We use changes in local income tax rates as a demand shifter allowing us to identify supply responses. In turn, cross-sectional differences in topographical constraints and administrative efficiency are taken as supply shifters, allowing us to identify the price elasticity of housing demand, which will be key to calibrating the idiosyncratic location preference dispersion parameter \( \lambda \) (see footnote 50).

A.2.1 A simultaneous-equation IV framework

Our starting point is the following simultaneous-equation model for a cross-section of municipalities \( j \):

\[
\Delta \ln P_j = \frac{1}{\eta^p} \Delta \ln H_j + \eta^p \Delta \ln \tau_j + \mu_j + \phi_c + \epsilon^d_j, \quad (A.1a)
\]

\text{and}

\[
\Delta \ln P_j = \frac{1}{\eta^p} \Delta \ln H_j + \beta_1 SDL_j + \beta_2 TTP_j + \mu_j + \phi_c + \epsilon^s_j, \quad (A.1b)
\]

where \( \Delta \) represents long first differences. \( P \) denotes residual housing prices, \( H \) the residential housing stock, \( \tau \) the personal income tax rate, \( X \) is a vector of local amenities, \( SDL \) is the share of developed land, \( TTP \) (“time to permit”) is a proxy for local administrative efficiency, and \( \phi_c \) are canton fixed effects.

The model described by equations (A.1a) and (A.1b) identifies the elasticities of housing demand \((\eta^d, p)\) and of housing supply \((\eta^s, p)\), contingent on a set of exclusion restrictions and validity conditions.

The exclusion restrictions we impose are that housing demand shifters do not affect housing supply, and that housing supply shifters do not affect housing demand. That is, we need that \( \text{cov}(\Delta \ln \tau, \epsilon^s) = \text{cov}(SDL, \epsilon^d) = \text{cov}(TTP, \epsilon^d) = 0 \).

One concern is that changes in local income tax rates \( \Delta \ln \tau \) could also lead to shifts in the supply curve. The atomistic absentee landlord described in Section 1.2 differs from our empirical setting insofar as rental income in Switzerland is taxed by the jurisdiction where the dwelling is located. We show in Online Appendix W.2 that the supply side of the model is independent of changes in income taxes if landlords’ running costs are tax deductible or taxed at the same rate as income. While mortgage interest, property tax payments and maintenance costs can be deducted from income taxes in Switzerland, transaction taxes are not deductible, and capital gains are in some places taxed at a different rate than the income tax. We exploit the heterogeneity in tax laws across Swiss cantons to filter out jurisdictions where changes in income tax rates are statutorily linked to changes in taxes that affect supply. Specifically, we replace \( \Delta \ln \tau_j \) in (A.1a) by a vector \( \Delta \ln \tau_j = [\Delta \ln \tau_j \quad \Delta \ln \tau_j \times NCM \quad \Delta \ln \tau_j \times PT \quad \Delta \ln \tau_j \times TT] \) and \( \eta^p = [\eta^{dp} \quad \eta^{dp \times NCM} \quad \eta^{dp \times PT} \quad \eta^{dp \times TT}] \). Dummy variables indicate the cantons in which municipalities are not restricted to use the same multiplier for capital gains and personal income taxes (\( NCM \), for no common multiplier), and cases in which municipalities have autonomy to set property tax rates (\( PT \) and transaction tax rates...
(TT). The main effect $\Delta \ln \tau_j$ then measures the effect of local income taxes as measured in jurisdictions where changes in these taxes directly affect housing demand but not housing supply.

The second concern is that topographical constraints and local administrative efficiency, our two supply shifters, are correlated with location-specific features, such as nice views, sunlight or accessibility, that also affect housing demand growth (see, e.g., Davidoff, 2016). We address this concern by controlling extensively for municipality-level differences in amenities $X$. Moreover, our supply shifters are measured in 1985 for the share of developed land and 1997-2003 for the building permits, several years prior to our sample period for housing prices and housing stocks.

Valid identification furthermore requires that the shifters be exogenous to the system of equations, i.e. $\text{cov}(\Delta \ln \tau, \epsilon^d) = \text{cov}(\text{SDL}, \epsilon^s) = \text{cov}(\text{TT}, \epsilon^s) = 0$. Topography is used to define the amount of ‘developable’ land, and the amount of developed land is measured several years prior to our sample period, as is $\text{TT}$. Therefore $\text{SDL}$ and $\text{TT}$ are considered plausibly exogenous.

We however expect that local tax rates are endogenous with respect to local housing demand, in first-differences as well as in levels (i.e. $\text{cov}(\Delta \ln \tau, \epsilon^s) \neq 0$). To address the endogeneity of the tax rate, we turn to three separate estimation steps on our sample of border municipalities.

First, we take cross-cantonal spatial differences of the demand equation (A.1a), yielding

$$\nabla \Delta \ln P_{jk} = \frac{1}{\eta^{dp}} \nabla \Delta \ln H_{jk} + \eta^p \nabla \Delta \ln \tau_{jk} + \mu \nabla X_{jk} + \phi_c + \epsilon_{jk},$$

(A.2)

where $\nabla$ indicates the cross-cantonal spatial difference within pairs of municipalities $jk$ in two neighboring cantons, $c$ and $d$, with $(j \in c) \neq (k \in d \neq c)$. We instrument $\nabla \Delta \ln H_{jk}$ with $\nabla \text{SDL}_{jk}$ and $\nabla \text{TT}_{Pjk}$ and the vector $\nabla \Delta \ln \tau_{jk}$ with the equivalent vector of canton tax rate differentials $\nabla \Delta \ln \tau_{cd}$ (including interaction terms). Standard errors are two-way clustered at origin and destination municipality level.

Second, we back out the implied housing supply elasticity by estimating the following reduced-form equations separately,

$$\nabla \Delta \ln H_j = \eta^s \nabla \Delta \ln \tau_{jk} + \beta_1 \nabla \text{SDL}_{jk} + \beta_2 \nabla \text{TT}_{Pjk} + \mu \nabla X_{jk} + \phi_c + \epsilon_{jk},$$

(A.3)

$$\nabla \Delta \ln P_{jk} = \eta^p \nabla \Delta \ln \tau_{jk} + \beta_1 \nabla \text{SDL}_{jk} + \beta_2 \nabla \text{TT}_{Pjk} + \mu \nabla X_{jk} + \phi_c + \epsilon_{jk}$$

(A.4)

The vector $\nabla \Delta \ln \tau_{jk}$ is instrumented with the vector $\nabla \Delta \ln \tau_{cd}$. The parameter vectors are $\eta^{s} = [\eta^{s}, \eta^{s \times \text{nem}}, \eta^{s \times \text{pt}}, \eta^{s \times \text{lt}}]$, $\eta^{p} = [\eta^{p}, \eta^{p \times \text{nem}}, \eta^{p \times \text{pt}}, \eta^{p \times \text{lt}}]$ and coefficients of interest

---

64We constructed the variables $\text{nem}, \text{pt}$ and $\text{tt}$ in the cross-cantonal setting as follows. The binary variable $\text{nem}$ is equal to one if both cantons in a pair require that capital gain taxes vary with personal income taxes, and zero otherwise. The binary variables $\text{pt}$ is equal to one if both cantons in a pair do not allow property taxes to be levied, and zero otherwise. The binary variable $\text{tt}$ is equal to one if municipalities, in the canton pair, can levy a transaction tax or if the transaction tax is set at the cantonal level, and zero otherwise.
Table A2.1: Demand equation IV estimates

<table>
<thead>
<tr>
<th></th>
<th>Spatial difference of rental price growth rate</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
</tr>
<tr>
<td>Housing stock ($\hat{\eta}_d$)</td>
<td>-1.056** (0.519)</td>
</tr>
<tr>
<td>Amenity controls</td>
<td>YES</td>
</tr>
<tr>
<td>Fiscal controls</td>
<td>YES</td>
</tr>
<tr>
<td>Origin canton FE</td>
<td>YES</td>
</tr>
<tr>
<td># of observations</td>
<td>2,004</td>
</tr>
<tr>
<td># of origin clusters</td>
<td>524</td>
</tr>
<tr>
<td># of dest. clusters</td>
<td>524</td>
</tr>
<tr>
<td>Instrument</td>
<td>SDL, TTP</td>
</tr>
<tr>
<td>Kleibergen-Paap F Stat</td>
<td>7.37</td>
</tr>
<tr>
<td>Estimator</td>
<td>2SLS</td>
</tr>
</tbody>
</table>

Notes: Two-way cluster robust standard errors at origin and destination municipality level. The sample consist of cross-canton pairs of municipalities with a pairing road distance of 10 km. Regressions are weighted by the log population in 2000 of the smallest municipality in the pair. Housing demand elasticity coefficient has already been transformed for direct interpretation. Amenity controls include indices of accessibility, exposure to natural risks, architectural heritage, and hours of sunlight. Fiscal controls include the interactions between the income tax rate and dummy variables NCM, PT, and TT. ***p<0.01, **p<0.05, *p<0.1.

are $\eta^s$ and $\eta^p$, respectively. The implied housing supply elasticity is given by

$$\hat{\eta}_s^{\delta p} = \frac{\hat{\eta}_s}{\hat{\eta}_p},$$

where standard errors can be calculated using the delta method.

A.2.2 Results

The results of estimating equation (A.2) are presented in Table A2.1. Columns (1) and (2) report the estimation first without and then with canton-level tax rates as instruments. Our main object of interest, the estimated price elasticity of housing demand, equals around -1, irrespective of whether tax rates are instrumented or not.\(^{65}\) Note that a unit elasticity implies zero residential mobility in a model with Cobb-Douglas preferences over housing and non-housing goods.

Table A2.2 presents our estimates of the housing supply elasticity. Columns (1) and (2) show the OLS and 2SLS estimations of equation (A.3), while columns (3) and (4) show the OLS and 2SLS estimations of equation (A.4). Taking the ratio of the point estimates of columns (2) and (4) yields an implied IV estimate of the housing price elasticity of housing supply.

\(^{65}\) First-stage Kleibergen-Paap F-statistics are low, indicating a potential weak first-stage issue. Estimating the simultaneous-equation model using the full set of municipalities leads to a housing demand price elasticity of -1.3, with a standard error of 0.61 (see Appendix Table A3.3).
Table A2.2: Supply equation IV estimates

<table>
<thead>
<tr>
<th></th>
<th>Spatial difference of dwelling space growth rate</th>
<th>Spatial difference of rental price residual growth rate</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
</tr>
<tr>
<td>Dwelling space elasticity of</td>
<td>-1.240***</td>
<td>-0.695</td>
</tr>
<tr>
<td>income taxes ($\hat{\eta}_s$)</td>
<td>(0.314)</td>
<td>(0.429)</td>
</tr>
<tr>
<td>Rental price elasticity of</td>
<td></td>
<td></td>
</tr>
<tr>
<td>income taxes ($\hat{\eta}_p$)</td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>-1.405***</td>
<td>-0.951*</td>
</tr>
<tr>
<td></td>
<td>(0.388)</td>
<td>(0.557)</td>
</tr>
</tbody>
</table>

Implied housing supply elasticity ($\hat{\eta}_{s,p}$)$_{OLS}$: 0.883*** (0.331)
Implied housing supply elasticity ($\hat{\eta}_{s,p}$)$_{IV}$: 0.731 (0.622)

<table>
<thead>
<tr>
<th></th>
<th>YES</th>
<th>YES</th>
<th>YES</th>
<th>YES</th>
</tr>
</thead>
<tbody>
<tr>
<td>Amenity controls</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
</tr>
<tr>
<td>Fiscal controls</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
</tr>
<tr>
<td>Origin canton FE</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
<td>YES</td>
</tr>
<tr>
<td># of observations</td>
<td>2,004</td>
<td>2,004</td>
<td>2,004</td>
<td>2,004</td>
</tr>
<tr>
<td># of origin clusters</td>
<td>524</td>
<td>524</td>
<td>524</td>
<td>524</td>
</tr>
<tr>
<td># of dest. clusters</td>
<td>524</td>
<td>524</td>
<td>524</td>
<td>524</td>
</tr>
<tr>
<td>Instrument</td>
<td>Canton tax differential</td>
<td>Canton tax differential</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Kleibergen-Paap F Stat</td>
<td>8.51</td>
<td>8.99</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Estimator</td>
<td>OLS</td>
<td>2SLS</td>
<td>OLS</td>
<td>2SLS</td>
</tr>
</tbody>
</table>

Notes: Two-way cluster robust standard errors at origin and destination municipality level in parentheses. The sample consist of cross-canton pairs of municipalities with a pairing road distance of 10 km. Regressions are weighted by the log population in 2000 of the smallest municipality in the pair. Columns (1) and (2) come from the estimation of equation (A.3), while columns (3) and (4) come from the estimation of equation (A.4). The implied housing supply elasticity ($\hat{\eta}_{s,p}$)$_{OLS}$ comes from the ratio of point estimate in column (1) and column (3). The implied housing supply elasticity ($\hat{\eta}_{s,p}$)$_{IV}$ comes from the ratio of point estimate in column (2) and column (4). The corresponding standard errors are calculated using the delta method. Amenity controls include indices of accessibility, exposure to natural risks, architectural heritage, and hours of sunlight. Fiscal controls include the interactions between the income tax rate and dummy variables NCM, PT, and TT. ***p<0.01, **p<0.05, *p<0.1.

The OLS and IV estimates of the implied elasticities are again comparable and statistically significantly different from each other. The IV point estimate equals 0.7. We retain this value for our estimation of the structural parameters.
A.3 Supplementary tables

Table A3.3: Simultaneous equation estimates

<table>
<thead>
<tr>
<th></th>
<th>Rental price growth rate</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
</tr>
<tr>
<td>Demand equation (A.1a):</td>
<td></td>
</tr>
<tr>
<td>Housing stock ($\hat{\eta}^d_{p}$)</td>
<td>-1.558</td>
</tr>
<tr>
<td></td>
<td>(1.464)</td>
</tr>
<tr>
<td>Local income tax ($\hat{\eta}^{p,r}$)</td>
<td>-0.400***</td>
</tr>
<tr>
<td></td>
<td>(0.142)</td>
</tr>
<tr>
<td>Supply equation (A.1b):</td>
<td></td>
</tr>
<tr>
<td>Housing stock ($\hat{\eta}^{s,p}$)</td>
<td>0.594***</td>
</tr>
<tr>
<td></td>
<td>(0.211)</td>
</tr>
<tr>
<td>Share of developed land ($\hat{\beta}_1$)</td>
<td>0.082***</td>
</tr>
<tr>
<td></td>
<td>(0.029)</td>
</tr>
<tr>
<td>Time-to-permit ($\hat{\beta}_2$)</td>
<td>0.000</td>
</tr>
<tr>
<td></td>
<td>(0.007)</td>
</tr>
<tr>
<td>Canton FE</td>
<td>YES</td>
</tr>
<tr>
<td>Amenity controls</td>
<td>NO</td>
</tr>
<tr>
<td>Fiscal controls</td>
<td>NO</td>
</tr>
<tr>
<td># of observations</td>
<td>1,815</td>
</tr>
</tbody>
</table>

Notes: Standard errors in parentheses. Weighted by log municipal population in 2000. Housing demand and supply elasticities have already been transformed for direct interpretation. Amenity controls include indices of accessibility, exposure to natural risks, architectural heritage, and hours of sunlight. Fiscal controls include the interactions between the income tax rate and dummy variables NCM, PT, and TT. ***p<0.01, **p<0.05, *p<0.1.
### Table A3.4: Structural parameter and elasticity estimates: fully-fledged model

<table>
<thead>
<tr>
<th></th>
<th>Households without children</th>
<th></th>
<th>Households with children</th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Bottom 50 (1)</td>
<td>Next 25 (2)</td>
<td>Top 25 (3)</td>
<td>Bottom 50 (4)</td>
</tr>
<tr>
<td><strong>Panel A: Calibration using:</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Swiss Household Panel</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Housing tastes ( (\alpha) )</td>
<td>0.11</td>
<td>0.11</td>
<td>0.11</td>
<td>0.11</td>
</tr>
<tr>
<td>Minimal housing expenditure ( (\nu_{i}/h_{m}) )</td>
<td>0.75</td>
<td>0.69</td>
<td>0.57</td>
<td>0.80</td>
</tr>
<tr>
<td>Expenditure share on housing ( (\delta_{i}) )</td>
<td>0.38</td>
<td>0.24</td>
<td>0.17</td>
<td>0.36</td>
</tr>
<tr>
<td>Aggregate housing share ( (\tau_{i}) )</td>
<td>0.13</td>
<td>0.14</td>
<td>0.17</td>
<td>0.16</td>
</tr>
<tr>
<td>Municipal Tax Rate Database</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Income tax rates ( (\tau_{i}) )</td>
<td>0.12</td>
<td>0.16</td>
<td>0.24</td>
<td>0.05</td>
</tr>
<tr>
<td>Simultaneous Equation IV Estimates</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Idiosyncratic location preference dispersion parameter ( (\lambda) )</td>
<td>1.93</td>
<td>1.93</td>
<td>1.93</td>
<td>1.93</td>
</tr>
<tr>
<td>Housing supply price elasticity ( (\eta^{s}) )</td>
<td>0.73</td>
<td>0.73</td>
<td>0.73</td>
<td>0.73</td>
</tr>
<tr>
<td>Tax Base Database</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Taxpayer population share ( (s_{m}) )</td>
<td>0.44</td>
<td>0.17</td>
<td>0.13</td>
<td>0.06</td>
</tr>
<tr>
<td>Share of tax base ( (\gamma_{m}) )</td>
<td>0.17</td>
<td>0.19</td>
<td>0.30</td>
<td>0.02</td>
</tr>
<tr>
<td><strong>Panel B: Structural parameters</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Preference for public goods ( (\delta_{m}) )</td>
<td>0.291</td>
<td>0.246</td>
<td>0.111</td>
<td>0.014</td>
</tr>
<tr>
<td><strong>Panel C: Structural elasticities</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Tax base elasticities</td>
<td>0.286</td>
<td>-0.069</td>
<td>-0.689</td>
<td>0.011</td>
</tr>
<tr>
<td>Marginal willingness to pay rent</td>
<td>0.343</td>
<td>-0.110</td>
<td>-1.482</td>
<td>-0.131</td>
</tr>
<tr>
<td>Resident incidence</td>
<td>0.148</td>
<td>-0.036</td>
<td>-0.357</td>
<td>0.006</td>
</tr>
<tr>
<td>Landlord incidence ( (\eta_{p}^{r}) )</td>
<td>-0.146</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: Overall renter incidence is the weighted sum of the income class specific incidences using the corresponding relative population shares. We set \( \theta = 1 \) throughout.
### Table A3.5: Tax base elasticities for households without children: pensioners and non-pensioners

<table>
<thead>
<tr>
<th></th>
<th>Households without children</th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
<th></th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Non-pensioners</td>
<td>Pensioners</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td></td>
<td>(Bottom 50)</td>
<td>(Next 25)</td>
<td>(Top 25)</td>
<td>(Bottom 50)</td>
<td>(Next 25)</td>
<td>(Top 25)</td>
<td>(Bottom 50)</td>
</tr>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
<td>(4)</td>
<td>(5)</td>
<td>(6)</td>
<td>(7)</td>
</tr>
<tr>
<td><strong>Fixed effect panel model</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Income tax rate</td>
<td>-0.123***</td>
<td>-0.401***</td>
<td>-0.683***</td>
<td>0.028</td>
<td>-0.171**</td>
<td>-0.341***</td>
<td></td>
</tr>
<tr>
<td>(0.059)</td>
<td>(0.069)</td>
<td>(0.071)</td>
<td></td>
<td>(0.073)</td>
<td>(0.081)</td>
<td>(0.139)</td>
<td></td>
</tr>
<tr>
<td># of observations</td>
<td>17,283</td>
<td>17,283</td>
<td>17,283</td>
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<tr>
<td><strong>Panel B: OLS estimation on border municipalities</strong></td>
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<tr>
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<td>-0.597***</td>
<td>-0.072</td>
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<tr>
<td><strong>Panel C: OLS pairwise difference estimation on border municipalities</strong></td>
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<tr>
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<td>-0.659***</td>
<td>0.096**</td>
<td>-0.094</td>
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<td>(0.048)</td>
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<tr>
<td>Origin canton-year fixed effect</td>
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<td><strong>Panel D: IV pairwise difference estimation on border municipalities</strong></td>
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<td># of municipalities</td>
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<tr>
<td>Origin canton-year fixed effect</td>
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<tr>
<td>Instrument</td>
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<tr>
<td><strong>Panel E: IV pairwise difference estimation on border municipalities: distributed log model</strong></td>
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<tr>
<td>Income tax rate (cumulative effect)</td>
<td>0.143</td>
<td>-0.440***</td>
<td>-0.716***</td>
<td>0.043</td>
<td>-0.221*</td>
<td>-0.379*</td>
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<tr>
<td>(0.129)</td>
<td>(0.188)</td>
<td>(0.199)</td>
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<td>(0.070)</td>
<td>(0.113)</td>
<td>(0.200)</td>
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<tr>
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<tr>
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<tr>
<td>Instrument</td>
<td>Cantonal income tax rate differential</td>
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<td></td>
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**Notes:** Cluster robust standard errors reported in parentheses. In panels A and B, standard errors are two-way clustered at origin and destination municipality level. In the remaining panels, standard errors are two-way clustered at origin and destination municipality level. In municipalities with zero taxpayer in a given category, \( \ln(0) \) has been replaced by 0. Regressions in panel E employ a standard distributed log approach estimating \( \text{ln} \ u_{ijkt} = \eta \text{ln} \ u_{ijkt}^{\text{Top}} + \sum_{t=1}^{T} \delta_t (\text{ln} \ u_{ijkt} - \text{ln} \ u_{ijkt}^{\text{Top}}) + \xi_{ijkt} + \varepsilon_{ijkt} \), so that we may interpret \( \hat{\tau}_{ijkt} \) directly as the long-term effect. Controls in panel C include (time-invariant) indices of accessibility, exposure to natural risks, architectural heritage, and hours of sunlight. *** = p < 0.01; ** = p < 0.05; * = p < 0.1.
Table A3.6: Structural parameter and elasticity estimates: pensioners and non-pensioners

<table>
<thead>
<tr>
<th>Panel A: Calibration using:</th>
<th>Non-pensioners</th>
<th>Pensioners</th>
<th>Non-pensioners</th>
<th>Pensioners</th>
<th>Non-pensioners</th>
<th>Pensioners</th>
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<tr>
<td>Housing tastes ( (\alpha) )</td>
<td>0.12</td>
<td>0.12</td>
<td>0.12</td>
<td>0.12</td>
<td>0.12</td>
<td>0.12</td>
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<tr>
<td>Minimal housing expenditure ( (x_a / x_n) )</td>
<td>0.75</td>
<td>0.71</td>
<td>0.59</td>
<td>0.70</td>
<td>0.60</td>
<td>0.50</td>
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<td>Expenditure share on housing ( (x_n) )</td>
<td>1.37</td>
<td>1.04</td>
<td>0.17</td>
<td>0.40</td>
<td>1.24</td>
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<td>Aggregate housing share ( (x_n) )</td>
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<td>0.10</td>
<td>0.12</td>
<td>0.09</td>
<td>0.11</td>
<td>0.12</td>
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<td>Tax Rate Database</td>
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<tr>
<td>Income tax rates ( (\tau) )</td>
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<td>0.17</td>
<td>0.25</td>
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<td>Idiosyncratic location preference dispersion parameter (( \lambda ))</td>
<td>1.84</td>
<td>1.84</td>
<td>1.84</td>
<td>1.84</td>
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<tr>
<td>Housing supply price elasticity ( (\eta) )</td>
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<td>0.73</td>
<td>0.73</td>
<td>0.73</td>
<td>0.73</td>
<td>0.73</td>
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<td>Tax Base Database</td>
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<tr>
<td>Taxpayer population share ( (\eta) )</td>
<td>0.28</td>
<td>0.11</td>
<td>0.09</td>
<td>0.16</td>
<td>0.06</td>
<td>0.05</td>
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<td>Panel B: Structural parameters</td>
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<tr>
<td>Preference for public goods ( (\mu) )</td>
<td>0.128***</td>
<td>0.091**</td>
<td>0.027</td>
<td>0.144***</td>
<td>0.140***</td>
<td>0.136**</td>
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<td>(0.036)</td>
<td>(0.043)</td>
<td>(0.053)</td>
<td>(0.023)</td>
<td>(0.039)</td>
<td>(0.060)</td>
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<tr>
<td>Panel C: Structural elasticities</td>
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<tr>
<td>Tax base elasticities</td>
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<td>-0.312***</td>
<td>-0.884***</td>
<td>0.300***</td>
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<td>(0.059)</td>
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<td>(0.093)</td>
<td>(0.077)</td>
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<td>Marginal willingness to pay rent</td>
<td>-0.170***</td>
<td>-0.658***</td>
<td>-1.937***</td>
<td>0.142***</td>
<td>-0.106***</td>
<td>-0.178***</td>
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<td>(0.064)</td>
<td>(0.095)</td>
<td>(0.128)</td>
<td>(0.090)</td>
<td>(0.105)</td>
<td>(0.213)</td>
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<td>Resident incidence</td>
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<td>-0.188***</td>
<td>-0.486***</td>
<td>0.130***</td>
<td>-0.023</td>
<td>-0.209***</td>
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<td>(0.031)</td>
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<td>(0.021)</td>
<td>(0.032)</td>
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<tr>
<td>Landlord incidence ( (\rho^{**}) )</td>
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<td>(0.121)</td>
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</table>

Notes: Standard errors reported in parentheses. Overall renter incidence is the weighted sum of the income class specific incidences using the corresponding relative population shares. ***p < 0.01, **p < 0.05, *p < 0.1.
A.4 Supplementary figures

Figure A4.1: Stated preferences for different expenditure categories

Notes: This figure shows the percentage of respondents answering “Spend much more” or “Spend more” to the question: “Listed below are various areas of government spending. Please show whether you would like to see more or less government spending in each area. Remember that if you say “much more”, it might require a tax increase to pay for it.” Quintiles by net monthly household income. Source: International Social Survey Program (ISSP), Role of Government (survey responses for Switzerland, 1996).

Figure A4.2: Expenditure shares on housing

Notes: This figure reports the evolution of housing expenditure shares (defined as annual rent over disposable income) by income quartile. Source: Swiss Household Panel data.
Figure A4.3: The geography of housing prices in Switzerland

Notes: Panel (a) depicts the average rental prices in CHF per square meter, for the initial years 2004 and 2005. Panel (b) represents the difference in rental prices between the average of 2004, 2005 and 2013, 2014. The averages over the initial and final period serve to ensure the largest sample of municipalities. The white lines represent municipal administrative borders. The black lines represent cantonal administrative borders.
Figure A4.4: IV estimation sample at cantonal borders

Notes: Black lines are cantonal borders. Dark grey municipalities included used in the instrumental variable estimations. They are selected to be strictly adjacent to cantonal border and municipality pairs are based on a 10km road distance criteria.

Figure A4.5: Structural estimates: four income classes

Notes: In the left-hand-side panel, light gray bars represent the tax incidence experienced by households without children, dark gray bars represent the tax incidence experienced by households with children, and the black bar represents the group-size-weighted average of resident incidence. Incidence is the percent-change in resident welfare with respect to one percent increase in income taxes as defined by equation (11). Bar widths represent relative population sizes. Whiskers represent 95% confidence intervals.
Figure A4.6: Structural estimates with randomized starting values

Notes: Panel (a) and (b) present the sensitivity to starting values for the three-class and the four-class specifications, respectively. Bars represent the average incidence and preference estimates from 1,000 iterations of the structural estimation that randomly draws starting values for $\delta_m$ from a uniform distribution $\in [0,0.5]$. Bar widths represent relative population sizes. Dashed whiskers represent the average estimate +/- one standard deviation. 10% of the estimates with the highest chi-square statistic were removed.
Figure A4.7: Sensitivity to starting values of household preferences for the public good ($\delta_m$): three-class specification

Notes: The histograms plot the distributions of the estimated parameters $\delta_m$ for 1,000 iterations of the structural estimation that randomly draws starting values for $\delta_m$ from a uniform distribution $\in [0,0.5]$. The last panel reports the distribution of the Chi-square statistic related to the structural estimations. 10% of the estimates with the highest chi-square statistic were removed.
Figure A4.8: Sensitivity to starting values of households’ preference for the public good \( (\delta_m) \): four-class specification

Notes: The histograms plot the distributions of the estimated parameters \( \delta_m \) for 1,000 iterations of the structural estimation that randomly draws starting values for \( \delta_m \) from a uniform distribution \( \in [0,0.5] \). The last panel reports the distribution of the Chi-square statistic related to the structural estimations. 10% of the estimates with the highest chi-square statistic were removed.
Figure A4.9: Structural estimates: fully-fledged model

Notes: In the left-hand-side panel, light gray bars represent the tax incidence experienced by households without children, dark gray bars represent the tax incidence experienced by households with children, and the black bar represents the group-size-weighted average of resident incidence. Incidence is the percent-change in resident welfare with respect to one percent increase in income taxes as defined by equation (11). Bar widths represent relative population sizes. The starting values of $\delta_m$ are the same as those chosen for the estimation of the ‘myopic’ model presented in Figure 4.

Figure A4.10: Structural estimates by family status and age

Notes: Average estimates and standard deviations from 1,000 iterations of the structural estimation that randomly draws starting values for $\delta_m$ from a uniform distribution $\in [0, 0.5]$. The resident welfare panel reports four main elements across income classes. First, the light gray bars represent the income tax incidence borne by non-pensioner households without children. The gray bars represent the income tax incidence borne by pensioner households without children. Second, the darker gray bars represent the income tax incidence experienced by households with children. Last, the black bar represents the overall renter household welfare change measured by a weighted sum of household incidence using relative population shares as weights. Bar widths represent the relative population size. Whiskers represent intervals of one standard deviation of our 1,000 estimates per household type.
Figure A4.11: Sensitivity to starting values of household preferences for the public good ($\delta_m$) by family status and age

Households without children (non-pensioners)  
Households without children (pensioners)  
Households with children

Notes: The histograms plot the distributions of the estimated parameters $\delta_m$ for 1,000 iterations of the structural estimation that randomly draws starting values for $\delta_m$ from a uniform distribution $\in [0,0.5]$. The last panel reports the distribution of the Chi-square statistic related to the structural estimations. 10% of the estimates with the highest chi-square statistic were removed.
Figure A4.12: Structural estimates: differential mobility costs

Notes: Average estimates and standard deviations of the scale parameter $\lambda_m$ (for a constant $\delta$). The higher $\lambda$, the lower is residents’ attachment to location (hence the more mobile they are). Results are from 1,000 iterations of the structural estimation that randomly draws starting values for $\lambda_m$ from a uniform distribution $\in [0.5, 3.5]$. In the left-hand-side panel, light gray bars represent the tax incidence experienced by households without children, dark gray bars represent the tax incidence experienced by households with children, and the black bar represents the group-size-weighted average of resident incidence. Incidence is the percent-change in resident welfare with respect to one percent increase in income taxes as defined by equation (11). Bar widths represent relative population sizes. Dashed whiskers represent the average estimate +/- one standard deviation. 10% of the estimates with the highest chi-square statistic were removed.
Figure A4.13: Sensitivity to starting values sensitivity of households’ attachment to locations ($\lambda_m$).

Households without children

Households with children

Chi-square statistic

Notes: The histograms plot the distributions of the estimated parameters $\lambda_m$ (keeping $\delta$ constant) for 1,000 iterations of the structural estimation that randomly draws starting values for $\lambda_m$ from a uniform distribution $\in [0.5, 3.5]$. The higher $\lambda$, the lower is residents’ attachment to location (hence the more mobile they are). The last panel reports the distribution of the Chi-square statistic related to the structural estimations. 10% of the estimates with the highest chi-square statistic were removed.
Figure A4.14: Households’ and landlords’ incidence for different values of $\lambda$ across family status

Households without children
Households with children

Notes: Each graph is based on 10,000 different combinations of calibrated values for different $\lambda$ depending on the family status, in a range given by an interval of plus/minus one standard error of the estimated value presented in Appendix Section A.2.2. The intervals are truncated at 0.5 to avoid model instability. The black dots indicate baseline estimates. The z-axis scale is common to all panels.
Who Bears the Burden of Local Taxes?

– Online Appendix –

Marius Brülhart  Jayson Danton  Raphaël Parchet  Jörg Schläpfer

September 19, 2019

Contents

W.1 Technical derivations ................................................. 1
W.2 A modified housing supply .......................................... 5
W.1 Technical derivations

This section contains the detailed derivations of our baseline model. In Subsection W.1.1, we characterize the individual’s marginal willingness to pay rent (MWPR) for a small tax change. In Subsection W.1.2, we derive our system of equations characterizing the effect of a small change in the tax rate on the equilibrium number of residents in different income classes and on equilibrium housing prices. In Subsection W.1.3, we derive the incidence of a change in the tax rate on residents’ and landlords’ welfare.

W.1.1 The marginal willingness to pay rent (MWPR)

The optimization problem of household \( i \) in income class \( m \) and choosing a location \( j \) can be written as follows:

\[
\max_{h_{mj}, z_{mj}} U_{imj} = \alpha \ln(h_{mj} - \nu_h) + (1 - \alpha) \ln(z_{mj} - \nu_z) + \delta_m \ln(g_j - \nu_g) + A_{ij}
\]

s.t. \( z_{mj} + p_j h_{mj} = (1 - \tau_j) w_m \). \hspace{1cm} (W.1)

The individual Marshallian demands of this program are

\[
h^*_{mj} = \nu_h + \frac{\alpha [(1 - \tau_j) w_m - p_j \nu_h - \nu_z]}{p_j}, \text{ and}
\]
\[
z^*_{mj} = \nu_z + (1 - \alpha) [(1 - \tau_j) w_m - p_j \nu_h - \nu_z],
\]

where \( \nu_h \geq 0, \nu_z \geq 0 \) and \( \nu_g \geq 0 \) can be thought of, respectively, as existential needs for housing, the non-housing composite good and the public good. For simplicity, and without loss of generality, we assume \( \nu_z = 0 \) and \( \nu_g = 0 \). Unlike with a standard Cobb-Douglas utility, the elasticity of individual housing demand with respect to prices is not constant. It is given by

\[
\left| \frac{\partial h^*_mj}{\partial p_j} \right| = 1 - \frac{(1 - \alpha) \nu_h}{h^*_mj},
\]

which is equal to one only if \( \nu_h = 0 \) and less than one otherwise.

A household’s indirect utility given its choice of location \( j \), is

\[
V_{imj} = \kappa + \ln [(1 - \tau_j) w_m - p_j \nu_h] - \alpha \ln(p_j) + \delta_m \ln(g_j) + \ln(A_{ij}). \hspace{1cm} (W.2)
\]

We define as marginal willingness to pay rent the change in the housing price (‘bid-rent’ price change) a household with income \( m \) would require to be indifferent toward a given change in the local tax rate:
\[ dV_{imj} = \left[ \frac{\partial V_{imj}}{\partial p_j} dp_j + \frac{\partial V_{imj}}{\partial \tau_j} d\tau_j + \frac{\partial V_{imj}}{\partial g_j} dg_j \right] \]

\[ dV_{imj} = \left. \left[ -\alpha \left( \frac{h_{mj}^*}{h_{mj}^* - \nu_h} \right) \frac{dp_j}{p_j} - \alpha \frac{\tau_j}{(1 - \tau_j) S_{mj}} \left( \frac{h_{mj}^*}{h_{mj}^* - \nu_h} \right) \frac{d\tau_j}{\tau_j} + \delta_m \frac{dg_j}{g_j} \right] \right|_{dV_{imj}=0} \]

Hence,

\[ \left. \frac{dp_j}{dt} \frac{\tau_j}{p_j} \right|_{dV_{imj}=0} = - \left[ \frac{\tau_j}{(1 - \tau_j) S_{mj}} - \frac{\delta_m}{\alpha} \left( 1 - \frac{\nu_h}{h_{mj}^*} \right) \frac{dg_j}{d\tau_j} \right], \quad \text{(W. 3)} \]

where

\[ \frac{dg_j}{g_j} \left. \frac{\tau_j}{d\tau_j} \right| = 1 + \sum_m (\gamma_{mj} - \theta s_{mj}) \frac{dN_{mj}}{N_{mj}} \frac{\tau_j}{d\tau_j}, \quad \text{(W. 4)} \]

with \( \gamma_{mj} \equiv w_m N_{mj} / \sum_m w_m N_{mj} \) and \( s_{mj} \equiv N_{mj} / N_j \).

### W.1.2 Equilibrium

The model’s equilibrium is characterized by three main equations:

\[ N_j = \sum_m N_{mj} \quad \text{with} \quad N_{mj} = \frac{\exp(\lambda u_{mj})}{\sum_{j'} \exp(\lambda u_{mj'})} \quad \forall j \in J, \quad \text{(W. 5a)} \]

\[ H_j^d = H_j^* \quad \text{with} \quad H_j^d = \sum_m N_{mj} \cdot h_{mj}^* \quad \text{and} \quad H_j^* = B_j p_{j}^* \eta_j^p \quad \forall j \in J, \quad \text{(W. 5b)} \]

\[ g_j = \tau_j N_j^{-\theta} \sum_m w_m N_{mj} \quad \forall j \in J, \quad \text{(W. 5c)} \]

where (W. 5a) describes the population, (W. 5b) governs the housing market, and (W. 5c) is the government budget constraint for each jurisdiction \( j \).

Totally log-differentiating equation (W. 5c) and using the notation \( \dot{x} = dx / x \) yields:

\[ \dot{\ln g_j} = \frac{\partial \ln g_j}{\partial \ln \tau_j} d\tau_j + \sum_m \frac{\partial \ln g_j}{\partial \ln N_{mj}} dN_{mj} - \theta \sum_m \frac{\partial \ln g_j}{\partial N_{mj}} dN_{mj} \]

\[ \frac{dg_j}{g_j} = \frac{d\tau_j}{\tau_j} + \sum_m w_m dN_{mj} - \theta \sum_m \frac{N_{mj} dN_{mj}}{N_j N_{mj}} \]

\[ \dot{g}_j = \tau_j + \sum_m (\gamma_{mj} - \theta s_{mj}) \dot{N}_{mj} \]

\[ \dot{g}_j = \tau_j + \sum_m (\gamma_{mj} - \theta s_{mj}) \dot{N}_{mj}. \]
Totally differentiating equation (W.5b) yields:

\[
\sum_{m} \left[ \frac{\partial H^d_{mj}}{\partial N_{mj}} dN_{mj} + N_{mj} \frac{\partial h^*_{mj}}{\partial p_j} dp_j + N_{mj} \frac{\partial h^*_{mj}}{\partial \tau_j} d\tau_j \right] = \frac{dH^d_j}{dp_j} dp_j
\]

\[
\sum_{m} \left[ H^d_{mj} \dot{N}_{mj} + H^d_{mj} \left( \frac{\partial h^*_{mj}}{\partial p_j} \frac{p_j}{h^*_{mj}} \right) \dot{p}_j + H^d_{mj} \left( \frac{\partial h^*_{mj}}{\partial \tau_j} \frac{\tau_j}{h^*_{mj}} \right) \dot{\tau}_j \right] = \frac{1}{H^d_j} = \eta^s_j \dot{p}_j
\]

where \( \pi_{mj} \equiv H^d_{mj}/H^d_j \) is income class \( m \)'s share of aggregate housing demand and \( \rho_j \equiv \sum_m \pi_{mj} (1 - (1 - \alpha) \frac{\nu_h}{h^*_{mj}}) + \eta^s_j \dot{p}_j \) collects other parameters, notably the housing supply elasticity.

Finally, totally differentiating equation (W.5a) for a given municipality \( j \) and income class \( m \), yields:

\[
dN_{mj} = \lambda N_{mj} (1 - N_{mj}) \left[ \frac{\partial V_{mj}}{\partial p_j} dp_j + \frac{\partial V_{mj}}{\partial \tau_j} d\tau_j + \frac{\partial V_{mj}}{\partial \theta_j} d\theta_j \right]
\]

\[
\frac{1}{\lambda} \dot{N}_{mj} = -\alpha \left( \frac{h^*_{mj}}{h^*_{mj} - \nu_h} \right) \dot{p}_j - \frac{\alpha \tau_j}{(1 - \tau_j) S_{mj}} \left( \frac{h^*_{mj}}{h^*_{mj} - \nu_h} \right) \dot{\tau}_j + \delta_m \dot{\theta}_j
\]

\[
\frac{1 - \delta_m (\gamma_m - \theta_{s mj}) \lambda}{\alpha \lambda} \left( \frac{1 - \nu_h}{h^*_{mj}} \right) N_{mj} - O + \dot{p}_j = \left[ \frac{\delta_m}{\alpha} \left( 1 - \frac{\nu_h}{h^*_{mj}} \right) - \frac{\tau_j}{(1 - \tau_j) S_{mj}} \right] \dot{\tau}_j
\]

where \( O \equiv \frac{\delta_m}{\alpha} \sum_{m' \neq m} (\gamma_{m'j} - \theta_{s m'j}) N_{m'j} \).

Stacking the \( M \) population equations and the equilibrium rental price solution into a system of equations yields

\[
A_j \ddot{y}_j = B_j \dot{\tau}_j,
\]

where

\[
A_j = \begin{bmatrix}
-\delta_1 (\gamma_{1j} - \theta_{s1j}) \lambda & 1 - \frac{\nu_h}{h^*_{1j}} & \cdots & 1 - \frac{\nu_h}{h^*_{Mj}} & 1 \\
-\delta_2 (\gamma_{2j} - \theta_{s2j}) \lambda & 1 - \frac{\nu_h}{h^*_{2j}} & \cdots & 1 - \frac{\nu_h}{h^*_{Mj}} & 1 \\
\vdots & \vdots & \ddots & \vdots & \vdots \\
-\delta_M (\gamma_{Mj} - \theta_{sMj}) \lambda & 1 - \frac{\nu_h}{h^*_{Mj}} & \cdots & 1 - \delta_M (\gamma_{Mj} - \theta_{sMj}) \lambda & 1 \\
\pi_{1j} & \cdots & \cdots & \pi_{Mj} & -\rho_j
\end{bmatrix}
\]
and

\[
B_j = \begin{bmatrix}
\frac{\delta \ln \eta}{\alpha} \left(1 - \frac{\mu_{m}}{h_{mj}}\right) - \frac{\tau_j}{(1-\tau_j)S_{mj}} \\
\vdots \\
\frac{\delta \ln \eta}{\alpha} \left(1 - \frac{\mu_{m}}{h_{mj}}\right) - \frac{\tau_j}{(1-\tau_j)S_{mj}} \\
\alpha \left(\frac{\tau_j}{1-\tau_j} \sum_m \frac{\pi_{mj}}{S_{mj}}\right)
\end{bmatrix} .
\]

**W.1.3 Incidence**

Overall renter household welfare is given by

\[
W^R = \sum_m s_m \cdot \frac{1}{\lambda} \log \left( \sum_j \exp(\lambda u_{mj}) \right) .
\]

The effect of a change in the tax rate of municipality \(j\) on household welfare in income class \(m\) is given by

\[
dW^R_m = N_{mj} \left[ \frac{\partial u_{mj}}{\partial p_j} dp_j + \frac{\partial u_{mj}}{\partial \tau_j} d\tau_j + \frac{\partial u_{mj}}{\partial g_{mj}} dg_{mj} \right]
\]

\[
dW^R_m = \alpha N_{mj} \left\{ \frac{\delta \ln \eta}{\alpha} \left(1 - \frac{\mu_{m}}{h_{mj}}\right) \frac{dm_j}{p_j} - \frac{\tau_j}{(1-\tau_j)S_{mj}} \left( \frac{h_{mj}^{*}}{h_{mj}^{*} - \nu_h} \right) \frac{d\tau_j}{\tau_j} + \frac{\delta m_{j}}{\alpha} \frac{dm_{j}}{g_{j}} \right\}
\]

\[
dW^R_m = \alpha N_{mj} \left( 1 - \frac{\nu_h}{h_{mj}^{*}} \right)^{-1} \left\{ \frac{\tau_j}{(1-\tau_j)S_{mj}} \frac{dm_j}{p_j} - \frac{\delta m_{j}}{\alpha} \left(1 - \frac{\nu_h}{h_{mj}^{*}}\right) \left( \frac{dg_{mj}}{\tau_j} \right) - \frac{\delta m_{j}}{\alpha} \frac{dm_{j}}{g_{j}} \right\} , \quad \tag{W.7}
\]

where \(\eta^{p,\tau}\) is the observed change in the equilibrium housing price. The overall change in household welfare is then \(\frac{dW^R_m}{d\ln \tau_j} = \sum_m s_m \cdot \frac{dW^R_m}{d\ln \tau_j}\).

The producer surplus is given by

\[
W^L = \int_0^{H^{*}} \left( p_j^* - \left( \frac{x}{B_j} \right)^{1/\eta_{h,p}} \right) dx = \frac{p_H^*}{1 + \eta_{h,p}} .
\]

The change in the landlord’s welfare after a change in the local tax is then

\[
dW^L = \left( \frac{1}{1 + \eta_{h,p}} \right) \left( \frac{\partial W^L}{\partial p_j^*} dp_j^* + \frac{\partial W^L}{\partial H_j^*} dH_j^* \right)
\]

\[
dW^L = \left( \frac{1}{1 + \eta_{h,p}} \right) (H_j^* p_j^* + p_j^* dH_j^*)
\]

\[
dW^L_{\tau_j} = \left( \frac{p_j^* H_j^*}{1 + \eta_{h,p}} \right) \left( \frac{dp_j^*}{\tau_j} \frac{\tau_j}{p_j^*} + \frac{dH_j^*}{\tau_j} \frac{p_j^*}{p_j^*} \right) \left( \frac{dH_j^*}{dp_j^*} \right)
\]

\[
dW^L_{\ln \tau_j} = p_j^* H_j^* \left( \frac{dp_j^*}{\tau_j} \frac{\tau_j}{p_j^*} \right) . \quad \tag{W.8}
\]
As a result, landlords’ welfare is fully determined by changes in equilibrium rental prices.

### W.2 A modified housing supply

In this section, we present a modified version of the setting proposed in (Brueckner, 2011, Ch.6). Atomistic absentee landlords own a stock of dwelling space with a net-of-tax rental revenue of \((1 - \tau_j)p_jH(k, l_j)\), where \(p_j\) is the rental price per square meter, which is considered as given.\(^1\) \(H(k, l_j)\) represents a concave constant returns to scale housing production function, using non-land capital \(k\) and land \(l_j\) as inputs.\(^2\) Non-land capital is rented at price \(r_k\) and land rent per unit is \(r_l\).\(^3\) The cost of housing is given by \(C(k, l_j) = r_kk + r_ll_j\), which is financed entirely with mortgage debt. Non-land capital is assumed to be supplied perfectly elastically, making \(r_k\) an exogenously fixed parameter.

Landlords need to cover running costs when supplying the rental market. To simplify notation, let \(x_j\) denote the capital-land ratio \(k/l_j\), which can be interpreted as building density or height. Substituting \(x_j\), a landlord’s profit maximization program is

\[
\max_{x_j} l_j[(1 - \tau_j)p_jh(x_j) - ac(x_j)],
\]

where \(h(x_j) \equiv H(x_j, 1)\) and \(c(x_j) \equiv C(x_j, 1)\) denote the dwelling space and housing cost per unit of land, and \(\gamma\) represents the fraction costs effectively borne by landlords after consideration of tax deductions. The \(h\) function satisfies \(h'(x_j) \equiv H_1(x_j, 1) > 0\) and \(h''(x_j) \equiv H_{11}(x_j, 1) < 0\).

Given a fixed parcel of land, \(l_j\), the landlord chooses \(x_j\) to maximize profit per unit of land, given by equation (W.9), while land prices adjust until profits per unit of land are zero. Due to the fact that profits are zero, for any value of \(l_j\), the scale of the landlord’s building is indeterminate. The maximization of (W.9) with respect to \(x_j\) and the zero profit condition are

\[
(1 - \tau_j)p_jh'(x_j^*) = ar_k, \quad (W.10a)
\]
\[
(1 - \tau_j)p_jh(x_j^*) - ar_kx_j^* = ar_l^*.
\]

The landlord’s profit-maximizing dwelling stock per unit of land is given by \(h(x_j^*)\), with \(x_j^*\) being the optimal structural density determined by (W.10a).

The total dwelling stock in municipality \(j\), using a Cobb-Douglas production function, is equal to

\[
H^j_s = l_j \cdot h(x_j^*) = l_j \left[ \frac{B(1 - \tau_j)p_j}{ar_k} \right] \eta_s^{\alpha_p}, \quad \forall j \in J,
\]

where \(\eta_s^{\alpha_p} \equiv B/(1 - B)\) represents the housing supply elasticity of rental prices and \(B \in \)...

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\(^1\)While landlords are absentee, to be consistent with our simplifying assumption, we assume that they pay a tax on rental income in the jurisdiction in which their dwelling is located, to be consistent with our empirical setting.

\(^2\)As in Brueckner (1987), the building is implicitly being split up into housing units (apartments) that consumers can then rent from the landlord.

\(^3\)Factor prices are assumed to be strictly positive.
is the Cobb-Douglas share of capital expenditure in housing production. Furthermore, the endogenous price of land is $r^*_l$, which is determined by replacing $x^*_j$ into (W.10b). Finally, movements along the supply curve are measured by the housing supply elasticity with respect to rental prices. The partial derivative of equation (W.11) with respect to $p_j$ yields

$$\frac{\partial H^j}{\partial p_j} \frac{p_j}{H^j} = \eta^{sp} \geq 0.$$  \hspace{1cm} (W.12)

**Tax deductibility assumptions** In the model, the landlord incurs running costs when supplying the rental market with dwelling space. Mortgage interest payments are equal to $i \cdot C(k, l_j)$, where $i$ denotes the interest rate. Over time, the landlord’s housing stock depreciates at rate $\dd$, representing a cost of $d \cdot C(k, l_j)$. In addition, property taxes need to be paid, which amount to $b_1 \cdot C(k, l_j)$, where $b_1$ is the property tax rate. Furthermore, transaction taxes $b_2$ are due if housing stock is sold and amount to $b_2 \cdot C(k, l_j)$, in the event of a sale. Finally, capital gains through appreciation of housing prices reduce costs by $g \cdot C(k, l_j)$. If all costs are deductible and capital gains are taxed at the same rate as income, we can collect the various running costs and define the fraction of the housing stock’s value allocated to running costs as

$$a = (1 - \tau_j) \dd,$$

with $\dd = (i + d + b_1 + b_2 - g)$.

In this case, equation W.11 becomes

$$H^j = l_j \left[ \frac{B \cdot p_j}{\dd \cdot r_k} \right]^{\eta^{sp}},$$

and housing supply is independent of changes in income tax rates.

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4See Combes, Duranton and Gobillon (2016) for a discussion on the choice of the production function.

5Capital gains $g$ represent the expected rate of increase in housing value. Bear in mind that this source of revenue is only realized at the sale of the housing stock, but is still anticipated.