Residential exodus from Dublin circa 1900: Municipal annexation and preferences for local government*

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Online appendix

1 Literature

The theoretical Tiebout model was first empirically tested by Oates (1969), who showed that inter-municipal differences in local government taxes and spending were reflected in house prices - often referred to as "fiscal capitalization". Virtually all studies, looking for tax capitalization and public service capitalization in housing prices, find a statistically significant negative impact of property taxes and positive impact of local public goods quality on house values (see Ross & Yinger 1999, p 2032). Many subsequent studies call Tiebout models "sorting models", because they explain voluntary intermunicipal sorting by demographic group according to household demand for local government services (amongst other local characteristics). The result of sorting can be perfect stratification - segregation - by group across communities. Hoyt & Rosenthal (1997) provide an innovative empirical test for the existence of Tiebout sorting; their findings are consistent with Tiebout and also satisfy a necessary, but not sufficient, condition for the efficiency of residential sorting across areas. Implicit in the sorting phenomenon are issues of endogenous selection, where residents of a community are both consumers of and inputs into

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the local services in their jurisdiction; consequently efficiency in consumption and in production become potentially inseparable problems.

Following Ellickson (1971), a strand of this literature is devoted to developing theoretical sorting models which characterize equilibrium when both the populations of local jurisdictions and the policies of those jurisdictions are endogenous (see, also, Epple et al. 1984; Epple & Sieg 1999; Nechyba 2000; Calabrese et al. 2002; Calabrese et al. 2006). Recent empirical work often makes use of natural experiments or other forms of applied causal analysis to deal with this endogeneity issue. In a highly cited study, Black (1999) uses boundary discontinuity design to show that high public-school test-scores are capitalized into the value of owner-occupied housing within school districts. Boustan (2008 and 2012) also utilizes administrative/municipal boundaries in order to identify effects of local government on residential location decisions. Boustan (2012) uses urban-suburban boundaries to uncover the effects of municipally ordered school desegregation (a change in the bundle of local public goods available to city residents) in the 1970s for American cities. She finds that desegregation reduced the demand for urban housing and consequently led to a reduction in the urban tax base, imposing a fiscal externality on the remaining city residents. Even more directly related to our study, Boustan (2008) uses urban-suburban boundaries to identify the effects of local government composition on residential demand, and consequently suburbanization, in the United States between 1960 and 1980. She finds that "white flight" from American urban centers was not necessarily solely driven by a demand for racial homogeneity, but also by a desire for political autonomy, as residents were seeking to escape the changing city electorate. Baum-Snow & Lutz (2011) use quasi-experimental data to look at how the residential sorting equilibrium (racial) in a metropolitan area changes with an exogenous shock to a local public good (school quality) in just one jurisdiction. Particularly, they use the fact that the timing of court-ordered desegregation, which is equivalent to a decrease in quality of public schools for whites, varies across a metropolitan area. They find that white enrollment in public schools dropped by 6 to 12 percent due to desegregation, and that this manifested itself in white movement to the suburbs in the South and private school attendance by whites (without moving residence) in other areas of the nation.

Empirical tests of Tiebout's theory generally do not attempt to directly identify the predicted migration patterns in response to marginal changes in taxation or local public goods. Most instead focus on the result of such migration, like the implicit capitalization in house prices. However, Schmidheiny (2006), Banzhaf & Walsh (2008) and Dahlberg et al. (2012), do empirically test the actual Tiebout mechanism. Schmidheiny uses a 1997 household level cross-section of migrating households within metropolitan Basel to test a model of income segregation using a multinomial response framework (individuals choosing to live in one of a number of jurisdictions within a single metropolitan area). Only a handful of earlier works had utilized the discrete choice framework to focus explicitly on the impact of local fiscal variables on the community choice problem (see Friedman 1981, Quigley 1985 and Nechyba & Strauss 1998). This framework avoids some of the endogeneity problems inherent in capitalization/hedonic models, as from the perspective of the individual choosing to live in a given area, community characteristics can be taken as given at that moment. He finds that households choose their residential jurisdiction based upon local taxation, and that such preferences lead to sorting by income group. Following a similar econometric framework, Dahlberg et al. (2012) study detailed Swedish data and find that households choose municipalities in the Stockholm metropolitan area that offer more attractive levels of public goods (in this case, child care financing). Banzhaf & Walsh (2008) use a difference-in-difference approach (taking advantage of exogenous marginal changes in a local public good) with data at the community level, and show that people *do* indeed "vote with their feet" to choose residential communities with prefered bundles of taxes and public goods (in this case, represented by environmental quality).

Scholarly work on annexation similarly builds on Tiebout (1956). Fragmentation, having multiple local governments in a single municipal area, is found to be a better, more efficient, local government organizational strategy than municipal conglomeration. Fragmentation promotes competition amongst governments for residents and thus annexation reduces this competition. Moreover, since many people have chosen to live in unincorporated municipal fringe areas specifically to avoid paying higher city taxes, annexation effectively "thwarts the preferences that these residents have revealed in their choice of where to live" (Liner & McGregor (2002, p 1477)). The argued benefits provided by annexation: scale economies, coordination and uniformity of municipal service provision, orderly land-use planning and overall reduction of political chaos are often shown to be outweighed by the negative "fiscal deficit" (Hamilton (1976)) borne by certain groups due to the impact of annexation on demographic composition, tax base, electorate, and indirectly through its eventual capitalization in housing prices. A fiscal deficit occurs when a household's cost in local taxes exceeds its return in local public services (see, for example, Bradford & Oates 1974, Wagner & Weber 1975, Epple & Zelenitz 1981, and Bollens 1986 for evidence of this phenomenon).

Annexation is a process that is often argued to strictly benefit annexing cities (poorer) at the expense of the annexed suburbs (wealthier). Filer & Kenny (1980) use a median voter model to empirically test and confirm the common perception that annexation (consolidation) is a vehicle for the redistribution of wealth: enabling poorer city communities to expropriate wealth, in the form of collectively distributed local public goods, from richer former suburban residents. However, Garasky & Haurin (1997) find evidence suggesting that pivotal-voter behavior on municipal referenda does not always follow the expected logic shown in Filer & Kenny (1980) (i.e. voters are not rational). In their study, central city voters who would be left behind after a deannexation of a wealthy neighborhood, overwhelmingly supported it even though they would be made worse off. See Wildasin (1986) for a literature review of related models employing the median-voter approach, and see Ross & Yinger (1999) for a discussion of the limitations of this approach. Heim (2011) confirms that local government competition for tax revenues is a major driving factor behind annexation (specifically, promoting the "border wars" which developed in the Phoenix metropolitan area, from 1970 onwards). She finds that municipalities acted strategically in the timing of annexation, as particularly evident in the association between border expansion and infrastructure development (highway construction). The same occurred in Dublin, as the installation of the main drainage scheme for the city coincided with its push for boundary expansion.

Recent work has become increasingly complex in its modeling of metropolitan equilibrium; allowing populations and local government policies to endogenously evolve. Calabrese et al. (2002) develop a computational multiple jurisdiction model of local government and use simulations to confirm that annexation generally increases the welfare of residents in the pre-existing city, while residents in the annexed suburbs lose out (as do the residents of the suburbs that remain independent, but to a lesser extent). Alesina et al. (2004) use a similar multiple jurisdiction model and show that there is a premium for racial political autonomy. Residents are willing to pay more in order to maintain racial homogeneity versus to maintain income homogeneity in their local jurisdiction (by avoiding municipal conglomeration).

Not all scholarly evidence sides against annexation (municipal conglomeration), or supports the hypothesis that its main motive and outcome is (inefficient) redistribution of wealth. Using a median-voter model and census tract level data from large U.S. cities annexed in the 1950s, Austin (1999) finds that city desire to expropriate suburban tax base did not motivate annexation. Cities were sensitive to other economic factors, such as the desire to offset the political and racial effects of urban migration. Liner & McGregor (2002) find evidence that there is an "optimal level of annexation" for which the growth in per capita taxation and the growth in per capita spending are minimized: that is, there are service delivery efficiencies that "at least for a time" more than offset the fiscal losses attributable to municipal monopoly. Many studies questioning the benefits of jurisdictional fragmentation and local government competition rest on the belief that Tiebout sorting is itself inefficient. This can be due to either an inefficient allocation of people into communities, or an inefficient selection of public services by local governments. See Ross & Yinger (1999) (particularly Section 4) for a detailed review of earlier work on sources of inefficiencies that might present themselves in decentralized local governments with Tiebout sorting. Dur & Staal (2008) find that jurisdictional fragmentation results in a sub-optimal outcome when positive cross-border spillovers of local public goods exist. Cross-border spillovers occur when residents in townships that are located nearby a city take advantage of the services provided by that city. Many township dwellers work in cities and often find the city a good place to spend their free time. The opposite scenario is not so often true. In their two-district (city and village) theoretical model accounting for spillovers, they show that jurisdictional autonomy leads to an under-provision of local public goods in both jurisdictions. However, they determine that annexation is not the ideal solution to the problem, as it would in fact make the distribution of local public goods between the two areas highly skewed towards the original city. They show that a socially optimal scheme of national transfers to local governments is a good solution to the inefficiencies inherent in jurisdictional fragmentation. Calabrese et al. (2012) also find large inefficiencies with jurisdictional fragmentation when *ad valorem* property taxation (rather than head taxation - which has been shown to be efficient) is the method used to finance local public goods. This was not the first paper to note the differential effects of head taxes versus property taxes in a sorting model. See Mieszkowski & Zodrow (1989) for a survey of earlier work. Using a computational model they find that in large part these inefficiencies can be attributed to the notion of the "poor chasing the rich": when relatively richer communities (suburbs) come into existence, that have high expenditures on local public goods (the Tiebout equilibrium), poorer households then crowd into these communities and free ride; consuming relatively little housing in that area (i.e. paying little by way of property tax on their low-valued properties). They argue that this effect would dissipate most, if not all, of the potential welfare gains of decentralization over centralization. Hamilton (1975) was first in a long string of literature to model the idea of zoning into the Tiebout framework, and he found that this practice in its "perfect form" (under strong assumptions) would make it possible for high income households to completely prevent subsidization of the type depicted in Calabrese et al. (2012); eliminating the problem of the poor chasing the rich.

The current literature argues that residents do have preferences over local government which will lead them (perhaps efficiently) to sort by demographic group into preferred areas, and that the exogenous (unanticipated) shock of annexation should generate a reaction whereby residents who were instantaneously made worse off will "vote with their feet" to choose a new best option.

2 Simulations: Long Term Implications of Estimated Preferences on Area Composition

The purpose of the simulation exercise in this extension is to provide an image of what the preferences (marginal effects) estimated in the main body of the paper imply about the demographic composition of an area over time (here, across the twentieth century); with a particular focus on what would have happened to area three (the independent townships) had it been annexed to the city like originally proposed by the Dublin city council. We first proceed with a description of the calculation of odds ratios for each area, which will determine who replaces whom in the simulations that follow. Using odds ratios implicitly surmounts the issue that the estimated coefficients are normalized with different error variances for each area (allowing us to compare results across areas).

We calculate odds ratios for residential replacement on both the religion and income proxy attributes. To do so, we utilize the estimated coefficients on these variables for areas one (the original city) and two (the independent townships), but for area three, we impose the departure rule for area two (and thus use the estimated coefficients for area two) on the existing population of area three in order to simulate the annexation, as described below. Following the calculation of odds ratios, we describe the actual simulations. Results of the simulations should be interpreted cautiously, as they implicitly assume that the projected exit process (for both religion and income proxy groups) is perpetuated constantly over the entire period.

2.1 Setting up Simulations: Religion Variable

In the notation in this section on religion we assume that the vector of household characteristics x includes all household characteristics, but religion.

2.1.1 Area one

The estimated coefficient for the Protestant dummy gained from the probit regression is $\hat{\beta}_{prot} = -0.359$ and the average proportion Protestant among movers is $\overline{Prot} = 0.146$. The estimated utility for the average household of inward and outward movers is $\left(\overline{x}\hat{\beta}+\hat{\beta}_{Prot}\overline{Prot}\right)=0.096$. From that value we net out the estimated average religion effect on the utility value, i.e. $\hat{\beta}_{Prot}\overline{Prot}=-0.051$, and obtain the predicted utility value for the average Catholic household, $\overline{x}\hat{\beta}=0.147$. The predicted marginal effect that a Protestant household moves into the area is calculated (from the probit regression) as the difference:

$$\begin{aligned} P\left(Y=1|\overline{x},Prot;\hat{\theta}\right)-P\left(Y=1|\overline{x};\hat{\beta}\right) &= \Phi\left(\overline{x}\hat{\beta}+\hat{\beta}_{prot}\right)-\Phi\left(\overline{x}\hat{\beta}\right) \\ &= \Phi\left(-0.212\right)-\Phi\left(0.147\right) \\ &= 0.416-0.558=-0.142, \end{aligned} \tag{1}$$

where we have denoted with $\hat{\theta}\equiv\{\hat{\beta},\hat{\beta}_{prot}\}$. And for the marginal effect that a household moves out of the area we have: $P\left(Y=0|\overline{x},Prot;\hat{\theta}\right)=\left[1-\Phi\left(\overline{x}\hat{\beta}+\hat{\beta}_{prot}\right)\right]=0.584$ and $P\left(Y=0|\overline{x};\hat{\beta}\right)=\left[1-\Phi\left(\overline{x}\hat{\beta}\right)\right]=0.442$; leading to the obvious mirroring difference in probability $P\left(Y=0|\overline{x},Prot;\hat{\beta}\right)-P\left(Y=0|\overline{x};\hat{\beta}\right)=0.142$.

We employ the above probabilities computed for the average household (with the average religion effect removed) and simulate the dynamics for religion in the area. From above, the probability that a Protestant family leaves the area is 58.40% and that a Catholic family leaves the area is 44.20%. The probability that a Protestant family moves into the area is 41.60% and the probability that a Catholic family moves into the area is 55.80%. Given that the probabilities sum to one (100%) over inward and outward movers, we find it convenient to express the religion relation with odds and odds ratios. The odds that a Protestant family replaces a Protestant family is $\frac{0.416}{0.584}=0.712$ and that a Catholic family replaces a Catholic family is $\frac{0.558}{0.442}=1.262$. Hence, the odds ratio that a Protestant family replaces a Catholic family is $\frac{0.712}{1.262}=0.564$, which is about one in two, and vice versa is the reciprocal $\frac{1.262}{0.712}=1.772$ (almost two in one).

We utilize these odds and odds ratios to compute the dynamics of the religion area composition. We begin with a hypothetical period one (think of it as year 1901) and a sample of 1,000 households living in an area inhabited by a certain proportion of Protestants and Catholics.

Between period one and period two (which can be thought as year 1911) 40% of the households (a realistic but arbitrary percentage) leave the area at random. The odds and odds ratios are employed to identify who replaces whom. If the (average) household leaving the area is Catholic, it will face a $\frac{1.262}{1.262+0.564} = 0.691$ probability to be replaced by another (average) Catholic family, and $(1 - 0.691) = 0.309$ probability to be replaced by an (average) Protestant family. If the (average) household leaving the area is Protestant, it will face a $\frac{1.772}{1.772+0.712} = 0.713$ probability to be replaced by a (average) Catholic family, and $(1 - 0.713) = 0.287$ probability to be replaced by another (average) Protestant family. We study such dynamics over 11 periods, so as to cover the time span (1901-2001), and repeat the procedure 100 times to gain insight on the distribution of the dynamics.

While interpreting the results the reader should be aware that the exercise has assumed that the religion exit process (that we estimated) is perpetuated constantly over time.

Below we will repeat the same exercise for areas two and three, but to economize space and tedious repetitions, we avoid restating much of the text, presenting only the different values and results. A full description of the simulations is presented in Subsection 2.3.

2.1.2 Area two

The estimated coefficient for the Protestant dummy is $\hat{\beta}_{prot} = -0.260$ and the average proportion of Protestants among movers is $\overline{Prot} = 0.352$. From the estimated utility for the average household of inward and outward movers, $\left(\overline{x}\hat{\beta} + \hat{\beta}_{Prot}\overline{Prot}\right) = 0.116$, we net out the estimated average religion utility value, $\hat{\beta}_{Prot}\overline{Prot} = -0.092$, and obtain the predicted utility value for the average Catholic household, $\overline{x}\hat{\beta} = 0.208$. The predicted marginal effect that a Protestant household moves into the area is computed from the probit regression as the difference:

$$
\begin{aligned}
P\left(Y = 1|\overline{x}, Prot; \hat{\theta}\right) - P\left(Y = 1|\overline{x}; \hat{\beta}\right) &= \Phi\left(\overline{x}\hat{\beta} + \hat{\beta}_{prot}\right) - \Phi\left(\overline{x}\hat{\beta}\right) \\
&= \Phi\left(-0.052\right) - \Phi\left(0.208\right) \\
&= 0.479 - 0.582 = -0.103,
\end{aligned}
\tag{2}
$$

and for the households that leave the area we have: $P\left(Y=0|\overline{x}, Prot;\hat{\theta}\right)=$ $=\left[1-\Phi\left(\overline{x}\hat{\beta}+\hat{\beta}_{prot}\right)\right]=0.521$ and $P\left(Y=0|\overline{x};\hat{\beta}\right)=\left[1-\Phi\left(\overline{x}\hat{\beta}\right)\right]=0.418$; leading to the mirroring difference in probability $P\left(Y=0|\overline{x}, Prot;\hat{\beta}\right)-P\left(Y=0|\overline{x};\hat{\beta}\right)=0.103$.

From above: the probability that a Protestant family leaves the area is 52.10% and that a Catholic family leaves the area is 41.80%. The probability that a Protestant family moves into the area is 47.90% and the probability that a Catholic family moves into the area is 58.20%. The odds that a Protestant family replaces a Protestant family is $\frac{0.479}{0.521}=0.919$ and that of a Catholic family replacing a Catholic family is $\frac{0.582}{0.418}=1.392$. Hence the odds ratio of a Protestant family replacing a Catholic family is $\frac{0.919}{1.392}=0.660$, about one in two, and vice versa it is $\frac{1.392}{0.919}=1.515$.

If the (average) household leaving the area is Catholic, it will face a $\frac{1.392}{1.392+0.660}=0.678$ probability of being replaced by another (average) Catholic family and a $(1-0.678)=0.322$ probability of being replaced by an (average) Protestant family. If the (average) household leaving the area is Protestant, it will face a $\frac{1.515}{1.515+0.919}=0.622$ probability of being replaced by a (average) Catholic family and $(1-0.622)=0.378$ probability of being replaced by another (average) Protestant family.

2.1.3 Area three if annexed

The estimated coefficient for the Protestant dummy is $\hat{\beta}_{prot}=-0.143$ and the average proportion of Protestants among movers is $\overline{Prot}=0.455$. From the estimated utility of the average household of inward and outward movers, $\left(\overline{x}\hat{\beta}+\hat{\beta}_{Prot}\overline{Prot}\right)=0.112$, we net out the estimated average religion utility value, $\hat{\beta}_{Prot}\overline{Prot}=-0.065$, and obtain the predicted utility value for the average Catholic household, $\overline{x}\hat{\beta}=0.177$. The effects on religion for this area are, however, statistically insignificant, indicating that in fact residential replacement was at random in the area that remained independent: Protestants were replacing Protestants and Catholics were replacing Catholics. We now consider what would have happened to area three had it been annexed to the city, as originally proposed by the Dublin municipal government. We look at the hypothetical effect of the annexation, by replacing $\hat{\beta}_{prot}=-0.143$ with $\hat{\beta}_{prot}=-0.260$

(the coefficient for area two, where the annexation did occur). The predicted marginal effect that a Protestant household moves into the area is computed from the probit regression as the difference:

$$\begin{aligned} P\left(Y=1|\overline{x}, Prot; \hat{\theta}\right) - P\left(Y=1|\overline{x}; \hat{\beta}\right) &= \Phi\left(\overline{x}\hat{\beta} + \hat{\beta}_{prot}\right) - \Phi\left(\overline{x}\hat{\beta}\right) \\ &= \Phi\left(-0.083\right) - \Phi\left(0.177\right) \\ &= 0.467 - 0.570 = -0.103, \end{aligned} \tag{3}$$

And for the households that leave the area we have: $P\left(Y=0|\overline{x}, Prot; \hat{\theta}\right) =$
$= \left[1 - \Phi\left(\overline{x}\hat{\beta} + \hat{\beta}_{prot}\right)\right] = 0.533$ and $P\left(Y=0|\overline{x}; \hat{\beta}\right) = \left[1 - \Phi\left(\overline{x}\hat{\beta}\right)\right] = 0.430$, leading to the obvious mirroring difference in probability $P\left(Y=0|\overline{x}, Prot; \hat{\beta}\right) - P\left(Y=0|\overline{x}; \hat{\beta}\right) = 0.103$.

From above: the probability that a Protestant family leaves the area is 53.30% and that a Catholic family leaves the area is 43.00%. The probability that a Protestant family moves into the area is 46.70% and the probability that a Catholic family moves into the area is 57.00%. The odds that a Protestant family replaces a Protestant family is $\frac{0.467}{0.533} = 0.876$ and that of Catholic family replacing a Catholic family is $\frac{0.570}{0.430} = 1.326$. Hence the odds ratio of a Protestant family replacing a Catholic family is $\frac{0.876}{1.326} = 0.661$, and vice versa is $\frac{1.326}{0.876} = 1.514$.

If the (average) household leaving the area is Catholic, it will face a $\frac{1.326}{1.326+0.661} = 0.667$ probability of being replaced by another (average) Catholic family and $(1 - 0.667) = 0.333$ probability of being replaced by an (average) Protestant family. If the (average) household leaving the area is Protestant, it will face a $\frac{1.514}{1.514+0.876} = 0.633$ probability of being replaced by a (average) Catholic family and $(1 - 0.633) = 0.367$ probability of being replaced by another (average) Protestant family.

2.2 Setting up Simulations: Income Proxy Variable

In the notation in this section on income proxy We assume that the vector of household characteristics x includes all household characteristics, but the income proxy.

2.2.1 Area one

First, we identify five equal size intervals from the income proxy domain [30,100]. The estimated utility of the average household is $\left(\overline{x}\hat{\beta} + \hat{\beta}_{Hisc}\overline{Hisc}\right) = 0.096$. From that value we subtract the difference $\left(\overline{Hisc} - \overline{Hisc_1}\right)\hat{\beta}_{Hisc} = (61.776 - 39.590) * (-0.010) = -.222$ - so as to calibrate the average utility to the mean point of the first household HIS-CAM income group in the area in period 1 (39.590), and obtain the new mean average utility $\left(\overline{x}\hat{\beta} + \hat{\beta}_{Hisc}\overline{Hisc_1}\right) = 0.318$. The probability computed at the new average utility is 0.625. We report below the partial effect estimated in the probit regression:

$$\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}; \hat{\theta}\right)}{\partial \overline{Hisc}} = -0.004. \tag{4}$$

We use the above partial effect to approximate the partial derivatives for each group, computed relative to the midpoint (mean value) of the first quintile. We report, in increasing order, the mean value of the HIS-CAM variable in each group: 39.590, 54.580, 63.280, 77.750, 96.940. We denote the groups with $q = (1, 2, \cdots, 5)$, where $q = 1$ indicates the poorest group and $q = 5$ the wealthiest one. We then compute a linear approximation of the partial derivatives for each group:

$$\begin{aligned}
\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc_2}; \hat{\theta}\right)}{\partial \overline{Hisc_2}} - \frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc_1}; \hat{\theta}\right)}{\partial \overline{Hisc_1}} &= -0.060 \\
\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc_3}; \hat{\theta}\right)}{\partial \overline{Hisc_3}} - \frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc_1}; \hat{\theta}\right)}{\partial \overline{Hisc_1}} &= -0.095 \\
\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc_4}; \hat{\theta}\right)}{\partial \overline{Hisc_4}} - \frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc_1}; \hat{\theta}\right)}{\partial \overline{Hisc_1}} &= -0.153 \\
\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc_5}; \hat{\theta}\right)}{\partial \overline{Hisc_5}} - \frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc_1}; \hat{\theta}\right)}{\partial \overline{Hisc_1}} &= -0.229.
\end{aligned} \tag{5}$$

Needless to say, the partial derivatives for the zero outcomes (households that leave the area) have the same values as those above but inverted signs, as we have seen for the religion effect. The odds that an average household moves into the area relative to a household leaving the area is, in increasing order of income quintile: $\frac{0.625}{0.375} = 1.667$, $\frac{0.565}{0.435} = 1.299$, $\frac{0.530}{0.470} = 1.129$, $\frac{0.472}{0.528} = 0.895$, $\frac{0.396}{0.604} = 0.655$. The odds ratios with respect to the lowest quintile are: $\frac{1.299}{1.667} = 0.779$, $\frac{1.186}{1.667} = 0.677$, $\frac{1.042}{1.667} = 0.537$, $\frac{0.835}{1.667} = 0.393$. The full table of income quintile odds ratios is:

	1	2	3	4	5
1	1.000	1.283	1.477	1.862	2.546
2	0.779	1.000	1.151	1.451	1.985
3	0.677	0.869	1.000	1.261	1.725
4	0.537	0.689	0.793	1.000	1.368
5	0.393	0.504	0.580	0.731	1.000

2.2.2 Area two

The estimated utility of the average household is $\left(\overline{x}\hat{\beta} + \hat{\beta}_{Hisc}\overline{Hisc}\right) = 0.116$. From that value we subtract the difference $\left(\overline{Hisc} - \overline{Hisc}_1\right)\hat{\beta}_{Hisc} = (67.336 - 38.800) * (-0.007) = -.200$ - so as to calibrate the average utility to the mean point of the first household HIS-CAM income group in the area in period one (38.800), and obtain the new mean average utility $\left(\overline{x}\hat{\beta} + \hat{\beta}_{Hisc}\overline{Hisc_1}\right) = 0.316$. The probability computed at the new average utility is 0.624. We report below the partial effect estimated in the probit regression:

$$\frac{\partial P\left(Y = 1 | \overline{x}, \overline{Hisc}; \hat{\theta}\right)}{\partial \overline{Hisc}} = -0.003. \tag{6}$$

We use the above partial effect to approximate the partial derivatives of each group, computed relative to the mid point (mean value) of the first quintile. We report, in increasing order, the mean value of the HIS-CAM variable in each group: 38.80, 54.55, 63.91, 76.89, 95.12. We then compute a linear approximation of the partial derivatives for each group:

$$
\begin{aligned}
\frac{\partial P\left(Y=1|\overline{x},\overline{Hisc}_2;\hat{\theta}\right)}{\partial\overline{Hisc}_2}-\frac{\partial P\left(Y=1|\overline{x},\overline{Hisc}_1;\hat{\theta}\right)}{\partial\overline{Hisc}_1}&=-0.047\\
\frac{\partial P\left(Y=1|\overline{x},\overline{Hisc}_3;\hat{\theta}\right)}{\partial\overline{Hisc}_3}-\frac{\partial P\left(Y=1|\overline{x},\overline{Hisc}_1;\hat{\theta}\right)}{\partial\overline{Hisc}_1}&=-0.075\\
\frac{\partial P\left(Y=1|\overline{x},\overline{Hisc}_4;\hat{\theta}\right)}{\partial\overline{Hisc}_4}-\frac{\partial P\left(Y=1|\overline{x},\overline{Hisc}_1;\hat{\theta}\right)}{\partial\overline{Hisc}_1}&=-0.114\\
\frac{\partial P\left(Y=1|\overline{x},\overline{Hisc}_5;\hat{\theta}\right)}{\partial\overline{Hisc}_5}-\frac{\partial P\left(Y=1|\overline{x},\overline{Hisc}_1;\hat{\theta}\right)}{\partial\overline{Hisc}_1}&=-0.169.
\end{aligned}
\tag{7}
$$

Again, the partial derivatives for the zero outcomes (households that leave the area) have the same values as those above but inverted signs, as we have seen for the religion effect. The odds that an average household moves into the area relative to a household leaving the area is, in increasing order of income quintile: $\frac{0.624}{0.376}=1.660$, $\frac{0.577}{0.423}=1.363$, $\frac{0.549}{0.451}=1.216$, $\frac{0.510}{0.490}=1.040$, $\frac{0.455}{0.545}=0.835$. The odds ratios with respect to the lowest quintile are: $\frac{1.363}{1.660}=0.821$, $\frac{1.216}{1.660}=0.733$, $\frac{1.040}{1.660}=0.626$, $\frac{0.835}{1.660}=0.503$. The full table of income quintile odds ratios is:

	1	2	3	4	5
1	1.000	1.218	1.365	1.596	1.988
2	0.821	1.000	1.121	1.311	1.632
3	0.733	0.892	1.000	1.169	1.456
4	0.626	0.763	0.855	1.000	1.245
5	0.503	0.613	0.687	0.803	1.000

2.2.3 Area three if annexed

The effects on income for this area are statistically insignificant, indicating that in fact residential replacement was at random in the area that remained independent: households in a particular income group which exited the area were being replaced by the same types of households. In a similar vein to that done for the religion variable, we now look at what would have happened to area three had it been annexed to the city, as originally proposed by the Dublin municipal government. To do so, we impose the departure rule from area two (via the estimated coefficients) and consider how the existing population in area three would have evolved by income group.

The estimated utility of the average household is $\left(\overline{x}\hat{\beta} + \hat{\beta}_{Hisc}\overline{Hisc}\right) = 0.112$. From that value we subtract the difference $\left(\overline{Hisc} - \overline{Hisc}_1\right)\hat{\beta}_{Hisc} = (70.087 - 41.620) * (-0.007) = -.199$ - so as to calibrate the average utility to the mean point of the first household HIS-CAM income group in the area in period one (41.620), and obtain the new mean average utility $\left(\overline{x}\hat{\beta} + \hat{\beta}_{Hisc}\overline{Hisc}_1\right) = 0.311$. The probability computed at the new average utility is 0.622. We report below the partial effect estimated in the probit regression:

$$\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}; \hat{\theta}\right)}{\partial \overline{Hisc}} = -0.003. \tag{8}$$

We use the above partial effect to approximate the partial derivatives for each group, computed relative to the mid point (mean value) of the first quintile. We report, in increasing order, the mean value of the HIS-CAM variable in each group: 41.62, 54.49, 63.24, 78.23, 96.36. We then compute a linear approximation of the partial derivatives for each group:

$$\begin{aligned}
\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}_2; \hat{\theta}\right)}{\partial \overline{Hisc}_2} - \frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}_1; \hat{\theta}\right)}{\partial \overline{Hisc}_1} &= -0.039 \\
\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}_3; \hat{\theta}\right)}{\partial \overline{Hisc}_3} - \frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}_1; \hat{\theta}\right)}{\partial \overline{Hisc}_1} &= -0.065 \\
\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}_4; \hat{\theta}\right)}{\partial \overline{Hisc}_4} - \frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}_1; \hat{\theta}\right)}{\partial \overline{Hisc}_1} &= -0.110 \\
\frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}_5; \hat{\theta}\right)}{\partial \overline{Hisc}_5} - \frac{\partial P\left(Y = 1|\overline{x}, \overline{Hisc}_1; \hat{\theta}\right)}{\partial \overline{Hisc}_1} &= -0.164.
\end{aligned} \tag{9}$$

The odds that an average household moves into the area relative to a household leaving the area is, in increasing order of income quintile: $\frac{0.622}{0.378} = 1.646$, $\frac{0.583}{0.417} = 1.400$, $\frac{0.557}{0.443} = 1.258$, $\frac{0.512}{0.488} =$

1.050, $\frac{0.458}{0.542} = 0.844$. The odds ratios with respect to the lowest quintile are: $\frac{1.400}{1.646} = 0.851$, $\frac{1.258}{1.646} = 0.765$, $\frac{1.050}{1.646} = 0.638$, $\frac{0.844}{1.646} = 0.513$. The full table of income quintile odds ratios is:

	1	2	3	4	5
1	1.000	1.175	1.308	1.567	1.949
2	0.851	1.000	1.113	1.334	1.659
3	0.765	0.898	1.000	1.198	1.490
4	0.638	0.750	0.835	1.000	1.244
5	0.513	0.603	0.671	0.804	1.000

2.3 Simulations

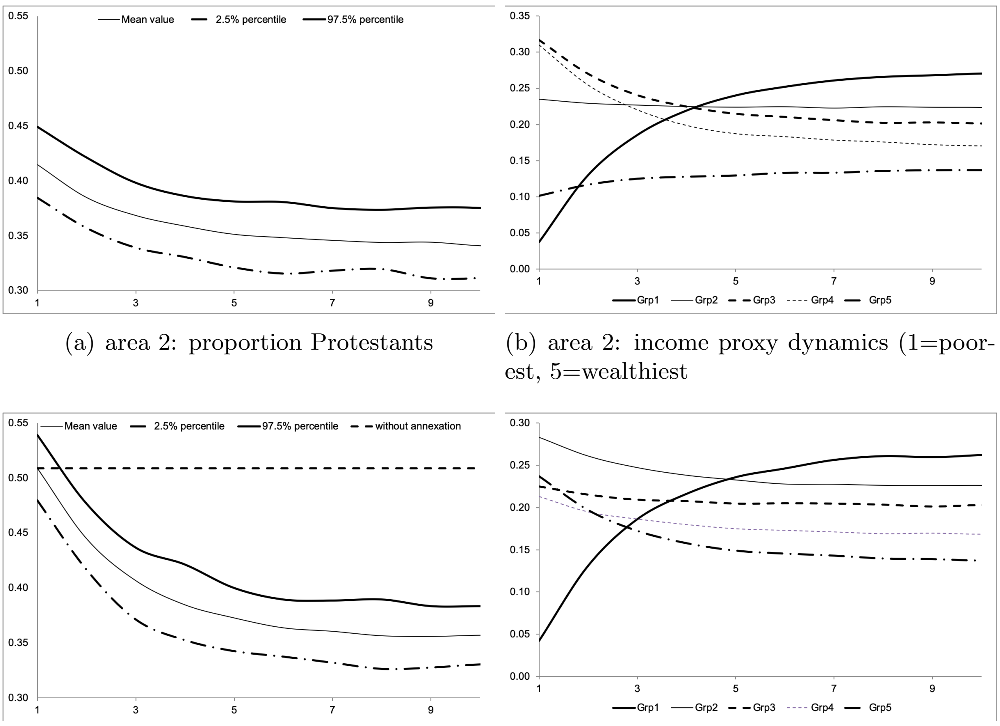
2.3.1 Religion

The following procedure is done for each area. We begin with a hypothetical period one (think of it as year 1901) and a sample of 1,000 households. We calibrate the original area composition to the actual proportion of Protestants in the area in 1901. We draw 1,000 values from the uniform distribution [0,1] and if a drawn value is below that of the calibrated proportion of Protestants in the area. We determine that household to be Protestant (and vice versa to be Catholic). Then we assume that between period one and period two 40% of the households leave the area at random. We determine the departure process in the following way: We have drawn 1,000 new random values from the uniform distribution [0,1] and denote a departure whenever the value is below 0.4.

The odds and odds ratios are employed to identify who replaces whom. If the household leaving the area is Catholic, we will make use of the odds and odds ratios to determine the probability that it is replaced by another Catholic or instead by a Protestant family - as explained above.

To complete the determination of the replacement, we again draw 1,000 new values from the uniform distribution [0,1]: We use the aforementioned information on the religion of the departing household and the associated probabilities, and we determine the religion of the replacing household (whether Catholic or Protestant) based upon the value of each corresponding draw. Of course, for those households who were determined not to have departed from the area in the previous draw, the corresponding value of this last draw will not be used. We repeat the procedure for nine additional periods, so as to cover a hypothetical time span (1901-2001). We then renew the same steps 100 times to gain insight on the distribution of Protestant dynamics. In the graph on the top left panel of Figure 1, we have plotted the mean value and the 95% confidence interval for the Protestant dynamics for area two. The top right panel repeats the same exercise for area three, where we have imposed the same departure rule as in the annexed area two. The actual value of the (non annexed) area three is assumed to be constant over time, for the coefficient was not significant.

Figure 1: Simulated dynamics (deciles from 1901(1) to 2001(10))



(a) area 2: proportion Protestants

(b) area 2: income proxy dynamics (1=poorest, 5=wealthiest

(c) area 3: proportion Protestants)

(d) area 3: income proxy dynamics (1=poorest, 5=wealthiest)

References

Alesina, A., Baqir, R. & Hoxby, C. (2004), 'Political Jurisdictions in Heterogeneous Communities', *Journal of Political Economy* **112**(2). 5

Austin, D. (1999), 'Politics vs Economics: Evidence from Municipal Annexation', *Journal of Urban Economics* **45**(3), 501–532. 5

Banzhaf, S. & Walsh, R. (2008), 'Do People Vote with Their Feet? An Empirical Test of Tiebout's Mechanism', *The American Economic Review* **98**(3), 843–863. 3

Baum-Snow, N. & Lutz, B. F. (2011), 'School desegregation, school choice, and changes in residential location patterns by race', *The American Economic Review* **101**(7), 3019–3046. 2

Black, S. (1999), 'Do Better Schools Matter? Parental Valuation of Elementary Education', *Quarterly Journal of Economics* **114**(2), 577–599. 2

Bollens, S. (1986), 'A Political-Ecological Analysis of Income Inequality in the Metropolitan Area', *Urban Affairs Review* **22**(2), 221–241. 4

Boustan, L. (2008), 'Escape from the City? The Role of Income and Local Public Goods in Post-War Suburbanization', *NBER Working Paper* **13301**. 2

Boustan, L. (2012), 'School Desegregation and Urban Change: Evidence from City Boundaries', *American Economic Journal: Applied Economics* **4**(1), 85–108. 2

Bradford, D. & Oates, W. (1974), "Suburban Exploitation of Central Cities and Governmental Structure", *in* H. Hochman & G. Peterson, eds, 'Redistribution Through Public Choice', Columbia University Press, New York. 4

Calabrese, S., Cassidy, G. & Epple, D. (2002), 'Local Government Fiscal Structure and Metropolitan Consolidation', *Brookings-Wharton Papers on Urban Affairs* pp. 1–43. 2, 5

Calabrese, S., Epple, D. & Romano, R. (2012), 'Inefficiencies from Metropolitan Political and

Fiscal Decentralization: Failures of Tiebout Competition', *The Review of Economic Studies* **Advance Access Online**, 1–31. 6

Calabrese, S., Epple, D., Romer, T. & Sieg, H. (2006), 'Local Public Good Provision: Voting, Peer Effects, and Mobility', *Journal of Public Economics* **90**(6), 959–981. 2

Dahlberg, M., Eklöf, M., Fredriksson, P. & Jofre-Monseny, J. (2012), 'Estimating Preferences for Local Public Services Using Migration Data', *Urban Studies* **49**(2), 319–336. 3

Dur, R. & Staal, K. (2008), 'Local Public Good Provision, Municipal Consolidation, and National Transfers', *Regional Science and Urban Economics* **38**(2), 160–173. 5

Ellickson, B. (1971), 'Jurisdictional Fragmentation and Residential Choice', *The American Economic Review* **61**(2), 334–339. 2

Epple, D., Filimon, R. & Romer, T. (1984), 'Equilibrium Among Local Jurisdictions: Toward an Integrated Treatment of Voting and Residential Choice', *Journal of Public Economics* **24**(3), 281–308. 2

Epple, D. & Sieg, H. (1999), 'Estimating equilibrium models of local jurisdictions', *Journal of Political Economy* **107**(4), 645–681. 2

Epple, D. & Zelenitz, A. (1981), 'The Implications of Competition among Jurisdictions: Does Tiebout Need Politics?', *Journal of Political Economy* **89**, 1197–1217. 4

Filer, J. & Kenny, L. (1980), 'Voter Reaction to City-County Consolidation Referenda', *Journal of Law and Economics* **23**(1), 179–190. 4

Friedman, J. (1981), 'A Conditional Logit Model of the Role of Local Public Services in Residential Choice', *Urban Studies* **18**(3), 347–358. 3

Garasky, S. & Haurin, D. (1997), 'Tiebout Revisited: Redrawing Jurisdictional Boundaries', *Journal of Urban Economics* **42**(3), 366–376. 4

Hamilton, B. (1975), 'Zoning and Property Taxation in a System of Local Governments', *Urban Studies* **12**(2), 205–211. 6

Hamilton, B. (1976), 'Capitalization of Intrajurisdictional Differences in Local Tax Prices', *The American Economic Review* **66**(5), 743–753. 4

Heim, C. (2011), 'Border Wars: Tax Revenues, Annexation, and Urban Growth in Phoenix', *International Journal of Urban and Regional Research* . 4

Hoyt, W. & Rosenthal, S. (1997), 'Household Location and Tiebout: Do Families Sort According to Preferences for Locational Amenities?', *Journal of Urban Economics* **42**(2), 159–178. 1

Liner, G. & McGregor, R. (2002), 'Optimal Annexation', *Applied Economics* **34**(12), 1477–1485. 4, 5

Mieszkowski, P. & Zodrow, G. (1989), 'Taxation and the Tiebout Model: The Differential Effects of Head taxes, Taxes on Land Rents, and Property Taxes', *Journal of Economic Literature* **27**(3), 1098–1146. 6

Nechyba, T. (2000), 'Mobility, Targeting, and Private-School Vouchers', *The American Economic Review* **90**(1), 130–146. 2

Nechyba, T. & Strauss, R. (1998), 'Community Choice and Local Public Services: A Discrete Choice Approach', *Regional Science and Urban Economics* **28**(1), 51–73. 3

Oates, W. (1969), 'The Effects of Property Taxes and Local Public Spending on Property Values: An Empirical Study of Tax Capitalization and the Tiebout Hypothesis', *Journal of Political Economy* **77**(6), 957–971. 1

Quigley, J. (1985), 'Consumer Choice of Dwelling, Neighborhood and Public Services', *Regional Science and Urban Economics* **15**(1), 41–63. 3

Ross, S. & Yinger, J. (1999), "Sorting and Voting: A Review of the Literature on Urban Public

Finance", *in* E. Mills & P. Cheshire, eds, 'Handbook of Regional and Urban Economics', Elsevier Science, chapter 47, pp. 2001–2060. 1, 4, 5

Schmidheiny, K. (2006), 'Income Segregation and Local Progressive Taxation: Empirical Evidence from Switzerland', *Journal of Public Economics* **90**(3), 429–458. 3

Tiebout, C. (1956), 'A pure theory of local expenditures', *Journal of Political Economy* **64**(5), 416–424. 3

Wagner, R. & Weber, W. (1975), 'Competition, Monopoly, and the Organization of Government in Metropolitan Areas', *Journal of Law and Economics* **18**(3), 661–684. 4

Wildasin, D. (1986), *Urban public finance*, Routledge. 4