Decline or Surge?
Tipping Reactions and Tipping Point Models in European Context*

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Abstract

It is often argued that Schelling-type tipping point models are fitted for U.S. dynamics, and that tipping reactions in Europe are markedly more difficult to identify. Here, I apply the Schelling setting to a large selection of Census data and newly published series, finding reactions which are not markedly different from the U.S. figures of [Card, Mas, and Rothstein 2008]. I ponder the interpretation of this result.

Keywords: Migration, Segregation, Racial Discrimination, Mobility

JEL Codes: D70, H50, J11, J15, J61

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1 Introduction

The common wisdom about racial segregation in Europe may be summarized by two tenets: 1. The situation is less severe in Europe than in the U.S., and less severe in continental Europe than in Britain; 2. This situation is improving or at least remains stable in most European countries.

To the first hypothesis, there is no obvious answer, and comparisons vary greatly depending on the source used and the measure of interest. Descriptive statistics about school segregation usually rank France and Germany below the U.S. and Britain in terms of concentration of children of migrants [OECD 2019], but with such low margins that definitive conclusions would be presumptuous. Available comparative data on residential segregation [OECD 2021; Commission et al. 2017] does not suggest that continental Europe performs better overall; if we use traditional segregation indices, Britain exhibits a greater number of 'extreme' case districts where migrant concentration is two or three times higher than in the most extreme continental cases, but the nationwide average of this indices does not suggest a comparative under-performance of the U.K.

As to the second hypothesis, the few existing comparative data at the OECD level does not leave way for much optimism; [OECD 2019] ranks Britain dead last on the evolution of migration-based segregation at school, and Germany is not far behind. There is a growing body of evidence, in European context, about segregation strategies of native parents, for instance in France [Beffy and Davezies 2013], Denmark [Rangvid 2009], Sweden [Andersson, Malmberg, and Östh 2012] and Spain [Farre, Ortega, and Tanaka 2018] among others.

When it comes to residential segregation, attempts at replicating items of the American literature focused on the post-desegregation-era *white flight* have been relatively successful, especially in the context of Nordic countries [Bråmå 2006; Müller, Grund, and Koskinen 2018; Andersen 2017; Bolt and Kempen 2010].

In France, on the opposite, there seems to be a relative consensus, both in the public discourse, in the academic [Rathelot and Safi 2014] and administrative literature [Botton et al. 2020], about the idea that the country performs better on that spot than most European partners, and that the situation is stable or slightly improving [Préteceille 2009].

Actually, conclusions vary greatly depending on the index chosen. [Botton et al. 2020] indeed conclude that, over the INSEE's Census data 1990-2015, the dissimilarity index of extra-European migrants is going *down* if we take the district (IRIS) as the infra-unit; but it is going *up* if we take the city (*commune*) as the infra-unit. Such a pattern is strikingly indicative of a U.S.-type ghetto expansion mechanism : migrant population is rising in residential districts that lie just adjacent to a 'priority district' (QPV, former ZUS) and middle-class populations tend to leave these districts for a city farther away; that is why migrant segregation *within* the communes of French metropolises is declining, but the segregation *between* the communes is on the rise. Also consistent with the expanding ghetto hypothesis is the finding that, at the IRIS level, if the dissimilarity index of extra-European migrants is going *down*, while their isolation index is going *up* [Pan Ké Shon and Verdugo 2014]. This spatial concentration of segregation effects is a differential feature compared with the American estimates. It might suggest that native families are tipping, not necessarily to avoid special groups, but rather to avoid specific impoverished districts.

Within the economic literature, there is a simple model to formalise such cumulative processes: the tipping point model of Thomas Schelling [Schelling 1971]. It combines a spatial residential framework and a micro model where households of the dominant or privileged group have a distaste for residing near families of minority groups, with the intuition that, even if the disutility of the lower-group-proximity is very low, complete residential separation is the only Nash equilibrium of the setting[1]. A key feature is that the utility function of the members of the upper group is non-monotonous; it brutally jumps down when the share of the lower-group population in a district exceeds a certain level l^*, i.e. the upper members are indifferent to the presence of the other group till a certain level: subjectively and at a micro level, this level is a *threshold of tolerance*, but at the aggregate macro level, it becomes a *tipping point*, i.e. a massive flight reaction of the upper group. In this setting, one household moving away for a mixed district creates considerable negative externatilities for those remaining, generating a chain reaction which stops when the system reaches absolute segregation. I.e. even if mixity is a social optimum, and even if agents are

[1] A detailed review of these early models is provided by [Goffette-Nagot, Jensen, and Grauwin 2009]

aware of it, individual choices are set in a way that separation is the only Nash equilibrium.

2 Empirical transcription of the Schelling model

2.1 A focus on tolerance thresholds

One of the most well known empirical transcriptions of this model in the U.S. context is to be ascribed to [Card, Mas, and Rothstein 2008]. Their original intuition might be summarised as such. If tipping patterns do exist, we should be able to detect structural breaks in the variation of the share of the upper group within each district, when the share of the lower group evolves over time. However, a model which would use these two variations as dependent and explanatory variables would be faced with an obvious problem of colinearity. Hence the idea to compare the share of the lower population at the beginning of a decade, and the evolution of the upper population over that decade. We assume that over the decade, the upper group households will move to the most privileged districts which have the lowest shares of minority population.

2.2 Main specification

Formally, the explanatory variable is defined as the share of the lower group, in district i of region c at time $t-10$ (i.e. at the beginning of the decade):

$$l_{ic,t-10} = L_{ic,t-10}/N_{ic,t-10}$$

Where l and L go for the share and population of the lower group, and N the total population of the referenced district.

The dependent variable is the evolution of the share of the upper group, defined similarly as:

$$\Delta u_{ic,t} = (U_{ic,t} - U_{ic,t-10})/N_{ic,t-10}$$

For each region c, two specifically built algorithms (which are detailed in the annex) determine the tipping point, i.e. the initial share $l^*_{c,t-10}$ above which there is a brutal drop in the upper group population. The first algorithm (our *method 1*) is based on a structural break research device; the second one (our *method 2*) seeks a fixed point, modelising the variations of the dependent as a quartic polynominal and finding its roots. The points are identified over a randomly drawn 2/3 subsample of the data, and the whole model is estimated on the remaing 1/3. Variable δ is defined as the distance between the tipping point for the whole region and the lower group share in district i of that region:

$$\delta_{ic,t-10} = l_{ic,t-10} - l^*_{c,t-10}$$

The final model is specified as:

$$\Delta u_{ic,t} = p(\delta_{ic,t-10}) + d\mathbb{I}[\delta_{ic,t-10} > 0] + \tau_c + X_{ic,t-10}\beta + \varepsilon_{ic,t} \tag{1}$$

Where p() is fourth-order polynomial, τ_c a region-specific fixed effect, and X a matrix of controls. d is the coefficient of interest, which shall provide the magnitude and significance of the decline in the upper-group population over the tipping point.

The original article relies on data from the U.S. census. The district or *infra*-unit is the census tract, the *supra*-unit, the *Metropolitan Statistical Area*. In our setting, we'll use the INSEE's Census and the RFL-Filosofi database, with the district (IRIS) as the *infra*-unit, and the *Zone d'emploi* as the *supra*-unit[2].

2.3 Choice of the lower group

As emphasised by [Goffette-Nagot, Jensen, and Grauwin 2009], tipping models have been built to incorporate any definition of the privileged and underprivileged population, be it based on income, ethnicity, education level or

[2]Contrary to exclusively urban *supra*-units like the AAV (*Aire d'attraction de ville*) or the UU (*Unité urbaine*), the ZE cover the entirety of the French territory; it is convenient for rural areas; for provincial metropolises, choosing the ZE or the agglomeration makes very little difference (it provides a definition of the metropolis which is generally wider than the UU but narrower than the AAV). One main difference between choosing the ZE or the UU-AAV lies in the treatment of the Paris metropolis; it is one huge AAV-UU, split in more than a dozen ZEs; such a division seems more natural, since a detailed analysis shows that tipping points are highly heterogeneous over the capital metropolis, with thresholds generally going down as we move from the center to the periphery. Yet as we show in our annex, the significance and magnitude of the estimations are not fundamentally altered by alternative choices.

another variable. In the original article, the authors focus on the white population, as opposed to the minority (African-American and Hispanic) population, over three decades between 1970 and 2000. They find consistent and significant tipping around their estimated zone-specific points. Typically, in Los Angeles 1990-2000, around a threshold of 15% minority population, there is a −7pp drop in the evolution of the white share. Over the decade 1970-1980, which just follows the banning of segregationist policies, tipping is happening for lower shares of minority population, and results in much more violent drops (with some extrema over −30pp).

Replicating that setting over the INSEE's data leaves us with a dual choice: we can pick either the migration status or the nationality. We shall use the latter one, which has the advantage to be stable over time (an individual might be granted French nationality over the period, while being born within or without French territory is a stable feature). We shall therefore define the majority group as the *natives* (people born on French metropolitan and oversee territory) and the minority group as the *migrants* (people born in a foreign country[3]). The census allows us to test other definitions of the upper and lower group, based on diploma, SES, or SES interacted with country of origin, but these alternative definitions perform but poorly (a wide array of categories have been tested, some major results being reported in our annex).

We develop in the corresponding annex a framework in which we find some evidence of tipping based on incomes, middle-class families reacting to a rise of the local population earning less than the national first decile of the income distribution, but we lack the data that would be necessary to ensure that these estimates are perfectly robust.

3 Tipping patterns

3.1 Estimated level

We present first our estimates of *origin*-based-tipping points. In table 1, we report the average value of the tipping thresholds estimated by method 1 (structural break) and method 2 (fixed point). Compared to the original article, we find very similar results, but our thresholds are generally a bit lower than Card and alii's recent values (which are respectively 14.5% and 13.9% for methods 1 and 2 over their last decade, 1990-2000).

Table 1: *Origin*-based-tipping - Estimated tipping points

	1999-2010		2010-217	
	Structural break (1)	Fixed point (2)	Structural break (3)	Fixed point (4)
Mean	10.39%	10.87%	7.78%	10.61%
SE	10.84	8.95	10.49	9.66
Without identified threshold	0	0	0	0
Correlations				
1999-2010 Structural break	1.00			
1999-2010 Fixed point	0.21	1.00		
2010-2017 Structural break	0.32	0.18	1.00	
2010-2017 Fixed point	0.17	0.22	0.19	1.00

Points are expressed in share of migrant pop. in district. Summary stats are unweighted.

[3]The INSEE has a narrower definition of what a *migrant* is, namely a person which is born in a foreign country, did not have the French nationality at birth, and has been living on the French territory for more than two years.

3.2 Estimated reaction around the threshold

Table 2 provides estimates for the tipping behaviour around *Origin*-based estimated thresholds. For each method, we provide the average drop in native population when the district-IRIS lies beyond the tipping point of the zone.

In the original article, Card and his coauthours found, for their last decade 1990-2000, a coefficient d of -7.1 and -9.3pp (for methods 1 and 2 respectively), meaning that the magnitude of the effect in our data is approximately between one third and one half of the American figures.

Table 2: *Origin*-based-tipping - Regression discontinuity model for change of native share around the tipping point

	Dependent var.: Change in native population in the district from $t-10$ to t					
	Method 1 - Structural break			Method 2 - Fixed point		
	Base	F.E.	Full	Base	F.E.	Full
	(1)	(2)	(3)	(4)	(5)	(6)
1999-2010 decade						
Beyond tipping point (coef. d)	$-.96$pp	-2.88pp***	-2.49pp***	-3.93pp***	-3.82pp***	-3.04pp***
SE	(1.11)	(0.81)	(0.74)	(1.34)	(0.91)	(0.78)
Observations	16271	16271	16271	16271	16271	16271
R^2	1.5%	18.8%	22.1%	1.3%	18.7%	21.9%
F-stat	16.63	3.91	4.6	8.97	4.12	4.78
2010-2017 decade						
Beyond tipping point (coef. d)	.83pp	-1.45pp**	-1.83pp***	$-.89$pp	-2.93pp***	-2.58pp***
SE	(0.69)	(0.54)	(0.64)	(0.71)	(0.72)	(0.66)
Observations	16281	16281	16281	16281	16281	16281
R^2	4.2%	7.3%	8.4%	4.2%	7.1%	8.5%
F-stat	14.61	5.85	3.44	7.82	5.69	3.32
Zone fixed effects		X	X		X	X
Controls			X			X

Sign. thr. : *p<0.1; **p<0.05; ***p<0.01

Note: The unit of interest is the district (IRIS). Data are drawn from the INSEE's Census. We report the estimates of model 1, specifically the estimation of coefficient d. The dependent variable is the growth of the native population (defined as these persons who are born on French territory) within the district from t to $t-10$, as percentage of the base population at date $t-10$. The main explanatory variable is a dummy equal to one if the share of migrant population (those persons who are not born on French territory) is beyond the ZE-specific estimated tipping threshold. All specification also include a quartic polynomial in the deviation from the district's migrant share distance from the local tipping point, plus Z.E. fixed effects. The vector of controls, drawn from the INSEE's census, includes unemployment rate, share of working-class people, share of persons with no diploma, and share of vacant accommodation. Standard errors are clustered at the ZE level.

One important remark: we failed to replicate that setting over former issues of the INSEE's Census (1962, 1968, 1975, 1982, 1990). Actually, replication of the exact same specification is not possible before 1990. IRIS-level data were not provided before that time[4] and information about the country of origin are scarcely reported. For issues prior to 1990, we must rely on a very poor proxy, i.e. a strategy at the commune level[5], using nationality to define the upper and lower groups (French nationals / Foreigners). That strategy fails to identify tipping points: actually, the population of French citizens tends to rise between two issues of the Census in communes with the highest shares of foreigners. Using the share of persons repatriated from Algeria after the end of the war, or the share of Algerian Muslim population living on metropolitan territory (provided for the 1962 issue) leads to the very same result. In issues of the Census in which the share of SES categories can be matched over time, we find no tipping

[4] A special subdivision of the communes, the *Ilôt*, did exist, but over these, information about nationality and country of origin are provided for one issue only - 1982 - and cannot therefore be matched over time.

[5] There, the commune is the *infra*-unit; as to the *supra*-unit, we tested different options: the 1994 ZEs, the 2010 ZEs, the *départements*; that choice does not change results fundamentally.

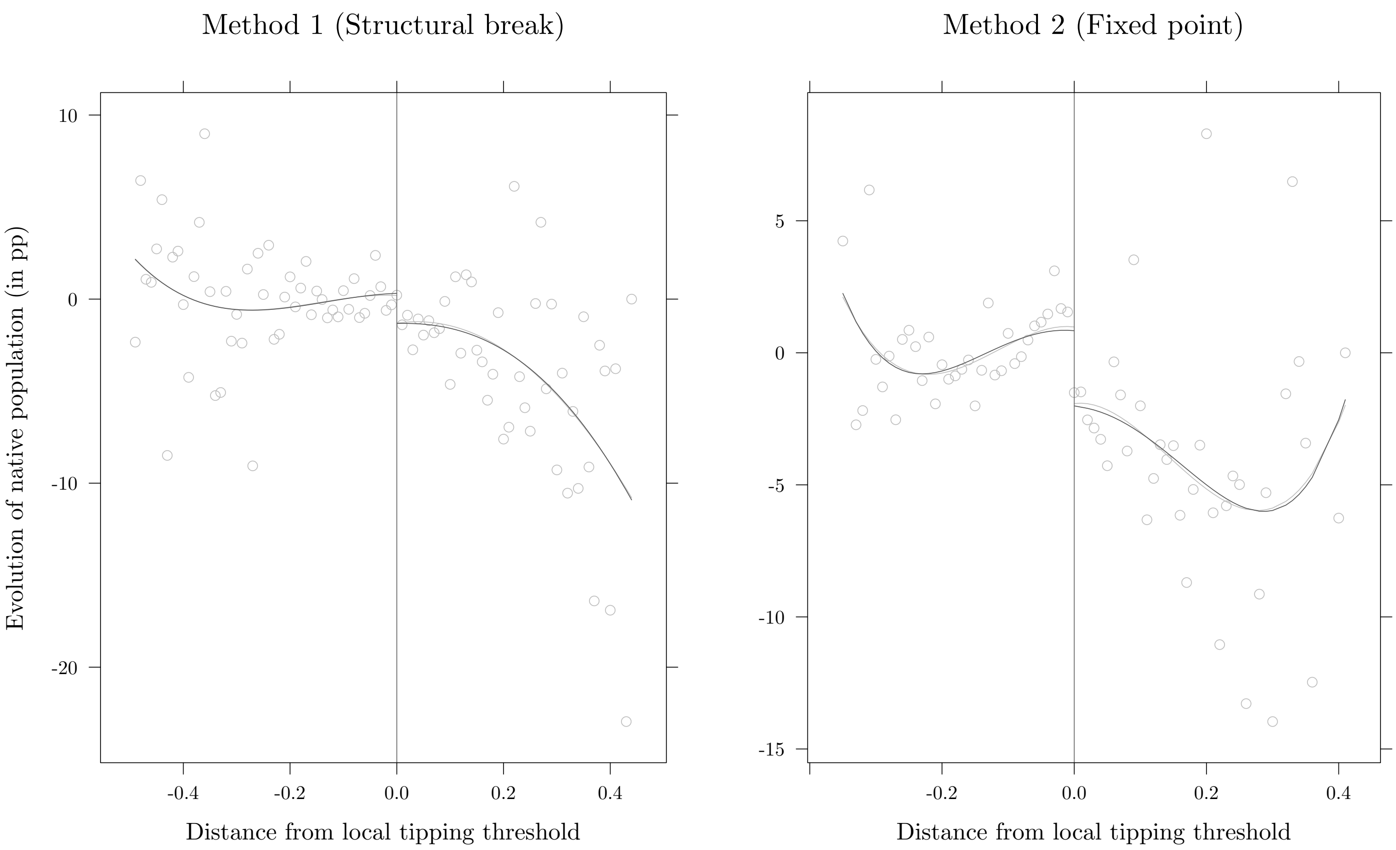
reaction from the upper and middle-class to the local shares of foreign population[6].

Yet we also fail to replicate our *origin*-based strategy on the 1990-1999 decade, while we have under hand everything we need to apply the very same specification. The algorithms estimate consistent tipping points, tipping reaction of the native population around these points is negative, but not significant, or significant at a 10% level in one specification only. This result is however consistent with data analyses [Préteceille 2009] and qualitative evidence [Beaud and Pialoux 2003] which suggest that ethnic segregation, after a slow decline during the 1980s, has been deterioriating since the late 1990s onward.

3.3 Graphical evidence

Figure 1 plots the results of the pooled analysis displayed in table 2 for methods 1 and 2. We put, on the x-axis, the share of migrant population (centred around the local tipping point), and on the y-axis, the evolution of the native population (centred around the zone's mean).

Figure 1: *Origin*-based-tipping - Pooled estimation



Note: Results displayed come from the specification of table 2. x-axis is the share of migrant population within each district at the beginning of the decade minus the zone tipping point ($\delta_{ic,t-10}$). y-axis is the evolution of the native population as percentage of the base population at the beginning of the decade, centred around its mean ($\Delta u_{ic,t} - \overline{\Delta u_{ic,t}}$). Dots give averages in 1-percentage-point bins. We use two different fits: a 4th-order polynomial with an intercept shift at zero (light-grey) and a 3rd-order polynomial on the two subsamples right and left of zero.

Figure 2 gives an idea of what tipping looks like at a local level, taking the example of 4 zones. For this local approach, we do not center our variables: the x-axis reports the brute district migrant share, the y-axis, the brute evolution of native population share. We represent method 1 thresholds only.

The specific case of the Parisian metropolis is probably the most interesting one. The classical social geography of Paris is polarised over a East vs. West, bourgeois vs popular, axis. However, the analysis shall focus there, not on the extreme cases, but on these cities which lie just above or just below the threshold of tolerance.

These neighbourhoods share one pivotal common characteristics. Most of them are located adjacent to a priority district, but a district which does not belong to the worst cases, and is generally considered by the administra-

[6]Great care must be taken in interpreting these results: contrary to country of birth, nationality is not a stable individual characteristic and can change over the life course. The variations we observe might well be an upward trend in the naturalization process of foreigners.

tion as “low-priority”[7]. Most of these special suburbs are found at the frontier between the East-West poles, around a North-South axis. On the South axis, we might mention Boissy-St-Léger, Longjumeau, Vigneux-sur-Seine/Montgeron, Dammarie-les-Lys/Le Mée-sur-Seine; that last example is paradigmatic : it is a relatively calm low-priority district, but provides the sharpest decline of native population among the whole data (-78.5pp over 1999-2010 in the IRIS immediately adjacent to the Fontainebleau forest, where l_{1999} was just 1.5pp below the threshold; the effect is surely driven by the Plaine-du-Lys QPV; the IRIS at the heart of that suburb exhibits a tipping reaction of -29.2pp, starting from a far superior initial migrant share of 22%)[8]. On the North axis, compelling examples include Franconville, Sannois, Tremblay-en-France, Asnières-sur-Seine[9], Mantes-la-Ville (the IRIS adjacent to Mantes-la-Jolie tips at -28.1pp over 1999-2010 for $l_{1999} = 9.8\%$ migrant population share). One important final remark. Almost every single aforementioned city is either a low-priority district, or a district which was not considered as a QPV before 2015 ; tipping behaviour seems to predate the official classification[10].

4 Towards an economic interpretation

4.1 Significance of tipping models in their original U.S. context

Residential segregation and tipping reactions should not be considered as a universal and timeless feature of hierarchical societies. As emphasised by [Cutler, Glaeser, and Vigdor 1999], in late 19th c. America, all indexes of spatial segregation were paradoxically lower than they are today, and even lower in the South than in the North [Kellogg 1977][11]; in a sense, the racial hierarchy was so stringently enforced at the time that no isolation or distinction was necessary for the dominant group. The real story of American segregation dawns in the 1910s. It is the half-century of the American Apartheid best described by Denton and Masey [Denton and Massey 1992]: flocks of black immigrants move from the South to the major Northern cities. As a reaction, white residents, afraid at the idea of being submerged, build a comprehensive complex of institutions to ensure that their neighbourhood will remain segregated, exclusively-white. Hence these ill-famed legal devices of that time like contracts with restrictive clauses prohibiting resale to black persons (*restrictive covenants*), quotas on mortgage lending to black residents (*redlining*), corrupt promoters threatening locals to sell to a black person in order to make extra-profits (the so-called *blockbusters*). These restrictions were stringently enforced : in 1950 Chicago, 80% of real estate transactions had a racial clause [Clark and Perlman 1948]. Quantitative evidence is striking : the 1950s-1960s are the all-time maxima of segregation indexes[12]. In 1970, the average black person was living in a tract with an average 68% Black share. Concurrently, and maybe paradoxically, in the academic fields, it was a time of great quantitative inventiveness, with the rise of the main segregation measures like the dissimilarity[13] and isolation[14] indexes [O. Duncan and B. Duncan 1955], but also a time of heated debate on whether or not segregation was a

[7] In the technical jargon of the administration, such low-priority districts can be distinguished by the fact that they are not covered by the 2015 initiative known as the Nouveau programme national de renouvellement urbain (NPNRU), or by the fact that the NPNRU is considered of “regional” and not “national” interest.

[8] When it comes to *income*-based tipping, that commune provides also among the sharpest decline of the share of the middle 40% over 2000-2010 in the Paris zone ($\Delta u = -2.56pp$) and it lies just below the theoretical threshold for the area ($l_{2001} = 7.1\%$ vs. $l^*_{2000,Paris} = 7.2\%$)

[9] It is one of the most paradigmatic examples for the middle 40 vs bottom 10 analysis : $\Delta u_{2001,2011} = -1.8pp$ for $l_{2001} = 7.3\%$

[10] The same pattern can be identified in many provincial metropolises. In Toulouse, there is tipping West of the oldest and most well known QPV, Le Mirail, but the sharpest declines of the middle-class population are observed in recently created QPV units: Blagnac-Baradel, Colommiers-Val d'Aran, Toulouse-Soupetard. In Marseille, the sharpest drops of Δu are happening, not in the popular North, but in Southern (La Cravache - Le Trioulet) and Eastern priority districts (in M40vsB10 analysis, -16.2pp over 2001-2011 in the district just West of La Rouguière, where the initial l was relatively low, at 8.5%). In the zone of Saint-Etienne also, tipping is observed in peripheral cities relatively preserved till now: in Le Chambon-Feugerolles (in M40vsB10 analysis, -8.3pp, $l_{2001} = 10.2\%$) or at the heart of the ideal city planned by Le Corbusier at Firminy-Vert (in M40vsB10 analysis, -6.9pp, $l_{2001} = 12.6\%$).

[11] As shown by [Cutler, Glaeser, and Vigdor 1999], this will remain a permanent feature of Southern cities; throughout the 20th c., their mean dissimilarity index is systematically around 0.2 lower than the Northern average.

[12] In 1960, we get $D_c = 0.8$ and $I_c = 0.6$ [Cutler, Glaeser, and Vigdor 1999]

[13] The main segregation indexes of the American quantitative tradition were defined very early [O. Duncan and B. Duncan 1955], though attempts at axiomating their structure are very recent [Frankel and Volij 2011; Echenique and Fryer 2007]. The two most well known indexes are the dissimilarity and the isolation index. The dissimilarity index D_c is usually defined as:

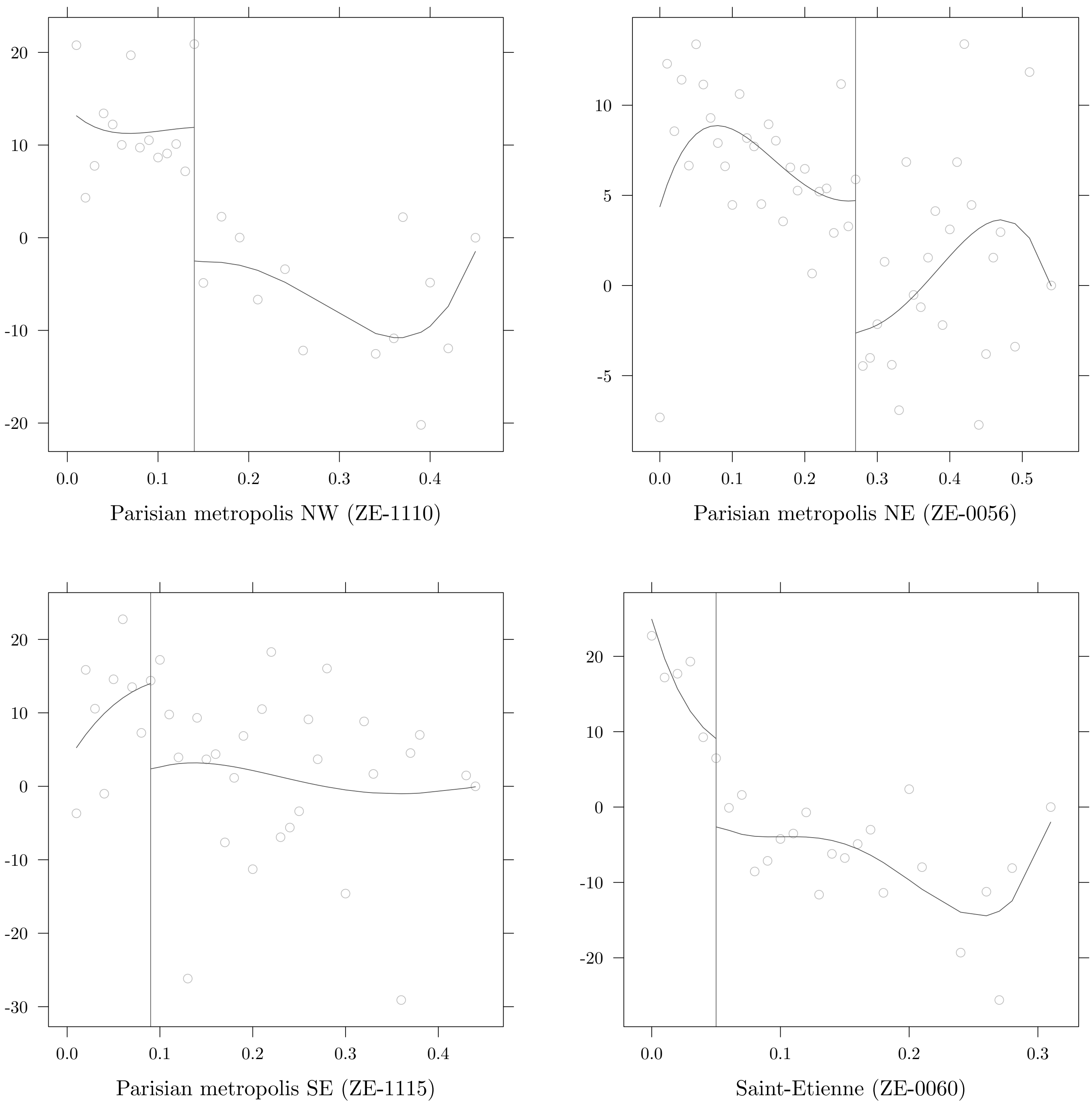
$$D_c = \frac{1}{2} \sum_i \left| \frac{b_{c,i}}{B_c} - \frac{w_{c,i}}{W_c} \right|$$

Which averages the ratio of the different census tracts i within a city c, $b_{c,i}$ being the black population of the tract, $w_{c,i}$ the non-black population of the tract, B_c and W_c, the black and non-black population of the whole city. D_c is interpreted as the share of the district's population which should move out to ensure spatial homogeneity.

[14] Using the same notations as in previous footnote, the isolation index I_c is defined as:

$$I_c = \sum_i \left(\frac{b_{c,i}}{B_c} \times \frac{b_{c,i}}{b_{c,i} + w_{c,i}} \right)$$

Figure 2: *Origin*-based-tipping - Some illustrative zones



Note: x-axis is the share of migrant population within each district at the beginning of the decade ($l_{ic,t-10}$). y-axis is the evolution of the native population as percentage of the base native population at the beginning of the decade, ($\Delta u_{ic,t}$). Dots give averages in 1-percentage-point bins. We use two different fits: a 4th-order polynomial with an intercept shift at zero (light-blue) and a 3rd-order polynomial on the two subsamples right and left of zero. The tipping point estimated through method 1 is shown as a dashed grey line. Results come from four different zones (from top left to bottom right): ZE1110 (Parisian metropolis North-West / Mantes-la-Jolie), ZE0056 (Parisian metropolis North / Roissy), ZE1115 (Parisian metropolis South-East / Evry) and ZE0060 (Saint-Etienne metropolis).

social evil in itself; in the Interwar period, Chicago School scholars had emphasised the fact that ghettos of the great metropolises were not necessarily places of social anomy [Whyte 1943] and could even act as springboards for economic opportunities [Halbwachs 1932], as a buffer for a faster integration within the nationwide labour market[15].

The Civil Rights policy of the Kennedy-Johnson administration put an end to this era of institutionalised segregation. Tipping models emerged by the same time, in the early 1970s; it was a paradoxical moment: the great Civil Rights Acts of 1964 and 1968 had curbed the tide of spatial segregation for the first time since a half century; yet the political coalition which had allowed the Democratic party to pass much of these policies fractured very early in the 1970s [Gethin, Martinez-Toledano, and Piketty 2021]; as a consequence, if the legal framework of the great Civil Rights Acts survived, many longer term provisions designed by the Johnson administration to foster residential desegregation were never applied; this is most notably the case of the 1968 Housing Act (one of the "three most important pieces of legislation of [his] presidency" according to president Johnson himself) which required, in each new housing project, a 50% quota for public or moderate rents housing. That scheme was terminated by the Nixon administration, which replaced it with two types of measures, suffused with a *self-help* narrative, which would define much of the policy path of the next decades: 1. Voucher programs to help minority families leave impoverished districts (the direct ancestor of what *Moving to Opportunity* will be for the 1990s) ; 2. A dilution policy meant to dissolve ghettos by relocating minority people to a myriad of little public housing projects built among middle-class suburbs.

However elusive their results[16], these policy items were widely imitated in Northern America and in Europe, being in harmony with the intellectual atmosphere of the time: the primacy of individual-level approaches of inequality; the conjecture that, great political aggregates (nation, class, religion) being on the decline, social identity had become more volatile and self-defined, that the social destiny of a person now depended less on its background than on its environment; it was thought that manipulating this environment and the type of people with whom that person was interacting with on a daily basis could fundamentally change its social fortune. Critical and acritical transcriptions in social sciences spawned a surge of phenomenological and interactionnist approaches, the econometrical equivalent of which were the peer-effects literature spawned by the seminal model of [Manski 1993]. Within this vein, tipping models might be construed as a formal attempt to nuance the tenets of the time, with the paradoxical conclusion that, even a purely interactionnist framework, with rational and benevolent agents, can spawn a social structure as thoroughly segregated as it was at the time of the American Apartheid, a phenomenon against which interactionist policies are doomed to remain powerless. Very early, the peer-effects literature itself provided intuitions that network impacts where either spontaneous representations of social life unidentifiable within a cautious research program [Abdulkadiroglu, Angrist, and Pathak 2011] or that they were encapsulating more structural underlying forces [Card and Giuliano 2016]. Unsurprisingly, this pervasive intuition, common to other academic fields, was doomed to generate a dialectic reaction of equal magnitude in favour of structural and holistic approaches.

4.2 Fleeing districts rather than people

There is evidence that tipping reactions are not limited to the *origin* dimension. In our annex F, we manage to identify tipping thresholds based on *income*, with the families belonging to the first decile of the national fiscal income distribution as the lower-group. Since many districts with high concentrations of non-native population are plagued with an accumulation of social and economic disadvantages [Algan, Hemet, and Laitin 2016], we could fear that our main specification might be unable to disentangle the origin-based tipping reaction from other driving variables. It is all the more true if we heed to the well-identified feature of the French case that the social perception of these phenomena has crystallised over the official State-sponsored proto-*Affirmative Action* labels: the ZUS-QPV labels for cities, the ZFU label for firms, and the ZEP-REP labels for primary education. If the impact

The isolation accounts for the fact that, even in an ideal city where minority people are not segregated at all (i.e. $D_c = 0$), if that minority represents a very low percentage of the population, it is improbable for a majority person to meet a minority person.

[15] Recent research in economic history tends to inform that thesis. [Pérez 2019] reviews a wide range of primary sources and recent studies about Italians in the U.S., urging the fact that economic integration of migrants was generally smoother in the West than in the great ghettos of Northwestern cities. Consistently, when [Abramitzky, Boustan, and Eriksson 2014] apply their equation main specification to a restriction on urban centres, the initial wage premia become wage penalties; i.e. *ceteris paribus*, urban interstitial districts were not the best places to assimilate economically for a newcomer migrant.

[16] As to the flaws of *Moving to Opportunity*-like projects, we might quote [Ludwig, G. Duncan, and Hirschfield 2001] or [Chyn 2018] ; for the failure of relocation policies, [Oreopoulos 2003].

of the two latter ones is discussed [Lafourcade and Mayneris 2017], school labels are known to trigger sizeable tipping reactions; [Beffy and Davezies 2013] estimate that, when a school becomes eligible for such a label, over 50% white-collar children and more than 80% of children of teachers leave instantly, whatever the initial social conditions in the school or in the surrounding district.

Hence the idea to check whether or not the tipping reactions we identify are concentrated around the most impoverished suburbs of the country as defined by official district-level labels, formerly the list of the *Zones urbaines sensibles* (ZUS) and since 2015, the new list of the *Quartiers prioritaires de la politique de la ville* (QPV).

In order to do so, we adapt one of the robustness checks of the original article[17]. Using interacted dummies, we shall see whether our main tipping coefficient is significantly different depending on the distance to the nearest priority district (with three intervals: less than 1km, between 1 and 3, more than 3). The results of that exercise for *origin*-based tipping points are displayed in table 3, those for *income*-based tipping points in the annex. These tables might be read as such: 1. The main effect provides an idea of the tipping reaction in the immediate vicinity of poor suburbs; 2. Line 3 provides an idea of that reaction in urban context; 3. Line 5 gives an estimate for a rural or semi-rural context. In column (7) and (8), we also test for the existence of spillover effects between districts[18]:

In the original article, all estimates of d were robust to an interaction with the distance to the nearest ghetto, i.e. racial tipping was nationwide in the late 20th c. United States, not concentrated on urban areas or around ghettos. Here on the contrary, we often fail to identify significant tipping reactions more than 3km away from a priority district; it seems that the aggregate effect we identified might be considerably driven by a priority-district-stigma[19]. However, we find some indications of a slow transition to the nationwide framework characteristic of the U.S. case: in the 2000s, tipping seems highly concentrated around poor suburbs ; away from urban centres, we even find nonsignificant *positive* values of d for the *origin*-based strategy. The early 2010s display quite a different picture: tipping reactions have now become more systematic, uniformly distributed over the territory.

The interpretation of this concentration around QPVs is more challenging. One line of argument that the tipping framework itself can discard is the idea that isolation would be driven by non-natives themselves, a thesis [Cutler, Glaeser, and Vigdor 1999] label the *port of entry theory* which, in the U.S. context, owes much to the theses of the Chicago School mentioned above.

Its main basis in contemporary research is that paradoxical finding of the literature that, even if higher segregation within a city [Cutler and Glaeser 1997] or equivalently higher diversity [Hémet and Malgouyres 2018] might hurt the labour prospects of minority workers, these workers seem to draw sizeable benefits from local hiring networks [Bayer, Ross, and Topa 2008], which are known to be racially stratified (i.e. neighbours are more likely to help each other find a job if they are of the same race [Hellerstein, McInerney, and Neumark 2011]) and relatively

[17] Card, Mas & Rothstein hypothesised that tipping might be triggered by some missing variable, one of them being the proximity of a ghetto or minority-dominant district. Over their 1990-2000 decade, it seems that tipping more than 2 miles away from a ghetto is not 5% significant. They also gauge the impact of distance from the CBD to isolate the impact of the great postwar white flight from the center of U.S. cities, but that did not seem relevant for our setting.

[18] In the original article, Card and his coauthors found a very large negative coefficient on this interacted variable Beyond TP $\times$ None of neighbours with $l > l^*$, of far greater magnitude that the main effect (-32pp vs -3pp for the 1980-1990). I.e. when a district located in a preserved or privileged neighbourhood tips, tipping reaction is much more violent. The authors interpret that result as evidence for the existence of considerable spillover effects between districts. The share of lower-group population in the district of residence might actually be only a proxy of the real variable driving departures; it seems like to some white families, there's a huge preference, not only for non-mixed districts, but also for non-mixed districts with non-mixed neighbours. That peculiarity is absent in our replication. The coefficient on the interacted variable Beyond TP $\times$ None of neighbours with $l > l^*$ is positive-significant, and total tipping reaction in districts with preserved neighbourhood, generally not significant. I.e., in the French context, the decision to leave is driven by the situation in the immediate vicinity, not by the situation of neighbouring districts, and at the margin, in the most preserved districts, spillover effects between IRISes might even be positive (middle-class family accept to remain in a district which has tipped, because neighbouring districts are still segregated (it is important to heed to the fact that the definition of what a preserved neighbourhood means varies considerably between zones.In provincial context, it generally means that all neighbouring districts have $l < 0.05$; within areas of the Paris metropolis, it is rather $l < 0.15$.).

[19] We must be extremely careful when comparing these results with the original article. Card and his coauthors defined a *ghetto* as a district with more than 60% minority population. If we use that definition over the INSEE's data, even with a far lower ceiling of 30% migrant share, distance to the nearest ghetto has no significant impact on the tipping behaviour. We find significant impacts when we use the distance to a priority district, and it is important to keep in mind that the QPV list is a very wide definition of what a poor suburb is (there are almost no districts in Paris or Lyon which are more than 3km away from a QPV). It is quite possible that our estimates simply capture the fact that tipping is exclusive to urban areas.

Table 3: Tipping reaction by distance from nearest priority district (*origin*-base tipping)

	Dist. to ZUS		Dist. to QPV		Dist. to new QPV		By nearby spillovers	
	1999-2010	2010-2017	1999-2010	2010-2017	1999-2010	2010-2017	1999-2010	2010-2017
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Main effect: beyond tipping point	-7.03^{***}	-2.97^{***}	-6.65^{***}	-3.51^{***}	-4.96^{***}	-2.82^{***}	-3.23^{***}	-1.84^{***}
	(0.75)	(0.65)	(0.71)	(0.63)	(0.85)	(0.63)	(0.55)	(0.45)
Interacted: Beyond TP × Nearest priority district is 1-3 km away	2.33^{***}	0.32	2.19^{***}	1.29^{**}	1.25	0.81		
	(0.65)	(0.56)	(0.59)	(0.51)	(0.75)	(0.51)		
Total tipping effect	-4.69^{***}	-2.64^{***}	-4.47^{***}	-2.22^{***}	-3.71^{***}	-2.02^{*}		
	(0.65)	(0.52)	(0.65)	(0.51)	(0.64)	(0.54)		
Interacted: Beyond TP × Nearest priority district is >3 km away	5.61^{***}	1.93^{***}	6.03^{***}	2.76^{***}	2.54^{***}	1.36^{**}		
	(0.67)	(0.61)	(0.64)	(0.67)	(0.76)	(0.56)		
Total tipping effect	-1.43^{**}	-1.04^{*}	-0.62	-0.76	-2.42^{***}	-1.45^{***}		
	(0.58)	(0.52)	(0.61)	(0.57)	(0.58)	(0.48)		
Interacted: Beyond TP × None of neighbours with $l > l^*$							2.56^{**}	1.32
							(1.13)	(1.41)
Total tipping effect							-0.66	-0.53
							(1.21)	(1.44)

Sign. thr. : *p<0.1; **p<0.05; ***p<0.01

Note: The unit of interest is the district (IRIS). Data are drawn from the INSEE's Census. We report the estimates of model 1, specifically the estimation of coefficient d. The dependent variable is the growth of the native population (defined as these persons who are born on French territory) within the district from t to $t-10$, as percentage of the base population at date $t-10$. The main explanatory variable is a dummy equal to one if the share of migrant population (those persons who are not born on French territory) is beyond the ZE-specific estimated tipping threshold. All specification also include a quartic polynomial in the deviation from the district's migrant share distance from the local tipping point, plus Z.E. fixed effects. The vector of controls, drawn from the INSEE's census, includes unemployment rate, share of working-class people, share of persons with no diploma, and share of vacant accommodation. The specification is still 1, now fully interacted with the indicated tract characteristic. We report results using the tipping points estimated with the fixed point method. In column 7 and 8, *neighbours* are the four closest IRISes (computed in terms of distance from centroid to centroid). Observations are weighted by total Census population, and clustered at the level of the region (pre-2015 geography).

efficient to help non-native workers to find a job in a labour market with hiring discrimination[20] [Dustmann, Glitz, et al. 2016[21]; Aslund, Hensvik, and Skans 2014; Giuliano, Levine, and Leonard 2009[22]; Bandiera, Barankay, and Rasul 2009[23]]. Even though some evidence suggests that these informal racial networks might lock-up minority workers in low-skilled, bad paying jobs [Hellerstein, Kutzbach, and Neumark 2014], where pay raises are slower [Dustmann, Glitz, et al. 2016][24], these local networks might be a second-best for many non-native people. Yet on

[20]It is an important feature of the works of [Aslund, Hensvik, and Skans 2014] or [Brown, Setren, and Topa 2012] that they seem to confirm the theory of statistical discrimination : for low-skilled jobs, in which a C.V. does not provide much information about the quality of a person, an extra source of information, coming from a referral for instance, might overcome the prejudice of managers against some types of persons.

[21][Dustmann, Glitz, et al. 2016] use a German dataset about newly hired employees ; for each employee, the share of workers of the same ethnicity that this person is used as a proxy to determine whether or not that person was hired externally, or through an informal job search ethnic-based network. The authors estimate that employees hired through these informal channels earn higher wages, are less likely to leave the firm, but experience a slower rise in wage.

[22][Aslund, Hensvik, and Skans 2014] use a Swedish linked employer-employee database to show that non-native managers significantly recruit more frequently non-native workers ; however, when native managers recruit within a pool of former co-workers, the bias against non-native employees disappears, indicating that the effect is driven, not by taste-based discrimination, but by the fact that non-natives do not have access to sufficiently large informational networks. [Giuliano, Levine, and Leonard 2009] make a similar point on a US dataset.

[23][Bandiera, Barankay, and Rasul 2009] show in a field experiment that when managers are paid with fixed wages, they will disproportionately recruit persons from their ethnicity, and that these matchs are generally sub-optimal for the firm. Performance pay for managers reduces that bias.

[24]These findings however are not always replicated in the literature : [Brown, Setren, and Topa 2012] is a good counter-example (in their setting, racial networks allow people to find jobs faster, and to be hired at a higher pay).

existing French data, it is difficult to substantiate such a thesis; in sheer descriptive statistics, we find no evidence that non-native families belonging to different SES groups tend to conglomerate together; the few significant correlation we found over Census data point to the opposite effect.

A more compelling interpretation involves the residential strategies of native middle-class families determined to avoid the reputational penalties attached to these districts, penalties which seems to have a strong impact on labour market outcomes. We allude to this special vein of the *Spatial Mismatch* literature [Gobillon, Selod, and Zenou 2005] which investigates the reputational effect of the place of residence indicated in a CV; in French context, there's evidence that indicating an address in a priority districts on one's CV is a considerable detrimental feature in a candidacy [Bunel, L'Horty, and Petit 2016], even once controlled for the ethnic origin of the applicant [Duguet, Leandri, et al. 2010]; more surprising, it seems like the burden of these penalties is borne by natives rather than by non-natives [Duguet, Gray, et al. 2020], i.e., employers who have overcome the racial prejudice will likely easily overcome the place-of-residence one, while prejudiced employers might overreact to the mention of a QPV in the CV of a native person.

4.3 Tipping reactions and the conservative vote

Another feature which vindicates against an identitarian interpretation of our findings is the fact that, as opposed to the original article, we fail to correlate the level of tipping thresholds with indexes of attitude towards race, especially here, with the FN-RN vote. Actually, it is much easier to isolate a correlation with the centre-right than with the far-right vote, a finding consistent with recent European estimates of the effect of migrations on local political outcomes [Barone et al. 2016; Dustmann, Vasiljeva, and Damm 2019].

In the original article, Card and his coauthors hypothesised that hostility towards contact between races would push the tipping point down. They find a significant correlation between the level of the threshold in each MSA and a specifically built Race attitude index. To take two extreme examples, in a city exemplifying the South like Memphis, the tipping point is 7 points lower than in a liberal city like San Diego.

We take a similar approach in the model of table 4, trying to find some correlation between the level of the tipping point and voting behaviour. We aggregated series from the CDSP-Sciences Po about city-level vote shares, to obtain vote shares at the zone level for each presidential election. We create five political aggregates (from far-right to far-left) and use the whole left-wing vote share as the base category. Empirically, the best predictors are the results of the election which happened next to the beginning of the corresponding decade (2002 for the 1999-2010 decade for instance).

Results suggest that tipping behaviour is polarised over the Left-Right axis. We find a strong correlation between the level of the tipping point and the conservative and centre-right vote, which consistently pushes the threshold down. Conversely, when estimated alone, the centre-left and far-left vote shares significantly push that threshold up[25]. The magnitude of the effect is sizeable : for the 2002 election, one extra percentage point for conservative candidates, as compared to the left base, is estimated to drive the tipping point 1.22pp down[26]. The negative marginal impact for the conservative total is not primarily driven by the Gaullist coalition: when we estimate the centre-right and Gaullist parties separately, we often find slightly superior coefficients for the latter ones (even in 2017, the coefficient on the LREM vote is negative, but nonsignificant).

More surprising, we fail to identify a significant impact of the Le Pen vote on tipping behaviour, even when we try a first-difference specification (to test the impact of the evolution of the FN-RN vote on the change in the level of the threshold). This might be explained by the heterogeneity of the far-right vote: we find very low tipping points in the old FN bastions of the South-East (3.9pp for Orange), while more recently conquered areas can exhibit extremely high thresholds (20.3pp for Béziers)[27].

[25] Estimated alone, the vote shares of L. Jospin in 2002 and F. Hollande in 2012 are among the strongest predictors which are positive and 5% significant

[26] Thresholds are indeed considerably lower in relatively affluent areas which are centuries-old bastions of conservative parties (2.7pp for Les Herbiers, 5.1pp for Rambouillet).

[27] Note however that the correlation might indicate a reverse causality: there is a suspicion, from the descriptive data, that areas with the clearest hikes in the far-right vote between 2002 and 2012 were characterised, over 1999-2010, by low tipping thresholds (+6.6pp in

Consistent with this interpretation, it is possible to obtain a significant negative far-right-coefficient in some specifications of table 4 with the addition of regional fixed effects, with the general picture that tipping points tend to be structurally lower (by a 10pp margin) in Provence and Nord regions (in which many old bastion cities of the party are found), but relatively high in other regions of the East where the FN scored recent victories.

Table 4: Determinants of the *origin*-based tipping point

	Dep. var.: Level of the tipping threshold							
Time period	Decade 1999-2010				Decade 2010-2017			
Election year	*pres. 2002 rd1*		*pres. 2007 rd1*		*pres. 2012 rd1*		*pres. 2017 rd1*	
Method for tipping point est.	*Str. br.*	*Fix. pt*	*Str. br.*	*Fix. pt*	*Str. br.*	*Fix. pt*	*Str. br.*	*Fix. pt*
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Vote share in corresponding election								
Far-right	−0.07	−0.04	−0.01	−0.14	0.38	0.21	0.05	−0.002
	(0.16)	(0.22)	(0.29)	(0.32)	(0.25)	(0.24)	(0.13)	(0.08)
Right + centre-right	−1.22**	−0.18	−0.74**	−0.05	−0.58*	−0.31**	−0.46**	−0.17**
	(0.43)	(0.21)	(0.32)	(0.19)	(0.32)	(0.15)	(0.17)	(0.08)
Centre-left + far-left as base								
Controls	X	X	X	X	X	X	X	X
Obs.	304	304	304	304	304	304	304	304
R^2	0.18	0.14	0.09	0.15	0.23	0.05	0.31	0.05
F-stat	11.1***	8.3***	4.9***	8.4***	14.3***	2.7**	21.7***	2.7**

Sign. thr. : *p<0.1; **p<0.05; ***p<0.01

Note: The unit of observation is the ZE (*Zone d'emploi*, INSEE 2010 def.). The dependent variable is the level of the tipping threshold of natives, reacting to non-native pop. shares, estimated with both methods in tables 2, over the data of the Census, for the decade indicated in the first row. The main explanatory variables are the vote shares in the presidential election of interest. Political groups are defined in the annex. Controls are taken from the Census and include: the share of residents in the IRIS who live in a single-family home, the share of those who own their home, the share of those who live in a public housing unit (HLM), and the average number of persons per room in the IRIS. Coefficients on the control variables are always nonsignificant. Observations are weighted by the total Census population, clustered at the level of the INSEE's superzones.

5 Robustness checks

One issue might imperil the zero conditional mean assumption in this framework. The variations of our dependent variable might be driven, not by departures or arrivals, but by some inner phenomenon: maybe the local native population rises because of demographic dynamics (extra births for instance).

To that objection we can provide both a theoretical and an empirical answer: 1. From a theoretical standpoint, we can show than the growth rates of each local upper group are independently distributed dependent of the size of the upper group at the beginning of the period[28]; 2. As to the empirical dimension, one solution is provided by the grand mobility dataset of the INSEE's Census, which allows us to know, for each recorded individual, in which commune that person lives at time t, and in which commune that person was living at time $t-5$ or $t-10$

the Le Pen vote share between 2002 and 2012 for Béthune, where the tipping threshold is estimated at 1.1pp over the decade). Besides, in unreported results based on the *Mobilité* datasets of the INSEE, we found some evidence that a rise in the native population coming from a district that has *tipped* over the decade is associated with a rise in the FN-RN vote.

[28]This independence property is easily spoted on detailed data: to take the example of the *income*-based strategy, limiting ourselves to Paris, we see that destitute IRISes with very low values of $u_{i,2001}$ below 25% exhibit almost aleatory variations of the middle-class population (La Goutte d'Or III : $\Delta u_{i,2001,2011} = -1.11$ pp starting from $u_{i,2001} = 25.1\%$ over 2000-2010; La Courneuve - La Tour : +3.9pp starting from 20.1%). Conversely, the clearest tipping reactions are observed next to low-priority districts (with an extremum at −17.6pp in Asnières-sur-Seine, where in 2001, the deciles were almost mathematically drawn, with the bottom 10 making out 10.9% of the inhabitants, the middle 40, 39.93%).

(depending on the versions of the set). Over these datasets, it is possible to clearly isolate the flux of departures and arrivals from the inner dynamics. In the annex, we replicate our *origin*-based strategy on these mobility series; we find very similar tipping points, and coefficients d which are notably higher than the ones reported in table 2. If indeed there is a bias in our estimates because of the inner dynamics of the upper group, that bias leads us to underestimate the magnitude of d. Here's a quick overview of the other robustness checks provided in the corresponding annex:

- 1. Falsification exercises and placebos to check for biases in variables and selections:
  - 1.1. *Placebo tests on upper on lower groups* – We replicate our strategy on different alternative, closely related upper and lower groups. We find no consistent result on any other definition of the groups.
  - 1.2. *Flexible higher-order polynomials in controls* – We test the main specification with a quartic polynomial in each control variable, finding very similar estimates of d.
  - 1.3. *Alternative sources: Mobility series from the INSEE* – See above.
- 2. Robustness checks involving spatial interactions:
  - 2.1. *Alternative spatial levels* – There's a hotted debate in the spatial econometrics literature about a phenomenon known as MAUP (*Modifiable Areal Unit Problem*); i.e. in many empirical models, changing the spatial unit interest imperils the robustness of the estimates; coefficients which are sizeable and significant at the regional level might be nonsignificant at the county level. To check for such a risk, we tested different *supra*-units (UU, AAV) and different *infra*-units (the commune instead of the IRIS). Our estimates are robust to these changes;
  - 2.2. *Controlling for the proximity of a priority district* – See tab. 3;
  - 2.3. *Controlling for spillover effects between districts* – See tab. 3;
  - 2.4. *Checking spatial autocorrelation* – In the original article of Card and alii, there were suspicions of strong retroactive influence between districts (one upper group family leaves district i; families from neighbouring districts react with tipping; tipping from neighbours creates a retroactive effect in district i). We tested a Spatial Durbin model, very similar to the one reported in our annex H, to distinguish direct and retroactive effects on tipping; we found that these retroactive effects are most of the time nonsignificant: the decision to leave is influenced by the social situation in the immediate backyard only (a major difference with the American context).

6 Conclusion

Applying the Schelling tipping point model to a contemporary French data yields results which are widely consistent with the existing European literature, but tend to undermine the idea of a French or European exceptionalism on that spot. We investigated tipping points where shifts in migrant population share significantly impact native population dynamics. Our estimates reveal that even slight changes in migrant population proportions can trigger notable responses in native population trends. As to the location of these thresholds, and the intensity of the reaction around them, we find figures which overall lie below the U.S. levels, but by a slight margin only.

Our results are not consistent with an expanding ghetto pattern as does of [Card, Mas, and Rothstein 2008] overall, but a little number of very segregated areas exhibit emerging tendencies of that sort.

We argue that interpretations focused on political or identitarian polarization, which have dominated the U.S. literature, fare poorly in European context.

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Annex A – About the estimation strategy

Choosing data, finding relevant variables

Datasets - Card, Mas & Rothstein were relying on data from the U.S. Census. A natural French equivalent is the census of the INSEE (we used the issues of 1962, 1968, 1975, 1982, 1990, 1999 and the New annual Census from 2006 onward).

Geographical unit of interest - Applying our specification over actual data requires of choice of geographical level: one must pick a *supra*-unit, the city c, and an *infra*-unit, the district i. In the authors' original setting, the *supra*-unit is the MSA (Metropolitan Statistical Area), the *infra*-unit, the census tract. Datasets from the INSEE leave us with a wide range of options: the *infra*-unit might be the *commune* or the infracommunal unit (formally, the IRIS) ; for the *supra*-unit, we might take any definition of the urban agglomeration used by official statistics, from the narrowest, known as the UU (*unité urbaine*) to the widest, the AAV (aire d'attraction des villes), or a middle definition, the ZE (Zone d'emploi) which has the comparative advantage to cover the entirety of the territory:

- The main drawback of using communes lies in the fact that, contrary to the census tract, it is a highly uneven division (in our subsample, the tiniest communes do not surpass 1000 inhabitants, while the largest unit, Toulouse, is far above 300.000). However, communes are among the oldest administrative divisions; while the IRIS are barely known outside of a limited circle of experts, communes are the pivotal level at which citizens perceive the impact of public policies and social changes (symptomatically, municipal elections are among the few elections which have resisted the decline in the turnout rate over the last decades).
- Conversely, IRIS divisions are more homogeneous. Every commune with more than 10.000 inhabitants, and most communes over 5000 is divided into IRISs, and communes below 5000 act as an IRIS in the INSEE system. They are extremely useful when studying major metropolises like Paris or Lyons, but for the remainder of the French territory, they leave us with the same issue about the size of communes mentioned herein above.

All in all, IRISes in ZEs seems to be the most natural choice, at least the one which is the most similar to the original setting. To provide one comparative example, the city of Chicago has 2.7 million inhabitants and 866 census tracts, while the ZE of Lyons has 1.2 million inhabitants, 248 communes (in our subsample) and 538 IRISes. However, since there is wide controversy among spatial econometricians about how the choice of the geographical unit might bias the estimates, we will systematically report results at the AAV, UU and *commune* level in this annex.

Controls - Another issue is the choice of the control variables in $X_{ic,t-10}$. The authors have recourse to the following ones, always computed at the tract level: unemployment rate, log(mean family income), housing vacancy rate, renter share, fraction of homes in single-unit buildings, and fraction of workers who commute using public transit (used as a proxy for working-class population). At the IRIS level, we use data from the Census of the INSEE with the following controls: unemployment, share of persons with no diploma, and share of vacant accommodation; share of migrants (people not born in France) and share of blue-collar workers are included when they are not colinear with the explanatory.

Summary statistics

We'll use two major sources from the INSEE: the Census and the RFL-Filosofi database.

The Census provides highly reliable estimates with almost no selection issues. To provide but one example, the 1999 Census contains 50066 IRISes with a non-zero population (or 57.9 million inhabitants). Once we have ensured that we can match the IRISes over time, removed those units the INSEE recommends not to use in comparative analysis, and applied our selection rules[29], we are left with 47878 units, or 57.1 million inhabitants (87% of the national population).

The RFL-Filosofi sets, since they are dependent on statistical secrecy rules, provide far less information. Bracket breakdown of income is recorded for roughly one third of IRISes, mostly urban ones. This urban subsample has a

[29] The IRIS must be within a *supra*-unit, and within each area, we require a minimum of 12 IRISes.

more unequal income structure than the whole nation (in 2011, 11.9% of the subsample population lies in decile 1, 11.1% in decile 10, the population share of any other decile being below 10%; i.e. in our urban subsample there's a overrepresentation of the richest and poorest households). For the 2001-2011 restriction rules, once applied, leave us with 15615 units (or 51.4 million inhabitants, 20.6 million households in 2001). A detailed overview is provided in Table 5.

Table 5: Summary statistics

	Origin-based strategy		*Income-based strategy*	
	1999-2010	2010-2017	2001-2011	2011-2019
	(1)	(2)	(3)	(4)
Nb. of IRISes in original dataset	50066	50885	48404	48386
Nb. of IRISes preserved over time	49764	49281	48015	41978
Nb. of IRISes in final sample	48813	48848	15615	13326
Nb. of ZEs in final sample	305	305	199	199
Mean lower group share at beginning of decade	10.1	8.6	11.1	11.9
Growth in total population over the decade	7.1	3.8	13.9	6.57
Growth in upper-group population	7.8	2.5	10.9	6.54

Growths of ind. pop. for the Census, of the tax unit pop. for the RFL sets.

In the original article, there were 110 zones in the sample, and a mean 300 *infra*-units (census tracts) within each area. In our replication, we get 305 *supra*-units, and an average 160 *infra*-units within each one.

For the RC-IRCOM dataset, we used a special restriction which concatenates civil years 2000, 2010 and 2020 (which we called R4), which originally contains 4216 cities. Once restriction rules are applied[30], we are left with 2958 cities or 18.43 million tax units.

Structural break searching algorithm

In order to estimate the ZE-specific tipping point, we rely on two different methods.

Method 1 (Structural break) - In this setting, identification of the thresholds relies on the methodology of [Hansen 2000]. For each supra-region or ZE c, we define a search interval $[m_1, m_2]$ where:

$$m_1 = \min(l_{ic,t-10})$$
$$m_2 = \max(l_{ic,t-10}) \times \mathbb{I}[l_{ic,t-10} < 0.5]$$

Candidate tipping points are all the values of the sequence defined by $v_{n+1} = v_n + 1/1000, v_0 = 0$ that lie within $[m_1, m_2]$.

For each candidate point l^* within each area, the dependent variable $\Delta u_{ic,t}$ is regressed over a dummy equal to one if $l_{ic,t-10} > l^*_{ic,t-10}$, formally if $\delta_{ic,t-10} > 0$ (i.e. if the initial lower-group share is above the tipping threshold):

$$\Delta u_{ic,t} = a_c + \alpha \mathbb{I}[\delta_{ic,t-10} > 0] + \varepsilon_{ic,t}$$

That simple operation is reproduced over the whole range of hypothetical tipping points $l^*_{c,t-10}$ within the extreme values of $l_{c,t-10}$. The chosen area-specific tipping point is the value of $l^*_{c,t-10}$ which maximises the R-squared of the model (provided that coefficient α is negative)[31].

[30] I.e. being in a ZE, and having a minimum of 6 cities within each ZE

[31] Actually, in some parts of the replication STATA code provided by Jesse Rothstein, they are using the maximisation of the t-stat as the main criterion to choose the tipping point. We tested this criterion on some sets, but this alteration tends to make the threshold estimation less congruent across methods. Since their published article mentions the R-squared as the only criterion, consistent with the method described by Hansen 2000, we keep the original setting.

Method 2 (Fixed point) - Card, Mas & Rothstein propose a complementary, non-standard method to the classical approach of Hansen, which, according to them, performs better for lower-size cities. The intuition is that if we take the average of $\Delta u_{ic,t}$ over the zone, and center our dependent variable around this mean, the evolution of the upper-group population shall be positive in districts below the tipping point, and negative for districts which are beyond the tipping point. I.e., if we plotted our centred dependent variable vs. the lower group share, we should see a piece-wise function that is equal to 0 at l^*. More formally, Card and his coauthors rely on a smooth approximation of the differential variable:

$$\Delta u_{ic,t} - E[\Delta u_{ic,t}|c]$$

We fit that variable to a quartic polynomial in $l_{ic,t-10}$; to obtain a function $R(l_{t-10})$. The algorithm of method 2 then finds the roots of that polynomial, provided it is below 0.5. If multiple roots are identified, the algorithm picks the one point at which the first-order derivative of $R()$ is the lowest (most negative).

Method 1 algorithm returns a missing value when it cannot find any structural break associated with a negative drop. As to method 2 algorithm, it creates a *NA* when $R()$ has no root.

Exploratory results - Over the INSEE's census

One important limitation of empirical strategies tackling the issue of spatial segregation is the element of arbitrariness in the choice of the dominant and of the discriminated groups. There exists auto-aggregative algorithms which infer the relevant groups from the patterns of the spatial data themselves; we applied the one of [Louf and Barthelemy 2016] at the IRIS level for a fiscal-income-based strategy, yet, as in their own application, it leads us to use three main aggregates (the bottom 50% of the fiscal income distribution, the top 40%, and a little middle 10% in-between) which do not perform well in a tipping setting.

We therefore empirical tested several alternative definitions of the upper and lower group, comparing two issues of the INSEE's Census (1999 and 2010) and two issues of the RFL-Filosofi set (2001 and 2011).

Over the Census, group labels include the place of birth (for simplicity of language, we will improperly call *natives* those who are born on French territory, and *migrants* those born in a foreign country), the 8-classes SES scale of the INSEE (especially, the shares of white-collar and blue-collar workers, of employees and intermediate occupations ; we use the label *working-class* for the sum of blue-collar and employee population). Due to data limitations, the dependent variable must focus on each separate item : SES or place of birth (the interaction of the two being absent from the 2010 series). The explanatory on the contrary can rely on such an interaction (we can use as *explanans* the share of working-class migrants in 1999 for instance).

We always report the result of the full pooled model, with Δu as the dependent, and as explanatory variables, a quartic polynominal on δ, a dummy equal to one when $\delta > 0$, *supra*-units-specific fixed effects, and the full vector of controls. We require a minimum of 12 infra-units within the geographical supra-units (6 when the infra-unit is the commune). Controls in these specifications include: share of vacant accommodation, unemployment, share of people with no diploma, and when it is not redundant with the explanatory, share of migrants and share of working-class people. Over the RFL-Filosofi set, we add an extra control, the mean fiscal income of the household within the IRIS. Reported statistics are weighted by supra-unit (ZE) total population. Standard errors are clustered at the supra-unit level.

Table 6: Full model with alternative upper/lower groups and alternative spatial units (I)

Upper gr.	Lower gr.	M1/2 thr.	Method 1			Method 2			Obs.
			F-stat	d	t-stat	F-stat	d	t-stat	
—INSEE's Census - ZE / IRIS—									
Natives	Migrants	10.4/11.6	4.6	-2.29	2.44	4.78	-3.84	4.36	16271
Natives	Work.-clas. migr.	4.1/5.2	8.2	-2.2	3.7	8.3	-3.1	5.7	id.
White-col.	Work.-clas.	18.8/23.9	5.4	$+0.1$	0.67	7.8	-0.5	1.63	id.
White-col.	Work.-clas. migr.	3.1/4.2	4.9	-0.03	0.2	7.6	-0.09	0.2	id.
Intermediate	Work.-clas.	16.3/19.9	2.7	$+0.55$	1.37	2.6	-0.33	0.84	id.
Intermediate	Work.-clas. migr.	3.1/5.6	2.7	-0.67	3.9	2.8	-0.68	4.1	id.
—INSEE's Census - AAV / IRIS—									
Natives	Migrants	8.2/10.4	4.9	-3.05	4.8	5	-3.69	5.9	12903
Natives	Work.-clas. migr.	2/3.4	3.2	-5.1	6.5	3.1	-3.6	4.9	id.
White-col.	Work.-clas.	27.1/31.2	2.3	-2.1	3.4	2.2	$+0.67$	1.4	id.
White-col.	Work.-clas. migr.	2.3/6.1	3.7	-0.17	1.1	3.3	-0.26	1.7	id.
Intermediate	Work.-clas.	16.3/19.9	2.7	$+0.55$	1.37	2.6	-0.33	0.84	id.
Intermediate	Work.-clas. migr.	2.9/3.4	3.9	-1.1	5.13	3.3	-0.71	3.41	id.
—INSEE's Census - UU / IRIS—									
Natives	Migrants	7.3/11.3	4.1	-6.37	2.81	3.9	-6.51	3.61	5084
Natives	Work.-clas. migr.	2.5/5.2	2.1	-8.1	5.12	1.9	-1.68	1.32	id.
White-col.	Work. clas.	18.1/21.7	1.4	$+0.04$	0.13	1.4	$+0.07$	2.4	id.
White-col.	Work.-clas. migr.	2.2/24.1	4.2	-2.36	1.85	4.1	-1.3	1.45	id.
Intermediate	Work. clas.	19.1/26.3	1.8	-4.7	0.67	1.7	-26.2	4.7	id.
Intermediate	Work. clas. migr.	2.3/5.4	1.9	-1.19	1.1	1.9	-1.6	2.24	id.
—INSEE's Census - ZE / Commune—									
Natives	Migrants	7.52/9.66	7.4	-0.23	0.41	7.6	-3.39	5.43	11587

Over the RFL-Filosofi set, available group labels include the ten deciles of the national structure of fiscal income.

Table 7: Full model with alternative upper/lower groups and alternative spatial units (II)

Upper gr.	Lower gr.	M1/2 thr.	Method 1			Method 2			Obs.
			F-stat	d	t-stat	F-stat	d	t-stat	
—RFL-Filosofi - ZE / IRIS—									
Decile 2	Decile 1	14.4/15.81	8.7	−0.29	2.7	8.9	−0.56	4.4	5205
D3	D1	17.6/13.7	10.2	+0.1	0.93	10.7	0.04	0.38	id.
D4	D1	15.37/12.6	11.1	−0.46	3.7	10.8	−0.34	1.53	id.
D5	D1	12.3/14.8	8.7	−0.4	4.1	8.5	−0.59	5.2	id.
D6	D1	11.6/13.8	8.9	−0.68	6.2	8.8	−0.45	2.9	id.
D7	D1	10.6/14.1	8.5	−0.53	4.5	11.9	−0.69	9.2	id.
D8	D1	9.4/10.9	7.1	−0.64	4.1	13.7	−0.71	8.42	id.
D9	D1	11.31/11.7	6.9	−0.61	6.8	12.4	−0.62	7.46	id.
D10	D1	6.6/20.1	19.6	−0.41	1.41	20.9	−0.84	4.14	id.
D6 to 9	D2 to 5	40.1/37.5	17.12	−0.43	1.3	17.1	−0.63	1.81	id.
—RFL-Filosofi - AAV / IRIS—									
D6 to 9	D1	9.4/10.3	12.8	−1.26	3.1	12.9	−1.74	4.2	4746
D2 to 5	D1	13.8/19.4	4.2	−1.5	3.8	4.6	−0.6	1.15	id.
D10	D1	6.7/18.2	7.7	−0.65	4.3	7.6	−0.5	3.5	id.
D6 to 9	D2 to 5	44.1/46.2	8.1	−0.2	0.56	7.6	−0.5	3.4	id.
—RFL-Filosofi - UU / IRIS—									
D6 to 9	D1	9.33/9.4	12.8	−2.37	4.67	12.9	−0.94	2.28	3348
D2 to 5	D1	14.3/38.7	11.9	−0.23	0.35	11.5	−1.79	2.8	id.
D10	D1	14.4/22.5	6.4	−0.29	1.1	6.3	−1.41	2.3	id.
D6 to 9	D2 to 5	28.9/36.8	14.5	−1.53	-4.67	13.9	−0.03	0.04	id.
—IRC - AAV / Commune—									
D6 to 9	D1	7.6/8.1	4.8	−0.85	0.14	5.6	-1.56	2.51	1930
D2 to 5	D1	7.7/11.4	5.7	−0.33	1.1	5.8	-0.07	0.18	id.
D10	D1	7.3/7.6	6.2	−0.55	2.25	5.9	-0.59	2.9	id.
D6 to 9	D2 to 5	37.6/32.5	6.6	+0.15	0.34	6.7	-0.36	0.76	id.

Annex B – Towards an alternative strategy based on incomes

More interesting, the RFL-Filosofi allows us to search for tipping behaviour based on income. For these *income-based*-thresholds, we'll define the upper group as the *Middle 40%* (these households, the disposable income of which is comprised between decile 5 and 9 of the national distribution of income) and the lower group as the *Bottom 10%* (these households, the fiscal income of which is below decile 1 of the national distribution of income)[32]. We did not define those class arbitrarily: as we show in our annex, these two classes are the only one on which consistent tipping is identified[33]. The rationale behind this reaction seems relatively straightforward: To sum it up :

- The bottom 50% are trapped into an unwilling sedentarity. When the social conditions of their district starts to deteriorate, they are unable to move away. We detect some departure trends in districts with very high levels of underprivileged population, but no tipping or discontinuity. Constrained on the credit market, working-class families do not have the financial leverage to back an individualistic-optimizing residential strategy;

- The top 10% barely react to the rise of local migrant or underprivileged population. They are living in relatively preserved districts. Besides, these groups have close control over local politics, and are able to influence housing policies to maintain population homogeneity[34].

- The middle 40% are the only ones which react with tipping. They do not have enough control on local policies to prevent the social evolution of the neighbourhood, but sufficient leverage to leave when conditions deteriorate.

As we see in table 8, thresholds based on income are of very similar magnitude than thresholds based on origin, around the national average of each lower group. However, standard deviations are far lower for *income*-based estimates, and cross-correlation less robust; tipping based on income seems more systematic, less concentrated in certain areas (estimated *origin*-base-thresholds are often abnormally low in areas with high conservative or far-right vote shares):

Table 9 provides estimates for the tipping behaviour around *income*-based estimated thresholds. For each method, we provide the average drop in the middle-class population (those belonging to the national middle 40%) when the share of underclass tax units (those belonging to the national bottom 10%) is beyond a area-specific tipping threshold computed by our two methods.

The zone of Lyon provides a good example of income-based tipping dynamics. Among popular suburbs at the East of the city, the most deprived QPV still maintain a relatively stable middle-class, which seems unwilling to leave the district. At the heart of Bron-Parilly, the middle 40% account for 20.2% of the population in 2001, and there are only 8 departures coming from that group over the next decade. The sharpest declines in Δu are to be found further South, in individual houses districts which are adjacent to a well known QPV. One very symbolic example is found in Vénissieux - Les Minguettes, a priority district known to be the birthplace of one of the main antiracist activist movement of the late 1980s. In the IRIS lying at the heart of that suburb, the bottom 10% make out 29.4% of the population, and 33.1% of the population is not born in France ; but the sharpest tipping behaviour from the middle 40% are found in nearby single-home districts: (Charreard and Chassagnon with resp. -6.9 and -6.7pp over 2000-2010) which have remained relatively mixed neighborhoods (with l_{2001} at respectively 8.9 and 9.2%).

[32] The RFL-Filosofi dataset provides the deciles of fiscal income for the whole country, and for almost all IRISes which lie within a urban unit. It is then possible to use the interpolation method pioneered by [Blanchet, Fournier, and Piketty 2017] to determine the share of each national group within each district. To give but one example, over the whole zone of Lyon, the bottom 10% make out exactly 10.04% of the population, the middle 40%, 40.86. But in the most deprived priority districts, the figures are almost reversed: in Bron-Parilly, the bottom 10 provide 30% of the local population, the middle-class, 11%.

[33] Here again, we tested a wide range of specifications, the most important ones reported in our annex. The main outcome of that exercise is that tipping is exclusively observed coming from the middle 40% as they react to the bottom 10% population. The middle-class does not react with tipping to the dynamics of any other group (be it defined by income, SES, place of birth, or interaction between these last two); conversely, the bottom 10% population triggers no tipping from another group, whatever its definition.

[34] The list of communes which do not respect the ceilings of local public housing mandated by the 2000 SRU law, recently updated by [Ministère chargé du logement 2020], is almost comprised of communes with the highest of local population earning more than decile 10, if not centile 100 of the national structure of income

Table 8: *Income*-based-tipping - Estimated tipping points

	2001-2011		2011-2019	
	Structural break	Fixed point	Structural break	Fixed point
	(1)	(2)	(3)	(4)
Mean	11.52%	13.44%	11.85	13.71
SE	5.19	7.17	6.12	8.65
Without identified threshold	0	0	0	0
Correlations				
2001-2011 Structural break	1.00			
2001-2011 Fixed point	0.37	1.00		
2011-2019 Structural break	0.11	0.12	1.00	
2011-2019 Fixed point	0.06	0.11	0.26	1.00

Points are expressed in share of bottom 10% pop. in district. Summary stats are unweighted.

Table 9: *Income*-based-tipping - Regression discont. model for middle 40% pop. change around the tipping pt

	Dependent var.: Change in middle-class population in the district from $t-10$ to t					
	Method 1 - Structural break			Method 2 - Fixed point		
	Base	F.E.	Full	Base	F.E.	Full
	(1)	(2)	(3)	(4)	(5)	(6)
2001-2011 decade						
Beyond tipping point (coef. d)	-3.38pp***	-2.22pp***	-1.93pp***	-3.42pp***	-2.06pp***	-1.75pp***
SE	(0.63)	(0.54)	(0.64)	(0.49)	(0.54)	(0.64)
Observations	5205	5205	5205	5205	5205	5205
R^2	6.4%	25.7%	29.1%	7.6%	25.7%	29.1%
F-stat	70.6	8.6	9.8	8.97	8.57	9.81
2011-2019 decade						
Beyond tipping point (coef. d)	-2.85pp***	-3.46pp***	-2.46pp***	-2.81pp***	-3.93pp***	-2.39pp***
SE	(0.36)	(0.35)	(0.34)	(0.34)	(0.35)	(0.33)
Observations	4442	4442	4442	4442	4442	4442
R^2	8.7%	21.9%	33.1%	5.2%	22.3%	33%
F-stat	126.2	9.8	16.7	71.7	10.1	16.6
ZE fixed effects		X	X		X	X
Controls			X			X

Sign. thr. : *p<0.1; **p<0.05; ***p<0.01

Note: The un. of obs. is the IRIS. The dep. var. is the growth in middle 40% pop. (def. as these tax units, the fisc. inc. of which lies betwe. dec. 5 and 9 of the nat. struct.) within the IRIS over $t, t-10$, as perc. of the base pop. at $t-10$. The main exp. var. is a dummy equal to one if the share of bottom 10% pop. (simil. def.) is beyond the ZE-specif. estimated tipping pt. All spec. include a quartic polyn. in the devia. of lower-group shares from the tipping pt, plus ZE fixed eff. Controls include unempl. rate, blue-collar sh., sh. of migrants, no diploma sh., vacant accom. sh. and log mean fisc. inc. SE are clustered at the ZE level.

Annex C – Supplementary robustness checks

Higher-order polynomial in control variables

Table 10: Sensitivity of the tipping coefficient to flexible controls (*origin*-based tipping)

	(1)	(2)	(3)	(4)	(5)	(6)
1999-2010	-2.29^{**}	-2.19^{**}	-2.11^{**}	-2.18^{**}	-2.02^{***}	-1.83^{**}
	(0.94)	(0.88)	(0.86)	(0.93)	(0.92)	(0.91)
4-th order polynominal in:						
Unemployment		X				X
Share with no diploma			X			X
working-class share				X		X
Vacant accommodations rate					X	X

Sign. thr. : *p<0.1; **p<0.05; ***p<0.01

Note: The specification is still 1, plus quartic polynomials in the indicated variable. We report exclusively the results from the first method of estimation of the tipping points, results for method 2 being similar. Standard errors are in parentheses.

Table 11: Sensitivity of the tipping coefficient to flexible controls (*income*-based tipping)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
2001-2011	-1.93^{***}	-1.31^{***}	-1.39^{***}	-1.29^{***}	-1.61^{***}	-1.45^{***}	-1.47^{***}	-0.82^{**}
	(0.64)	(0.41)	(0.38)	(0.38)	(0.41)	(0.39)	(0.39)	(0.39)
4-th order polynominal in:								
Unemployment		X						X
Logged mean income			X					X
Migrant share				X				X
Share with no diploma					X			X
working-class share						X		X
Vacant accommodations rate							X	X

Sign. thr. : *p<0.1; **p<0.05; ***p<0.01

Note: See fig. 10

Distance to a priority district

Table 12: Tipping reaction by distance from nearest priority district (*income*-base tipping)

	Dist. to ZUS		Dist. to QPV		Dist. to new QPV		By nearby spillovers	
	2001-2011	2011-2019	2001-2011	2011-2019	2001-2011	2011-2019	2001-2011	2011-2019
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Main effect: beyond tipping point	-2.73^{**}	-3.68^{***}	-3.21^{***}	-3.32^{***}	-2.27^{***}	-2.35^{***}	-1.63^{***}	-2.83^{***}
	(1.01)	(0.44)	(0.87)	(0.41)	(0.77)	(0.44)	(0.54)	(0.34)
Interacted: Beyond TP × Nearest priority district is 1-3 km away	$+0.92^{*}$	$+1.43^{***}$	$+1.28^{**}$	$+0.76^{**}$	$+0.69$	-0.28		
	(0.48)	(0.34)	(0.55)	(0.33)	(0.45)	(0.44)		
Total tipping effect	-1.81^{***}	-2.25^{***}	-1.93^{***}	-2.56^{***}	-1.57^{***}	-2.64^{***}		
	(0.47)	(0.37)	(0.48)	(0.38)	(0.46)	(0.38)		
Interacted: Beyond TP × Nearest priority district is >3 km away	$+1.79$	$+0.81^{*}$	$+3.61^{***}$	$+0.91^{**}$	$+0.93^{*}$	-0.5		
	(1.15)	(0.41)	(1.05)	(0.44)	(0.55)	(0.39)		
Total tipping effect	-0.94^{*}	-2.87^{***}	$+0.41$	-2.41^{***}	-1.34^{**}	-2.85^{***}		
	(0.49)	(0.39)	(0.54)	(0.44)	(0.46)	(0.37)		
Interacted: Beyond TP × None of neighbours with $l > l^*$							$+1.29^{*}$	$+1.59^{**}$
							(0.73)	(0.66)
Total tipping effect							-0.33	-1.24^{*}
							(1.18)	(0.69)

Sign. thr. : *p<0.1; **p<0.05; ***p<0.01

Note: The specification is still 1, now fully interacted with the indicated tract characteristic. In column 7 and 8, *neighbours* are the four closest IRISes (computed in terms of distance from centroid to centroid).

Consistently lower tipping thresholds in trade-competition-exposed zones

Another candidate hypothesis for the interpretation of such tipping reactions involves labor rather than identity issues.

The underlying argument has been developed several times in the qualitative literature [Beaud and Pialoux 2003]; shrinking manufacturing employment does not only cripple unionism and political organising at the firm level; it creates a more profound rift within working families; in the aftermath of major layoffs, there's a rise of social anomy within the families of lower-paid-lower-skilled workers; searchers notice growing concerns, among households of more qualified blue-collars, about the declining social atmosphere of working-class suburbs, the families with sufficient financial leverage even starting to leave the neighborhood, which becomes more and more segregated.

We test this hypothesis there with a 2SLS specification similar to model where the location of the tipping point is the *explanandum*, and the explanatory variable is an index of local exposure to the import competition from emerging economies, similar to [Autor, Dorn, and Hanson 2013]. Results of this estimation are reported in table 13; a rise in import exposure indeed seems to push the tipping threshold down.

Table 13: Simple model for the location of the *origin*-based tipping point

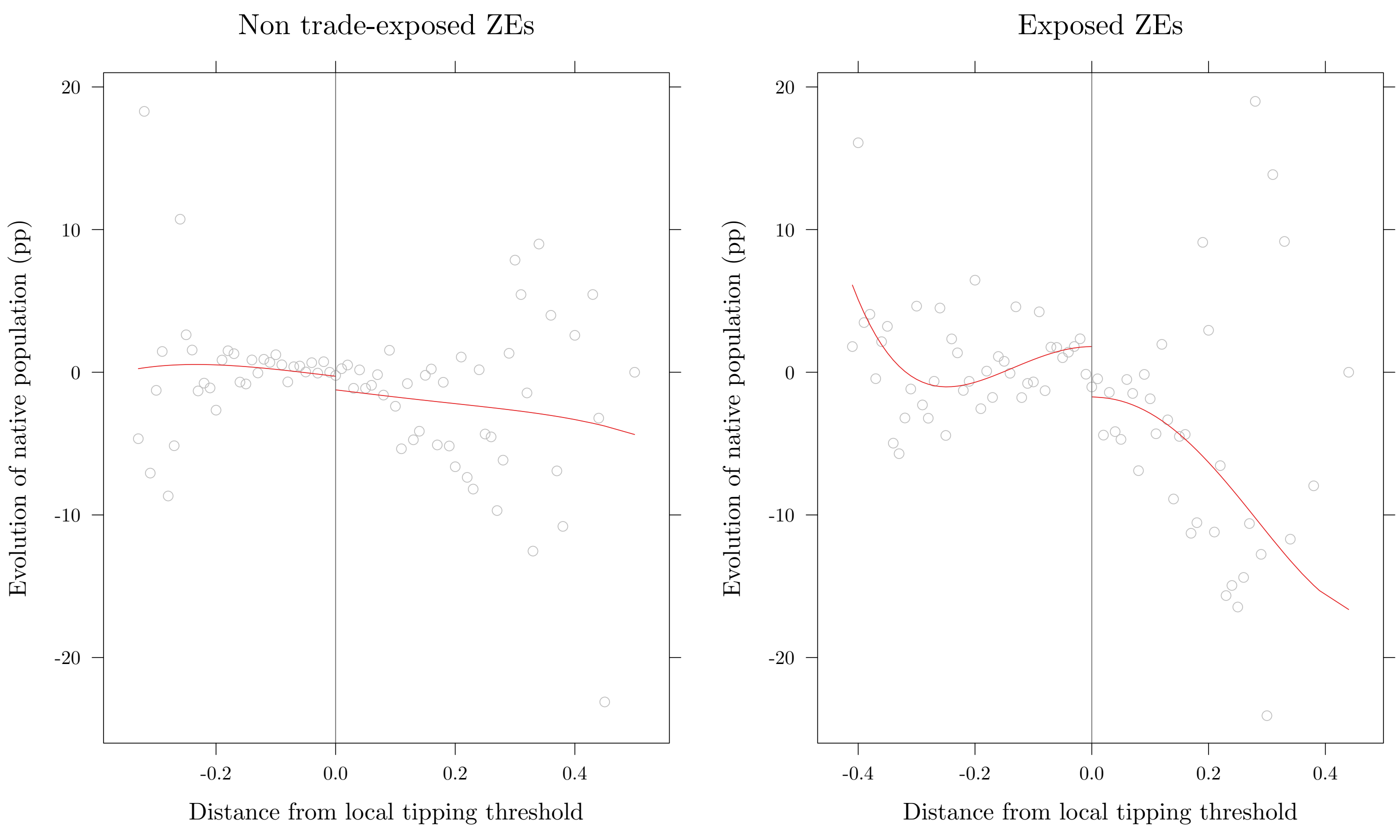
Dep. var.: Evolution of the tipping threshold (1999-2010 versus 2010-2017)				
	Origin-based T.P.			
	(1)	(2)	(3)	(4)
Rise in import compet. exposure				
Panel A. 1999-2008 exposure				
	-2.21^{**}	-4.36^{**}		
	(1.01)	(1.21)		
Panel B. 2008-2018 exposure				
			-1.69	-1.93
			(5.42)	(9.31)
Controls		X		X
Obs.	304	304	304	304
R^2	0.18	0.29	0.02	0.31
F-stat	3.1^{***}	3.9^{***}	2.6^{*}	3.9^{***}

Sign. thr. : *p<0.1; **p<0.05; ***p<0.01

Note: The unit of observation is the ZE (*Zone d'emploi*). The dependent variable is the difference between the tipping thresholds of each ZE for the second decade (2010-2017) and for the first decade (1999-2010) as estimated with the fixed point method, the estimation of which is reported in table 2. The main explanatory variable is the index ΔIPW described in [Autor, Dorn, and Hanson 2013]. Observations are weighted by the total population reported in the 1999 Census. Standard errors reported in parentheses are clustered at the level of the INSEE's superzone.

Not only is the tipping point lower in more exposed regions, but the tipping reaction itself seems more pronounced. If we distinguish the winning and losing ZEs from trade competition exposure in the sense of the intermediate scenario of [Fournel 2023], and reestimate our main tipping model over a newly drawn sample for the first decade (1999-2010), we get a nonsignificant d among winning ZEs (-0.03, $t = 0.07$) versus -4.11 ($t = 5.87$) for losing ZEs. Figure 3 plots the results of that exercise.

Figure 3: Tipping reaction around the tolerance threshold depending on the type of ZEs



Note: The unit of interest is the district (IRIS). Data are from the INSEE's Census. The tipping threshold specific to each *Zone d'emploi* estimated through the structural break method is displayed as a vertical grey line. x-axis gives the share of migrant population within each IRIS at the beginning of the decade ($l_{ic,t-10}$) minus the tipping point of the ZE to which it belongs. y-axis is the evolution of the native population as percentage of the base native population at the beginning of the decade, ($\Delta u_{ic,t}$). Dots give averages in 1-percentage-point bins. We plot in red a 3rd-order polynomial on the two subsamples right and left of zero. We estimated the main tipping model separately on the pool of ZEs deemed winners and losers of trade competition exposure as defined in section 3 herein above (intermediate scenario based on gains of trade estimates by [Borusyak and Jaravel 2021]). Observations are weighted by the start-of-the-decade total population. S.E. are clustered at the INSEE superzones' level.

That such tipping behaviours are correlated with industrial decline caused by exposure to import competition is a novel finding, though it is perfectly consistent with the political effects of trade competition exposure general found in recent literature [Dorn and Levell 2021].

Replication on the mobility series of the INSEE

One critical assumption of our identification strategy revolves around the evolution of the upper group share $\Delta u_{i,t-10,t}$:

- We assume that this evolution is driven exclusively by departures and arrivals in district i;
- Yet it might also be explained through the inner dynamics of the upper group. Maybe native population recedes because of its demographic dynamics; maybe local middle-class population recedes because it has been impoverished over the decade after a negative chock specific to the region.

Actually, the INSEE's Census provides one way to determine which effects predominates. The *Mobilité* series allow us to know, for a very large sample of the national population, in which *commune* one person was living at the beginning of a reference period, and in which *commune* that person lives at the end of that period. That reference interval has changed over time: the 1999 Census recorded mobility over 10 years (i.e. between 1989 and 1999) ; the 2006,2007,2008 Censuses, over 5 years, and the most recent series, since 2013, over one year (there was a break in the series between 2008 and 2013). These sets offer no information on individual incomes, but data about the origin of the person is provided since 2006 onward.

We therefore picked the 2006 issue and replicated the very same *origin*-based tipping identification strategy, with two special differences :

- 1. We are know working on a representative subsample of 19.8 million people (roughly one third of the national population);

- 2. The *supra*-unit is still the ZE, but the *infra*-unit is the *commune*[35]

Except for these two points, there are no other differences with the original strategy. $l_{ic,t}$ is still defined as the local share of the commune's population who is of foreign origin (not born on French territory). We estimate the tipping points with the very same algorithms, and the explanatory is still a dummy equal to one if $l_{ic,t}$ is above the ZE-specific tipping threshold. Our dependent is still the evolution of the native population (i.e. the number of arrivals, minus the number of departures of native individuals, as a share of the base population at the beginning of the reference period). Our reference interval is 2001-2006. Controls are not computed on the subsample of the mobility set; we matched each communal observation with the corresponding value of the control variable in the 1999 Census[36].

The outcome of that robustness exercise is displayed in table 14.

Table 14: *Origin*-based-tipping - Regression discontinuity model for change of native share around the tipping point

	Dependent var.: Change in native population in the district					
	Method 1 - Structural break			Method 2 - Fixed point		
	Base	F.E.	Full	Base	F.E.	Full
	(1)	(2)	(3)	(4)	(5)	(6)
2001-2006 semi-decade						
Beyond tipping point (coef. d)	-2.94^{***}	-3.04pp^{***}	-2.95pp^{***}	-3.62pp^{***}	-4.59pp^{***}	-4.48pp^{***}
SE	(0.44)	(0.48)	(0.48)	(0.43)	(0.49)	(0.49)
Observations	17378	17378	17378	17378	17378	17378
R^2	0.5%	4.3%	4.4%	0.6%	4.6%	4.7%
F-stat	18.1	2.51	2.56	21.1	2.71	2.76
ZE fixed effects		X	X		X	X
Controls			X			X
Average estimated T.P.		7.95%			11.03%	

Sign. thr. : *p<0.1; **p<0.05; ***p<0.01

Note: The unit observation is the commune. The dependent variable is the growth of the native population (defined as these persons who are born on French territory) within the commune from t to $t-10$, as percentage of the base population at date $t-10$. The main explanatory variable is a dummy equal to one if the share of migrant population (those persons who are not born on French territory) is beyond the ZE specific estimated tipping threshold. All specifications also include a quartic polynomial in the deviation from the district's migrant share from the local tipping point, plus zone fixed effects. The vector of controls, drawn from the INSEE's Census, includes unemployment rate, share of working-class people, share of persons with no diploma, and share of vacant accommodation. Reported statistics are weighted by total population. Standard errors are clustered at the ZE level.

[35] Restriction rules are not relevant there since, even with our 12 infra-units minimum per supra-unit, we can still preserve all communes.

[36] It is a unsatisfactory strategy, but a second best in this context. The Mobility series record individual characteristics, not at the beginning, but at the end of the reference period, meaning if we took the control variables from that set, we would have the values of 2006, not of 2001. Besides, some control variables, like the share of vacant accommodation, cannot be extracted from the set.