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Native Population Turnover & Emerging Segregation: The Role of Amenities, Crime and Housing

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Native Population Turnover & Emerging Segregation: The Role of Amenities, Crime and Housing

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Abstract: Using geo-coded full-population grid-level data for the three largest metropolitan areas in Sweden, 1993-2016, this paper *i)* estimates the level and pace of ethnic segregation, *ii)* examines possible tipping points in this development, and *iii)* gauges the importance of several mitigating or exacerbating factors (such as the mix of housing area tenure type, different types of amenities, and crime). We use OLS and 2SLS to estimate outcomes at two different geographic levels; 250 x 250 square meter grids and SAMS areas (roughly equivalent to US census tracts), respectively. On average, we find that for every 1 percentage point increase in immigration, native growth is reduced by around -0.3 percentage points. Crime levels exacerbate developments and factors such as housing area tenure-type mix and access to various amenities slows it down, but only marginally so. Using repeated and single random sampling for cross-validation, and the twin common methodological approaches as suggested in the literature, we estimate possible tipping points in these segregation developments. In contrast to most other studies in the literature, none of our potential tipping points are however statistically significant when probing their relevance in explaining factual population developments, suggesting a rather more continuous – albeit steeply so – segregation process rather than a structural brake. In terms of tipping point methodology, our main findings are that fixed-point estimation is less robust than R-square maximization for small geographical units, and that the former consistently selects for lower tipping-point candidates than the latter.

JEL-codes: C26, J15, R23.

Keywords: Tipping point, ethnic enclaves, segregation, urban amenities, geo-coded data, micro-data

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1. Introduction

A large body of literature has documented general white flight and changes of demographic profiles at the neighborhood level in the US (see Ellis, Wright, Fiorio, & Holloway, 2018; Reibel & Regelson, 2011), and continuing positive net-migration inflows to the US and Europe has motivated substantial research efforts, both as to how immigrants arrange themselves in the cities to which they migrate (e.g. Bartel, 1989; Borjas, 1999; Catney, 2016; Åslund, 2005) and the persistence and consequences of ethnic enclaves (e.g. Bayer, McMillan, & Rueben, 2004; Cutler, Glaeser, & Vigdor, 1999; Edin, Fredriksson, & Åslund, 2003; Patacchini & Zenou, 2012; Saiz & Wachter, 2011, among many others).

Building on Schelling (1971) and earlier work testing its implications (for example Clark, 1992), a more contemporaneous literature also specifically addresses the *tipping point* phenomenon, i.e. whether neighborhoods follow a pattern in which at a threshold point of some proportion of new settlers, the outflow of native-born residents is such that the neighborhood "tips" from native born to a majority of new migrant settlers. An important paper by Card, Mas and Rosenstein (2008) identifies tipping points across US cities as occurring within a range of 5-20 percent neighborhood minority share. For Swedish data using similar methodology, Aldén, Hammarstedt & Neuman (2015) and Böhlmark and Willén (2020) also find tipping point thresholds of comparable magnitude for the residential share of non-European foreign-born. In contrast to these results, recent work – also using Swedish data – by Andersson, Berg & Dahlberg (2021) finds increasing ethnic segregation over time but with no clearly defined thresholds, and studies by Fernández-Huertas Moraga, Ferrer-i-Carbonell, & Saiz (2019) for Spain, and Ong (2017) for the Netherlands, have also come to very similar conclusions. For the US, Easterly (2009) also suggest a rather linear (i.e. non-accelerating) development over time.

However, two problems in these studies are *i)* a use of a low level of geographical resolution when calculating both residential ethnic segregation and the potentially confounding neighborhood level controls. For example, the most commonly used US and Swedish area-based delineators are census tracts and SAMS areas, respectively. While the latter of these two is more geographically detailed than the former, both encompass rather large local populations and potentially display substantial within-variation in terms of the share of ethnicities, local amenities as well as types of housing and tenure. Another problem *ii)* is lack of detail and to some extent relevance as regards these neighborhood level controls, an example of which is that – to the best of our knowledge – no study includes controls for a crucial variable such as local crime rates, or include any measure of the extent to which an area is regarded as safe by residents to reside in. This is an important methodological issue that has also been raised in the literature (in other words; is it ethnic background per se that drive segregation developments, or different types of social problems that the majority population view as associated with minority population areas?).[8] This lack of good quality neighborhood level controls also goes for data on different types of local amenities, a factor that can potentially also affect the pace and extent of ethnic segregation. We argue that these gaps in the literature may help explain previous inconclusive results.

The purpose of our paper is therefore to estimate the relationship between growth in the number of non-European immigrants and growth in the native-born population, both its pace and potential non-linear development, while addressing these two methodological concerns. Our empirical and theoretical contributions to the literature are as follows: Firstly, in terms of data, we use a higher geographical resolution than what has previously been used (250 x 250 square meter grids), controlling for information on crime as well as rich data set on local amenities (at grid level). Second, we employ an empirical design that allows us to assess the relative importance of the qualitative nature of the neighborhoods for native flight. In other words, we do not merely control for neighborhood level characteristics but rather study how the same level of exposure to minority inflow interacts with factors such as natural amenities, public transport, access to services and different types of housing (tenure mix). Thirdly, in terms of theoretical contributions, when testing for structural brakes (tipping points) in developments by way of the twin common estimators as suggested in Card et. al. (2008), we compare and probe

[8] For earlier US debates, see Farley et al. (1994) and Zubrinsky and Bobo (1996) for an example of the former view (that white flight and aversion is driven by racial prejudice), and Harris (1999, 2001) taking the latter standpoint (that associated poverty and social problems is the main driver).

the accuracy of using single random sampling methods (the go-to method in virtually all previous papers) with repeated sampling techniques.

Our empirical analysis suggests that the native-born outflow as related to residential inflow of non-western immigrants is rather substantial: Using OLS and instrumenting for the share of non-western foreign born, while controlling for municipality fixed effects and grid-level heterogeneity, we find on average that for every 1 percentage point increase in immigration, native growth is reduced by around -0.3 percentage point, over a 5-year period. Our results also suggest that crime and different types of amenities have significant independent effects but that these are not very large, suggesting that policies, which tangibly address these factors will also affect ethnic segregation, but that these efforts will most likely have to be substantial to affect outcomes.

As for structural brakes and non-linearities in developments, for our three metropolitan areas and two decadal time periods, we find candidate tipping points that range between 13-24 percent non-western immigrants (17 percent on average). This average estimate is similar to what was found in Card et. al. (2008) and is in line with one of the previous studies also using Swedish data (Böhlmark and Willén, 2020). However, in contrast to these outcomes, none of our tipping-point candidates significantly help explain factual population developments, which suggests a rather more continuous – albeit steeply so – segregation process rather than a structural brake. In terms of theory and tipping-point methodology, our main findings are that *i)* fixed-point estimation is less robust than R-square maximization using small geographical units, and *ii)* that the former methodology consistently selects for lower tipping point candidates than the latter, at higher levels of geographical resolution. Comparing different sampling techniques when searching and testing tipping-point candidates, *iii)* we also find that the use of repeated sampling when estimating tipping point location reveals a substantial heterogeneity, which in turn may reflect back upon the wide range of tipping point estimates previously found in the literature wherein single random sampling has been the norm.

Finally, a note on causality in these outcomes. A common problem in the segregation literature is the endogenous nature of the outcomes, i.e., the residential change that over time leads to ethnic segregation is a simultaneous process where we do not beforehand know the causal direction. In our paper, we employ two separate approaches to address this problem. Firstly, following Card, Mas and Rothstein (2008) and virtually all subsequent papers addressing this problem, we employ an IV-estimator where the predicted number of minority residents in a neighborhood is used as instrument for the minority share within each grid (the so-called shift-share instrument). Second, as a robustness test of these estimates, we use a (quasi) natural experiment research design and exploit the fact that Sweden in 2006 experienced a sudden and pronounced increase in Middle Eastern and African immigration (mainly from Iraq, Syria and Somalia). Since this increase in immigration could only to a limited extent have been predicted beforehand, we can use this increase as an exogenous event and estimate outcomes before and after this particular point in time.

In what follows, section 2 provides a further literature review and motivates the research. Section 3 details our specific contribution while sections 4 and 5 outline empirical design and data description, respectively. Section 5 provides the empirical analysis and in section 6 we summarize our findings.

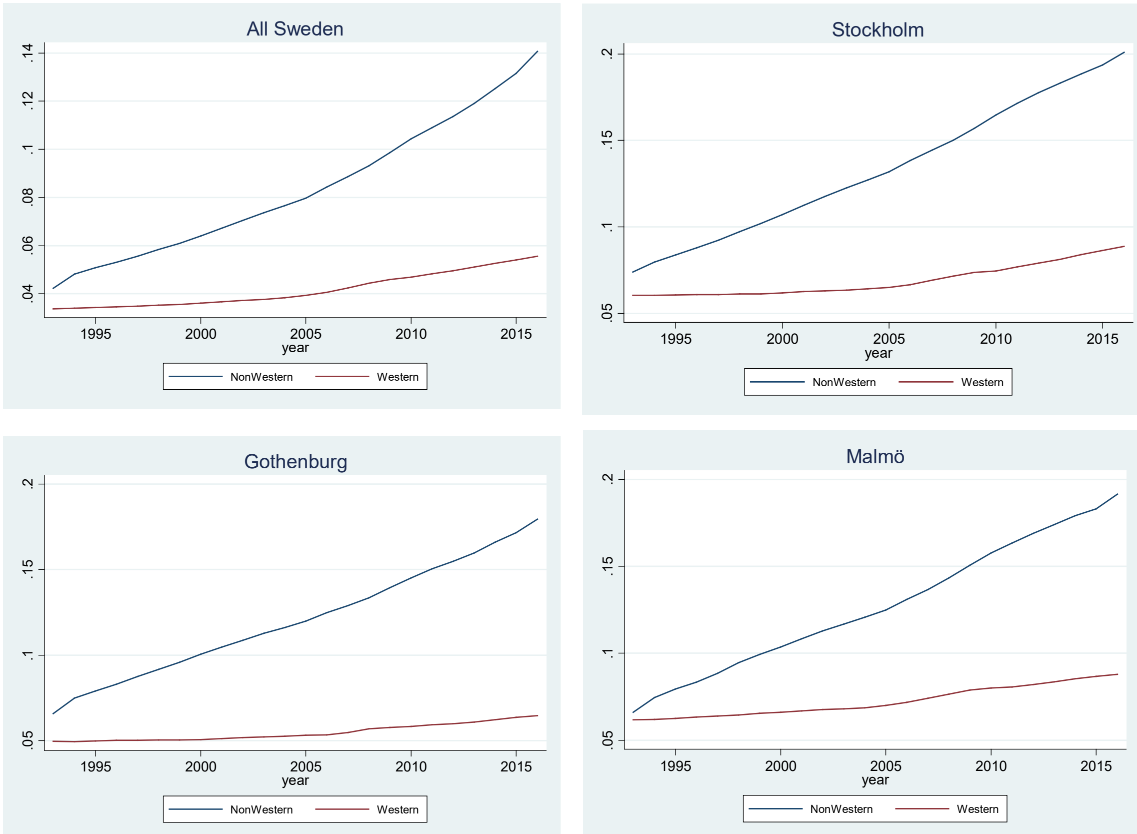
2. Background; what makes the Swedish case interesting?

During a historically short time span, Sweden has moved from being relatively homogenous in terms of the ethnic background of its population to being much more diverse. While around 95 percent of the population had a Swedish- or western-country background at the start of our studied time-period in 1993, this figure had dropped to 85 percent in 2016, with total population increasing by around 14 percent, or 1.25 million people. Currently, Sweden has around 19 percent foreign born (see Figure 1, panel a), which puts it in the top among western countries in the world, with the highest share of immigrants, both compared to the rest of Europe as well as the United States (Eurostat, 2020). As seen in Figure 1, these average numbers also display regional variation with relatively larger shares of first- and second-generation immigrants found in the major metropolitan regions (Figure 1, panel b, c & d).

Immigrant source countries have also varied greatly over time; whereas the 1970s and 1980s were characterized by immigration from Latin America and Iran and Iraq, the early 1990s saw large inflows from former Yugoslavia following the Balkan war, and major subsequent source countries have for example been Iraq, Somalia and Syria (SCB, 2020). Although not necessary for our theoretical framework, it is relevant as a background to note that in particular this latter development (i.e. immigration from Africa and the Middle East) has also increased both educational and cultural disparities between the native and immigrant population, and in terms of cultural disparities, much more so than – for example – as compared the past decades of immigration to the US.[9]

Figure 1. The share of western and non-western first- and second-generation immigrants. All of Sweden (a) and the larger metropolitan areas of Stockholm (b), Gothenburg (c) and Malmö (d), 1993-2016.

[9] This statement follows from results in the World Values Survey and the fact that the US in comparison has a much smaller share of immigrants from the Middle East and Africa. In a two-dimensional figure, with "traditional values vs. secular-rational values" on the y-axis, and "survival values vs. self-expression" on the x-axis, Sweden places itself in the top right-hand corner whereas most countries in the Middle East and Africa are placed bottom left (see e.g. Ingelhart & Welzel, 2015). The extent to which these differences persist among immigrants residing in Sweden is highlighted in Puranen (2019)

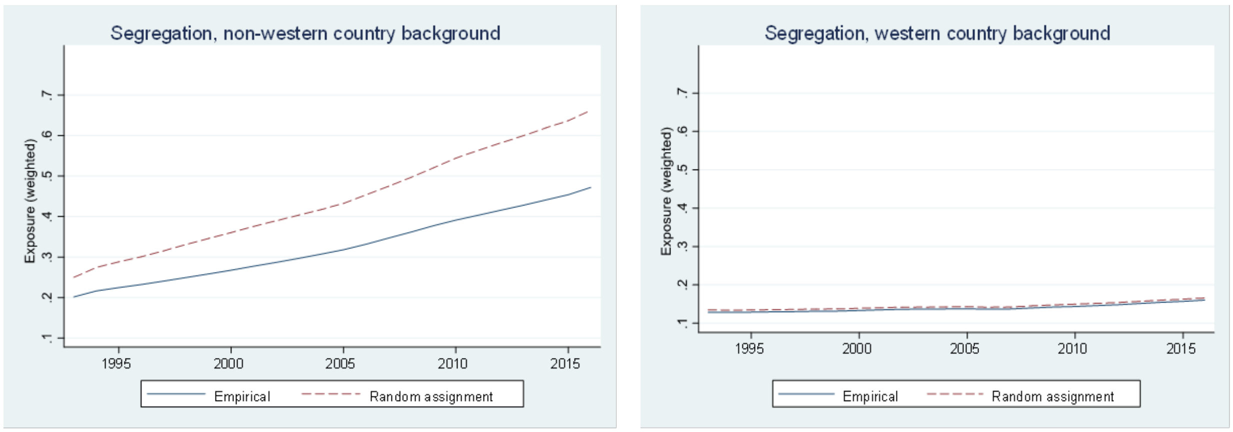


Note: "Non-western" defined as first- and second-generation immigrants born in either Asia, Middle-East, Africa or Eastern-Europe. "Western" defined as all other. Source; Statistics Sweden (Mona database).

The immigration inflow is also paralleled by increasing residential segregation, the dynamics of which are the focus of our study. In Figure 2 below, we follow Nordström-Skans & Åslund (2010), and Hedström et al (2017), and highlight the change in segregation over time as measured by the so-called exposure index, for the three largest metropolitan areas in Sweden.

The index estimates the likelihood of a native person (defined here as being Swedish born with at least one Swedish-born parent) of meeting someone in their residential neighborhood that is of foreign descent, defined as either a foreign-born person or someone with two foreign born parents. This *factual* likelihood is then compared with the equivalent *theoretical* likelihood had all first- and second-generation immigrants randomly sorted themselves into neighborhoods (that is, were they represented within all neighborhoods in equal proportion as to the metropolitan region at large). The resulting gap between these two curves over time can then be interpreted as a measure of the change in segregation.[10]

Figure 2. Development of segregation as captured by the exposure index, non-western and western country immigrants



In Figure 2, the lines show developments of segregation in the form of the exposure index from 1993 to 2016, for immigrants from non-western and western (non-Swedish) countries, respectively (panels *a* and *b*). The fully drawn line in both panels corresponds to the factual (empirical) development across grids in Stockholm, Malmö and Gothenburg, whereas the dashed line shows the theoretical exposure from a random allocation. Comparing the vertical distance between the two curves, we see that for immigrants of non-Western descent, this distance has increased from a fairly modest separation in 1993 to almost three times as large in 2016. For immigrants with western country background (panel *b*), the two curves have, however, developed in tandem, indicating an almost negligible change in the levels of segregation between these two groups.

[10] The segregation index *I* here refers to the probability weighted sum of the immigration share at the grid level, where the probability is given by the share of native residents in the grid and the native population at the metropolitan level. Specifically, $I_t = \sum_{i=1}^{n}\left[\left(\frac{X_{it}}{P_{it}}\right) * \left(\frac{N_{it}}{N_{mt}}\right)\right]$, where X_{it} and P_{it} is the number of non-natives and the total number of residents, respectively in the grid, whereas N_{it} and N_{mt} is the number of natives in the grid and in the metropolitan area in the year t.

3. Theoretical background and earlier studies

As noted in the introduction, the literature on ethnic segregation and demographic sorting in urban areas is vast and continually evolving. The principal methodology of much of the earlier research within the fields of geography, demography and sociology was more descriptive in nature, focusing on how whites in US cities over time left neighborhoods as blacks moved in (Farley, Fielding, & Krysan, 1997; Crowder 2000; Krysan & Bader, 2007; Lewis et al., 2011 Bader & Krysan, 2015), as well as how these changes were related to factors influencing people's residential choices beyond race and ethnicity, such as education, income and wealth.[11] While the tipping phenomena served a background to these studies, they did in general not engage with the specific mechanism or attempt to model the tipping process. At the heart of these analyses (most often based on survey and interview data) were instead the potential drivers of changes that we see in the residential mosaic, with much discussion focused on the relative role of preferences vis-a-vis discrimination and prejudice (for an overview of this literature, see Clark, 2007).

An important general conclusion from these research efforts is that ethnicity and race seem to play a separate although not overarching role when it comes to residential choice. For example, in Clark (1992), depending on race or ethnic background, at least 50% of respondents express choices of own race characteristics in the ethnicity of a presumptive neighborhood, but other types of factors such as affordability, access to schools and communications were deemed as equally important. Own race selection as related to residential movements has also been documented in more recent studies such as Fosset (2006a, 2006b) and Quillian (2015). A second noteworthy result, which relates to the question of the extent to which the earlier US segregation literature is relevant in a European context, is that there seemed to be a rank order of sorts as related to ethnic preferences and neighborhood residential choice. The most and least preferable neighbors for white residents are Anglo-Saxon (white) and African American, respectively, with Hispanic- and Asian neighbors placed somewhere in-between, a

[11] Within this literature there is a debate about the way in which neighborhood characteristics are defined, where these are often imputed by the very people that live there (Bader, 2004). Defining neighborhood characteristics only by way of demographic composition is problematic to the extent that such characteristics are determined endogenously by the very transition, or the change, that one tries to estimate.

preference order which to some extent also reflects the socio-economic positioning of these groups in US society (Denton and Massey, 1991; Clark, 1992; Zubrinsky and Bobo, 1996; Rosenbaum and Schill, 1999; Zubrinsky, 2000; Krysan, 2002). Some studies also suggest that the full mixed neighborhood is deemed as being the least desirable alternative (Clark, 1992; Pais, 2020).

In comparison to the US, European immigrant-dense neighborhoods are very diverse, indeed as we saw in the preceding section, the main segregation patterns that we find is that between the native (or residents of western descent) as opposed to residents from non-western countries (see also Hedström, 2019; le Grand & Szulkin, 2002). To our knowledge there have been no comparable studies of residential preferences in a European context, but to the extent that these US results are applicable, this would suggest that ethnic/cultural preferences as related to native residential choice are even more relevant here, as are then behavioral segregation models such as those developed by Schelling and Card.

In the remainder of this section, we will primarily focus on studies dealing directly with the tipping phenomenon as they pertain to our study. Following the seminal work of Thomas Schelling, these studies all assume and point to individual preferences and the resulting social distance dynamics as an important factor when trying to understand segregation developments over time (Clark and Fossett, 2008). Within much of other social science research in the field this theoretical and methodological stance is deemed controversial, or at the very least highly debatable (Fossett, 2006, p.187). Some prominent sociologists for example argue that the role of individual preferences in understanding segregation is not at all important, and rather highlight factors such as public policy and discrimination (see e.g. Massey & Denton, 1988; Yinger, 1995).

In this context it is however important to keep in mind that Schelling (1978, p.138) explicitly limited the scope of his work, stating that

> "..at least two main processes of segregation are outside this analysis. One is organized action-legal or illegal, coercive or merely exclusionary, subtle or flagrant, open or covert, kindly or malicious, moralistic or pragmatic. The other is the process, largely but not entirely economic,

by which the poor get separated from the rich, the less educated from the more educated, the unskilled from the skilled… in where they work and live and eat and play…"

We also adopt Schelling's caveat for our analysis; i.e., we do not claim to capture everything in terms of the dynamics of this process, but rather, we test for the implications of the framework while – as far as our data allows – controlling for other aspects that may affect segregation.[12] An apparent strength in Schelling's theoretical framework is that it offers a surprising yet logical explanation for an evident gap between results found in research on individual residential ethnic/cultural preferences, on the one hand, and the factual segregation seen in the urban landscape on the other. For example, a recent large survey by the Global Village, a Swedish non-profit working with integration issues, found that about 20 percent of the Swedish native-born population have a preference for non-Muslim neighbours (The Global Village, 2020), preferences which only to a very limited degree are mirrored in factual urban residential patterns.

This discrepancy between stated preferences and sorting-outcomes would have come to no surprise to Thomas Schelling, however. Indeed, his main intellectual endeavour in this context was to analyse the effect of own-group preferences in generating segregated neighborhoods, and to demonstrate how often well-intentioned individual preferences ("micromotives") may lead to outcomes in the aggregate which are highly inconsistent with these preferences ("macrobehaviour", see Schelling, 1969; Schelling, 1971, 1978).

His original work encompasses two basic models: The Spatial Proximity model and the Bounded Neighborhood Model, both of which are sometimes referred to as the "Schelling segregation model".

[12] The issue of ethnic- as opposed to economic sorting, and the relative importance of the two, is a complex issue which is still largely unsolved. For example, Card et. al. (2008), in their tipping analyses for US cities, find no concomitant structural brake in either average income or rents in neighborhoods with minority shares that over time lead to tipping into all minority neighborhoods, and argue that this speaks in favor of a strong separate role for majority preferences in tipping dynamics. For Sweden, Andersson et al (2020) find that second generation non-western immigrants display similar flight- and avoidance behavior as do natives, arguing that this speaks to the importance of income in segregation developments. In line with this reasoning, Malmberg and Clark (2020) find that income-based sorting to some extent counteracts the ethnic sorting taking place across Swedish neighborhoods with the result that complete segregation as predicted by Schelling type models will in fact not materialize. Still, the research cannot as yet provide an estimate of the relative contribution of income versus preferences.

In the first, we assume a society with two groups of people belonging to either a majority or a minority, both of which have preferences for a certain share of own-group individuals in their neighborhood (an average "tolerance level" of the other group). All resident locations are assumed to be occupied by either a majority- or minority person or left vacant. Further, individuals continually (in each consecutive time-period) evaluate their housing situation and make residential moves to maximize their utility based on these preferences, and a stable equilibrium is reached when all individuals have satisfied their preferences. In social interaction models such as this, each individual's choice in one period affects the available options and therefore the choices for other individuals in the consecutive time-period, i.e., one individual move generates a feedback effect that subsequently affects the choices of others. Starting with a level playing field and a high level of initial integration, the interesting dynamic outcome in this model is that random individual movements can over time result in amplified effects generating considerable clustering of each of the two groups, even when assuming so-called "weak minority preferences" – such as most individuals preferring not to be in the minority (a tolerance level up to 50 percent of the other group).[13]

In the second model, the behavioural assumptions are similar but instead of being average, individual tolerance levels (preferences) now vary from very low to very high. The reasoning in the model is centered on a mixed neighborhood that large shares of individuals from both groups prefer to reside in, given that their own-group preferences are satisfied, once an individual's other-group tolerance level is exceeded, he or she will move out of the neighborhood. The mixed neighborhood (staring point) is one of three possible equilibria. It is however an instable one since anything that causes the share of either group to change only slightly can trigger a chain reaction, with further movement of individuals whose tolerance levels are now exceeded which in turn results in further change to the neighborhood group-mix and subsequent out-movement. This chain of events that are hereby set off eventually results in a either full majority- or full minority

[13] This surprising and somewhat counter-intuitive theoretical outcome is dubbed the "paradox of weak minority preferences", the paradox being that the existence of extreme preferences, such as either tolerating no minority population whatsoever or being completely indifferent (tolerating all), will lead to an integrated outcome, whereas tolerant preferences, say 20-70% own-group preferences, will lead to segregated outcomes.

neighborhood, the two remaining possible equilibria in the model. These two segregated equilibria are, in contrast, deemed stable since individuals are assumed as preferring to not be the sole minority person in a neighborhood (i.e., regardless of ethnicity, which survey data suggests). Therefore, save for a large and coordinated movement by like-minded residents, no one will take the first step and move into the other group's neighborhood, even though many actually prefer living in somewhat mixed rather than fully segregated neighborhoods.

The main insight from the Spatial Proximity model is that mild preferences for like-neighbors may over time lead to segregation, and that the process can arise endogenously, or by individuals merely acting according to their preferences. The key insight from the Bounded Neighborhood model is that it suggest *i)* the phenomena of "tipping" and self-reinforcing developments that lead to full segregation outcomes once a threshold corresponding to the average other-group tolerance levels of those residing in a mixed neighborhood is exceeded, and *ii)* that mixed neighborhoods are likely to be unstable in terms of their population dynamics, and can be sensitive to external events that "trigger the movement of a system from one equilibrium to another" (Zhang, 2011, p.172).

Empirical estimates of tipping dynamics

Social interaction models such as Schelling's are often referred to as "two-sided" models, since the interaction that arises depends on preferences of both population groups and the outcome hinges on the initial distribution and the speed of the reaction of each respective group. On the basis of a formalisation of Schelling's work, Card, Mas and Rothstein (2008) examine tipping dynamics in the US major metropolitan areas using data on census tracts, 1970-2000, and estimate tipping points that vary from 5 to 20 percent minority population (with an average of 13 percent). Beyond these thresholds the majority population (white) start to leave, and developments subsequently evolve towards all minority neighborhoods. Comparing their findings against survey data on racial preferences (Cutler et al., 1999), the authors also find tipping estimates to correlate with the majority tolerance level of minorities, where for example San Diego and Memphis are located at opposite sides in terms of both the tolerance- and tipping estimates (highest and lowest, respectively). The estimated discontinuities in majority population development are however not found for either rents or housing prices around these thresholds, which the authors argue further points to preferences as an important factor driving developments (see also Becker & Murphy, 2009, for general social interaction models).[14]

Building on their 2008 paper, Card, Mas and Rothstein (2011) develop a "one-sided" tipping model where the tolerance level of the majority population matters first and foremost, and integrated mixed neighborhoods remain stable as long as the share of the minority within the neighborhood remains below a share that corresponds to the average tolerance level of the neighborhood majority population. In comparison to Schelling and their earlier work (2008), the stability of the outcome is further enhanced due to additional assumptions as regards amenity preferences of the majority population (e.g. natural or cultural amenities), which in turn makes them less sensitive in their preferences as regards the minority group. Using the same US census data as in their previous paper, the authors evaluate these two models in Card et al. (2011), finding that occurrences of neighborhood tipping is predominantly one-sided; that neighborhoods with minority shares below their estimated tipping points attract both white and minority residents, and that there are few indications that minorities tend to leave the neighborhoods where the minority shares are below the previously estimated tipping points. They also find that tipping-point thresholds on average have risen over time, particularly so as viewed in through the lens of a few select mid-western large cities where census tract data is available from the 1940s.

As mentioned, studies estimating tipping dynamics using Swedish data and similar methodologies as in Card et. al. (2008), include Aldén, Hammarstedt & Neuman (2015), Andersson, Berg and Dahlberg (2021) and Böhlmark and Willén (2020). These three

[14] We should note that there is a considerable debate revolving around the theoretical aspects of Schelling's work. For example, Bruch and Mare (2006) argue that high levels of segregation occur only when "...individuals' preferences follow a threshold function" (p. 667) which they argue is not an empirically plausible assumption. They instead develop a simulation model on the basis of linear preference alternatives to Schelling's threshold function and find that under such circumstances segregation outcomes largely disappears. However, these ideas are countered by the findings of Van de Rijt, Siegel, and Macy (2009) who show that Schelling rather understated the tendency to segregation which emerges regardless of whether a majority population tolerates diversity or even seeks diversity as long as actors are also sensitive to small changes in ethnic composition.

studies all find tipping dynamics to be an important feature in residential segregation developments but vary considerably in terms of the size of the estimates, and also – as we shall see – as regards important variable definitions and model specifications.

Aldén, Hammarstedt & Neuman (2015) estimate changes in native population growth as related to increasing immigration inflow to Sweden's 12 largest municipalities, 1990-2000 and 2000 to 2007. The study identfies tipping points in these developments using the structural brake method (as in Card et al, 2008a, see further discussion below), as occuring at around 9.5 and 3.5 percent share of European and non-European immigrants, respectively, with somewhat higher estimates for both immigrant categories in the latter time-period. The authors estimate both native inflow and outflow and point to outflow as the main cause of hightened residential segregation during both time-periods, rather than reduced native inflow ("avoidence") into segregated neighbohoods. Natives are definend in the study as being born in Sweden, with or without Swedish-born parents, while immigrants are categorized as being born inside and outside of Europe, respectivly.

Using neighborhood data for all of Sweden, Andersson, Berg and Dahlberg (2021) estimate separate native inflow- and outflow effects of refugee immigration from non-OECD countries, 1997-2010. By way of a fixed-effects estimator controlling for all neighborhood unobserved heterogeneity, outcomes are measured within one year of additional foreign immigration with estimates of long-term effects included as robustness tests. Natives are here either categorized as all persons born in Sweden, or as one of three categories depending on parents' country of birth; native-born with native-born parents or those born in Sweden to parents with either a western- or non-western immigration background. Further, models are estimated for renters and homeowners separately, where the authors find no significant displacement effects (outflow) for their full sample, but a relatively small effect using a sub-sample of homeowners. The estimated effect is however equally strong regardless of native category and on that basis the authors argue that displacement is likely driven by changes in the socio-economic character of the neighborhood, rather than ethnic preferences per se.

Although tipping dynamics are not a specific focus in their paper, outflow coefficients are estimated separately for initial immigrant shares of 4-25 percent, finding a discrete increase in native outflow at around 18 percent. Again, this finding is very similar across the different native population categories used in the paper. The confidence intervals for these different tipping estimates are however overlapping and need to be addressed in further research efforts to be fully conclusive.

Finally, Böhlmark and Willén (2020) assess individual level consequences of growing up in Swedish segregated neighborhoods, defining such neighborhoods as being either above or below their estimated tipping points as related to non-western immigration, 1990-2000. Using the "fixed-point" estimation method of Card, Mas and Rothstein (2008, see below) applied to neighborhoods within the three largest municipalities, they find a 9-14 percent drop in native population growth as occurring around 18 percent immigrant share (the same level found in Andersson, Berg and Dahlberg, 2021), his drop being almost exclusively driven by native avoidance of high non-western immigration areas. Natives are categorized as those born in Sweden to native born parents, plus first- and second-generation immigrants from western countries (select high- income OECD countries). Immigrants from non-western countries and natives born in Sweden with two non-western immigrant parents constitute their immigrant category.

In contrast to these outcomes, Ong (2017) and Fernández-Huertas Moraga et. al. (2019) both gauge tipping-dynamics with similar methods but finding little evidence in support of such discontinuous developments. Ong (2017) estimates the decadal neighborhood growth rate of the native Dutch (including Western minority) population developments in the three major metropolitan areas 1998-2008, finding a structural brake for only one of them, which however does not hold when adding other types of neighborhood level controls. The study does not rule out tipping points in earlier developments (much segregation was already present prior to 1998) but suggest that the result may be either due to lack of statistical power (the methods suggested in Card et al 2008 are highly data intensive, see discussion below) but also that the Dutch social housing- and area-based policies may have functioned as mitigating factors.

For Spain, using grid-based micro neighborhood definitions (500 x 500 meters) and detailed amenity data similar to what we use in our study, Fernández-Huertas Moraga et. al. (2019) estimate the change in native population as related to the large immigrant inflow to Barcelona and Madrid, 2000-2008. Calculating neighborhood population change between these two points in time, their estimates suggest that one native resident moves out for every three immigrants arriving from a developing country, an "outflow-inflow paramter" corresponding to about -0.3. However, based on conventional methods for estmating tipping points (a version of the "structural brake" method, as in Card et al 2008), they find no conclusive evidence for any type of such a break in these developments, cautioning that this result may also be due to lack of statistical power (very few neighborhoods had any significant number of developing country migration prior to the starting-year of their study). Intersting also, in a separate analysis of newly built housing areas, the native and immigrant population inflow was found to be very similar in magnitude. In other words, the two were positively related, leading to the emergence of intergrated neighborhoods.

Finally, Easterly (2009) argues that the assumptions underlying the interpreation of the Schelling model by Card et al (2008) are too restrictive in that they only take into account the necesseary local conditions of the model, but fails to specify the sufficient global properties, or the global dynamics of the model. In practice this implies paying more attention to where residents that leave a neighborhood actually end up living, something which he argues amounts to more of a general equilibrium approach to the dynamics implied by the Schellling model, rather than focusing on a partial equilibrium and the specific dynamics around the a possible tipping point. Using the same Census data and changes over three decades as in Card et al (2008), Easterly indeed finds that about 10 percent of the sample of neighborhoods move from being all white to all black, but that this neighborhood change did not follow the non-linear dynamics of tipping. Since he finds a certain degree of "white flight" happening in all types of neighborhoods, even those with high initial white shares, he argues that the data better fits a main story of continous white suburbanization over the studied time-period.

To summarize, although some studies point to tipping points in the development of ethnic segregation, the available literature is still somewhat inconclusive. On the basis of the methodologies suggested in Card et al (2008), estimates for the US and Sweden range between 13 percent neighborhood share of non-white population as an average tipping estimate (Card et al, 2008, corresponding to 5-20 percent depending on metropolitan region) to around 3-4 and 17-18 non-white for Sweden, outcomes essentially depending on categorization of immigrant population (as in Card et. al. 2008, Aldén et. al., 2015, Andersson et. al, 2021, and Böhlmark and Willén, 2020, where estimates found in the latter two studies are more in line with one another). However, by employing the same data as in Card et. al. but using a different methodological approach, Easterly (2009) finds no tipping whatsoever. This latter approach has been much less explored in the subsequent literature. For continental Europe, building on Card et al (2008), the two available studies do not show any convincing structural brakes in developments, although the authors in both these articles point to lack of statistical power as a potential problem (initial tipping estimates are not robust to adding different neighborhood-level controls). The need for additional probing of tipping dynamics is thus clear; both as regards factual tipping outcomes and to what extent estimates vary when applying different types of methods to single data sets.

4. Data and empirical methods

Aside from data on neighborhood crime levels and local amenities, all variables used in our study are based on full-population register data from Statistics Sweden (the so-called Mona database), containing detailed information on for example place of residence, sources of income, education, and country of birth of both individuals and their parents. Also, based on their latest registered address, all individuals are geocoded as related to a system in which all of Sweden is divided into 250 x 250 square meter grids. On the basis of these geocodes, we can hereby aggregate our individual level data so as to construct variables that capture change over time in the characteristics of the local population living within these grids.

It is important to note that our grids are consistently defined both across geographic space and the time frame, and that this consistency allows us to avoid some common methodological problems raised in the segregation literature. Chief of these problems arise from the fact that segregation measures are often based on administratively defined neighborhoods, such as census tracts in the US and SAMS-areas in the Swedish context (Small Areas for Market Statistics). If such units are defined differently across metropolitan regions, or if definitions are differently implemented, an equal dispersion of the population can result in completely different levels of segregation.[15] Perhaps more importantly, our use of fine-grained data as neighborhood definitions allows us to identify different types of population enclaves *within* administratively defined areas. Such empirical advantage is explored only in a handful of other papers, none of which relate to ethnic population dispersion or tipping behavior (see e.g. Andersson, Klaesson, & Larsson, 2016)

Our crime data are provided and compiled by the Swedish National Council for Crime Prevention (www.bra.se) and consist of all criminal convictions in Sweden during the time-period in focus. Criminal convictions include court decisions, summary sanction orders and prosecution waivers. Conviction data cover only the tip of the iceberg of criminal offending. However, from an international viewpoint, Swedish conviction data are considered to have a relatively broad coverage of offences (cf. von Hofer, 2014). The data gives us the three- and five-year average numbers of residents that are either charged with or convicted of crimes as related to violence and drug offences (so-called *brottsbalksbrott*), and crime deemed as less serious (such as petty theft). As our main crime variable, we chose the former category since convictions as related to the latter category (non-violent crime) is perceived as less important as regards the extent to which a neighborhood is viewed as safe to reside in. In qualitative studies, heightening levels of violence is also viewed as the main problem by those living in immigrant dense neighborhoods (see e.g. Esaiasson, 2019). We do however also test our models using the more encompassing crime category.

[15] These types of methodological problems are often discussed in the segregation literature as the "modifiable area unit problem", or MAUP. For a much useful summery and illustration, see Hennerdal and Nielsen (2017)

Finally, our data on amenities is provided to us by the Institute for Housing and Urban Research, Uppsala University, and contains grid-level information on distance to both private- and public amenities such as restaurants, bars and shopping areas as well as distance to parks, lakes, streams, and various forms of public transportation. Summary statistics of all variables included in the analysis are provided in the appendix, Table A1.

Empirical approach: Outflow-inflow analysis

In terms of estimation and statistical modelling, our analysis starts by addressing the so-called outflow-inflow relationship, i.e., while controlling for a range of covariates, we estimate the response in native population growth as related to changes in immigrant population growth. Growth is here captured by the change in natives and immigrants relative the initial population size for each given time-period expressed in percent. For this purpose, we consider for grid i at time $t = 2001, 2006, 2011, 2016$ the following model:

$$Dn_{i,t} = \alpha + \beta_1 Dim_{i,t}^{DG} + \beta_2 Dim_{i,t}^{DD} + x_{i,t-1}\gamma + z_i\theta + \epsilon_{i,t}, \qquad (1)$$

where $Dn_{i,t} = 100 \times (Native_{i,t} - Native_{i,t-1})/P_{i,t-1}$ is the 5-year growth in the native population, defined as individuals with at least one native-born parent, and where $P_{i,t-1}$ is the population size in the grid at the start of the period. We define our variable for immigrant from developing countries ($im_{\mathrm{it}}^{\mathrm{DG}}$) as individuals with two parents born in the Middle East, Africa, Latin America, Eastern Europe or Asia, and those from developed countries ($im_{\mathrm{it}}^{\mathrm{DD}}$) as all other foreign-born.[16] This categorization of immigrants into two rather broadly defined groups largely follows previous studies ￼￼ as well as from theory: In Schelling's account, the majority population is assumed to have preferences for "likeness", and therefore to a varying extent as reacting to any number of residents from visible minorities. In this context, second-generation immigrants with both parents

[16] For individuals with parents that have different ancestries, classification is made based on ancestry of the mother. If for some reason this information is missing, classification is instead based on the father. In case information on both parents' ancestral home region is missing, the individual is not included in any of the groups. However, this does not imply that these individuals are purged from the data, as they are still part of the population, it only means that they are not part of the numerator in the region specific population varibles.

from visible minorities will likely be sorted into this category as well, including those born in Sweden with both parents belonging to this group. However, to further gauge to what extent our results are dependent upon immigrant categorization, we also provide results of model (1) ￼ is made between the various sub-groups that make up our developing country ￼.

Returning to the model, $\mathbf{x}_{i,t-1}$ is a vector capturing lagged grid-level characteristics such as population size, average wage income, crime and the share of those with university level education. As for z_i, it denotes a vector of time-invariant variables capturing the Euclidian distance from each grid center to different types of amenities, such as public transportation, natural and cultural amenities, shopping areas and bars and restaurants (for a full list of controls and variable description, see Table A1 in the appendix).

Our model as outlined in equation (1) is akin to many previous papers in the segregation literature, but particularly to Fernández-Huertas Moraga, Ferrer-i-Carbonell and Saiz (2019), which also uses extensive geo-coded grid level registry data while incorporating the qualitative nature of neighborhoods into their analysis. An important difference with their study, however, is that, as our main approach we define Dn_{it} as as well as our other population variables as the relative change as in most of the previous literature, rather than as population differences over the period

Our data restrictions mirror those employed in Fernández-Huertas Moraga et. al. (2019). We drop the 1st and 99th percentile in the weighted population growth distribution for each consecutive time-period, where initial population size is used as weight. This implies that for our main analysis we remove the fastest growing and as well as the fastest decreasing grids which can otherwise skew average outflow-inflow estimates considerably (all in all comprising 2 percent of the total initial population).[17] For the remaining grids, the percentage growth rate is thereby confined to $-100\% <$

[17] Grids with zero residents in the initial year can arguably show infinite growth, yet their grow rate isn't defined which means that they are thus not affected by this procedure. Since our regressions are weighted by initial population size these grids however do not affect our outcomes substantially. As for grids lacking residents in the end of the analyzed time-period their percentage growth rate is -100. These types of cases make up most of the grids in bottom 1 percentile and are thus excluded from the analysis. In total, we exclude 8866 grid-year observation (2216 per 5-year period), which across our three metropolitan areas corresponds to 58 684 residents (2 percent of our population).

$100 \times \left(P_{i,t}/P_{i,t-1} - 1\right) \leq 134\%$, which implies a reduction mainly of the upper tail in the percentage growth-rate distribution.[18] These measures are in line with previous Swedish segregation studies based on higher level tracts (SAMS), which drop all observations showing population growth that extends 5 standard deviations above mean growth, and neighborhoods showing more than 500 percent overall native population growth.[19]

As highlighted by way of introduction, the process of neighborhood population turnover is an inherently endogenous process where identification is not immediately straightforward. To account for this problem, we follow much of the previous literature and use the so-called shift-share instrument, as first proposed and tested in Altonji & Card (1991, see also Card 2001). The basic idea behind this instrument is to use previous immigration patterns to predict the future immigrant outflow/inflow of those who share a similar immigration background as the current immigrant population residing in a certain neighborhood, and then to use this predicted inflow as instrument for the factual observed migrant inflow. Since immigrants (as do natives) often seek to reside among compatriots, or at least in proximity to people to some extent sharing their cultural background (as documented empirically in e.g. Kasy, 2015; Saiz & Wachter, 2011), this instrument hereby controls for the non-random sorting of the immigrants into neighborhoods.

Adopting the above reasoning, we follow Fernández-Huertas Moraga, Ferrer-i-Carbonell and Saiz (2019) and consider the following instrument for our immigrant growth variables. Focusing on the growth if developing country immigrants ($Dim_{\text{it}}^{\text{DG}}$), it is given by:

[18] We have also run the analysis by only excluding grids with percentage growth rates below the 0.5th and above the 99.5th percentile, with little differences in the main analysis. When including all grids in the analysis, however, our data surely violate basic OLS assumptions. Our data restrictions alter little by way of the qualitative interpretation of the basic analysis, except in the case of our IV-estimates which without restrictions suffer from much weaker identification.

[19] In addition, as a complement to our main analysis using (1), we run separate regressions using initially empty grids as well as grids with only very low shares of immigrants in the initial year. This also serves as a robust test of our main modelling approach and is motivated since significant segregation had already taken place at the start of our studied time-period, and simultaneously modelling potentially disparate developments in older and newly established neighborhoods can pose a significant challenge.

$$\widetilde{Dim}_{i,t}^{DG} = \frac{1}{P_{i,t-1}} \sum_{g=1}^{\#g} im_{i,k,m,t-1}^{g} \frac{IM_{i,m,t}^{g}}{IM_{i,m,t-1}^{g}} \qquad (2)$$

It corresponds to the predicted total number of developing country immigrants within each grid in time t, as a share of the lagged grid population size ($P_{i,t-1}$). It is given by the interaction of $im_{k,m,t-1}^{g}$, which represents the lagged number of immigrants from each group *g* of developing-country immigrants within each grid, and $IM_{m,t}^{g}/IM_{m,t-1}^{g}$, i.e., the change from the previous period of that particular subset of immigrants within each municipality. Each of these (predicted) sub-group totals are then summarized to arrive at the predicted total number within each grid, and divided by the lagged population size gives the predicted share. Put differently, on this basis we project the municipality growth rate of each ethnic group and use this projection to forecast the expansion of these immigrant groups within each neighborhood (grid) – an expansion that is unlikely to be correlated with the concurrent change in native residents.

Empirical approach: Tipping-point analysis

The outflow/inflow analysis assumes that the sorting taking place across neighborhoods happens in a more-or-less linear fashion. Since both theory and earlier research suggest the existence of non-linarites and tipping points (i.e. structural breaks) in such developments, we proceed in the analysis by relaxing the linear assumption. Specifically, we aim to test for the presence of tipping-point dynamics in the segregation process.

A tipping point in this context is simply defined as the share of visible minority within a neighborhood where, if exceeded, the native population growth rate is expected to decrease rapidly, either because of natives not moving in (avoidance) or because of natives moving out. In what follows, we go about finding these tipping points using the two basic methods as suggested in the literature: "R^2-maximization" method, as first employed by Card et al. (2008) and used in e.g. Aldén et. al. (2015) and Fernández-Huertas Moraga et al. (2019), and the so-called "fixed-point" method, also first developed in Card et al. (2008), and subsequently used in Ong (2017) and Böhlmark and Willén (2020).

Before we further delve into the intuition of these two methodologies, we should note that we henceforth follow Böhlmark and Willén (2020) and study the potential tipping point behavior among individuals from western countries as related to changes in the non-western population (i.e. in addition to native born individuals in our dependent variable we also include all individuals from our developed-country category). As before, non-western is here synonymous with our developing country variable (that is, having been born in a developing country, or having both parents that were born in a developing country). The reason is that an important outcome of our outflow-inflow analysis is that we find a strong positive correlation between the movements of native-born and immigrants from developed countries, and no indication whatsoever that their localization patterns are any different than those observed for the native born (a finding also in line with e.g. Hedström et. al. 2019). In a second departure from the previously used framework, we now also follow the tipping-point literature and analyze 10-year intervals instead of 5-year intervals used in our outflow-inflow analysis.

In the first of our two chosen approaches (the R^2-max method), potential tipping point "candidates" are searched for by way of running a series of repeated regressions estimating the neighborhood rate of western population growth solely as a dichotomous function of whether or not a neighborhood share of non-western immigrants at base year exceeds different candidate tipping points (ranging from 1 to 50 percent share non-western). Following these 50 trial regressions, we then select the candidate tipping point (share of non-western immigrants) that produces the highest singular R^2.

Formally, for $i = 1,\dots,N_m$ grids in metropolitan area m = Stockholm, Gothenburg, Malmö, where N_m gives the number of grids in metropolitan area m, let the percentage share of non-wester immigrants be given by $s_{i,t-1} = 100(P_{i,t-1} - west_{it-1})/P_{i,t-1}$. Then, for each metropolitan area and time-period (1996 and 2006), we repeatedly estimate the following model

$$Dwest_{i,m,t} = \alpha(k)_{m,t} + \beta_{m,t} d(k)_{i,m,t} + \epsilon_{i,m,k}, \quad (3)$$

over $k = 1,\dots,50$ where $d(k)_{imt}$ corresponds to an indicator variable defined by $d(k)_i = 0$ for all grids i if $s_{i,(t-1)} \leq k$ and $d(k)_i = 1$ for all grids i for which $s_{i,(t-1)} > k$. Thus, for a given k, $\beta_{(k)}$ gives the difference in mean western population growth ($Dwest_{i,m,t}$) between grids with an immigrant share lower than k and grids with an immigrant share higher than k. Following the literature, this stage of the analysis does not include any covariates.

As noted above, from the set of 50 trial regressions for each metropolitan area and time-period the candidate tipping point (henceforth referred to as $cTP_{i,t-1}$) is then inferred from the regression k producing the largest R^2, which we can write as $cTP_{i,t-1} = \arg max_k R^2(k)$. With two 10-year periods and three metropolitan areas, this hereby gives us a total of six candidate tipping points.[20]

Alternatively, a candidate tipping point can be located by comparing western growth rates in each neighborhood (grid) with the average western population growth rate in the larger city or metropolitan area (the so-called "fixed-point" analysis, as first suggested by Card et. al., 2008). For neighborhoods where western growth over the analyzed period is *lower* compared to the metropolitan average, the initial share of non-western residents is (on average) expected to be higher, and vice-versa (i.e., higher western population growth) for neighborhoods where the initial share of non-western residents is lower. With this type of population dynamic, a potential candidate tipping point can be inferred from the neighborhood share of non-western population for which the rate of western population growth coincides with the rate of western population growth at the metropolitan level (i.e., where the difference between the two rates is equal to zero).

In practical terms, and following the previous literature, finding potential tipping point candidates using this methodology is commonly done by way of a two-step analysis. Firstly, we regress the difference between the two growth rates on a non-linear function (a quadratic polynomial) of the share of non-western residents at base year. Each of the factors contained in this polynomial equation (see below) are included to capture the potential functional form of the relationship between the two population growth rates (e.g., to capture a steep drop or sharp increase in the difference around a certain share of non-western population). Secondly, after fitting the equation to the data, we locate the

[20] Instead of choosing the k (i.e. candidate non-western population share) that gives the highest R2, we have also tried using the Akaike and Bayesian information criteria as alternative, but with the same result.

candidate tipping points (i.e., the shares for which the difference between these two growth rates is zero) by calculating the roots to the estimated polynomial function.[21]

We use the following model specification,

$$Dwest_{i,m,t} - Dwest_{m,t} = \alpha + f(s_{i,m,t-1}) + \epsilon_{i,m,k}, \quad (4)$$

The first two elements of the equation capture the difference between the change in neighborhood and metropolitan western population growth, measured as a fraction of total neighborhood and metropolitan population, and f is a 4th degree polynomial of the non-western neighborhood share at the beginning of the time-period (i.e., $f(s_{i,j,t-1}) \equiv \sum_{p=1}^{4} \beta s_{i,m,t-1}^{p}$). To then locate the roots of the estimated polynomial (i.e. the tipping point candidates), we utilize the Mata routine ("polyroots()") in Stata. In case there are multiple roots, following the previous literature, we compute the analytical derivative of the estimated polynomial function evaluated at the roots, and choose for candidate tipping point the root for which the said (negative) derivative is the steepest. Once a candidate tipping point is located, we repeat the procedure in a second step restricting the data to only include 10 percentage points around the tipping point.[22]

As with the R-square method discussed above, we here use Stata's analytical weight function with initial population size, and since global polynomial methods are sensitive to outliers, we only include grids for which the initial share of non-western is below 50 percent, i.e., $s_{i,t-1} < 50$.

Testing the tipping-point candidates

In the final step of the analysis, our different tipping-point candidates are tested for potential explanative power using the inflow/outflow model delineated above but

[21] Note however, in a departure from the framework of our outflow/inflow analysis, we cannot include grids with initial zero population because the immigrant share is not defined for zero.

[22] To avoid that this second stage produces a less valid candidate, we require *(i)* that the number of remaining observations is at least 100, and *(ii)* that the p-value of the F-statistic is larger than 0.01. If any of these conditions fail or if there is no candidate (identified root) in the 0 to 50 range, we instead opt for the TP candidate located in the first step.

substituting the immigration growth variable with a dummy variable corresponding to the candidate tipping point. If statistically significant and large enough to correspond to a structural brake in western population developments, we can presume that factual tipping has occurred and that our dummy variable also captures something of factual importance.

For external validity, also following Card et. al. (2008), we use separate random subsets of the data for "testing" (i.e., searching for candidate tipping points as described in the previous section) and "replication" (gauging their explanative power). For the former part of the analysis, we draw a 60 percent share of the full sample for each metropolitan area and period), and the remaining 40 percent is used for replication.

For our remaining 40 percent of observations, we then generate a new dummy variable *d*(0) that corresponds to the shifted share of the initial number of non-western residents, $s^*_{i,t-1} = s_{i,t-1} - (cTP_{j,t-1})$, as estimated using either of our two methodologies (R^2 max or fixed-point estimation). The definition of *d*(0) is thus given by *d*(0) = 1 for all grids where $s^*_{i,t-1} > 0$, and *d*(0) = 0, for all grids where $s^*_{i,t-1} \leq 0$ (i.e., the dummy is equal to 1 for all grids with an initial non-western population share above our estimated TP, and zero otherwise). In the testing regression we also control for a 4th degree polynomial around s^*_i, defined by $\mathbf{P}_i = \sum_{p=1}^{4} (s^*_i)^p$. Our final model used for replication can thus be written as

$$\Delta Western_{it} = \alpha + \gamma d(0)_i + P_i \Gamma + X_{i,t}\beta + \epsilon_{i,t} \,, \tag{5}$$

which we estimate with robust standard errors clustered on the running variable s^*, rounded to the nearest integer. If there is a tipping point in the Schelling sense, we expect $\hat{\gamma}_j > 0$. If $\hat{\gamma}_j > 0$ while also controlling for non-linear deviations around s^*, this result implies that grids with an initial share of non-westerners higher than the candidate tipping point experienced a subsequent lower net inflow of westerners, as compared to grids for which the share of non-westerners was lower than the candidate tipping point. Moreover, by controlling for non-linear deviations, the difference in averages is due to a vertical shift (structural break) around the tipping point.[23]

To summarize, as regards both searching for and testing potential tipping points, our approach thus encompasses both the two methodologies suggested in Card et. al. (2008), which are also the main strategies employed in the subsequent literature. In addition, for completeness and methodological parsimony, we also use two sampling procedures for both tipping estimators; simple random sampling (SRS), following virtually all previous studies, and a repeated Monte Carlo sampling procedure (MC) where our tipping estimates and subsequent results represent the averages from 100 randomly drawn samples of our data.

Empirical approach: Amenities, crime & housing

To address the potential of role of neighborhood characteristics that go beyond general socio-economic covariates such as average income level and the educational level of an area, we specifically address amenities (public and private), crime level and different types of housing tenure. The purpose of this is straightforward; as much as possible, we are interested in estimating the extent to which ethnic segregation occurs while at the same time accounting for the role played by area characteristics that also affect residential location decisions but are not necessarily correlated with the ethnic mix of a neighborhood. We address these factors by, firstly, including them as control variables in our outflow/inflow model (1) as well as when testing our candidate tipping points in (5). Second, to further gauge potential effects that go beyond average estimates, we separately plot and analyze the interaction between these variables and the marginal effect of our developing-country population variable.

A note on the quality of our data. Our data on public- and private amenities represent averages for the years 2012-2016, and we thus project these values backwards to the start of our analyzed time-period. This is a weakness of our data, but since our broadly defined

[23] In this rather strict test, we follow the previous literature. We should however note that in theory a tipping-point does not have to correspond to such a sharp shift in developments (e.g., it can also merely be a steep drop in population development, i.e., a non-linearity). For completeness, in subsequent tests we therefore also test tipping developments without our controls for non-linearities.

housing area characteristics (such as access to shopping, libraries and areas with restaurants and bars) are not subject quick change, we argue that excluding this information when estimating our outcomes poses more of a risk (omitted variable bias) than the potential errors that arise through using the data. We should however keep these data limitations in mind when interpreting our results.

For the role of tenure type, our dummy variable captures neighborhoods that had mixed tenure all through the years 2012-2016 (i.e., the category excludes areas with either only public or only owner-occupied forms of housing). Since Swedish metropolitan areas have been subject to privatization of formerly rent-controlled housing during our studied time-period (see e.g., Andersson and Turner, 2014), this variable therefore contains two sources of measurement error.

Firstly, it excludes some neighborhoods that have moved from being characterized as rent-only housing areas to owner-occupied only. However, such areas are unlikely to constitute a large share of our grid sample: Far from all residents living in formerly rent-controlled areas subject to privatization had neither the stable income nor the finances required for the bank mortgage to purchase their formerly rented apartments. Most such apartment buildings therefore still have a certain number of residents on rent-controlled forms of tenure (and as such appear as mixed-tenure areas in our data). Secondly, our variable also includes some areas that were 100 percent rent controlled at the start of our studied time-period. We argue however that this does not constitute a big problem since these initially rent-only housing areas (subsequently subject to privatization) contain the very same type of mixture of higher- and lower income residents which the policies aiming to promote mixed-tenure residential areas factually aim to achieve (and the potential effect of which is what we estimate).

5. Results

5.1. Outflow-inflow analysis

Starting with our outflow-inflow estimates, Table 1 below shows coefficients of model no. (1) as estimated for each of our 5-year time-periods. The estimates show that neighborhood level changes in the native-born population as related to immigrants with western-country background (i.e., foreign born; developed country) is positive and closely related (at 0.43, column 1), and it is negative as related to residents with a background in developing countries (-0.21). Both estimates hold for adding our different controls throughout models 2-6, and each of the two estimates increase in size when adding instruments to account for endogeneity of the outcomes (column 6). Thus, when taking account for endogeneity, we end up with final population movement estimates that correspond to a positive ~0.90 percentage point change in residents of western descent for every 1 percentage point change in residents with native decent (i.e., people of western decent and the native-born move to similar residential areas). The corresponding estimate for those with developing country background is around -0.3, which corresponds to a change of around one native resident for every three individuals of non-western descent moving in.

As noted, these population change estimates also hold for taking account of socio-economic covariates as well as taking account of crime, tenure mix within the area and different types of local amenities. As for socio-economic factors, neighborhood average income seems to be of most importance (strongly positive) all through regression 1-5, whereas our controls for residential area education- and employment levels (share high education and share unemployed) are both smaller in coefficient size and in the former case barely statistically significant). On the whole, these socio-economic characteristics reduce the initial value of our developing country variable from a negative .21 to negative .14 (a 34 percent reduction), meaning that socio-economic area characteristics certainly affect these developments but are to no means the

Table 1. Estimates of the change in native born population as related to the inflow of those with background in either western or non-western countries, 1993-2016

VARIABLES	(1) OLS	(2) OLS	(3) OLS	(4) OLS	(5) OLS	(6) 2SLS
1. Foreign born: Developed country	0.55***	0.52***	0.51***	0.48***	0.48***	1.02***
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.15)
2. Foreign born: Developing country	-0.21***	-0.14***	-0.12***	-0.11***	-0.11***	-0.28***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)
3. Share high education		0.04***	0.03***	-0.01***	-0.01	-0.01**
		(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
4. Share unemployed		-0.14***	-0.09***	-0.16***	-0.15***	-0.05***
		(0.01)	(0.02)	(0.01)	(0.02)	(0.02)
5. Average income (log)		1.48***	1.07***	1.30***	1.08***	0.24
		(0.25)	(0.25)	(0.24)	(0.24)	(0.29)
6. Number of crimes per 1000 residents (Sams)			-2.62***	-1.74***	-1.76***	-0.88***
			(0.20)	(0.20)	(0.20)	(0.21)
7. Tenure mixed area				1.77***	2.02***	1.64***
				(0.11)	(0.11)	(0.13)
8. Resturant and bars (10km)					0.50***	0.43***
					(0.12)	(0.12)
9. Nearby water (0.2km)					0.35***	0.28***
					(0.07)	(0.07)
10. Nearby lake (1km)					0.20***	0.16***
					(0.06)	(0.06)
11. Nearby ocean (1km)					1.01***	0.90***
					(0.09)	(0.09)
12. Nearby park (0.3km)					-0.17***	-0.14***
					(0.05)	(0.05)
Constant	8.85***	2.41**	4.92***	4.13***	4.20***	4.56***
	(0.09)	(1.18)	(1.20)	(1.16)	(1.19)	(1.38)
Observations	77,231	77,231	77,231	77,231	77,231	77,231
R-squared	0.09	0.10	0.10	0.11	0.12	0.10
Instrumented variables	No	No	No	No	No	1.2
Kleibergen-Paap Wald F statistic						99.34

whole story.

As we would expect, our added crime variable (Column 3) is negative and statistically significant. It further reduces the estimate of our developing country variable, from negative .19 to negative .16 (around minus 15 percent), and also holds throughout for subsequently adding our remaining variables for area characteristics. Out of these added covariates, our tenure mixed area dummy (column 4) is significant and positively related to change in native population; only one of our private- and public amenity controls (column 5) ends up as significant – access to restaurants, dining and pubs – which also goes for a minority of our remaining variables related to natural amenities, the majority of which are positive in terms of coefficient sign. We should note however that these added covariates neither affect our segregation estimates nor explained variation (R-squares) to any significant degree.

To what extent does our model and coefficient estimates capture actual developments? Since significant segregation had already taken place at the start of our studied time-period there is the possibility that our approach of simultaneously modelling (possibly disparate) developments in older and newly established neighborhoods is potentially problematic. As a robust test of our main modelling approach above, Figure A1 in the appendix therefore shows results of running our outflow/inflow model (model no. 1) on both the full sample (panel 1) as well as a limited sample of grids which displayed only very low shares of developing country immigrants – below 4 percent – in the initial year of our analysis (panel 2). As seen in the figure, both panels depict a pattern of positive estimates of changes in developing country immigrants up to shares of 7-8 percent, after which they decrease and become negative. Since the estimates are largely similar in the two panels, this suggests that covariates included in model no. 1 sufficiently controls for prior segregation developments.[24]

[24] As mentioned by way of introduction, we can also address causality of our estimates by way of a quasi-natural experiment research design and exploit the fact that Sweden in 2006 experienced a sudden and pronounced increase in Middle Eastern and African immigration (mainly from Iraq, Syria and Somalia). Since this increase in immigration can only to a limited extent have been predicted beforehand, we can use this increase in immigration as an exogenous event and, as a robustness test, estimate outcomes during a few years following this particular event. In Table A3 in the appendix, we therefore include inflow/outflow estimates using model no. 1 for population changes 2006-2011. Coefficient estimates are here similar as viewed for the whole time-period, suggesting that endogeneity is not a big cause of concern.

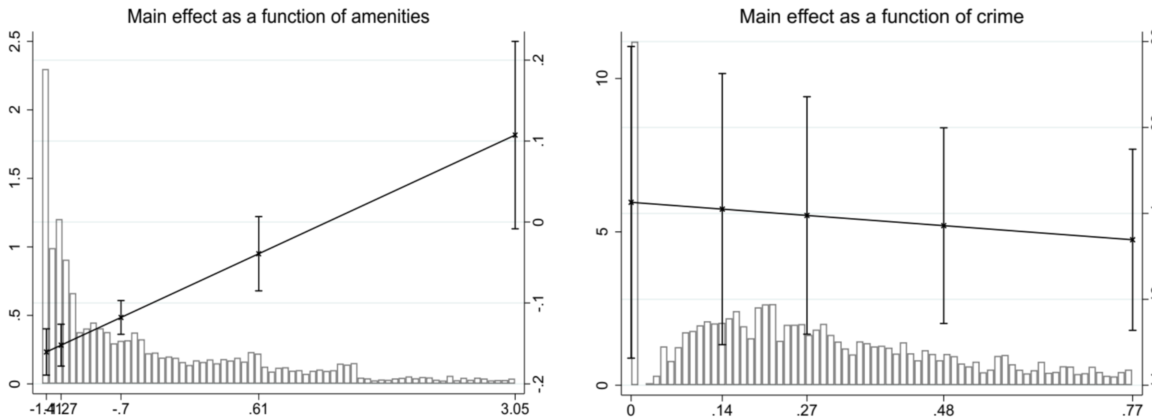
5.2. The role of public and private amenities, crime levels and housing.

In Figures 1a and 1b below, we provide an additional take on the results presented in Table 1 and specifically address the role of public- and private amenities, crime and housing (tenure type) in these developments. As in the main analysis above, we utilize model no. 1 but add variables for the potential interaction effects between either of these variables and the change in number of residents with a background in developing countries, our main variable of interest. Our focus here is to further gauge whether and to what extent each of these separate factors moderate the response of an inflow of immigrants from developing countries. I.e., to what extent the net change in the native-born population as related to developing country immigration is dependent upon these three factors.

Starting with public- and private amenities, to simplify the analysis, we utilize Principal Component Analysis (PCA). As with all PCA, the aim is to get a better grasp of the most important dimensions of the variation in the data, as well as to standardize the amenity variables (kernel-density estimates) and thereby make them easier to interpret in relation to the other covariates used in our main analysis. In Figure 1a below, the coefficient for the change in residents with a background in developing countries is negative for low values of principal component no. 1 but turns positive for those neighborhoods which are ranked higher in this regard. The results highlight that emerging segregation – as measured by our outflow-inflow estimates – happens at a slower pace or not at all in neighborhoods that higher in different types of private-, private- and natural amenities. We should note however that these high amenity neighborhoods are in the minority in our sample; most neighborhoods cluster around lower values of PCA factor no. 1 where the change in natives as related to an inflow of residents with a non-western background is still in the negative.

Figure 1b in turn shows the corresponding marginal effects of our developing country variable for different crime rates. Contrary to our amenity coefficient estimates, this analysis suggests that crime is less of an important factor than amenities in segregation developments; even for the very low crime rates where most of our neighborhoods are clustered, the coefficient for developing country residents is firmly in the negative, moving from only slightly above to slightly below -0.1, for neighborhoods with the lowest and highest crime rates, respectively. The interaction term is however not

Figure 1a & 1b. The marginal effects of the change in local population with a background in developing countries at different values of a) PCA factor no. 1., b) crime.



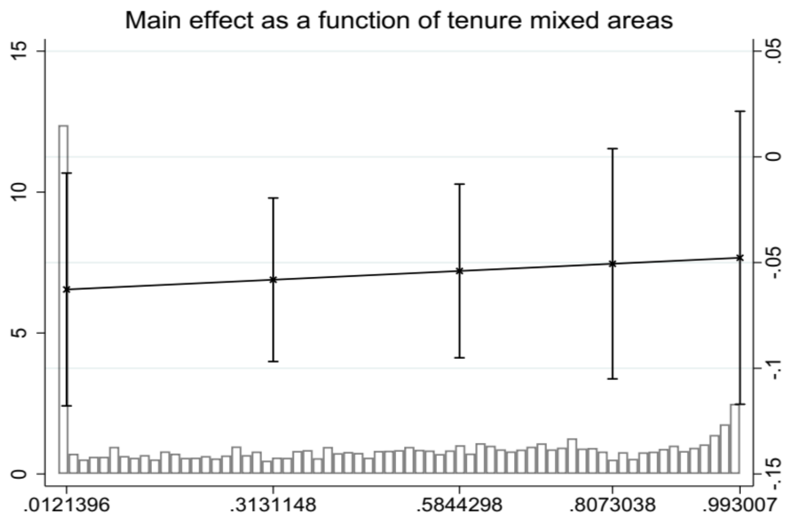
statistically significant.[25]

As with amenities and crime, in Figure 2 below we highlight the interaction between developing country immigration and housing tenure mix. The full drawn line leans slightly upward indicating that the larger the share of private housing, the more stable is the area in terms of segregation developments. As we can see on the right-hand side of the figure, the positive slope of the full drawn line is however mostly driven by the areas which are close to 100 percent private housing (but where standard errors around the estimate are large). Note also that the estimate is in the negative all through the increasing shares of private housing, indicating that areas with mixed tenure housing experience only marginally less segregation as compared to other areas. [26]

[25] Results of model no. 1 estimated with interaction effects are available from the authors upon request but cannot be included here.

[26] As an alternate analysis of the role of housing, Table A2 in the appendix shows the outcome when estimating model no. 1 only using mixed tenure type housing areas. As compared to using the full data set, coefficient estimates of both developed and developing country migrants are similar to results in Table 1. The effect of our developing country migrant variable is however weak in terms of statistical significance and is only significant using the full model and controlling for endogeneity (in Column 6). Note that we should however place less confidence in this estimate since its test statistic for the instrument (Kleibergen-Paap F-statistic) is much weaker than previously.

Figure 2. The marginal effect of the change in local population with a background in developing countries for different shares of private housing.



Source: Mona, Statistics Sweden

5.3. Estimating the candidate tipping point

For the sake of parsimony and as to give the reader a sense of the impact on the outcome that arises from using different types of estimators when gauging candidate tipping points, we employ two distinct methods to test for the tipping point location. First, R-square maximization (the most common type of estimator in the literature) by which the researcher estimates possible tipping points by way of regressing western population growth on a range of different initial shares of immigrant population (1 to 50 percent), and then chooses as candidate tipping point the one share that maximizes the R-square. Second, as outlined in the methodology section, the so-called fixed-point method where the candidate tipping is instead arrived at by way of comparing neighborhood and metropolitan western population growth, and as candidate tipping point choosing the one where the two growth rates are equal. In both these methodological approaches we use percent change in local population as basis for the analyses. For additional methodological parsimony, as mentioned, we also employ two sampling procedures for both estimators: simple random sampling (SRS) and repeated Monte Carlo sampling (MC).

In Table 2 below, we show tipping-point estimates separately for Stockholm, Gothenburg and Malmö and for two subsequent time-periods, 1993-2006 and 2006-2016. Using R2-maximisation we are able to locate candidate tipping points for all metropolitan areas and time-periods, ranging between 12 and 24 percent local population of developing country background. For both Stockholm and Gothenburg these estimates increase over time whereas for Malmö the estimate is lower for the second time-period as compared to the first. On average for all three cities, however, there is an increase: From 16 percent for the first time-period to 18.3 percent for the second.

In terms of variation in results depending on sampling methods, we find no strong discernable pattern in the variation, neither over time nor between cities. In other words, at this point in the analysis, it does not seem to be the case that the use of simple random sampling (SRS) methods results in outcomes that in any way are less consistent than using the average estimates from many different samples, as in the case when using Monte Carlo sampling. This is to us a somewhat surprising result. In Figure A2 in the appendix, we can however readily see that – in all likelihood – we should place more faith in the Monte Carlo estimates as opposed to those arrived at by way of SRS, a discussion which we return to below.

Turning to our tipping estimates using the fixed-point estimator, in contrast to R2-maximization, we fail to produce any valid results. The standard polynomial fixed-point method fits a 4th degree polynomial to the shifted grid-level percentage growth rate of non-western residents (i.e., it fits the polynomial to the difference between the metropolitan and grid level western population growth rates). At the grid level, the fixed-point estimator does however not produce any real roots anywhere in the 0 to 50 percent range of initial share of immigrants, regardless of city or time-period. This result has likely to do with there being too much variance in the data points when comparing grid- and metropolitan level growth rates. In other words, when it comes to using fixed point estimation, our conclusion is that our grid-level data is likely a too high geographical resolution. As we shall return to below, using larger local population areas as basis for the analysis – SAMS areas in our case – produces consistent results using both R-square and fixed-point estimation methods.

Table 2: Candidate tipping point estimates for Swedish metropolitan areas based on two types of estimation and sampling methods, simple random sampling (SRS) and Monte Carlo sampling (MC). Numbers represent local shares of developing country population.

Method:	R2 max:		Fixed point:	
Metropolitan Area \ Time-period	2006	2016	2006	2016
Stockholm (SRS)	13.00	14.00	.	.
Stockholm (MC)	12.26	18.84	.	.
(Sd)	(1.15)	(5.52)	.	.
Gothenburg (SRS)	15.00	23.00	.	.
Gothenburg (MC)	13.00	22.11	.	.
(Sd)	(2.05)	(2.00)	.	.
Malmö (SRS)	24.00	16.00	.	.
Malmö (MC)	18.92	15.81	.	.
(Sd)	(5.03)	(4.99)	.	.

NOTE: The table shows the results from separate tipping point estimates for Stockholm, Gothenburg and Malmö, for time-periods 1993-2006 and 2006-2016, respectively. The results are based on a 60 percent randomly drawn sample, using two sampling procedures; simple random sampling (SRS) (i.e. a single random sample) and Monte Carlo (MC), where the estimate represents the average from 100 randomly drawn samples. Parenthesis corresponds to the standard deviation in the Monte Carlo results.

5.4. Estimating the size and significance of the candidate tipping point

Table 3 below shows our coefficient estimates of the size in the drop of native-born population variable (western countries) that occurs around these shares of non-western residents, first using an empty model and then subsequently adding controls as per model no. 5. Since our "replication" data samples are smaller than our "testing" samples, to gain additional power, these estimates correspond to the average size of tipping points of all three regions and time-periods (the regression has thus been normalized to 0).

As seen in Table 3, when testing these candidate tipping points along with our other determinants of local population growth (using model no. 5) we find no significant result that lend support to any substantial role of these candidate tipping points in explaining subsequent local population developments once the share of non-western population surpasses these presumed tipping point shares. Using tipping point candidates based on simple random sampling (SRS) and Monte Carlo sampling (MC) gives us similar results in terms of the effect on population growth (negative in the magnitude of around 30 percent, see column 1, Table 3). Standard errors for both estimates are however high, indicating that the results are less reliable. Both estimates are significant at 90 percent level of confidence using our empty model, but T-statistics for the estimates quickly drop when adding other types of controls all through columns 2-4.

The results of this analysis can of course be due to us using a geographical resolution that is too high to be suitable for this type of analysis. In other words, when choosing when to move out of or into a residential area, each presumptive mover considers the ethnic background and character of a neighborhood area which is larger than the 250 x 250 meters that we use here, something which may have consequences for the validity of our analysis. To test this possibility, we conduct an additional analysis using larger SAMS-areas as the basic geographical unit, equivalent to that which has been used in the previous literature as concerns developments in Sweden. As seen in Tables A4 in the appendix, however, neither of these alternate analyses produces any significant tipping point estimates. Our conclusion is therefore that the process by which ethnic segregation occurs is mostly linear, albeit steeply so, and is not subject to structural brakes.

We should note however that our test for existence of tipping point developments follows the prevailing literature, and one can argue that the methodology as concerns these tests set a rather high bar for what we should view as a tipping point. In these tests we control for all non-linear developments and as we can readily see in figure 3 below, the R2 values from the regressions in many cases increase rater sharply beyond a certain share of non-western neighborhood population. A possibility is of course that this in fact is what a tipping point development actually looks like (i.e., a steep non-linearity) rather than the structural brake that the current methodology tests for.

Table 3: Estimating the size of the tipping point for all three metropolitan areas

R2 max: Growth	No controls (1)	Initial crime rate (2)	Social variables (3)	Full model (4)

Estimate (SRS)	-3.445 (-1.860)	-3.115 (-1.880)	-2.523 (-1.586)	-2.041 (-1.335)
Estimate (MC)	-3.323 (-1.767)	-2.759 (-1.521)	-2.345 (-1.290)	-1.876 (-1.000)

NOTE: T-statistics are in parenthesis. The table presents estimates of the size of the candidate tipping point as evaluated on a random sample that comprises 40 percent of the analyzed sample (i.e. consisting of data points that were not used to estimate the candidate tipping points). Estimates correspond to the average size of the tipping points of all three regions, which in the regression have been normalized to 0. The regressions are identical to that of model no. 1 but uses western population growth rate as dependent variable, controlling for a 4th degree polynomial of the difference between the estimated tipping point growth of non-western population and grid population growth. The models are also weighted using initial population size.

Finally, our estimates and tests of tipping points using different estimators and sampling techniques at the highly detailed grid level as well as the larger SAMS areas point to three noteworthy methodological findings.

Firstly, in terms of estimators, fixed-point estimation is less robust than R-square maximization when our unit of analysis is relatively small. As seen in Table A4 in the appendix, when we use the larger SAMS areas, finding the candidate tipping points and roots to our fourth-degree polynomial equation (i.e., the instances when the difference between metropolitan and local native population growth is zero) is not a problem.

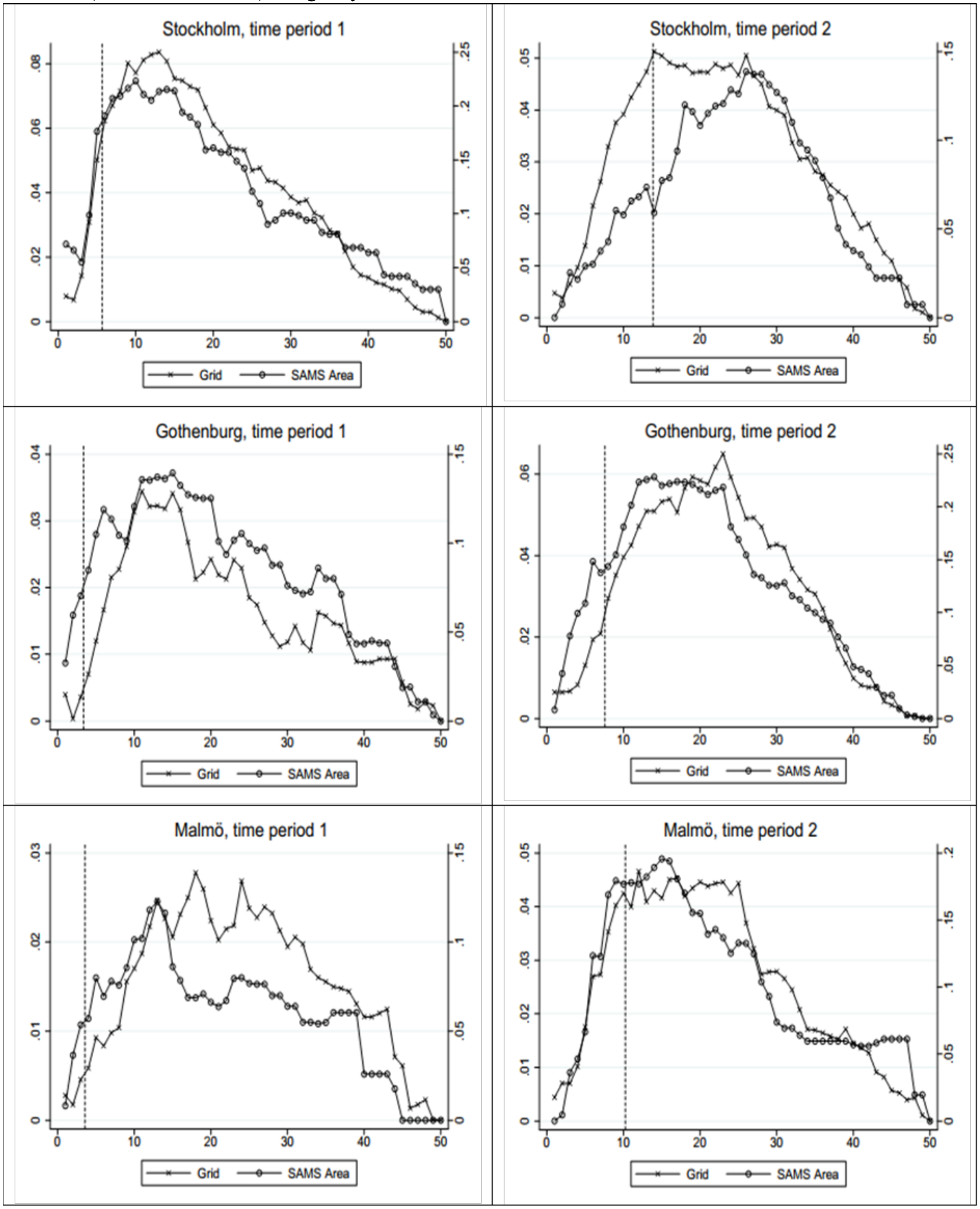
Second, as we highlight in Figure 3 below, when running the R-square max and fixed-point estimators on the same testing samples (each using 60 percent of all observations) at the SAMS level, we readily see that the tipping point candidates produced by these different estimators are for most part *not* equivalent to one another. Importantly, fixed-point estimation consistently selects for lower shares of non-western population as candidate tipping points than what is the case using R-square maximization (compare the doted vertical line – the fixed-point estimate – to the highest point on R-square curve, i.e., the R-square maximum). This is the case in all cases but one; Malmö, time-period 2, where they are close to equal (See Figure 3, panel 3).

Thirdly, our access to full population data allows for a comparison between the R-square maximization *parameter* on the one hand, and the estimate of that parameter arrived at using samples of the very same population data on the other. The full drawn line in Figure

A2 (see appendix) shows the amount of explained variation using local grid-level shares of non-western population as the only independent variable. The different bars on the x-axis shows the frequency distribution of the R-square max estimates arrived at using the 100 randomly selected Monte Carlo samples. The figure allows us to draw three tentative conclusions. *i)* in only two of our six studied cases most of the samples select for a tipping point corresponding to the "true" value (i.e., in Stockholm time-period 1 and 2). In two other cases they are somewhat close to that true estimate (Gothenburg and Malmö, time-period 1 and 2, respectively), and in the remaining two cases (Gothenburg and Malmö, time-period 2 and 1) they are wildly off mark. Further, *ii)* sampling as a way of choosing candidate tipping points is especially problematic when the factual R-square distribution displays either multiple local peaks or a plateau-like shape (as in the cases of Gothenburg and Malmö). Finally, iii) given the quite large spread in the distribution of these displayed 100 samples, and that a single SRS could correspond to any of them, using Monte Carlo sampling when gauging potential candidate tipping points is likely the most cautious and correct way to go.

Figure 3. R-square and fixed-point estimates of tipping points for Stockholm, Gothenburg and Malmö, time-periods 1 and 2 (1993-2006, 2006-2016). R-square tipping points estimated for both grid- and SAMS level data, fixed-point

estimates (dotted vertical line) using only SAMS data.



6. Conclusions and discussion

Our aim has been twofold when conducting this study. Firstly, we wanted to extend the current literature by estimating the pace in segregation outcomes while controlling for – and delving into in detail – area characteristics that are not commonly explored and used in the literature (such as crime levels, public and natural amenities, different types of housing). Second, we wanted to gauge the role of the geographical resolution level used when estimating outcomes, employing information at a more detailed grid level than what has – to our knowledge – hitherto been used. In terms of tipping point estimation and methodology, our aim has also been to fully exploit our comparatively rich panel data to test and compare both twin methodologies suggested in Card et. al. (2008), rather than choosing just one of them, as well as to extend the literature by introducing Monte Carlo sampling when searching for and testing candidate tipping points (as opposed to using single random sampling).

Our conclusions are as follows:

The outflow-inflow estimates highlighted in Table 1 indicate that the pace of residential segregation happens at a rate of around one native-born resident leaving a grid for every three immigrants from developing countries moving in. This pace or speed of development is similar to what has been estimated in a recently published paper for Spain which also exploits high resolution geographical data (in their case 500 by 500 square meter grids, see Fernandez et. al., 2019).

Somewhat surprisingly perhaps, this estimate is very robust to subsequently including a range of controls in the regression. In other words, adding residential grid level characteristics such as those mentioned above only marginally affects this outflow-inflow coefficient. This suggest that the social mechanisms and individual preferences underlying this sorting mechanism are strong, and in terms of consequences for policy, it suggests that measures aimed at slowing or alleviating segregation developments need to be very substantive in order to have an effect.

Our specific analysis of the interaction between our developing country variable and crime, a commensurate measure of amenities (PCA) and tenure type, as shown in Figure 1a-b and 2 all strengthen this conclusion. For example, to the extent that our crime variable accurately captures factual local developments, the effect on the rate of outflow-inflow is weak at best. Likewise, the coefficient estimate of change in the native born as related to an inflow of developing country immigrants is negative for most of the 0-100 range of increasing shares of privately owned housing, and only turns positive for grids where the share of privately owned housing is at about 85 percent. We should note however that the coefficient is less negative than for other housing areas, and that the dummy variable for tenure mixed housing areas is strongly positive and significant all through columns 4-7 in Table 1.

As for tipping points in these developments, on the basis of R2 maximation and fixed-point estimation we find candidate tipping points that range from 13 to 24 percent (for out three metropolitan areas and two time-periods) using the former but not the latter method (which likely does not allow for such high-level resolution data as we employ here). However, when testing these candidate tipping points using our ordinary population growth model (model no. 5) we find no significant result that lends support to any substantial role of these candidate tipping points in explaining subsequent local population developments once the share of non-western population surpasses these presumed tipping point shares. While the effect on the difference in westerners over the period is found to be negative in the magnitude of 3-4 percent, standard errors are high, and the results therefore statistically insignificant.

We also find that this outcome is not dependent on the level of geographical resolution. That is, when estimating and testing for tipping points at the level of larger SAMS areas this does not change this outcome. Our tentative conclusion is therefore that the process by which ethnic segregation occurs is mostly linear and is not subject to structural breaks (i.e. a tipping point development under current definitions). As discussed above, we should however note that our test for the existence of tipping point developments follows the literature, and that prevailing state-of-the-art methodology sets a rather high bar for what we should view as a tipping point (i.e., we control for all non-linear development around the tipping point share which is put to the test). As highlighted in Figure 3, R-square values when regressing native population changes on the share of non-western neighborhood population in many cases increase rater sharply beyond a certain point. A possibility is that this is what a tipping point development actually looks like, i.e., a steep non-linearity rather than the structural brake that the current methodology tests for.

Finally, our study points to three noteworthy methodological findings as related to tipping point methodology, as first employed by Card et. al. (2008).

Firstly, in terms of estimators, fixed-point estimation is less robust than R-square maximization when our unit of analysis is relatively small. As we highlight in Table A4, finding the candidate tipping points and roots to our fourth-degree polynomial equation is readily done when we use the larger and more commonly utilized SAMS areas as geographical unit of analysis.

Second, when running both the R-square max and fixed-point estimators on the same testing samples (each using 60 percent of all observations) at the SAMS level, we find readily see that the tipping point candidates produced by these different estimators are for most part *not* equivalent to one another. Importantly, fixed-point estimation consistently selects for lower shares of non-western population as candidate tipping points than what is the case using R-square maximization.

Thirdly, our access to full population data allows for a comparison between the R-square max *parameter* on the one hand, and the estimate of that parameter arrived at using samples of the very same population data on the other (see Figure A2, appendix 1). Comparing the amount of explained variation using local grid-level shares of non-western population as the only independent variable, and the frequency distribution of the R-square max estimates arrived at using the 100 randomly selected Monte Carlo samples allows us to draw three tentative final conclusions from our study:

i) The majority of samples select for a tipping point corresponding to the “true” value in only two of our six studied cases. In two additional cases they are somewhat close to that true estimate and in the remaining two cases they are wildly off mark (see Figure 3).

ii) Sampling as a way of choosing candidate tipping points is especially problematic when the factual R-square distribution displays either multiple local peaks or a plateau-like shape (as in the cases of Gothenburg and Malmö).

iii) Finally, given the quite large spread in the distribution of our 100 samples and that a single tipping point candidate as selected using simple random sampling could correspond to any of them (SRS – hitherto the most commonly employed approach in the literature), using Monte Carlo sampling when gauging potential candidate tipping points is likely the most cautious and correct way to go.

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Appendix 1.

Table A1. Definitions and summary statistics for the main neighborhood level variables included in the analysis (1993–2016).

Local population varibles	**Mean**	**Sd.**	**Min**	**Max**
Native born growth: Born in Sweden, w. at least 1 Swedish born parent	6.522	16.16	-71.11	121.7
Immigrant, developing = 1st or 2nd gen. immigrant (both parents) from either Africa, Middle East or Asia	2.710	6.857	-75.68	111.1
Immigrant, developed = all remaining 1st or 2nd gen. immigrants (both parents), e.g. Europe, North America, Australia	0.399	5.050	-50	63.64
Residential area characteristics				
Crime = No. individuals below age 25 with a minimum of 2 convictions (3 yr. average)	0.292	0.281	0	1.471
Average wage inc. = Average yearly income from employment (logged)	4.920	0.438	-0.916	6.716
Unemployment = share with at least one month/year of unemployment	6.090	5.541	0	54.55
Gridpopsize = Population size	.	.	.	.
University education = Share with three-year university-level education	34.63	16.30	0	100
Housing variables				
Tenure mixed area = Grids containing both privately owned and rent controlled housing (categorical)	0.299	0.458	0	1
Amenities				
Restaurants and bars = number of restaurants and bars within five kilometers (kernel density estimates)	0.619	1.115	0	6.269
Lakes & streams = Distance to water course (kernel density)	0.245	0.937	0	8.440
Distance commuter railways = Average distance to subway and commuter trains (distance kilometer)	0.295	1.058	0	9.752
Services = Services within 5 km (supermarkets, convenience stores, etc.) (kernel density)	0.433	1.132	0	10.40

Table A2. Estimates of the gg in native born population as related to inflow of foreign born, 1993-2916. Areas with mixed tenure type only.

VARIABLES	OLS	OLS	OLS	OLS	OLS	2SLS
1. Foreign born: Developed country	0.79***	0.75***	0.75***	0.75***	0.74***	1.11***
	(0.04)	(0.04)	(0.04)	(0.04)	(0.04)	(0.18)
2. Foreign born: Developing country	-0.14***	-0.09***	-0.07***	-0.07***	-0.06***	-0.21***
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)
3. Share high education		0.01**	0.00	0.00	-0.02***	-0.03***
		(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
4. Unemployment		-0.18***	-0.13***	-0.13***	-0.16***	-0.03
		(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
5. Average income (log)		0.97**	0.56	0.56	0.08	-1.46***
		(0.45)	(0.45)	(0.45)	(0.44)	(0.48)
6. Number of crimes per 1000 residents (Sams)			-2.68***	-2.68***	-2.26***	-1.13***
			(0.32)	(0.32)	(0.32)	(0.33)
8. Restaurants and bars (10km)					0.35**	0.28**
					(0.14)	(0.13)
12. Nearby water (0.2km)					0.26***	0.18*
					(0.10)	(0.10)
13. Nearby lake (1km)					0.06	0.02
					(0.07)	(0.08)
14. Nearby ocean (1km)					0.64***	0.43***
					(0.13)	(0.13)
15. Nearby park (0.3km)					-0.03	-0.03
					(0.06)	(0.06)
Constant	11.57***	8.71***	11.24***	11.24***	12.83***	16.20***
	(0.21)	(2.16)	(2.17)	(2.17)	(2.14)	(2.30)
Observations	23,130	23,130	23,130	23,130	23,130	23,130
R-squared	0.15	0.16	0.16	0.16	0.17	0.17
Instrumented variables						1.2
Kleibergen-Paap Wald F-statistic						81.98

NOTE: Robust standard errors in parenthesis, *** p<0.01, ** p<0.05, * p<0.1. Dependent variable is the 5-year net change in native born residents over the period 1993 to 2016 (where the first period is 5+3 years). All estimates are weighted by the grid's initial population size and corresponds to the restricted sample of 250 m. by 250 m. grids in Stockholm, Malmö, and Gothenburg metropolitan area. List of amenity controls not complete but includes all statistically significant parameter estimates.

Table A3. Estimates of the net change in native born population as related to inflow of foreign born, 2006-2010.

VARIABLES	OLS	OLS	OLS	OLS	OLS	2SLS
1. Foreign born: Developed country	0.49***	0.46***	0.46***	0.46***	0.45***	0.84
	(0.07)	(0.07)	(0.07)	(0.07)	(0.06)	(0.59)
2. Foreign born: Developing country	-0.28***	-0.23***	-0.22***	-0.22***	-0.20***	-0.39***
	0.49***	0.46***	0.46***	0.46***	0.45***	0.84
3. Share high education		-0.01	-0.01	-0.01	-0.02*	-0.02*
		(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
4. Unemployment		-0.25***	-0.21***	-0.21***	-0.24***	-0.14***
		(0.04)	(0.04)	(0.04)	(0.05)	(0.05)
5. Average income (log)		0.29	-0.01	-0.01	-0.40	-1.76*
		(0.90)	(0.91)	(0.91)	(0.89)	(0.93)
6. Number of crimes per 1000 residents (Sams)			-1.92***	-1.92***	-1.54**	-0.70
			(0.61)	(0.61)	(0.61)	(0.66)
8. Restaurants and bars (10km)					0.39**	0.34**
					(0.15)	(0.16)
12. Nearby water (0.2km)					0.20	0.17
					(0.18)	(0.17)
13. Nearby lake (1km)					0.46***	0.43***
					(0.14)	(0.15)
14. Nearby ocean (1km)					0.60***	0.44*
					(0.22)	(0.22)
15. Nearby park (0.3km)					0.10	0.10
					(0.11)	(0.11)
Constant					4.12***	4.14
					(0.18)	(4.40)
Observations	5,759	5,759	5,759	5,759	5,759	5,759
R-squared	0.05	0.06	0.06	0.06	0.07	0.07
Instrumented variables						1.2
Kleibergen-Paap Wald F-statistic						8.626

NOTE: Robust standard errors in parenthesis, *** p<0.01, ** p<0.05, * p<0.1. Dependent variable is the 5-year net change in native born residents over the period 2006 to 2010. All estimates are weighted by the grid's initial population size and corresponds to the restricted sample of 250 m. by 250 m. grids in Stockholm, Malmö, and Gothenburg metropolitan area. List of amenity controls not complete but includes all statistically significant parameter estimates.

Table A4: Estimating the size of the tipping point for all three metropolitan regions using SAMS areas

	No controls (1)	Initial crime rate (2)	Social variables (3)	Full model (4)
R2 max: Growth				
Estimate (SRS)	196.16 (1.18)	196.1 (1.18)	265.05 1.59	256.92 (1.61)
Estimate (MC)	10.95 (0.10)	11.20 (0.10)	-11.22 (-0.12)	-5.75 (-0.06)
Fixed-point:				
Estimate (SRC)	-20.34 (-0.25)	-21.31 (-0.26)	-3.33 (-0.04)	-11.81 (-0.18)
Estimate (MC)	-62.81 (-0.83)	-63.23 (-0.84)	-35.36 (-0.51)	-37.18 (-0.56)

NOTE: T-statistics are in parenthesis. The table presents estimates of the size of the candidate tipping point as evaluated on a random sample that comprises 40 percent of the analyzed sample (i.e. consisting of data points that were not used to estimate the candidate tipping points). Estimates correspond to the average size of the tipping points of all three regions, which in the regression have been normalized to 0. The regressions are identical to that of Table 3 that uses the difference in, here, western residents controlling for a 4th degree polynomial of the initial population size. Additionally, the models also include a forth degree polynomial of the non-western share. Since All models are weighted using the initial population size. The table includes estimates for three different methods, by which the candidate tipping points have been estimated: (i) R2 max: growth, which gives the candidate tipping i.e. initial non-western share that maximize the R2 from regressing western percentage growth on tipping points of 1…50 percent of non-western population. (ii) Fixed point instead gives the root to a polynomial that corresponds to the non-western share for which the percent change in western residents is equal to the change in western residents at the metropolitan level.

Figure A1. Estimates of the change in native population and developing country immigration using the full sample (panel 1) and only grids with at most 4 percent developing country immigrants in the initial year, 1993 (panel no. 2). Estimated using model no. 1.

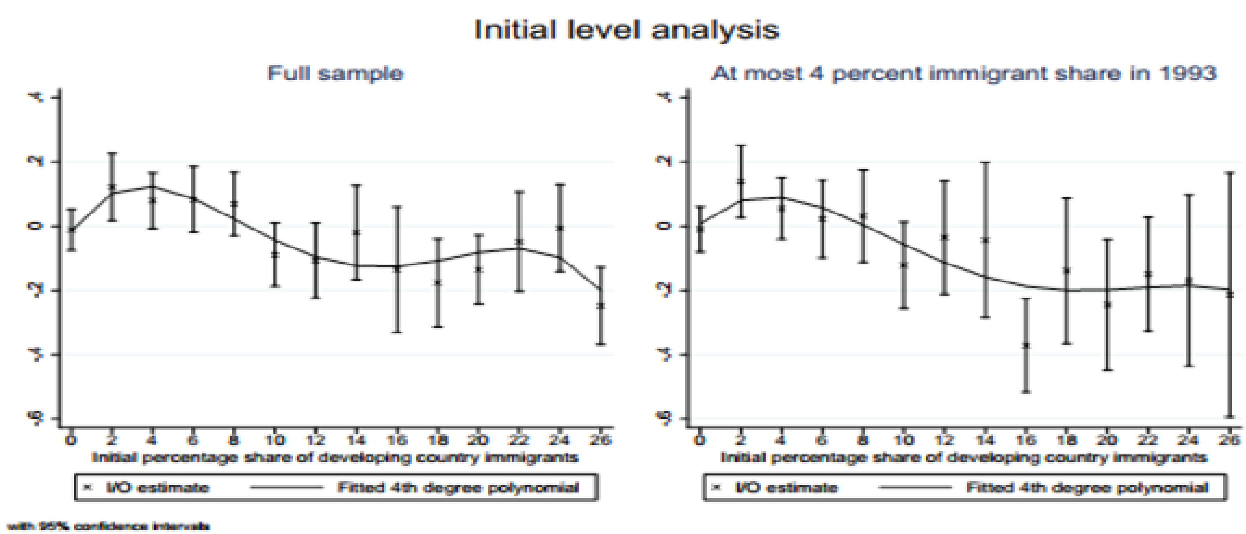

Figure A2. Tipping point estimates using R2 maximization methods, full population estimates and Monte Carlo sampling distribution for Stockholm, Gothenburg & Malmö.