UNU-MERIT Working Paper Series

#2014-008

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UNU-MERIT Working Papers

ISSN 1871-9872

Maastricht Economic and social Research Institute on Innovation and Technology, UNU-MERIT

Maastricht Graduate School of Governance MGSoG

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School choice, segregation, and forced school closure[1]

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February 2014

Abstract

We exploit the forced closure of three segregated primary schools in Amsterdam to establish the determinants of school choice of ethnic minority pupils. The schools were closed due to mismanagement and poor assessment from the Education Inspectorate. Most of the affected students were of socially disadvantaged and non-western migrant background. Our analysis contrasts the respective school choice decisions of the 'early movers' who had voluntarily changed schools within two years before the forced closure and the 'forced movers' who had to move to other schools after the closure. Using a conditional logit model and a nested logit framework, we find that: (i) students of segregated schools tend to re-concentrate into the same schools rather than disperse into different schools; (ii) primary school choice is nested upon school type; and (iii) the 'forced movers' prefer schools with more peers of own (non-western and low socioeconomic) background, less peer truancy, and shorter residence-to-school distance.

Keywords: School choice; Ethnic segregation; School closure; School mobility, Nested logit

JEL-classification: I20, I28, R28

[1] We would like to thank Henriëtte Maassen van den Brink for the very useful comments on an earlier version of this paper and the Municipality of Amsterdam for providing the dataset.

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1. Introduction

How do students select themselves into schools? While this question appears trivial and answerable via a simple analysis of the observed school choices in a system of free school choice, it is not due to three reasons. First, the observed, aggregated student characteristics of the chosen schools do not reveal parental school choice but rather the school composition. While the latter influences and is influenced by school choice, it is not equivalent. School composition here is a static concept, while school choice is dynamic, subject to student mobility or transitions during the course of education (Cameron and Heckman 2001; Declerq and Verboven 2013). Second, focusing on the school choices of new students (e.g. students enrolling in the first grade) is not representative. Even in a system of free school choice, most schools have an admission policy that allocates the potentially scarce places, e.g. based on proximity to residence, sibling's enrolment, religion, and lottery. This makes it less about school choice but more about matching principles (c.f. the student assignment mechanism literature by Abdulkadiroğlu and Sönmez 2003; Hastings, Kane, and Staiger 2005). Third, students do not choose schools randomly. They follow an implicit nested decision-making structure in which the school denomination is preferred above other observed school characteristics. Ignoring this nested structure would result in biased (or at least, incomplete) evidence.

Exploiting a forced school closure setting, this paper is able to focus on school choice as exogenously determined by a policy intervention. In particular, it examines the school choice of students in three Amsterdam primary schools that were forced to close in July 2007. The three schools had operated under the same school board, the Foundation for Islamic Primary Schools in Amsterdam (*Stichting Islamitische Basisscholen Amsterdam*, SIBA). The primary reason for the closure was the consecutive and deemed to be irreversible poor performance assessments by the Education Inspectorate on the schools' finances, management, and educational attainment of students (De Witte and Van Klaveren 2012; Dijksma 2007). The policy intervention was sudden with most of the affected students de-enrolling from the schools immediately after the closure and not before. School closures are very rare in the Netherlands and in this case, it is the indirect result of the discontinued public funding without which the schools were no longer financially viable. Consequently, our study augments the growing research field on school closures and its effects (De Witte and Van Klaveren 2012; Egelund and Laustsen 2006; Engberg et al. 2012).

Through school enrolment records, we also extract a sample of 'early movers' who had changed schools within two years prior to the forced closure to contrast their choice determinants with that of the forced movers. The former are students who have self-selected themselves into the three schools which were later forced to closed but have exhibited potential 'Tiebout mobility', i.e. by leaving the weakly assessed schools early for a presumably better school (Tiebout 1956; Hanushek, Kain, and Rivkin 2004; Allen, Burgess, and Key 2010; Brunner and Imazeki 2008). Our observations comprise of interschool mobility within the municipality so it is reasonable to assume that the school changes for these voluntary movers were more due to choice than to circumstance (e.g. a move out of the city). Through the two group comparison, we hope to isolate the actual determinants of school choice given an external shock such as forced school closure. The paper is novel in the sense that it studies the school choices of involuntary movers which is expected to differ significantly from voluntary school choice.

Since the three primary schools are Islamic schools with predominantly students of Moroccan and Turkish background, our paper contributes to the under-researched school choice literature for students of ethnic minority or migrant background. The increasing ethnic diversification of European cities due to immigration has led to substantial research interest in school choice and segregation (Allen 2007; Burgess, Wilson, and Lupton 2005; Karsten et al. 2006; Rangvid 2007; Cantillon 2009; Denessen, Driessena, and Sleegers 2005; Söderström and Uusitalo 2010). Ethnically diverse, the city of Amsterdam makes an interesting case study for the determinants of primary school choice. First, with the Netherlands being a relatively new migrant-receiving country, there is significant difference in school choice between native Dutch students and those of 'non-western' migrant origin. The latter typically refers to the four largest 'non-western' ethnic groups – Aruban and Dutch Antillean, Turkish, Moroccan, and Surinamese – with ethnicity defined by the parents' country of birth. Next, without school catchment conditions and with per capita public funding for almost all schools, the Dutch school system approximates a universal voucher system (Friedman 1955; De Haan, Leuven, and Oosterbeek 2011) with clear parental choice-driven sorting into schools. Mediating the economic factor in school choice, non-socioeconomic school segregation has been sustained as parents choose according to other considerations such as religious denomination, educational philosophy, and student ethnic composition. The latter's salience in school choice has been exacerbated by secularisation and the growing population of inhabitants with a foreign background (*allochtonen*) since the 1960s.

More than half of the primary school-attending children in Amsterdam are of non-western origin – 55.8 per cent in 2007 – yet ethnic composition in schools reflects a far more segregated reality. According to the dissimilarity index score in *Table 1*, 56.6 per cent of the non-western minority students in 2007 will need to change schools in order to achieve ethnic evenness at the city-level. The dissimilarity index is measured as the cumulative mean deviation in each school between the number of non-western pupils, weighted by its city-level population, and the number of native Dutch and western pupils, weighted by its city-level population (Duncan and Duncan 1955). Likewise, the interactive or exposure index is the cumulative product between the probability of a student being of a native Dutch and western background in each school and the probability of a student being of non-western background in a city (Massey and Denton 1988). The isolation index as the converse of the exposure index is measured as the cumulative product between the probabilities of a student being of non-western background in the city and in each school. From both indices, we know that for the average non-western minority pupil in Amsterdam, 72.9 per cent of her school peers will be of non-western background while the remainder 27.1 per cent will be of native Dutch and western background. From a policy perspective, our paper's findings on disaggregated parental school preferences are useful for addressing aggregated issues such as school segregation.

Table 1: Ethnic segregation indices for Amsterdam primary schools

	2005	2006	2007	2008
Isolation Index	0.737	0.732	0.729	0.725
Exposure Index	0.263	0.268	0.271	0.275
Dissimilarity Index	0.574	0.571	0.566	0.560

Source: Municipality of Amsterdam (2005-2011) and CBS (2013). Authors' own calculation based on the Duncan and Duncan dissimilarity index (1955) and the isolation and exposure (or interaction) indices from Massey and Denton (1988).

We estimate a conditional logit model to exploit the large heterogeneity of primary schools in Amsterdam. It includes all school alternatives alongside their characteristics, including match-specific information such as residence-to-school distance. Then, we relax the conditional logit's independence of irrelevant alternative assumption by estimating a nested logit model so that the school-specific error terms within a 'nest' can be correlated with one another. This is crucial given the profile of our student sample that have all enrolled in an Islamic school – many of which still exhibited parental preference for staying within the Islamic denominational schools after the forced closure (ANP 2007a). Demand has always exceeded supply in Amsterdam as local studies have found the percentage of parents preferring an Islamic denomination school to be 2.5 times that of students enrolled in the school type (van Kessel 2000; van Kessel 2003; as quoted in Karsten et al. 2006)

As a summary, our paper's contributions to the existing literature on school choice are three-fold. First, we analyse the forced school closure and its effect on student mobility. Second, via the forced school closure and our second sample of 'early movers', we compare and contrast the school choices of those who had changed schools by choice and those who did so by circumstance, i.e. the forced closure. Lastly, we apply a nested logit model to analyse primary school choice as a decision nested upon school type or religious denomination.

In the following sections, we review the key literature on the determinants of school choice (Section 2) and the primary school system in the Netherlands (Section 3), before describing our data (Section 4), methodology (Section 5) and results (Section 6). A final section concludes with several key lessons and a brief policy discussion.

2. Literature review

The following section summarises three main strands of literature relevant to school choice analysis: (i) the primary determinants of school choice; (ii) the use of stated preference versus revealed preference choice data; and (iii) the various discrete choice methods applied to school choice analysis. For brevity of this paper and given our application on the Dutch school system, our review of school choice determinants is largely centred on the Netherlands.

One of the main determinants of primary school choice is school type or religious denomination (Lankford and Wyckoff 1992; Driessen and Merry 2006; Allen and West 2009). For example, it was the main factor behind the 1917 Dutch constitutional reform that finally led to public funding of private religious schools (Ritzen, van Dommelen, and De Vijlder 1997). Besides school type, proximity or distance between school and residence is also a robust predictor for school choice (Ruijs and Oosterbeek 2012; Long 2004; Kelchtermans and Verboven 2010a; Kelchtermans and Verboven 2010b; Hastings, Kane, and Staiger 2005).

Ethnicity matters (Ladd and Fiske 2009; Ladd, Fiske, and Ruijs 2009; Karsten et al. 2003; Karsten et al. 2006; Burgess, Wilson, and Lupton 2005; Clotfelter 1999) – much more than most parents are willing to admit (Schneider and Buckley 2002; Schneider, Elacqua, and Buckley 2006). The distaste for schools with disproportionately more non-western minority students is apparent for both native Dutch and non-western parents themselves. Karsten and colleagues (2003) observe that parents regardless of ethnic background find predominantly 'non-white' schools in the neighbourhood – i.e. schools with 23 per cent disproportionately less native Dutch students compared to the four-digit postcode area – to be 'unsuitable'. It is also equally possible that ethnicity functions merely as a heuristic to parents for student ability, parental resources, and even teacher quality. With regards to the latter, schools with more disadvantaged or minority pupils could have difficulty attracting high quality teachers (Boyd et al. 2005; Clotfelter, Ladd, and Vigdor 2005; Jacob and Lefgren 2007). The disproportionately higher vacancy rates in schools with high share of non-western minority pupils coincides with the previous observation (Karsten et al. 2006; Ladd and Fiske 2009). In addition, the strong correlation between non-western background and low parental education could lead parents to proxy school peers' parental resources with share of non-western pupils. There is also the 'familiarity' aspect as students have been found to 'herd' together with fellow primary school peers when selecting secondary schools (Ruijs and Oosterbeek 2012).

Exploiting discontinuity at school district boundaries, Black (1999) estimates a 2.1 per cent higher willingness-to-pay by parents for one standard deviation increase in school test scores. Lower school quality has also been found to increase the probability on parental decision to exit charter schools, more than regular public schools (Hanushek et al. 2007). School quality matters in the Netherlands as well (Koning and van der Wiel 2013) but one study in Amsterdam using first preference school choice exposed some inconsistencies of this predictor (Ruijs and Oosterbeek 2012), i.e. some of the school quality indicators were found to have either no effect or negative effect on school choice. Guided by previous literature, we aim to reduce the potential omitted variable bias in our empirical analysis by incorporating all the possible determinants of school choice.

As a second line of earlier literature on school choice, stated preference studies tend to overestimate the importance of education quality over other factors such as racial and class peer composition (Schneider and Buckley 2002; Schneider, Elacqua, and Buckley 2006). In contrast, revealed preference research (see for instance Hastings, Kane, and Staiger 2005; Bayer, Ferreira, and McMillan 2007) accounts for: (i) the bias of survey respondents towards 'socially accepted' answers, (ii) the 'bundling' nature of school characteristics that needs to be trade-off with one another, and (iii) the constraints of regulation (e.g. catchment area) and school admission policies on parental choice. It is indeed a challenge for choice analysts to determine valid attributes that do influence choice selection because they are perceived subjectively by the decision-makers (Hensher, Rose, and Greene 2005). This is the main disadvantage of collecting revealed preference data from marketplace data such as school administrative records and not, for instance, from the parents as decision-makers.

A third line of literature considers school choice as a discrete choice problem. Random utility models (RUM), first introduced by McFadden (1973) with extensions by Ben-Akiva (1974), Williams (1977), Daly and Zachary (1978), and McFadden (1978) are popular in the school choice literature. Due to the higher computational demands of the conditional logit, multinomial logit models have been favoured in earlier studies (e.g. Manski and Wise 1983)[2]. Nonetheless, the latter can forsake key information by aggregating alternative-specific attributes while focusing more on the characteristics of the decision-maker that influences choice (Long 2004). Conditional logit models let the alternative-specific attributes to dictate choice. It also allows for pair-wise combined information between individual-specific attribute and alternative-specific attribute, such as residence-to-school distance, which is crucial for our study on primary school choice. As an extension of the conditional logit, nested logit models are common in the choice literature for tertiary education since it can account for the prior decision of whether or not to pursue post-compulsory education (Montgomery 2002; Kelchtermans and Verboven 2010a) – although the prevalence of the nested logit model with the 'not attend' option over the conditional logit model is contested by Long (2004). The mixed logit model is also an extension of the conditional logit model which, like the nested logit, relaxes the assumption of

[2] Random sampling of alternatives of the choice set (including the chosen alternative) is also a method to deal with the computational demands of the conditional logit estimations (Kelchtermans and Verboven 2010b; Kohn, Mansk, and Mundel 1976) and its nested logit extension (Montgomery 2002). See also the seminal paper by McFadden (1973).

independence of irrelevant alternatives (Hastings, Kane, and Staiger 2005; Ruijs and Oosterbeek 2012).

3. School choice in the Netherlands

The Dutch school system places a strong emphasis on free parental choice of schools (Dronkers 1995; Karsten et al. 2003; Ladd, Fiske, and Ruijs 2009). The decentralised school system in the Netherlands resembles a 'quasi-market' (Le Grand 1991) with public funding and private production of school services. Public funding of schools is allocated per capita with additional weights assigned based on parental education level. Schools with less than nine per cent students in need of additional weighting are not provided more than per capita funding.

Primary education is part of the compulsory education for all children between the ages of 5 and 16. Typically, a student begins primary school education from age 4 although the enrolment process starts much earlier, e.g. from the age of 2 in some Amsterdam schools (Gemeente Amsterdam 2013). Enrolment and admission policies are decentralised at the school-level although since 2009, there has been a trend for some schools in the same district towards harmonization.[3] Some religious schools have a religious admission criterion. While students are not bound to their residential neighbourhood schools – i.e. no 'catchment area' like in the United Kingdom – schools can give priority to children living in the school's neighbourhood. Other priority rules in school admission include having siblings who are enrolled in the school, children of the school's employees, and the attendance of similar pre-school education type. The more popular, oversubscribed schools also allocate placements by lottery. Lastly, unlike the United States, homeschooling is virtually non-existent in the Netherlands (Blok 2004).

While the Education Inspectorate reports and the national standardised test scores for schools are publically available, the municipality has been working towards greater transparency. Since 2011, the Municipality of Amsterdam has, in cooperation of the local school boards, published annual school quality indicator reports for all primary schools which include learning skill- and subject-based test scores, quality assessment from the Education Inspectorate, turnover rate, student socioeconomic composition, and type of secondary school stream recommendations (Gemeente Amsterdam 2011). The report allows for the direct comparison of schools and was aimed to improve the information asymmetry and efficiency of parental choice based on quality indicators. Similar information have been made publically available much earlier, from the 1990s, to guide secondary school choice (Ruijs and Oosterbeek 2012).

[3] After several successful pilot projects, a centralised system at the municipality-level is underway (Gemeente Amsterdam and Amsterdamse Schoolbesturen Primair Onderwijs 2013).

Choice within school closure

Primary school choice, however, is not limited to the first round of enrolment and is often a repeated decision-making event due to intentional or circumstantial school mobility (Alexander, Entwisle, and Dauber 1996; Hanushek, Kain, and Rivkin 2004; Allen, Burgess, and Key 2010). School mobility permits revealed preference choice analysis of new school choices and is more prevalent among non-western minority students in Amsterdam (see *Table 2*).

Table 2: Mobility between primary schools in column percentages (2000-2008), Amsterdam

# School Changes	Native Dutch	Moroccan	Antillean-Aruban	Surinamese	Turkish	Other Non-west	Western	Total
0	70.95	53.31	50.09	50.63	51.30	55.57	65.32	59.93
1	21.47	32.36	30.69	31.58	34.19	31.38	24.80	27.93
2	5.45	10.36	12.73	12.20	10.88	9.35	7.06	8.67
3	1.49	2.82	4.15	3.84	2.74	2.59	1.87	2.43
4 or more	0.64	1.15	2.35	1.75	0.89	1.10	0.95	1.04
Total	17,383	9,257	1,108	7,393	5,283	5,742	3,682	49,848

Source: Ong and De Witte (2013). 'Mobility' refers to move between schools (at the locational-level) offering standard primary education without correction for the merging, division or dissolution of schools.

Given the limitation of our data, we do not observe schools' admission policies, i.e. the supply-side constraints that affect school choice. Karsten et al. (2003), for instance, discuss gate-keeping measures taken by the schools during enrolment process. It is reasonable, however, to expect potential receiving schools to be open to small numbers of enrolment. More so with the mediation of the municipality that had assisted the displaced students with their relocation into other schools (personal communication, 2013). For the purpose of this paper, we make the strong but justifiable assumption that the students were unconstrained in school choice in contrast to the initial school enrolment when students were matched to schools based on the priority rules as previously discussed. Even so, approximately 500 out of the 600 students affected had chosen to move en bloc to one school under the same school board (Inspectie van het Onderwijs 2009). This was an unexpected outcome of the policy intervention – to the disapproval of the then State Secretary for Education, Culture and Science who was in favour of integrating the students into other, less insular schools (ANP 2007b). Temporarily, the students were allowed to stay in the same school buildings.[4] Nonetheless, there had been a complete change in school management, teachers, and other personnel.

4. Data description

For our study, we exploit school enrolment records provided by the Municipality of Amsterdam for the school years from 2005/2006 until 2010/2011. As many as 55,110 primary school students were enrolled in Amsterdam in July 2007. Between July 2005 and July 2007 when the forced closures took

[4] The larger two schools were renamed – *El Faroeq Omar* school became *IBS As-Siddieq (Zeeburg)* and *At Taqwa* school became *IBS As-Siddieq (Noord)* – and retained at original locations while students from the smaller *Abraham El Khaliel* school were absorbed by the other schools, including the original *IBS As-Siddieq (De Baarsjes)* school-location.

place, 665 students had been enrolled at least once in one of the three closed schools and had moved to another primary school within the municipality. For our final discrete choice analyses, observations with missing values in the battery of explanatory variables were excluded, leaving us with the final sample of 623 observations.

School alternatives in this paper refer to schools at the locational level, i.e. a school with two separate locations will be considered as two school alternatives. Altogether, there are 206 primary school alternatives across the years but the number varies between the years due to school closures and missing values in the covariates. The analysis were conducted separately for the two samples: the 'early movers' who had voluntarily changed schools within two years before the forced closure and the 'forced movers' who had to move to other schools after the closure. The latter group comprised a large subset of students who had collectively moved into one school that is managed by the same school board (Inspectie van het Onderwijs 2009). Although officially, 600 students were affected by the school closure, we only observe 486 of the 'forced movers' in our dataset as we focus exclusively on moves within the municipality and we exclude students who have since transitioned into secondary education. 'Early movers' comprise of 220 students who had changed schools within the municipality between July 2005 and June 2007.

In the selection of variables for the choice model, it is crucial to only include independent school attributes that actually differentiate the school alternatives from one another in the eyes of the parents as decision-makers (McFadden 1973). The dataset contains student-level covariates – ethnicity, residential postcode, and unauthorised absenteeism experience (truancy). Individual-level information is used to calculate aggregated school-level variables such as population size, peer ethnicity, and truancy rate. Meanwhile, Haversine-based distance from residence to potential schools is estimated via longitudinal and latitudinal information at the four-position postcode-level. Due to outliers, we cap our distance-to-school variable at 10 kilometres.[5] We were also provided test scores from the standardised national test (Cito) by the municipality but due to the lack of variation between schools (see *Table 3*), including this covariate did not alter much of our discrete choice analysis results.

Additionally, we have information on school type (i.e. teaching philosophy or religious denomination) and students' socio-economic status composition as measured by 2009[6] data on school funding

[5] Only 4.2 per cent of the students in our sample attend a school that is more than 10 kilometres away from their residence.
[6] Unfortunately, the pre-2009 student weight data is unsuitable for our purpose due to the Ministry of Education's gradual implementation of new student weight definitions between 2006 and 2009 (Ministerie van Onderwijs Cultuur en Wetenschap 2013). Up until 2006, additional weights were assigned by: 0.25 for native Dutch students with both parents having a maximum of lower vocational-level education; 0.40 for children of shipping crewmembers living away from the family; 0.70 for caravan-dwelling students; and 0.90 for first- and

weights for socially disadvantaged students. The latter pertains to public funding of schools which is allocated per capita and with additional weights assigned: 1.20 for students with at least one parent possessing primary-level education only; 0.30 for students with both parents (or the parent in-charge for primary care) having a maximum of lower vocational-level education (Rijksoverheid 2013). Given the large positive correlation between peer 'non-western' background and peer socioeconomic status as proxied by parental education (Pearson's product moment correlation = 0.86, p-value<0.000 for the very low education level, and 0.64, p-value<0.000 for the low education level), including both variables helps separate peer socioeconomic status from peer ethnicity when influencing school choice. Since additional funding is provided to schools with high proportion of socially disadvantaged pupils, we also include school personnel information to proxy for level of school resources. The earliest data provided by the Ministry of Education (*Dienst Uitvoering Onderwijs*, DUO) pertains to year 2008. We assume that both data on school funding weights and school staff (in fulltime equivalent) numbers did not fluctuate much between the years 2005 and 2008/2009. As a sensitivity analysis, the analysis was repeated with 2006 student weight data that only partially measures the proportion of low socioeconomic status students and the results were comparable[7] except for the variable for peers with very low socioeconomic status.

More than half of the students in our dataset are of Moroccan descent while a quarter of them are of Turkish descent. The non-significant chi-square test of association between student ethnicity and moving status indicates the relative similarity between the subgroups. Proximity is crucial as 90 per cent of them choose to attend a school that is within 500 metre away from their residences. In the bivariate table presented in *Table 3*, we compare the average student and receiving school characteristics of the two groups (and its subgroups) within our sample. It is apparent that students who chose to move en bloc to one school tend to live the closest to their new school while the potentially more selective early movers had moved to schools that are located further away. Truancy seems to be less of an issue for the early movers. Interestingly, truancy behaviour is more visible among those who did not move en bloc to the same school after the forced closure as nearly one out of five students have reported unauthorised absenteeism experience.

The vast majority (83 per cent) of the forced movers had moved into an Islamic school while the distribution across the three school denominations is roughly equal amongst the early movers. There are 10 percentage-points more non-western peers in the receiving school for the forced movers than for the 'early movers' – this difference is largely driven by the main receiving school for the forced movers with a high 93 per cent non-western student population. In contrast, forced movers who had

second-generation immigrants with at least one parent with a maximum lower vocational-level education or is unemployed, or the highest earning parent working in the manual or unskilled sector (Ladd and Fiske 2009).

[7] This set of results is available upon request.

moved to other schools ended up in schools with lower share of non-western peers and higher share of native Dutch and western peers. On the one hand, early movers appear more selective towards more native Dutch and western minority peers, smaller school size, lower truancy rate, and the higher level of school resources (as measured by the managerial and teaching staff-to-student ratio). On the other hand, forced movers emerge to be more selective towards schools with higher average test score and lower proportion of students with low parental education background (as measured by the 0.3 student funding weight).

With its relatively large student population, the main receiving school for the forced movers has a significantly smaller managerial and teaching staff-to-student ratio compared to the average receiving schools chosen by the other students. Potentially, the low teacher-to-student ratio is compensated by its relative high proportion of support and administrative personnel. Those who chose to move to this school have the average shortest residence-to-school distance which could indicate a potential trade-off between the school characteristics. The subsequent section with the conditional logit and nested logit explanatory models could shed light on this possibility.

Table 3: Distribution of student and receiving school characteristics by moving status

	Early Movers	Forced Movers		
	All	All	SIBA School	Other schools
Student characteristics				
Ethnicity (%)				
Native Dutch	2.3	2.7	2.5	3.6
Moroccan	60.0	58.9	61.0	48.8
Turkish	16.4	22.0	19.9	32.1
Surinamese	1.8	1.9	2.0	1.2
Antillean/Aruban	0.0	0.0	0.0	0.0
Other non-western	17.3	11.5	11.2	13.1
Western	2.3	2.7	3.0	1.2
With absenteeism experience (%)	**5.0**	**12.1**	**10.7**	**19.1**
Distance (in kilometre)	**0.6**	**0.2**	0.2	0.5
School characteristics				
School type (%)				
Public	**37.7**	**9.7**	**0.0**	**56.0**
Islamic	**27.7**	**82.9**	**100.0**	**1.2**
Christian	**34.6**	**7.4**	**0.0**	**42.9**
Peer ethnicity (%)				
Non-western	**80.5**	**90.4**	**93.0**	**78.0**
Moroccan	**39.9**	**46.2**	**49.5**	**29.7**
Turkish	**15.2**	**24.6**	**26.6**	**14.8**
Surinamese	**10.0**	**2.9**	**0.5**	**15.1**
Other non-western	**14.6**	**16.4**	16.4	16.6
Native Dutch	**14.5**	**4.2**	**1.6**	**17.0**
Western	**4.2**	**3.5**	**3.3**	**4.7**
Peer socioeconomic status (%)				
Low SES (0.3 weight)	**11.2**	**7.6**	**6.1**	**14.9**
Very low SES (1.2 weight)	35.2	37.3	37.8	34.6
Number of FTE staff per 100 student				
Management	**0.9**	**0.3**	**0.2**	**0.7**
Teaching	**7.5**	**6.1**	**6.0**	**6.8**
Support/Administrative	2.3	2.5	**2.7**	**1.8**
School size	**234.8**	**398.5**	**428.0**	**253.6**
Truancy (%)	**13.7**	**15.6**	15.7	15.6
Average test score	**534.9**	**535.5**	535.6	535.2
Sample size*	220	486	402	84

Source: Authors' own calculations with combined data from Municipality of Amsterdam (2005-2011), MINOCW (2009), DUO (2007-2008). Receiving school variables refer to 2005 with the exceptions of socioeconomic status (2009), test scores (2007/08), and school personnel (2008). Bold estimates indicate statistical significant difference at the 5% level between: (i) the early movers and (all) last movers; and (ii) among the forced movers, the SIBA-school and other schools. *Maximum sample size which does not account for missing values on covariates.

5. Methodology

5.1. Intuition behind the conditional and nested logit framework

Each school within the municipality of Amsterdam is considered to be a different school alternative. Given that we do not fully observe school mobility outside of Amsterdam, our analysis includes only intra-city mobility and does not include an 'out of Amsterdam' alternative. Due to the large set of differentiable school choice alternatives (i.e. 206 different schools), a parent's choice of primary school for their child[8] is estimated using a conditional logit model.

While its estimates report the likelihood of selecting a school with a given set of observed characteristics, the conditional logit model suffers from one major drawback – it does not allow for different substitutability or complementarity between alternatives with its assumption of independent and irrelevant alternatives (IIA). In our case, it is intuitive to see that when an Islamic school alternative is removed from or added to the choice set, the probability of selecting another Islamic school alternative changes disproportionally more than the probability of selecting a Public or Christian school alternative.

The mixed logit relaxes the assumption and allows for heterogeneous preferences for school characteristics but its lack of closed form solution creates computational limitations which we deem to be unnecessary for our relatively homogeneous student sample. Instead, the IIA assumption is relaxed in our study by means of a nested logit model which allows for the alternative-specific error terms within a 'branch' or 'nest' to be correlated with one another. The nested model does not imply a sequential decision-making process but it can be interpreted as such: parents as decision-makers prioritise the choice of school type or denomination before selecting schools within the school type based on a set of school attributes. To our best knowledge, our study is the first in applying a nested logit model to examine compulsory primary school choice that is nested within school type or religious denomination. Besides the composition of our unique sample of students enrolled in Islamic schools, we argue that parental choice for primary school is intuitively nested in the choice for school type. The failure to account for the decision's nesting structure would otherwise bias our choice estimates.

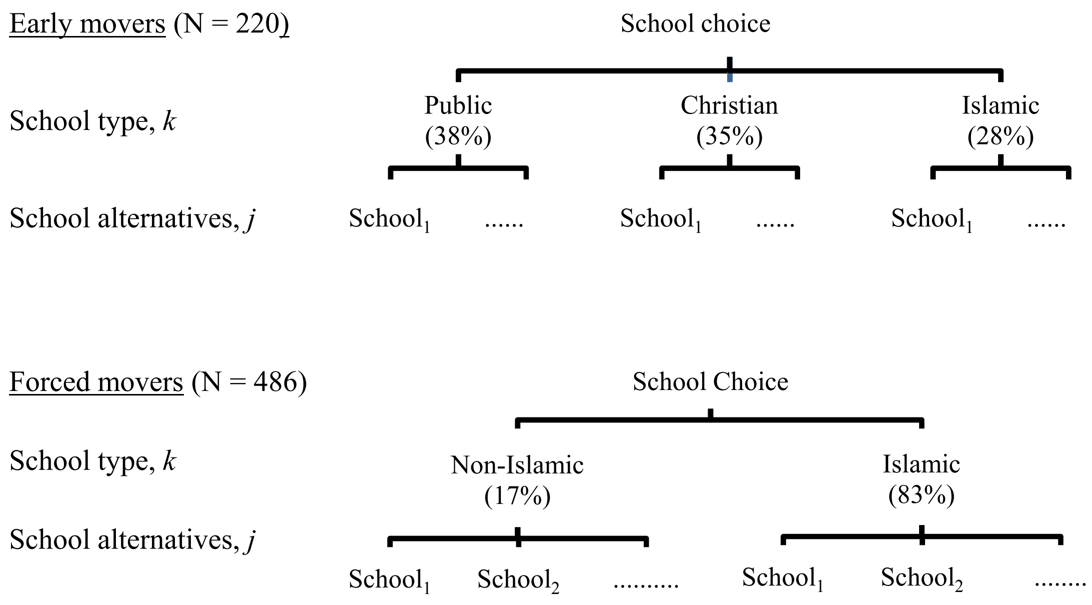
Even so, the main disadvantage of the nested logit model when compared to the conditional logit model is that the latter permits the estimation of the effect of school type as a choice determinant. It is also a useful comparison should either one of the nested logit models for our two sample groups be misspecified, i.e. that school choice is not nested on school type. Therefore, we estimate both the conditional logit and the nested logit probabilities of school choice for: (i) the 'early movers' who had

[8] In our paper, we assume that school choice is made by the parents of the student and/or the student and use both terms interchangeably.

changed schools within two years before the forced closure in July 2007; and (ii) the 'forced movers' who were forced to change schools after the forced closure (see *Figure 1*). Within a two-level nested logit model with school denomination as the nesting factor, the early movers can choose between three school types: public, Islamic, and Christian. We include school alternatives from other religious denominations in the last group, but since none of the non-Christian religious schools were selected by our sample, we can effectively consider this nest to be 'Christian'. In view of the small proportion of forced movers selecting non-Islamic schools (17 per cent), we only distinguish between those who had chosen the same Islamic school denomination, and those who had chosen 'non-Islamic' school types.

Except for one student, all of the forced movers who had chosen an Islamic receiving school had moved en bloc to the same school – a policy outcome that contradicts the initial aim and policy line of the Ministry of Education. Conceptually, we can interpret the nesting structure to be one of reconcentration-versus-dispersal behaviour by the students. The nesting structure is also consistent with the reality that most parents of the students from the closed schools were in favour of moving into an Islamic school instead of a non-Islamic school (as reported by ANP 2007a).

Figure 1: Nested logit model of school choice for the early movers and forced movers



Note: Number of observations includes those with missing values on some covariates.

5.2. School choice model

The intuitive description of *Subsection 4.1* is formalised below to give a more precise description of school choice.

Conditional logit model

Let the utility that a parent, *n*, derives from choosing school alternative*, j (1, ..., J)* to be U_{nj} (we follow the notations as in Train 2003). Since we do not observe all factors that can influence one's utility in choosing school j, we define V_{nj} as the part of utility derived by parent n that is observable to the researcher and ε_{nj} to be the random error term that captures the unobserved factors affecting utility:

$$U_{nj} = V_{nj} + \varepsilon_{nj} .$$

Under the assumption of utility maximisation, the parent will choose the school that will offer the highest utility. By choosing school i, the derived utility must be higher than the utility offered by all other school alternatives, $U_i > U_j \; \forall \, j \neq i$. By assuming the error term ε_{nj} to be independent and identically distributed (i.i.d.) with extreme value distribution and utility as linearly additive, the random utility model can be conveniently estimated with the standard logit model (c.f. McFadden 1973). Under the logit model, school choice is defined as a logit choice probability (expressed in its variance-scaled form) of parent *n* choosing school i,

$$P_{ni} = \frac{e^{\beta' x_{ni}}}{\sum_j e^{\beta' x_{nj}}}$$

where $V_{nj} = \beta' x_{nj}$ and x_{nj} represents the vector of observable characteristics of school alternative *j* that affect utility of parent *n*. The estimated β coefficients measure the actual effect of each observed variable scaled to the variance of the unobservables. And so the interpretation of individual β coefficients should account for the fact that a lower coefficient does not necessarily indicate a smaller effect as it could be due to larger variance of the unobservables (Train 2003).

Nested logit model

As elaborated previously, a nested logit model that relaxes the assumption of proportional substitution between school denominations seems more appropriate for our study. Analogous to the previous model, if we now assume school alternatives j to be separable into K number of non-overlapping 'nests' (or school denomination in our case), B_k, the utility of parent *n* choosing school *j* can be decomposed into three parts:

$$U_{nj} = W_{nk} + Y_{nj} + \varepsilon_{nj}$$

for $j \in B_k$ where W_{nk} is the observable utility component common to all school alternatives in nest B_k, Y_{nj} denotes the observable utility component specific to each school j within nest B_k and ε_{nj} is the idiosyncratic error term (Train 2003).

The probability for parent *n* to choose school i in nest B_k can be expressed as the product of the marginal probability of parent *n* choosing an alternative within nest B_k, P_{nB_k} and the conditional probability of parent *n* choosing school i in nest B_k conditional on choosing an alternative in nest B_k, $P_{ni|B_k}$:

$$P_{ni} = P_{ni|B_k} P_{nB_k}.$$

Employing the generalised extreme value (GEV) distribution of the error term ε_{nj} (c.f. McFadden 1978; Williams 1977; Daly and Zachary 1978), the probability for parent *n* to choose school *i* in nest B_k can estimated by a utility maximisation-consistent nested logit model. Under the nested logit, the i.i.d. condition still holds between the error terms of school alternatives in different nests:

$$cov(\varepsilon_{nj}, \varepsilon_{nm}) = 0 \text{ if } j \in B_k \text{ and } m \in B_l \text{ with } l \neq k.$$

But the error terms of different alternatives within the same nest are now allowed to be correlated:

$$cov(\varepsilon_{nj}, \varepsilon_{nh}) \neq 0 \text{ if } j, h \in B_k \,.$$

To estimate random utility model-consistent choice probabilities given the correlated error terms within a nest, the observable utility component specific to each school j within each nest, Y_{nj} needs to be normalised. This can be done via rescaling by the inverse of nest B_k's dissimilarity parameter, λ_k (we refer to the elaboration by Heiss 2002). Furthermore, this normalisation allows for the comparability across nests. Similar to the conditional logit before, the j school alternative-specific (rescaled) utility derived by parent *n* is estimated by a vector of observed school characteristics, x_{nj}: $Y_{nj}/\lambda_k = \beta' x_{nj}$. While the nest-specific utility, W_{nk} can be estimated by a vector of observable factors, z_{nk} common to all alternatives within each nest, B_k: $W_{nk} = \gamma' z_{nk}$. Hence, the corresponding marginal and conditional probability logit models:

$$P_{nB_k} = \frac{e^{\gamma' z_{nk} + \lambda_k I_{nk}}}{\sum_{l=1}^{K} e^{\gamma' z_{nl} + \lambda_k I_{nl}}}$$

$$P_{ni|B_k} = \frac{e^{\beta' x_{ni}}}{\sum_{j \in B_k} e^{\beta' x_{nj}}}$$

where the inclusive value, I_{nk} is the log of the denominator of the conditional logit probability model that links the two logit probabilities:

$$I_{nk} = \ln \textstyle\sum_{j \in B_k} e^{\beta' x_{nj}}.$$

Due to lack of information on the nest-specific factors, the z_{nk} vector is not estimated in this paper. Instead we only estimate, in addition to the alternative-specific utility, the overall utility a parent *n* derives from 'being able to choose the best alternative in the nest' which is equivalent to the inclusive value, $\lambda_k I_{nk}$ (Train 2003). In our application, the nested logit model estimates the following logit choice probability of parent n choosing primary school i given that they choose school denomination B_k:

$$P_{ni} = \left(\frac{e^{\beta' x_{ni}}}{\sum_{j \in B_k} e^{\beta' x_{nj}}}\right)\left(\frac{e^{\lambda_k I_{nk}}}{\sum_{l=1}^{K} e^{\lambda_l I_{nl}}}\right).$$

From this equation, it is apparent that the IIA condition with proportional substitution across alternatives still holds within each 'nest' or school type, but not between 'nests'.

Following Heiss (2002), we specify the coefficient of the inclusive value, λ_k (also known as dissimilarity parameter) to be equal to $\sqrt{1-\rho_k}$ with ρ_k measuring the correlation between the error terms of all alternatives within nest B_k. By this definition, a dissimilarity parameter value that is close to unity indicates more independence while a value close to zero suggests dependence or higher correlation between the unobserved utility components of the alternatives within nest B_k.

6. Determinants of primary school choice

The results of the conditional and nested logit estimates are presented in *Table 4*. As the results are complementary, we will discuss them simultaneously. The conditional logit and nested logit estimates concur with our expectation that, all things equal, peer ethnicity matters when it comes to primary school choice. Both the early movers and forced movers prefer schools with more non-western peers – every percentage point increase in non-western minority peers is associated with a 2 per cent increase in odds of school choice for an early mover and 7 per cent increase in odds of school choice for a forced mover (as derived from $e^{0.02}$ and $e^{0.07}$ respectively). It is also much more pronounced in the nested logit model for forced movers when school choice is first nested on school denomination. Here, a one percentage point increase in non-western peers increases the odds of school choice by 14 per cent, all else held constant. It is worth noting that our selected sample comprised of students who had enrolled themselves into ethnically segregated Islamic schools hence our estimates are not generalizable to the rest of the student population. We also cannot rule out the possibility that peer ethnicity is used merely as a heuristic device for parents to assess the school's overall student ability, parental resources, and teacher quality.

While peer ethnicity is an unambiguous determinant of school choice, peer socioeconomic composition has a mixed effect on school choice. First, early movers appear to be indifferent towards a school's proportion of students from socially disadvantaged background. Next, the effect of peer socioeconomic composition on the school choices of forced movers changes sign from negative to positive once school choice is defined to be nested on school type. The contradicting findings highlight the potential bias that arises when the appropriate nesting structure is not applied to a conditional logit model. All things equal, every percentage point increase in students of 'low' and ‘very low’ parental education background increases the likelihood of a forced mover selecting a school by 23 and 16 per cent, respectively.

For every one percentage-point increase in peer truancy rate, the likelihoods of selecting a school for the early movers and forced movers decrease by 4 per cent and 15 per cent respectively. Truancy here is measured by the proportion of students in the receiving school with unauthorised absenteeism experience. Distance between residence and school remains a robust determinant of school choice for both groups as parents prefer primary schools that are closer to home. When primary school choice is nested on school denomination, distance becomes a stronger determinant for the forced movers as every kilometre in additional distance reduces the odds of selecting a school by 76 per cent.

School size matters only to the forced movers as the likelihood of choosing a school increases by 2 per cent for every additional student. Controlling for school size and other factors, the early movers tend to choose schools with fewer teaching staff, while the forced movers appear to be indifferent towards it in the nested logit model. To explain the counter-intuitive finding, we allude to the fact that we have not controlled for teaching quality and other likely sources of unobserved heterogeneity (Boyd et al. 2005; Clotfelter, Ladd, and Vigdor 2005). In addition, the weighted student funding structure with its nine-per cent threshold has been acknowledged for its positive bias towards schools with more disadvantaged students by providing additional resources commonly used to hire additional personnel (Ladd and Fiske 2009). A survey of primary school principals in the Netherlands in 1992 positively correlates school size with the probability of having one full-time school director (De Haan, Leuven, and Oosterbeek 2011) – a factor that could improve overall school management and quality.

A consistent predictor for the forced movers, the number of support and administrative staff increases the likelihood of choosing a school by a factor of 3 for every additional personnel. Following our earlier hypothesis of students preferring larger schools for more managerial personnel, the number of managerial personnel has contrary effects on the early movers and forced movers. The former have a slight preference for schools with more managerial staff while the reverse is true for the latter. We posit two explanations for the forced movers’ ‘distaste’ in this: (i) parents trade off school characteristics in selecting primary schools and the benefits of being in, for instance, an Islamic school, overrides the cost of not having a full-time school director; (ii) although all school personnel information is publically available, school management is potentially less visible to parents and could be overlooked in their decision-making process.

From the conditional logit models, it is evident that both the early movers and forced movers do not distinguish between public and non-Islamic religious schools. This is not the case for Islamic schools – the likelihood of choosing one for the forced movers is 2.9 times the likelihood of choosing a public school, all else held constant. For the 'early mover' sample, importance of school type as a predictor depends on the model's battery of covariates since some schools – most notably, the three closed schools – are no longer observed in 2008/2009 on the school personnel and student weight funding variables (see *Table A* in the *Appendix*). There is substantial interschool mobility between the three closed schools prior to the forced closure in 2007: 50 out of 209 observations within two years before the closure. A school's Islamic denomination is a statistically significant determinant of school choice for the 'early mover' sample only when we account for mobility between the three Islamic schools (see *Table 4* and *Table A* in the *Appendix*).

Nevertheless, the rejection of the independence of irrelevant alternatives assumption within the school denomination branch(es) for both nested logit models justifies its use over the conditional logit. The null hypothesis of this likelihood ratio test defines all the dissimilarity parameters in the nested logit model to be equal to one, $\lambda_k = 1 \forall k$. If that were to be true, the model collapses into a standard conditional logit model. For the forced movers, the estimated dissimilarity parameter for the Islamic denomination schools of 0.143 corresponds to a very high correlation – approximately 0.9794 – between the error terms of the school alternatives[9].

[9] We note that for both models, at least one of the dissimilarity parameters had a value exceeding unity which violates the *global* utility maximization assumption under the additive random utility model (Hensher, Rose, and Greene 2005). In any case, large dissimilarity parameters are prevalent in the discrete choice literature and we argue that our model could still be locally consistent under utility maximization, i.e. for a subset of alternatives in each nest, or for a constrained range of all possible values of our covariates (for further discussion on the issue we refer to the works of Börsch-Supan 1990; Davis, Gallego, and Topaloglu 2012; Train 2003).

Table 4: Conditional and nested logit estimates by moving status

	Conditional Logit		Nested Logit	
	Early leavers	Forced movers	Early leavers	Forced movers
School denomination				
Public (reference)				
Islamic	0.55	1.04**		
	(0.42)	(0.43)		
Christian	0.26	0.03		
	(0.17)	(0.27)		
School size	0.00	0.00***	0.00	0.02**
	(0.00)	(0.00)	(0.00)	(0.01)
Truancy rate (%)	-0.03**	-0.05***	-0.04*	-0.16**
	(0.01)	(0.02)	(0.02)	(0.07)
Non-western peers (%)	0.02***	0.07***	0.02**	0.13***
	(0.01)	(0.01)	(0.01)	(0.05)
Low SES peers (%)	-0.02	-0.06***	-0.01	0.21**
	(0.01)	(0.01)	(0.02)	(0.10)
Very low SES peers (%)	0.02*	-0.03***	0.02	0.15**
	(0.01)	(0.01)	(0.01)	(0.07)
Number of managerial staff (FTE)	0.21	-0.23	0.34*	-3.55***
	(0.14)	(0.18)	(0.20)	(1.20)
Number of teaching staff (FTE)	-0.05***	0.07***	-0.07**	0.14
	(0.02)	(0.02)	(0.03)	(0.10)
Number of support staff (FTE)	0.05**	0.12***	0.03	1.10***
	(0.02)	(0.03)	(0.04)	(0.29)
Distance-to-residence (km)	-0.49***	-0.52***	-0.59***	-1.43***
	(0.09)	(0.08)	(0.21)	(0.29)
Log Likelihood	-761.563	-646.749	-758.673	-578.678
Akaike Information Criterion (AIC)	1545.126	1315.499	1541.346	1179.356
Reject IIA assumption within nests			yes	yes
Number of cases	160	463	160	463
Number of schools	195	196	195	196

Note: *** p<0.001, ** p<0.05, * p<0.10.Aggregated variables on receiving school refer to 2005 for the 'early movers' and 2006 for the 'forced movers'. For all samples, we use 2009 weighted student funding data on parental education, 2007/2008 CITO test scores, and 2008 school personnel data. Number of observations and school alternatives used in the analyses exclude those with missing values in any of the explanatory variables. Using the likelihood ratio chi-square test, we reject the null hypothesis of IIA within nests at the five per cent significance level with p=0.030 for the 'early movers' sample and p<0.000 for the 'forced movers' sample.

Robustness checks

Since the likelihood ratio test in the nested logit model could be susceptible to the tree structure, the alternative Hausman-McFadden test is usually performed first on the conditional logit model to determine the necessity of a nested logit model (Hensher, Rose, and Greene 2005). The test's null hypothesis of IIA is rejected for the conditional logit model of the 'forced mover' sample (χ^2= 139.35, p-value<0.000), i.e. removing one or a subset of alternatives does have a statistically considerable effect on the conditional logit estimates.

As mentioned before, there is substantial interschool mobility between the three schools prior to the forced closure. For the 'early mover' sample, the Hausman-McFadden null hypothesis is rejected for the 'early mover' conditional logit model when these interschool mobility observations are included (χ^2= 25.51, p-value=0.008) but not when they are excluded. To account for the effects of the additional school personnel covariates, we have re-estimated the model without these covariates for the 'early mover' sample with and without interschool mobility between the three closed schools (using 2006 student weight data that partially observes peer socioeconomic composition). The comparability of results leads us to conclude that school choice is indeed subject to school type or denomination although we cannot ascertain if the nested logit framework (especially under the current three-category nesting structure) is necessary for the 'early mover' sample.

7. Conclusion and policy implications

This paper exploits the forced closure of primary schools to examine the school choice of students in three segregated Amsterdam schools. While the school closure was caused by the schools' poor performance and administrative mismanagement, the selected schools comprised of students from predominantly non-western minority and disadvantaged background. The vast majority of the students who were forced to change schools chose to move en bloc to one school – an unintended and undesired policy outcome. Our study seeks to tease out the reasons underlying these school choices.

Despite the limited interpretation of our results to a specific student subpopulation, they contribute to the under-researched school choice literature for minority students in Western Europe. Quite intuitively given our select sample, school choice is nested in the choice for school type or religious denomination. Failure to account for the nesting structure in decision-making will lead to biased estimates, e.g. for peer socioeconomic composition. Besides the appropriate use of the school denomination nesting structure for the 'forced mover' sample, we also observe high student mobility between these three religious schools among the early movers. Like the forced movers, the voluntary movers prefer schools with more non-western students, less truancy behaviour among peers, and shorter distances from home. But unlike them, they do not retain a taste for peers of (most likely similar) low socioeconomic background after controlling for the additional school resources tied to the weighted student funding. They are also less concerned of school peers’ truancy behaviour – contrary to our earlier ‘Tiebout-mobility’ hypothesis – but seemed more selective towards schools with more managerial staff.

Our study provides some unique insights that could aid future policymaking namely, after an external shock of school closure: (i) students in segregated schools prefer re-concentration to dispersal; (ii) school choice of students is nested upon school type or religious denomination; and (iii) peer composition (in terms of ethnicity and socioeconomic status), distance, truancy, school size, and number of school personnel (managerial, support, and administrative staff) are relevant predictors for school choice. Similar policy interventions in the future should account for these school choice determinants, i.e. by ensuring the availability of desirable school substitutes to students, during the policymaking process.

APPENDIX

Table A: Conditional logit estimates by moving status

	Early Movers		Forced Movers		
Receiving school characteristics	(1)	(2)	(3)	(4)	(5)
School denomination					
Public (reference)					
Islamic	1.35	0.55	3.00	1.04	-3.12
	(0.20)	(0.42)	(0.21)	(0.43)	(1.19)
Christian	0.26	0.26	0.59	0.03	-0.38
	(0.17)	(0.17)	(0.25)	(0.27)	(0.27)
School size	0.00	0.00	0.01	0.00	0.01
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Truancy rate (%)	-0.02	-0.03	-0.06	-0.05	0.00
	(0.01)	(0.01)	(0.01)	(0.02)	(0.02)
Non-western peers (%)	0.04	0.02	0.12	0.07	0.00
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)
Low SES peers (%)	-0.03	-0.02	-0.11	-0.06	0.12
	(0.02)	(0.01)	(0.02)	(0.01)	(0.02)
Very low SES peers (%)	-0.08	0.02	-0.29	-0.03	0.03
	(0.02)	(0.01)	(0.03)	(0.01)	(0.01)
Number of managerial staff (FTE)		0.21		-0.23	0.30
		(0.14)		(0.18)	(0.21)
Number of teaching staff (FTE)		-0.05		0.07	-0.10
		(0.02)		(0.02)	(0.03)
Number of support staff (FTE)		0.05		0.12	0.02
		(0.02)		(0.03)	(0.05)
Distance-to-residence	-0.44	-0.49	-0.56	-0.52	-0.56
	(0.06)	(0.09)	(0.08)	(0.08)	(0.11)
Log Likelihood	-964.872	-758.149	-746.379	-623.712	-328.982
Akaike Information Criterion (AIC)	1945.745	1538.298	1508.759	1269.424	679.964
Pseudo R^2	0.1295	0.0957	0.6983	0.7448	0.1451
Number of cases	209	159	466	463	75
Number of schools	201	195	202	196	196

Note: Model (5) refers to the 'forced mover' sample excluding those who had chosen to move en bloc one school. Aggregated variables on receiving schools calculated in years 2005 for the 'early movers' and 2006 for the 'forced movers'. For all samples, we use 2006 weighted student funding data on parental education, 2007/2008 CITO test scores, and 2008 school personnel data. Number of observations and school alternatives used in the analyses exclude those with missing values in any of the explanatory variables. The difference between Models (1) and (2) is driven by students who were moving between the three closed schools prior to its closure.

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