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Gender, single-sex schooling and maths achievement

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ABSTRACT

This paper uses a distinctive feature of the Irish education system to examine the impact of single-sex education on the gender difference in mathematical achievement at the top of the distribution. The Irish primary school system is interesting both for the fact that many children attend single-sex schools, and because these single-sex schools are part of the general educational system, rather than serving a particular socio-economic group. In keeping with research on other countries, we find a significant gender gap in favour of boys, but contrary to suggestions in the literature, our results provide no evidence that single-sex schooling reduces the gap. If anything, the gender differential is larger for children educated in single-sex schools than in coeducational schools.

1. Introduction

The importance of mathematical skills has been well-documented. Schrøter Joensen and Skyt Nielsen (2010) provide evidence that maths skills have a causal effect on labour market outcomes and there is evidence that the individual returns to maths skills are higher than the returns to other skills (Buonanno & Pozzoli, 2009; Grogger & Eide, 1995; Koedel & Tyhurst, 2012; Paglin & Rufolo, 1990). The OECD (2010a) has emphasised the importance of mathematical proficiency for economic growth and there is concern in several countries that educational policy is not supporting children in attaining high levels of mathematical achievement. For these reasons, improving maths performance has been a key focus of educational policy for many governments.[1]

Although recent evidence suggests that females have surpassed males in general educational attainment in many industrialised countries (Pekkarinen, 2012), in maths, boys still tend to outperform girls (Bedard & Cho, 2010; Bharadwaj et al., 2012; Close & Shiel, 2009; Dickerson, McIntosh, & Valente; Fryer & Levitt, 2010; Hedges & Nowell, 1995; Husain & Milimet, 2009; OECD, 2010b).[2] Of the 65

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[1] Examples include projects aimed at addressing underachievement in maths by changing the attitudes and practices of schools, parents and children such as the Ocean Mathematics Project in London (Bernie & Lall, 2008); policies aimed at increasing the amount of time devoted to maths during school time, such as the 'numeracy hour' in the UK (Machin & McNally, 2005); the introduction in 2012 of bonus points for higher level maths when determining admission to university in Ireland; and the development of new assessment strategies for maths in the US (Stecher & Klein, 1996).

[2] Meta analyses conducted by Else-Quest, Hyde, and Linn (2010) and Lindberg, Hyde, Linn, and Petersen (2010) find negligible gender effects in average maths scores but note considerable heterogeneity across countries.

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countries participating in the 2009 Program for International Student Assessment (PISA), boys scored better in maths than girls in 54 countries. Since, in many countries the observed gender gap is largest at the top of the maths distribution, there is a growing concern that this is limiting the contribution of women in scientific and technical fields. As a result, strategies to raise the performance of girls have received particular attention in both educational policy discussions and in the media.[3]

In recent years, single-sex schooling has received increased attention. This is partly because several countries have been experimenting with single-sex classes within coeducational schools in an attempt to raise overall achievement. Title IX of the US Education Act was amended in 2006 to allow more flexibility to school districts to provide single-sex education, while in the UK, a 2007 government-backed review argued that boys should be taught separately, with more emphasis on 'competitive' lessons and the reading of non-fiction books (Department for Education and Skills, 2007). In addition, Fryer and Levitt (2010) have suggested that single-sex schooling might reduce the gender gap in maths. However, the effect of single-sex schooling on outcomes is usually difficult to test, as in most countries single-sex schools are selective and the numbers attending them are small (see Halpern et al., 2011). Ireland is unusual in that a sizable proportion of primary school children – about a quarter – are educated in single-sex, non-selective, state-funded schools. For this reason, the Irish educational system provides a valuable setting for examining the effects of single-sex schooling.

In this paper, we use data on Irish children to examine the determinants of the gender gap in maths at the age of nine, focussing on the impact of single-sex schooling. Concentrating on the early years of education is important because there is evidence that these are the years when the gender gap first opens up (Fryer & Levitt, 2010; LoGerfo et al., 2006; Robinson & Lubienski, 2011). Our results show that a gender gap in maths scores exists among Irish primary school children, particularly at the upper end of the distribution, but there is no evidence that single-sex schooling reduces the gap. In fact our results suggest that the gender gap may be larger for children attending single-sex schools.

The remainder of the paper is structured as follows: Section 2 provides a review of the literature. Section 3 describes the Irish educational system and Section 4 describes the data we use. Section 5 reports our results and Section 6 concludes.

2. Literature review

A number of recent studies have analysed the gender gap in mathematical performance among children. Halpern et al. (2007) provide a comprehensive summary of the literature, looking at the forces that contribute to this gender gap, including a discussion of biological theories related to the level and variance of innate spatial ability, and societal factors such as differences in the expectations of girls and boys. They conclude that the observed male advantage in maths is largest at the upper end of the ability distribution and that many variables work together in a complex way to determine this gender gap. Niederle and Vesterlund (2010) note that the differences in observed maths scores may not necessarily match gender differences in maths skills but may in part reflect how men and women respond differently to test-taking environments. However, they note that the observed gender gap in maths may have important implications irrespective of the underlying mechanism.

Fryer and Levitt (2010) use the Early Childhood Longitudinal Study Kindergarten Cohort (ECLS-K) and find that girls and boys in the US were observationally equivalent in both maths and reading when they entered kindergarten (normally at age 5 or 6), but by the end of fifth grade, at age 10 or 11, girls had fallen more than standard deviations behind their male counterparts. They test some societal theories of the gender gap by controlling for such variables as differential ratings by teachers, parental expectations, whether a child's mother worked in a maths-related occupation and the amount of time parents spent doing maths-related activities with children. They find that none of these has a substantial effect on the gender gap.

To gain further insight, Fryer and Levitt use two international datasets, the Trends in International Mathematics and Science Study (TIMSS) and PISA to investigate whether the size of the gender gap is related to gender inequality, using the World Economic Forum gender gap index. The authors find a strong positive relationship between gender equality and the relative performance of girls in maths using PISA[4] but not when using TIMSS. The difference in results is driven by the fact that TIMSS includes a large number of Middle Eastern countries that are not in PISA. Although these countries have a high degree of gender inequality, there is no gender gap in maths. The authors tentatively suggest that this may be due to the prevalence of single-sex schools in these countries.

Mael, Smith, Alonso, Rogers, and Gibson (2004), Roberson Hayes, Pahlke, and Bigler (2011) and Sullivan, Joshi, and Leonard (2010) all discuss mechanisms by which sex-segregated schools might affect relative academic performance across genders. Potential mechanisms include peer effects, differential attitudes to competition, and gender differences in approaches to learning.

One argument in favour of single-sex schooling is that peers' sexist attitudes interfere with girls' learning in coeducational schools. Sullivan et al. (2010) and Schneeweis and Zweimuller (2012) note that educational choices

[3] For example in June 2012 the European Commission launched a campaign challenging stereotypes on science and encouraging young people, especially girls, to study science. In another example, pop star <***>Will.i.am recently funded a prime time TV special called "i.am.first: Science is Rock and Roll" designed to get students, and young girls in particular, excited about learning STEM skills (science, technology, engineering and math).

[4] Guiso et al. (2008) report similar results, while Pope and Sydnor (2010) find that gender disparities in maths scores in the US are smaller in more gender-equal states.

may become more gender-stereotyped in a co-educational setting. In contrast, Halpern et al. (2011) note that there is evidence showing that sex segregation may increase gender stereotyping and, by doing so, legitimise sexism. Looking at attitudes to maths in particular, recent work by McCoy et al. (2012) reports that while in general boys are more positively disposed towards maths than girls, there is little difference in attitudes to maths between boys and girls educated in single-sex schools.

Hoxby (2000) and Lavy and Schlosser (2011) report evidence that the achievement of *both* boys and girls is increasing in the fraction of females in their peer group. Students who have more female peers report a lower level of classroom violence and better relationships with other students and teachers. In addition, their evidence suggests that teacher fatigue and burnout are inversely related to the proportion of girls in a class, in which case, single-sex schooling will increase the performance of girls, but at the cost of poorer achievement by boys. Roberson Hayes et al. (2011) cite evidence that boys tend to demand and receive more of teachers' attention than girls in coeducational classes, particularly in maths and science, indicating a further benefit for girls from single-sex schooling.

A second explanation as to why single-sex schooling may affect the achievement gap focuses on gender differences in response to competition. A number of papers investigate these differences using an experimental approach. For example, Gneezy, Niederle, and Rustichini (2003) test reactions to competition in maths puzzles by boys and girls and conclude that it is not that women are unwilling or unable to perform well in competitions *per se*, but rather that they do not compete well in competitions against men. Huguet and Régner (2007) find that this effect is particularly pronounced when girls are led to believe that a task measures maths ability rather than other abilities.[5]

Booth and Nolen (2012) examine the impact of school environment on attitudes to competition. They find that girls who are randomly assigned to all-girl groups are significantly more likely to be competitive. They also find that girls attending single-sex schools behave more competitively than girls in co-educational schools. For boys, they find that neither attendance at single-sex schools nor the gender make-up of boys' experimental peer group are important in explaining their attitudes to competition.

A third area of research on single-sex education considers the possibility that there are innate differences in how boys and girls learn. For example, Sax (2005) suggests that boys typically learn algebra better when introduced using numbers, whereas girls seem to be more engaged if the topic is introduced with a word-based problem. In this case, educational instruction is more effective when it is tailored to these differences, which may be easier in single-sex schools. However, Halpern et al. (2007) argue that there is no evidence that differences in brain function cause boys and girls to learn differently and maintain that this is not a valid justification for single-sex schooling.

Although single-sex schooling could affect academic performance through any of the mechanisms discussed above there is substantial debate in the empirical literature on the actual impact of single-sex schooling on performance.[6] Riordan (1985) found a significant advantage to single-sex education for girls but not for boys. Bryk, Lee, and Holl (1993) also found positive effects for girls' academic achievement as well as for social and personal development outcomes in girls' schools. However, Billger (2009) argues that much of the effect of single-sex schooling among private schools in the US accrues to students already likely to succeed and concludes that overall her results "do not provide a ringing endorsement of single-sex education." Halpern et al. (2011) go further and argue that there currently exists no well-designed research showing that single-sex education improves students' academic performance They quote Smithers and Robinson (2006) who conclude that "[t]he paradox of single-sex and co-education is that the beliefs are so strong and the evidence so weak."

The empirical evidence on the effect of single-sex schooling on mathematical achievement specifically is also mixed. Eisenkopf et al. (2012) analyse the impact of female-only classes on mathematical achievement, exploiting random assignment of girls into single-sex and co-educational classes in a single Swiss secondary school. They find that single-sex classes improve the performance of female students in maths and that this positive effect increases if the single-sex class is taught by a male teacher. However Jackson (2012), using random allocation of students to secondary schools in Trinidad and Tobago, finds that while girls with strong expressed preferences for single-sex schools had better exam performances, most students perform no better at single-sex schools. Furthermore he finds some evidence that girls in single-sex schools take fewer science courses.

Park et al. (2012) exploit the random assignment of children into single-sex and coeducational secondary schools in Korea to examine the impact of the gender composition of schools on maths scores, interest and confidence in maths and choice of STEM subjects in university. They find a significant causal effect of all-girls schools on maths scores, but no effect on other STEM outcomes; all-boys schools, on the other hand, have a strong positive effect on all of the outcomes studied. Booth et al. (2013) find that girls randomly assigned to a single-sex economics tutorial in a UK university do better than girls assigned to a coeducational tutorial, but there is no effect on the probability of choosing quantitative courses in their second year. Single-sex tutorials have no effect on either grades or subject choice for boys.

Almost all of the previous studies of single-sex schooling focus on education at either secondary or university level. None of the studies included in the comprehensive review of the impact of single-sex

[5] Further discussion of gender differences in competition can be found in Örs, Palomino, and Peyrache (2008), Gneezy et al. (2009), Günther, Ekinci, Schwieren, and Strobel (2010) and Cotton et al. (2010).

[6] For a recent survey, see Smyth (2010).

schooling on maths performance by Mael, Alonso, Gibson, Rogers, and Smith (2005) looked at primary school children. One recent paper by Dickerson et al. (2012) on the gender gap in maths in primary schools in a number of African countries finds mixed evidence on the impact of single-sex schooling; however, the authors note that the number of children enrolled in single-sex schools in the countries they study is very small.

This paper adds to the literature on single-sex schooling by studying its impact on the gender gap in maths in the formative primary school years. We exploit the distinctive features of the Irish education system, where single-sex schools are common and part of the general education system, rather than being restricted to higher socio-economic groups.

3. Description of Irish primary school system

Primary schooling in Ireland begins when children are aged four or five and lasts for eight years. The vast majority of primary schools in Ireland are state-aided parish schools, having been established under church patronage with the state giving explicit recognition to their denominational character. 92% of primary school children are educated at Catholic denominational schools, with the remainder attending non-Catholic denominational schools and multi-denominational schools. In some schools, called *gaelscoileanna*, instruction is carried out through the Irish language. These are typically also Catholic schools, but exist in addition to the local parish school, rather than replacing them. There are very few fee-paying schools at primary level in Ireland; less than 3% of students attend fee-paying primary schools.

Of the Catholic schools, a sizeable minority are single-sex schools simply because they were established at a time when separate schooling for boys and girls was the norm. These single-sex schools, often originally set up by religious congregations, continue to exist mainly in older residential areas in cities and towns where the school-going population is large enough to sustain at least two separate schools. Department of Education figures for 2007, the year to which our data refers, indicate that 27% of girls and 25% of boys of primary-school age were educated in single-sex schools in Ireland. This is unusually high by international standards. For example, in the UK just two single-sex state primary schools exist, accounting for fewer than 600 pupils.[7] Ireland is therefore unusual both for the fact that a sizeable proportion of children attend single-sex schools, and because these single-sex schools are part of the general educational system.

Irish parents can choose which schools they want their children to attend. Schools are not obliged to admit all children who apply but most parish schools operate a 'catchment area', whereby they accept all children of school-going age living within a particular area; however, parents can choose to apply to a parish school outside their own parish. Anecdotal evidence suggests that most parents send their children to their local schools[8] and, as noted above, whether that school is single-sex or coeducational is a matter of historical accident.

There is limited empirical evidence on how Irish parents choose schools for their children. However, some studies provide descriptive analyses that may be relevant. Fox and Buchanan (2008) conduct a qualitative study of school choice in Ireland, using in-depth interviews with parents, teachers and principals in nine Irish primary schools and focussing on the choice between local Catholic schools, *gaelscoileanna* and multi-denominational schools. Strikingly, they report that "[i]n not one single instance did parents speak to the academic reputation of schools in their neighborhoods as a factor in making their initial choice" (p. 12). Instead, parents emphasise convenience, class sizes and the school ethos in making their choices. The authors note in their conclusions that although schools differ in ethos, because they are all nationally funded, follow the same national curriculum, and are staffed by nationally certified teachers, "...[t]he generally accepted view in Ireland is that this ... results in schools of essentially equal quality" (p. 21).

O'Mahony (2008), in a survey of parents of children attending Irish Catholic primary schools, explores the factors that are important in choosing schools for their children. The author gave us access to this survey data, which allows us to investigate any differences between the parents of children attending single-sex schools and those attending coeducational schools, an issue that was not of central concern in the original report.[9]

Table 1 summarises the determinants of school choice available in the survey, distinguishing between parents of children attending single-sex and coeducational schools. Overall, 76% of the 443 parents surveyed report that they perceive the school that their child attends as their 'local school'. Importantly, this figure does not vary significantly across school types: 74% of single-sex school parents and 77% of coeducational school parents regard the school as their local one. As might be expected, the survey also shows that distance to school is important for most parents (59.0%), and again, there is no significant difference between parents according to whether their children attend single-sex or coeducational schools. As regards school choice, the vast majority of parents (89.1%) send their children to their school of first choice, again with no significant difference across school types. Moreover, consciously choosing schools along social class lines does not appear to be common: just 25.3% of the parents said that it was important that parents with a similar standard of living sent their children to the school, and this figure

[7] *Source*: Ofsted Inspection Reports for Winterbourne Junior Boys' school and Winterbourne Junior Girls' school.

[8] See for example the 'Moving to Ireland' website (http://www.movetoireland.com/movepag/schover.htm), which offers advice to people thinking of moving to Ireland. In response to queries about the quality of primary schools, they note "They're good. I can't think of a local elementary school within a 20 mile radius where I'd have any worries about sending my own kids. They are uniformly well thought of. [Parents] might dislike parts of the curriculum or the occasional teacher, but all speak well of their local school."

[9] Children living in Northern Ireland are excluded from the sample we use.

Table 1
Determinants of primary school choice.

	All	School type	
		Single-sex	Coeducational
School is 'local' school	.760	.743	.770
	(.428)	(.439)	(.422)
Distance to school (miles)	2.190	2.261	2.147
	(2.132)	(2.513)	(1.874)
Distance to school important	.590	.598	.585
	(.492)	(.492)	(.494)
School was first choice	.891	.874	.901
	(.312)	(.333)	(.299)
Parents with similar standard of living important	.253	.225	.271
	(.435)	(.419)	(.445)
School is single-sex	.381	1	0
	(.486)		
Single-sex important	–	.201	–
		(.402)	

Source: Survey of Catholic School Parents (O'Mahony, 2008).
Standard errors in parentheses.

does not vary significantly by school type (22.5% for single-sex school parents compared to 27.1% for coeducational school parents). Finally, parents were also asked whether the fact that the school their child attended was single-sex was important in choosing the school, and only 20% of the 169 parents of children attending such schools said that it was, thus reinforcing the notion that parents are not selecting schools on the basis of gender composition.

These descriptive analyses provide support for the view that whether a child attends a single-sex or coeducational school is effectively random; there is no evidence that parents believe that there are systematic differences in school quality between single-sex and coeducational schools. We exploit this feature of school choice when identifying the impact of single-sex schooling on the gender maths gap in Ireland.

4. Data

The Growing up in Ireland (GUI) survey tracks the development of a cohort of children born between November 1997 and October 1998. The data used for our analysis are from the first wave of interviews, which were carried out between August 2007 and May 2008.[10] The survey sample was generated through the national primary school system. 910 randomly selected schools participated in the study. Information was collected from the children, their parents, their teachers, the school principals, and their childminders (where relevant). Some administrative data were also provided with the Research Micro-File, including information on whether the school is co-educational or single-sex.

As part of the survey, each child took reading and maths tests, which were administered by GUI fieldworkers at the child's school. These tests, known nationally as 'Drumcondra' tests, have been used for many years in Irish schools and are linked to the national curriculum. The versions of the test used in our analysis were developed specifically for the GUI and had not been used or seen by schools, teachers or students prior to their use in the GUI. Since nine year old children in Ireland are distributed across three year grades in primary school (2nd, 3rd and 4th class) three different versions of the test (Levels 2, 3 and 4) were administered in the schools. The majority of children were in 3rd class and so took the Level 3 test. In this test children were given 35 min to complete a set of 25 questions. The questions tested a child's skills in problem solving, reasoning, and mathematical understanding. Questions were framed in a mixture of short answer and multiple choice formats and children were awarded one mark for each correct answer, giving a raw score range of 0–25.[11]

The Drumcondra test results, in the form of logit scores, provide the outcome variables of interest for this analysis. These logit scores are a transformation of the original raw test scores and are constructed using the principles of Item Response Theory (IRT) (Lord, 1952, 1980; Rasch, 1960). A key feature of the IRT approach is that it provides comparable estimates of ability even in situations where the test differs across individuals. The use of the logit scores therefore takes account of the fact that the children in GUI sat different levels of the test depending on which grade they were in.

In our analysis, we exclude children in 'special schools', which provide education to children whose needs cannot be accommodated in mainstream schools as well as children in private schools; together these account for 1.3% of children in the sample. This gives a base sample of 3674 girls and 3442 boys. In later sections we include a range of additional control variables, some of which are missing for some of the sample. The inclusion of these additional controls reduces the number of girls and boys to 3456 and 3220 respectively.

The administrative data provided with the survey allow us to distinguish among three school types – boys-only, girls-only and coeducational. 11% of schools in the sample are boys-only, 12% are girls-only and 77% are coeducational. Looking at student numbers we find that 28% of girls and 29% of boys in the sample are in single-sex schools.

Table 2 provides summary statistics by gender and school type for key variables. Looking first at the location indicators, it is clear that the geographical pattern of single-sex schools across the country reflects the historical evolution of the primary school system in Ireland, described in Section 3. Single-sex schools are significantly more likely to be attended by children living in urban[12] areas than coeducational schools are: 67% of girls attending single-sex schools live in urban areas, compared to 34% of girls attending coeducational schools, while the corresponding figures for boys are 62% and 35%. The regional variables included confirm this pattern – those areas where attendance at single-sex schools is particularly prevalent include Dublin, which is the largest city, and the South-West, which includes Cork, the second largest city.

Other variables included in the table are notable for the fact that differences according to attendance at single-sex

[10] Only the first wave is available to date.

[11] More details on the nature of the Drumcondra Tests used in the GUI analysis, including sample questions, are given in Murray et al. (2011).

[12] 'Urban' indicates living in a town of more than 10,000 people.

Table 2
Summary statistics.

	All		Single Sex		Coeducational	
	Boys	Girls	Boys	Girls	Boys	Girls
Urban	.427	.435	.618	.666	.349	.344
	(.495)	(.496)	(.486)	(.472)	(.477)	(.475)
Region: Border	.127	.131	.055	.082	.157	.150
	(.333)	(.337)	(.229)	(.274)	(.364)	(.357)
Region: Dublin	.221	.218	.290	.291	.192	.190
	(.415)	(.413)	(.454)	(.454)	(.394)	(.392)
Region: Mid-East	.114	.123	.115	.106	.113	.129
	(.317)	(.328)	(.320)	(.308)	(.317)	(.335)
Region: Midlands	.061	.066	.056	.050	.062	.072
	(.238)	(.248)	(.231)	(.217)	(.242)	(.259)
Region: Mid-West	.116	.101	.093	.053	.125	.120
	(.320)	(.301)	(.290)	(.223)	(.331)	(.325)
Region: South-East	.095	.116	.114	.168	.088	.095
	(.294)	(.320)	(.318)	(.374)	(.283)	(.293)
Region: South-West	.164	.149	.243	.187	.131	.134
	(.370)	(.356)	(.429)	(.390)	(.338)	(.341)
Region: West	.103	.097	.033	.065	.132	.110
	(.304)	(.296)	(.180)	(.246)	(.338)	(.312)
Maths score	−.704	−.805	−.671	−.838	−.718	−.792
	(.970)	(.880)	(.982)	(.898)	(.965)	(.872)
Maths top quartile	.283	.215	.306	.207	.273	.219
	(.450)	(.411)	(.461)	(.406)	(.446)	(.413)
Reading top quartile	.253	.240	.266	.237	.248	.242
	(.435)	(.427)	(.442)	(.426)	(.432)	(.428)
Single sex school	.292	.284	1	1	0	0
	(.455)	(.451)				
Parental education: degree	.276	.225	.251	.199	.287	.236
	(.447)	(.418)	(.434)	(.400)	(.452)	(.424)
Family income (000's)	50.52	48.52	51.44	46.21	50.13	49.44
	(31.00)	(30.94)	(33.82)	(27.92)	(29.75)	(32.02)
Father absent	.197	.223	.212	.299	.190	.193
	(.397)	(.416)	(.409)	(.458)	(.393)	(.395)
Teacher male	.184	.102	.267	.009	.150	.140
	(.388)	(.303)	(.443)	(.093)	(.357)	(.347)
Mother maths occupation	.052	.043	.058	.041	.050	.044
	(.223)	(.203)	(.233)	(.198)	(.218)	(.205)
Housework division unfair	.400	.428	.406	.449	.398	.420
	(.490)	(.495)	(.491)	(.498)	(.490)	(.494)
DEIS school	.179	.186	.225	.287	.161	.146
	(.384)	(.389)	(.417)	(.453)	(.367)	(.353)
Gaelscoil	.061	.058	0	0	.086	.080
	(.239)	(.233)			(.280)	(.272)
Multi-denominational School	.014	.010	0	0	.020	.014
	(.118)	(.100)			(.140)	(.118)
Child limited by language	.036	.025	.037	.027	.036	.024
	(.187)	(.156)	(.190)	(.162)	(.186)	(.154)
Mother's age	39.42	38.93	39.01	38.36	39.58	39.16
	(5.70)	(5.93)	(5.72)	(6.27)	(5.68)	(5.77)
Principal carer catholic	.861	.854	.875	.873	.856	.846
	(.346)	(.353)	(.330)	(.333)	(.351)	(.361)
Mixed-gender siblings	.629	.644	.617	.571	.633	.673
	(.483)	(.479)	(.486)	(.495)	(.482)	(.469)
Family size	2.88	2.86	2.82	2.57	2.90	2.98
	(1.18)	(1.23)	(1.16)	(1.12)	(1.19)	(1.25)
Class size	26.56	26.40	26.72	26.05	26.49	26.54
	(5.05)	(5.14)	(4.61)	(5.09)	(5.21)	(5.16)
Local area: % unemployed	4.29	4.47	4.52	4.92	4.20	4.29
	(2.44)	(2.43)	(2.88)	(2.53)	(2.23)	(2.37)
Maximum *N*	3442	3674	1009	1047	2433	2627

Standard errors in parentheses.

or coeducational schools are small. The variable means confirm that single-sex schools are not elitist in Ireland, unlike their counterparts in other countries. Exceptions to the similarity of the sample means across school types include whether the teacher is male (significantly more likely in boys-only schools and significantly less likely in girls-only schools) and whether the school is a *gaelscoil* or multi-denominational (there are no single-sex schools of this type in the sample, reflecting the fact that these school types have emerged only in recent years). However, across socio-economic indicators, the children attending single-sex schools come from similar backgrounds to those

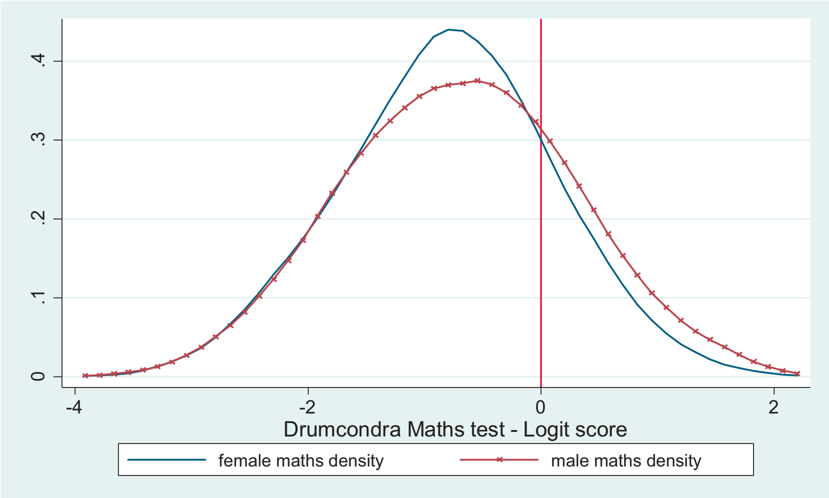

Fig. 1. Distribution of maths scores by gender.

attending coeducational schools. Where differences do arise – for example, in the absence of fathers (for which the difference between single-sex and coeducational schools is significant for girls) and the school's designation as a disadvantaged (DEIS) school[13] – they indicate that single-sex schools have *less* positive socio-economic characteristics. These differences are controlled for in our analysis.

5. Results on single-sex education and the gender gap

5.1. Establishing the gender gap in maths

Fig. 1 provides a Kernel density estimate of the distribution of maths scores for the boys and girls in the sample. A Kolmogorov–Smirnov test rejects the hypothesis that the two distributions are equal at the 1% significance level. Futhermore we can see that the gender difference in mean maths scores reported in Table 2 is driven by the fact that boys are outperforming girls at the top of the distribution. This finding is consistent with the literature reviewed by Halpern et al. (2007), as well as recent work by Ellison and Swanson (2010) who examine at the gender gap for very high achievers in maths in the US.

The vertical lines in Fig. 1 show the cut-off points for the top quartile of the maths distributions in our sample; 29% of males are in the top quartile compared to 22% of females. Since this is the point where the difference in maths scores between boys and girls becomes most pronounced, we focus on the top quartile in our analysis.[14]

5.2. The impact of single-sex schooling on the gender gap in maths

To examine the extent to which single-sex schooling affects the gender gap in maths, we begin by plotting the distributions of maths scores by gender and school type. The distributions in Fig. 2 are for boys attending coeducational schools, boys attending boys-only schools, girls attending coeducational schools and girls attending girls-only schools. The first point to note is that the differences occur primarily amongst those with high maths scores. Secondly, the ordering of the distributions indicates that boys perform better than girls, irrespective of school type, although the gap is most pronounced for boys attending single-sex schools. Finally, there appears to be little difference in the performance of girls at coeducational and single-sex schools.

To analyse these differences in more detail, we estimate probit models separately for males and females:

$$Q_{gi}^{75} = I(\alpha_g + \beta_g \text{Single-Sex}_i + \mathbf{X}i\boldsymbol{\gamma}_g + \varepsilon_i > 0),$$
$$\text{where } g = \text{male, female} \qquad (1)$$

Q_{gi}^{75} is a binary variable equal to one if the individual has a maths score in the top quartile and zero otherwise, Single-Sex is a dummy variable indicating the gender composition of the child's school, **X** is a vector of additional control variables, $I(.)$ is the indicator function and $\varepsilon \sim N(0, 1)$. We first estimate a model controlling only for the type of school attended. The parameter estimates from this model are presented in Columns 2 and 4 of Table 3. If single-sex schooling reduces the gender gap (as speculated by Fryer & Levitt, 2010), we expect girls to perform better in single-sex schools than in coeducational schools, and the girls' single-sex premium to be bigger than any boys' single-sex premium. In fact, the effect of attending an all-boys school is positive and the marginal effect for girls-only is negative, though neither are statistically significant in this specification.

To allow for possible differences in the observable characteristics of those attending single-sex and coeducational schools, we control for socio-economic factors in estimating the effect of single-sex schooling. To do this, we extend the previous model to include some additional explanatory variables. We control for parental education,

[13] The DEIS designation, which is assigned to schools on the basis of their socio-economic mix, is explained in greater detail in Section 5.

[14] A similar cut-off was used by Pope and Sydnor (2010) when looking at gender differences in test scores in the US.

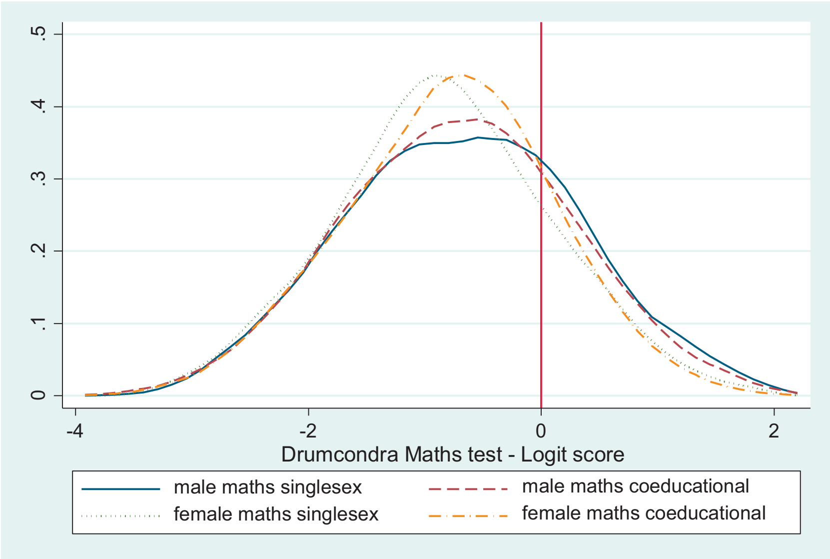

Fig. 2. Distribution of maths score by gender and school type.

Table 3
Marginal effects from probit models of scoring in top maths quartile.

	Boys		Girls	
Single sex school	.032	.046**	−.011	.014
	(.026)	(.023)	(.023)	(.022)
Parental education		.102**		.100**
		(.020)		(.018)
Family income		.001**		.0004*
		(.0003)		(.0002)
Father absent		−.048*		−.028
		(.029)		(.025)
DEIS school		−.056*		−.013
		(.031)		(.028)
Gaelscoil		.021		.035
		(.046)		(.036)
Multi denominational school		−.010		−.030
		(.078)		(.093)
Mother's age		.004**		.004**
		(.002)		(.002)
Local unemployment rate		−.014**		−.013**
		(.004)		(.004)
Child limited by language		−.188**		−.147**
		(.059)		(.052)
N	3442	3442	3674	3674

Standard errors in parentheses. Sampling weights are used in estimating the models; standard errors are adjusted for clustering at the school level.
* Significant at 10% level
** Significant at 5% level.

which is an indicator variable for whether either of the parents has a degree; family income, which is total household income measured in thousands of euro; mother's age measured in years; and the father being absent, which is a dummy variable taking the value one if the child's father does not live with the child.[15] We also include a dummy variable for whether the teacher assessed the child as having difficulties in class due to limited knowledge of the language of instruction. We also control for whether the school is a *gaelscoil* or a multi-denominational school. Since there are no single-sex *gaelscoileanna* or multi-denominational schools in our sample, it may be important to control for these school types in order to avoid confounding effects.

Finally, we use two variables to control for the socio-economic mix of the local area. The Department of Education classifies schools according to their socio-economic mix using a standardised system. The DEIS (Delivering Equality of opportunity In Schools) classification system is based on a range of factors such as parental unemployment, family type and size, and family income. Schools with a designated DEIS status receive extra resources. Approximately 20% of primary schools in Ireland are designated as DEIS schools.[16] As an additional control for the impact of the socio-economic environment, we also included the unemployment rate in the area in which the child lives, using Census Small Area Statistics provided with the data.

The marginal effects for this model are given in columns 3 and 5 of Table 3. Looking at the additional variables, we find that many are significant in determining whether a child's score is in the top quartile of the maths distribution, and all have the expected sign. The effect of single-sex schooling is now significant for boys at the 5% level, although it remains insignificant for girls.

Our main focus in this paper is the extent to which single-sex schooling reduces the gender gap in maths. However, this is not immediately obvious from the probit results reported in Table 3. To examine this we use the

[15] This includes cases where the child's father is deceased or living temporarily away from home.

[16] The Department of Education also provides a finer DEIS classification. Use of this finer classification made almost no difference to our final results.

parameter estimates to recover the implied probabilities of scoring in the top quartile for each of the four groups classified by gender and school type. Given the probit specification, the probability of scoring in the top quartile is

$$Pr(Q_{gi}^{75} = 1) = \Phi(\alpha_g + \beta_g \text{Single-Sex}_i + \mathbf{X}_i \boldsymbol{\gamma}_g),$$

where Φ is the standard normal cdf. To compare the probabilities of top quartile performance across genders and school types, it is necessary to equalise the values of the control variables across each group. In this way the differences in estimated probabilities take account of any underlying differences in observable characteristics across the groups. The difference between the probability of scoring in the top quartile for boys attending a single sex school and girls attending a single sex school is therefore given by

$$\Delta_{\text{mf:single-sex}} = \Phi(\alpha_m + \beta_m + \mathbf{X}^* \boldsymbol{\gamma}_m) - \Phi(\alpha_f + \beta_f + \mathbf{X}^* \boldsymbol{\gamma}_f)$$

and the difference among children attending a coeducational school is

$$\Delta_{\text{mf:coeducational}} = \Phi(\alpha_m + \mathbf{X}^* \boldsymbol{\gamma}_m) - \Phi(\alpha_f + \mathbf{X}^* \boldsymbol{\gamma}_f).$$

Of primary concern in this paper is the extent to which $\Delta_{\text{mf:single-sex}} < \Delta_{\text{mf:coeducational}}$.

The key findings are presented in Table 4. The first column simply reports the raw difference in predicted probabilities of high maths scores by gender with no controls. The results confirm that the gender gap is statistically significant; being male increases the probability of being in the top quartile by 6.7 percentage points. Column 2 reports the implied probabilities from a model including only a dummy for single-sex school, while Column 3 show the results from the model including the additional variables. To construct these probabilities we have to specify values for the control variables. The probabilities reported are for children with mean family income, mean mother's age and mean local unemployment rate, and where at least one of the parents has a degree, where the father is present, where the child does not have language limitations and where the school is not a *gaelscoil*, not multi-denominational and not a DEIS school. Since the results in Columns 2 and 3 are very similar, we focus on the results in Column 3.

When we control for these factors we see that boys in single-sex schools are 4.8 percentage points more likely to score in the top quartile than boys in co-educational schools, an effect that is significant at the 5% level. There is no corresponding effect of girls' single-sex education. As a result, single-sex schooling increases the gender gap; boys in co-educational schools are 5.9 percentage points more likely than girls to be in the top quartile in maths, whereas the gap is 9.4 percentage points for children in single-sex schools.

Although not the focus of this paper, it is worth noting that the impact of single-sex schools on reading scores is similar to the effect on maths scores. Boys attending single-sex schools are more likely to score in the top quartile than boys in co-educational schools. As before, there is no significant effect of single-sex education for girls. In contrast to the results for maths however, there is no gender gap in reading scores in either single-sex or coeducational schools. These results are reported in Tables A1 and A2 of the Appendix.

Table 4
Estimated probabilities of being in top maths quartile.

	Probit, no controls	Probit, control for school type only	Probit, socio-economic controls[a]
Male (1)	.282		
	(.011)		
Female (2)	.215		
	(.010)		
Gender gap (1) – (2)	.067		
	(.013)		
Male, coeducational school (3)		.273	.251
		(.012)	(.016)
Male single-sex school (4)		.306	.299
		(.024)	(.025)
Male school type gap (4) – (3)		.033	.048
		(.027)	(.025)
Female, coeducational school (5)		.219	.192
		(.011)	(.013)
Female, single-sex school (6)		.207	.205
		(.020)	(.022)
Female school type gap (6) – (5)		−.011	.013
		(.023)	(.023)
$\Delta_{\text{mf:coeducational}}$ (3) – (5)		.054	.059
		(.014)	(.019)
$\Delta_{\text{mf:single-sex}}$ (4) – (6)		.098	.094
		(.032)	(.032)

Bootstrap standard errors in parentheses.
[a] Probabilities are for children with mean income, mean mother age, mean local unemployment, no parental degree, father present, in non-DEIS, non-*gaelscoil*, non multi-denominational school and with no difficulty in the language of instruction.

5.3. Non-random selection and the single-sex effect

In the above analysis we have shown the existence of a significant gender gap in mathematical achievement at the top of the test distribution. We find no evidence that single-sex schools reduce this gap; if anything the gender gap is wider among students attending single-sex schools than among those in co-educational schools. Given the nature of the Irish educational system we believe that our results on single-sex schools are less likely to be affected by self-selection bias than is the case in other countries. In the final part of this section we examine the robustness of our findings to this exogeneity assumption.

The main concern in this literature is that populations attending single-sex schools may differ in unobservable ways from those attending coeducational schools. In this case, the earlier estimates cannot be interpreted as causal effects. In addition, if the decision to send children to single-sex schools differs for boys and girls then the gender comparison is compromised. A number of approaches have been suggested for dealing with endogenous selection into single-sex schools, which we consider below.

Table 5
Marginal effects from probit models of scoring in top maths quartile, controlling for birthweight.

	Boys	Girls
Single sex school	.044*	.018
	(.023)	(.022)
Parental education	.102**	.096**
	(.020)	(.018)
Family income	.001**	.0004**
	(.0003)	(.0002)
Father absent	−.0299	−.035
	(.029)	(.026)
DEIS school	−.051	−.018
	(.031)	(.028)
Gaelscoil	.026	.038
	(.047)	(.037)
Multi denominational school	−.005	−.024
	(.078)	(.095)
Mother's age	.004**	.005**
	(.002)	(.002)
Local unemployment rate	−.013**	−.013**
	(.004)	(.004)
Child limited by language	−.209**	−.131**
	(.061)	(.053)
Child's birthweight (kg)	.044**	.006
	(.014)	(.013)
N	3404	3629

Standard errors in parentheses. Sampling weights are used in estimating the models; standard errors are adjusted for clustering at the school level.
* Significant at 10% level.
** Significant at 5% level.

5.3.1. Omitted cognitive ability

Since there may be differences in the ability of children attending single-sex and coeducational schools, one approach to accounting for selection is to use a measure of the child's pre-school cognitive ability as an additional control variable. Ideally, we would like to have pre-school test scores for each child, but these are not available in our data. Instead we use information on the child's birthweight. There is a growing literature on the causal effect of birthweight on later outcomes (Black, Devereux, & Salvanes, 2007; Currie & Hyson, 1999; McGovern, 2013) and much of this research finds evidence linking low birthweight to cognitive deficits and behavioural problems among young children. Birthweight has also been used directly as an input when estimating the production function of cognitive achievement for young children (Todd & Wolpin, 2007).[17]

The results when birthweight is included as a control are given in Table 5.[18] The results suggest that birthweight has a significant positive effect on the maths achievement of boys at age nine, but has no effect for girls. A 1 kg decrease in birthweight decreases the likelihood of a boy scoring in the top quartile of the maths distribution by 4.4 percentage points. While this result is of interest in itself, what is important for our analysis is the impact of controlling for birthweight on the estimated gender gap in maths. Using the probit point estimates, we calculate the probability of scoring in the top quartile for children of average birthweight.[19] The gender gap among children attending single-sex schools is 7.9 percentage points and 5.2 percentage points among children at coeducational schools. Controlling for unobserved ability in this way does not alter our earlier conclusion; there is no evidence that the gender maths gap is smaller in single-sex schools.

5.3.2. Bivariate probit model

Our second approach to accounting for selection uses a bivariate probit model to isolate the causal effect of single-sex schooling. The model is given by

$$Q_{gi}^{75} = I(\alpha_g + \beta_g \text{Single-Sex}_i + \mathbf{X}_i\boldsymbol{\gamma}_g + \varepsilon_i > 0)$$
$$\text{Single-Sex}_{gi} = I(\eta_g + \mathbf{Z}_i\boldsymbol{\delta}_g + \upsilon_i > 0),$$

where $I(.)$ is the indicator function and $(\varepsilon, \nu) \sim N(0, 0, 1, 1, \rho)$. To aid identification, $\mathbf{Z}$ needs to contain at least one variable that is not in $\mathbf{X}$. In this analysis we use two instruments for school choice: urban location and the child's principal carer being Catholic.

The urban instrument attempts to capture geographical proximity to a single-sex school. In our data 43% of students living in an urban area attend a single-sex school compared to only 18% of students living in rural areas. Our identifying assumption is that, conditional on a range of family characteristics and the socio-economic mix of the school, the rural–urban indicator does not directly affect the mathematical performance of the child.[20]

The instrument based on the main carer (usually the mother) being Catholic exploits the fact that the majority of single-sex schools were founded by religious orders, and so religion may be more important to the ethos of these schools. As noted earlier, Fox and Buchanan (2008) found that school ethos was one of the factors mentioned by parents when making their school choices. We therefore construct a binary variable taking the value one if the main carer describes herself as Roman Catholic and zero otherwise. 86% percent of respondents were categorised as Catholic. Our identifying assumption in this case is that the religious denomination of the main carer has no effect on academic performance, conditional on the controls used in our analysis.[21] While the identifying restrictions on geographical proximity and religious denomination seem plausible, we first formally test the validity of our identifying assumptions making use of the fact that our model is over-identified.

Since no formal over-identification test exists for the bivariate probit that we are aware of, we follow Evans and Schwab (1995) and use a procedure based on an

[17] Including birthweight as an additional control is in keeping with the 'value added' approach to addressing endogeneity (Todd & Wolpin, 2003).
[18] In this and subsequent models where the inclusion of additional control variable leads to a slightly smaller sample, we have re-estimated our base model on the reduced sample; in all cases, this did not affect our main results.
[19] We set the other variables at the values discussed in Section 5.2.
[20] Some support for this assertion can be found in recent work by the OECD (2011).
[21] Evans and Schwab (1995) use a similar instrument in their analysis of public and Catholic schools in the U.S. and also argue that, once you control for other background characteristics, being from a Catholic family is not an important determinant of most economic outcomes.

Table 6
Marginal Effects from Bivariate Probit Models of Scoring in Top Maths Quartile and Choice of School Type.

	Boys		Girls	
	School type	Top quartile maths score	School type	Top quartile maths score
Single sex school		.273**		.133*
		(.099)		(.077)
Parental education	−.044*	.110**	−.013	.111**
	(.024)	(.022)	(.022)	(.021)
Family income	.000	.0008**	−.0002	.0004
	(.0003)	(.0003)	(.0003)	(.0003)
Father absent	−.015	−.048*	−.012	−.039
	(.028)	(.028)	(.038)	(.024)
DEIS school	.053	−.066**	.107	−.030
	(.072)	(.033)	(.071)	(.031)
Gaelscoil		.008		.029
		(.043)		(.041)
Multi denominational school		−.028		−.038
		(.084)		(.099)
Mother's age	−.0035*	.004**	−.0003	.005**
	(.002)	(.002)	(.0019)	(.002)
Local unemployment rate	−.004	−.015**	.0005	−.014**
	(.008)	(.005)	(.007)	(.004)
Child limited by language	−.020	−.167**	.008	−.138 **
	(.056)	(.041)	(.059)	(.033)
Urban	.222**		.243**	
	(.052)		(.044)	
Mother catholic	.038		.075**	
	(.031)		(.030)	
ρ		−.402**		−.237
		(.166)		(.147)
N	3442	3442	3674	3674

The reported marginal effects are derived from the marginal probabilities of test scores and single-sex education of the bivariate model. Sampling weights are used in estimating the model, and standard errors are adjusted for clustering at the school level.
* Significant at 10% level.
** Significant at 5% level.

Instrumental Variables estimation of a linear probability specification of our model.[22] In particular, we exploit the fact that we have two instruments for one endogenous variable to test the validity of the over-identifying restrictions using Hansen's J-statistic. The tests were run separately for boys and girls and the p-values of the over-identifying test were .32 and .74 for boys and girls respectively. Based on these results, we are clearly unable to reject the validity of our identifying assumptions. In addition, the F-statistic for the null hypothesis that the two instruments have no effect on school choice is 143 for boys and 164 for girls, well in excess of the value of 10 normally considered appropriate for instrumental relevance. These findings justify our use of these instruments in a bivariate probit model.

Table 6 presents the results of the model. The first and third columns show the results for the school choice decision for boys and girls respectively, while the second and fourth columns give the results for maths performance. The impact of geography on school choice is as expected, with boys and girls living in urban areas significantly more likely to attend a single-sex school. We also find that mothers who are Catholic are significantly more likely to send their girls to single-sex schools, although this effect is not significant for boys. The other estimates in the model for school choice support our view about the absence of positive selection into single-sex schools. For girls, neither the parental background characteristics, nor ρ, representing unobservables, are significant. For boys, only two variables – parental education and mother's age – are marginally significant. Moreover, to the extent that selection does occur for boys, it is in the opposite direction to that usually encountered in studies of single-sex education. The coefficients on parental education and mother's age are negative, as is the estimated value of ρ from the bivariate probit. This supports our assertion that Irish single-sex schools are not elitist.

The coefficients from the bivariate probit models can be used to estimate the marginal probabilities of scoring in the top quartile of the maths distribution for each of our gender-school composition groups, as before. The estimated gender gap in maths achievement for children attending coeducational schools is now smaller than that estimated using the univariate probit model, at 3.3 percentage points, and is not significantly different from zero. However, the gap among children at single-sex schools is larger than before: the point estimate is now 18 percentage points, although the standard error is also correspondingly larger, so that the effect is significantly

[22] The instrumental variable estimation of the linear probability model as an alternative to the bivariate probit approach was proposed by Angrist (1991). We have estimated our model both ways and obtained very similar marginal effects. Note that this is in contrast to Altonji, Elder, and Taber (2005a) who used differences between the 2SLS and bivariate probit estimates of the effect of Catholic schools in the US to cast doubt on the power of religion as an instrument in their analysis.

different from zero only at the 10% level. Even taking into account any possible selection effects, we fail to find any evidence that the gender maths gap is smaller in single-sex schools.

5.3.3. Final robustness checks

In this section we consider two final robustness checks of our analysis. Altonji, Elder, and Taber (2005b) provide a useful framework for gauging the likely impact of selection on estimated causal effects. The basic premise of their approach is to use the magnitude of selection on observables as a benchmark for the likely impact of selection on unobservables. A measure of the impact of selection on observables may be constructed by first re-estimating Eq. (1) under the restriction that $\beta_g = 0$ and calculating an index $\mathbf{X}_i\boldsymbol{\gamma}_g$. The amount of selection on observables is then measured by $E[\mathbf{X}\boldsymbol{\gamma}_g|\text{Single-Sex} = 1] - E[\mathbf{X}\boldsymbol{\gamma}_g|\text{Single-Sex} = 0]$. If the selection on unobservables needed to explain the estimated causal effect is substantially larger than the estimated selection on observables, then they argue that this casts doubt on selection as a possible explanation for the estimated effect.

The estimated selection on observables in our model is small and *negative*. Therefore in order for selection on unobservables to account for the fact that boys in single-sex schools do *better* than boys in coeducational schools, it would have to be the case that selection on unobservables operates in the opposite direction to selection on observables. This does not seem plausible and indeed is not consistent with the results from the bivariate probit analysis.

As an alternative way of assessing the role of selection on unobservables, Altonji et al. (2005b) also propose constraining the selection parameter in the bivariate probit model, ρ, to different values and then examining the impact on the remaining parameters. Following this approach, we consider the degree of selection on unobservables needed to eliminate the gender gap in maths for children attending single-sex schools. To do this, we keep the selection in the boys' equation fixed at its estimated value and then examine the value of ρ that would be required in the girls' equation in order for single-sex schooling to eliminate the gender gap. Experimenting with alternative values of ρ we find that the required selection on unobservables for girls would have to be more than twice as large as that actually estimated (ρ equal to $-.57$ compared to an estimated ρ equal to $-.24$). This is a substantial change and is further evidence against the role of selection on unobservables in explaining our findings.

Table 7 summarises our findings on the gender gap in maths by school type across the range of models we have estimated in this section. The results are consistent across each specification: there is no evidence that single-sex schools reduce the gender gap. The results indicate that, if anything, the gap in maths achievement is larger, not smaller for children attending single-sex schools.

Table 7
Summary of gender differentials in probability of scoring in top maths quartile.

Model, controls included	Raw gender gap	$\Delta_{\text{mf:coeducational}}$	$\Delta_{\text{mf:single-sex}}$
Probit, no controls	.067** (.013)		
Probit, school type only		.054** (.014)	.098** (.032)
Probit, socio-economic controls, excluding birthweight		.059** (.019)	.094** (.032)
Probit, socio-economic controls, including birthweight		.052** (.019)	.079** (.032)
Bivariate probit		.033 (.025)	.180* (.109)

Bootstrap standard errors in parentheses

6. Other explanations of the gender gap in maths

Since we find no evidence that single-sex schooling reduces the gender gap in maths, in the final section of the paper we consider alternative explanations of the gender gap, including teacher gender, mother's occupation, attitudes to equality and the gender mix within the household.

Firstly, there is evidence that teacher gender may affect students' outcomes. The evidence on the direction of the teacher gender effect is mixed. Dee (2007) finds that both boys and girls are adversely affected when taught by a female maths teacher, whereas Carrell, Page, and West (2010) find that girls, particularly those at the top of the distribution, perform better in maths and science when taught by a female professor, with no corresponding effect for boys. Winters, Haight, Swaim, and Pickering (2013) found no significant relationship between teacher gender and the achievement of either male or female students in the early years of schooling but a statistically significant relationship between teacher gender and student achievement in middle and high school grades. Finally, Cho (2012) found little evidence of any teacher gender effect in a recent study of academic achievement in fifteen OECD countries. Since in our data, boys are more likely to be taught by male teachers (18% of boys have a male teacher compared to only 11% of girls), then any teacher gender effect could in part explain the gender gap. To examine this, we include an additional control for the gender of the child's teacher.

Secondly, following Fryer and Levitt (2010), we include a control for whether the mother's occupation is maths-related.[23] This is intended to capture the possibility that

[23] Maths-related occupations are based on the 4 digit ISCO classification of occupations and are defined to include meteorologists, chemists, geologists and geophysicists, mathematicians and related professionals, statisticians, computer systems designers, analysts and programmers civil engineers, electrical engineers, electronics and telecommunications engineers, mechanical engineers, chemical engineers, accountants, economists, chemical and physical science technicians, civil engineering technicians, electrical engineering technicians, electronics and telecommunications engineering technicians, securities and finance dealers and brokers bookkeepers, statistical, mathematical and related associate professionals accounting and book-keeping clerks, statistical and finance clerks and mechanical engineering technicians. Where the mother is not currently working her most recent occupation is used.

the presence of a female role model in the family might counteract gender stereotyping, and so prevent maths being viewed as a male subject.

Thirdly, we include a control variable measuring whether or not the main care giver (almost always the mother) deems the division of housework between parents to be unfair. Guiso, Monter, Sapienza, and Zingales (2008), Fryer and Levitt (2010) and Pope and Sydnor (2010) argue that the gender gap is smaller in countries/states with a more gender-equal culture; the inclusion of this variable is intended to allow for the possibility of a similar effect between families.

Finally we also include a control variable for being in a mixed gender household. Just as girls' attitudes to maths may be affected by the presence of boys in coeducational schools, this may also be true if they grow up with brothers. Since the likelihood of growing up in a mixed gender household is higher in households with larger families, we also control for family size to assist with interpretation of the results.

The results of this extended analysis are given in Table 8. Looking at the coefficients on the new control variables, we find no evidence that the gender of the teacher or the division of housework within the family has any effect on maths performance for either boys or girls. We do find, however, that having a mother with a maths-related occupation has a significant positive effect on maths scores. This may be a causal effect, reflecting inherited innate ability or a more positive disposition to the subject. On the other hand, the variable may be picking up some measure of social class not captured by income or education. However, it is interesting that results from a separate analysis not reported here indicate that the mother having a maths occupation is insignificant when included as a control in a model of reading scores. This is not what we would expect if the mother having a maths occupation was simply proxying for social class. Finally, the results suggest that boys who grow up in a mixed sibling household tend to perform less well than boys in a single-sex household, while there is no corresponding effect for girls. It is interesting to note that the lower scores of boys in mixed-gender households is consistent with the poorer performance of boys in coeducational schools noted earlier.

While these individual coefficients are of interest in their own right, our main finding in this section is that the inclusion of these additional controls does not explain the gender gap. Using the probit point estimates, we again calculate the probability of scoring in the top quartile, as before setting the values of the variables at the mean for continuous variables and at the mode for dichotomous variables. Thus, in this case the probabilities are estimated for children with female teachers, with mixed gender siblings and a fair division of housework whose mothers do not have a maths occupation. The estimated male advantage in coeducational schools is 4.1 percentage points, which is significant at the 10% level; the male advantage in single-sex schools is 7.8 percentage points, which is significant at the 5% level. These additional variables do not explain the gender gap in maths in either type of school.

Table 8
Marginal effects from probit models of scoring in top maths quartile, extended model.

	Boys	Girls
Single-sex school	.033	−.005
	(.025)	(.024)
Parental education	.101**	.087**
	(.021)	(.018)
Family income	.001**	.0005**
	(.0003)	(.0002)
Father absent	−.051	−.031
	(.033)	(.028)
DEIS school	−.071**	−.025
	(.033)	(.030)
Gaelscoil	.012	.0200537
	(.049)	.0375571
Multi denominational school	.006	−.041
	(.080)	(.093)
Mother's age	.005**	.005**
	(.002)	(.002)
Local unemployment rate	−.010**	−.013**
	(.005)	(.004)
Child limited by language	−.149**	−.145**
	(.064)	(.056)
Teacher male	.029	−.002
	(.028)	(.026)
Mother has maths occupation	.161**	.114**
	(.044)	(.058)
Division of housework unfair	−.007	.017
	(.020)	(.017)
Mixed gender siblings	−.037*	.018
	(.021)	(.017)
Family size	.002	−.023**
	(.010)	(.007)
N	3220	3456

Standard errors in parentheses. Sampling weights are used in estimating the models; standard errors are adjusted for clustering at the school level.
* Significant at 10% level.
** Significant at 5% level.

7. Conclusion

In this paper, we examine the gender gap in maths scores using a representative sample of nine-year old Irish school children. In keeping with other studies, we find that boys perform better than girls, with the difference most pronounced at the top of the distribution. Examining the reasons for the gender maths gap is important because of the under-representation of women in STEM careers. Fryer and Levitt (2010) failed to find support for a range of explanations of the gender gap in their analysis. One alternative hypothesis tentatively suggested by these authors is that "mixed-gender classrooms are a necessary component for gender inequality to translate into poor female math performance". The distinctive characteristics of the Irish education system provide us with an opportunity to test this hypothesis at an important stage in children's development, and we find no evidence that the gender gap in maths is smaller in single-sex schools than in coeducational schools; if anything, it is larger. These findings are robust across a range of specifications, and support our view that the key results in the paper are not driven by selection effects.

While our analysis provides tentative evidence that boys in single-sex schools are more likely to achieve high maths scores than their counterparts in coeducational schools, we find little evidence of a similar effect for girls. Exploring the mechanisms underlying this finding is an interesting avenue for further research.

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Appendix

Tables A1 and A2.

Table A1
Estimated probabilities of being in top reading quartile.

	Probit, no controls	Probit, control for school type only	Probit, socio-economic controls[a]
Male (1)	.254		
	(.010)		
Female (2)	.240		
	(.010)		
Gender gap (1) – (2)	.014		
	(.014)		
Male, coeducational school (3)		.249	.205
		(.011)	(.013)
Male single-sex school (4)		.267	.243
		(.021)	(.023)
Male school type gap (4) – (3)		.018	.038
		(.024)	(.023)
Female, coeducational school (5)		.241	.198
		(.012)	(.013)
Female, single-sex school (6)		.237	.221
		(.019)	(.019)
Female school type gap (6) – (5)		–.005	.023
		(.022)	(.021)
$\Delta_{\text{mf:coeducational}}$ (3) – (5)		.008	.007
		(.015)	(.019)
$\Delta_{\text{mf:single-sex}}$ (4) – (6)		.030	.022
		(.029)	(.031)

Bootstrap standard errors in parentheses. Probabilities are for children with mean income, mean mother age, mean local unemployment, no parental degree, father present, in non-DEIS, non-*gaelscoil*, non multi-denominational school and with no difficulty in the language of instruction.

Table A2
Marginal effects from probit models of scoring in top reading quartile.

	Boys		Girls	
Single-sex school	.019	.038*	–.007	.023
	(.023)	(.023)	(.022)	(.021)
Parental education		.147**		.117**
		(.019)		(.019)
Family income		.0007**		.0005**
		(.0003)		(.0002)
Father absent		–.0008		–.002
		(.028)		(.027)
DEIS school		–.073**		.015
		(.032)		(.027)
Gaelscoil		.039		.083**
		(.038)		(.029)
Multi-denominational school		.064		.072
		(.056)		(.072)
Mother's age		.004**		.007**
		(.002)		(.002)
Local unemployment rate		–.006		–.013**
		(.004)		(.004)
Child limited by language		–.295**		–.580**
		(.072)		(.100)
N	3442		3674	

Standard errors in parentheses. Sampling weights are used in estimating the models; standard errors are adjusted for clustering at the school level.
* Significant at 10% level.
** Significant at 5% level.

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