UCD CENTRE FOR ECONOMIC RESEARCH

WORKING PAPER SERIES

2011

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WP11/13

July 2011

UCD SCHOOL OF ECONOMICS
UNIVERSITY COLLEGE DUBLIN
BELFIELD DUBLIN 4

Human Capital and the Quantity-Quality Trade-Off during the Demographic Transition: New Evidence from Ireland

Alan Fernihough*†

Abstract

In this article I measure the child quantity-quality relationship in 1911 Ireland. My analysis shows that sibship size had a strong impact on the probability of school enrollment in both Belfast and Dublin. However, the magnitude of the relationship varied considerably across different cohorts, most noticeably between the two cities. The existence of this relationship shows how the demographic transition played a vital role in the expansion of human capital and is highly consistent with the theoretical foundations of various long-run growth theories.

JEL Classifications: J10, N3, O10

Keywords: Quantity-Quality, Human Capital, Demographic Transition, Unified Growth Theory

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†I am grateful to Cormac Ó Gráda, Morgan Kelly, Neil Cummins, Paul Devereux and Oded Galor for helpful comments and advice.

1 Introduction

Unified growth theory argues that the child quantity-quality (Q-Q) trade-off was the key mechanism by which the progression from stagnation to modern economic growth was accelerated.[1] Before the 19th century, economic growth was slow and spasmodic, while fertility rates were persistently high. By the mid-20th century this situation (in the western world) had largely reversed. However, the demographic transition occurred prior to the advent of reliable data-series, and consequently, empirical evidence supporting the Q-Q mechanism during the transition is scarce. Furthermore, contemporary estimates, like Black et. al (2005) and Qian (2009), have yielded inconclusive results.

This paper is motivated by the apparent gap in empirical research evaluating the strength of the Q-Q trade-off during the demographic transition. This gap is understandable given the aforementioned dearth of available data. However, the recent digitization of the 1911 Irish census provides a rich source of individual level information for the entire Irish population, from which it is possible to measure the relationship linking fertility and child 'quality'. Using data for the Irish population introduces a number of caveats which need to be considered. It is widely accepted that Irish involvement in European fertility transition was limited (Ó Gráda, 1991). While I do not contest this view, I argue that there were two regions which exhibited a significant degree of purposeful marital restriction – Belfast and Dublin city.

To understand the extent of marital fertility control, I use the entire sample of Irish married women less than fifty years of age, and estimate a series of statistical models which respect the format of count data. The existing statistics on Irish marital fertility are aggregated at a county level and thus ignore Belfast, and the Dublin city environs.[2] My estimates represent the first attempt to compare marital fertility in Belfast and Dublin city with other Irish regions. The results are telling, as the level of marital fertility in Ireland's two major urban centers was dramatically lower when compared to the rest of Ireland.

Given that parents restricted child quantity, were household resources directed to-

[1]See Galor (2005) for an overview.
[2]Belfast rests between two counties: Antrim and Down.

wards child quality? In this article, preferences for child 'quality' are defined as the amount of extra resources a parent devotes to an offspring. Here, school enrollment provides an ideal proxy for 'quality' and permits me to model the effect of household fertility on so-called 'quality' in a reduced-form model. My reduced form Q-Q model tests whether having a larger number of siblings reduced the probability of school attendance for those aged 14-15 years. To estimate this model, I employ a number of statistical methodologies which address potential shortcomings in these data, including OLS, 2SLS, Multilevel and Probit.

I find strong evidence that the child Q-Q trade-off existed in Belfast. The marginal effect of increasing the number of siblings by one decreases the probability of remaining in school by about 2.7%. The relationship was weaker in Dublin, as the magnitude of the estimated Q-Q parameter is 1.0%. To understand how the trade-off varied between cohorts, I perform the model-based recursive partitioning scheme proposed by Hothorn et al. (2008). The results generated by this scheme are striking, and reveal a lot about the dynamics which encompassed the Q-Q trade-off in economic history. Firstly, the inter-city disparity cannot be reconciled by adjusting for social and economic groups. It seems that Belfast's modern industrial structure stimulated the Q-Q trade-off more than Dublin's proto-industrial landscape. Furthermore, I find that the relationship is weakest amongst the wealthiest in both cities, as the offspring of the elite went to school regardless of family size. There appears to be a wealth threshold, after which the Q-Q relationship vanishes.

The remainder of the paper is structured as follows. The following section provides the requisite overview of the existing literature concerning the Q-Q trade-off. Section 3 describes the historical setting. In this section, I pay particular attention to the demographic and educational aspects. The fourth section provides a brief overview of the data set and presents new estimates of marital fertility in Ireland. Section 5 presents the results testing the Q-Q trade-off, and section 6 provides an analysis exploring how the Q-Q mechanism varied across cohorts. Section 7 concludes.

2 Related Research

Trends in economic growth and fertility during the period 1850-1950 presented a puzzle to economic demographers. Classical theory, typically attributed to Malthus, assumes that children are a normal good, and thus the income elasticity of child demand is positive. However, the emergence of modern economic growth alongside the demographic transition was clearly at odds with this theory. To answer this puzzle, Becker (1960) proposed an extension to the classical economic model of fertility, and argued that parents made optimal (utility maximizing) child-rearing decisions across not one but two dimensions: quantity and quality. Quantity refers to the number of children, whereas the meaning of quality is less-precisely defined. For the purposes of this article, it is useful to think of child-resources devoted towards quality as being human-capital augmenting.

Since Becker's contribution, the theoretical framework of the Q-Q model has been expanded in various forms, such as Becker & Lewis (1973) and Becker & Tomes (1976). Indeed, several key contributions in the macroeconomic growth literature have cited this trade-off as the vital mechanism which fostered the transition to sustained economic growth. These citations include research by Galor & Weil (1999; 2000) and Galor & Moav (2002), who argued that an endogenous relationship between technological growth, the demand for human capital, and fertility emerged in the second phase of the Industrial Revolution. It is worth noting that a fundamental difference exists between the Becker-style Q-Q model and the more recent contributions (Galor, 2011). In Becker's model, increased levels in income stimulate a decline in fertility via a substitution effect. Alternatively, parental preferences are such that the income elasticity with respect to quality is higher than with respect to quantity for higher levels of income. This is not the mechanism proposed in Galor & Weil (1999; 2000) and Galor & Moav (2002), who argued that the future return on offspring human capital caused fertility to decline. In essence, technological growth stimulated an economic expansion which drove an increase in the demand for human capital, that in-turn decreased fertility via their Q-Q trade-off. Falling rates of fertility and consequent expansions in human capital reinforced technological and ultimately economic advances.

[FIGURE 1 ABOUT HERE]

Figure 1 presents a schematic diagram illustrating a simple Q-Q trade-off. Starting at point A, parents choose to have the number of children n_A and devote τ_A resources towards their quality. This consumption bundle maximizes utility, illustrated by the indifference curve U_A, subject to the budget constraint in the dotted line. Imagine two simultaneous changes which cause the budget line to change. Firstly, the real wage for women increases in a way which raises the relative price of children, since the opportunity cost of having and rearing children is time away from work. Secondly, the government introduces a new law abolishing school fees and thus reducing the relative price of quality. Envisaging these changes leads parents to reduce their fertility, to n_B, and increase the amount of resources devoted towards child development, τ_B, thus attaining a new level of utility U_B.

The close timing and sequential nature of the second phase of the Industrial Revolution, demographic transition and emergence of modern economic growth, presents a narrative consistent with the above hypothesis. Nevertheless, there is a deficit of hard empirical evidence supporting the Q-Q trade-off during the demographic transition. Furthermore, empirical studies which test the empirical validity of the Q-Q model using more contemporary data samples have found mixed results.

The relationship between sibship size and child 'quality' is endogenous. Consequently, researchers wishing to measure the Q-Q effect must employ empirical methodologies which account for endogeneity bias. Typically, this requires the use of relatively large samples of reliable data containing a number of potentially confounding variables and/or instrumental variables (IV). Thus, the lack of evidence on the Q-Q trade-off in historical demography is understandable considering the relative scarcity of suitable data meeting the aforementioned criteria. Recently, however, a few studies have sought to overcome this issue. Becker et al. (2010) present a case in point. Using aggregated regional data for mid-19th century Prussia, they found that areas with higher fertility also had lower levels of school enrollment. Interestingly, the authors also found evidence in favor of reverse-causality. In other words, areas where the level of education was higher also had lower levels of fertility. This result echoes the findings of Bleakley & Lange (2009), who use the eradication of hookworm disease in Southern US states as a form of natural experiment. Bleakley & Lange argued that the eradication of this disease reduced the cost of child 'quality', and the subsequent increase in education and decrease in fertility were consistent with the Q-Q framework and also unified growth theory.

Hatton & Martin (2010) use a unique individual level sample of British children in 1937-39 to measure the relationship between family size and height, which serves as a proxy for human capital (health). Their results are also consistent with the Q-Q hypothesis, as they show how family size was a key determinant of height. More recently, research from Van Bavel et. al (2011) shed light on the nature of the Q-Q trade-off during the demographic transition. In this study, the authors use life-course data from the Belgian city of Antwerp to measure the relationship between sibship size and socioeconomic mobility. The results of this study are inconclusive, since the Q-Q relationship seemed only to apply to the wealthy as a defensive strategy for maintaining status quo. For the poorest, family-size appeared to have no bearing on socioeconomic advancement of the child.

Several studies have attempted to measure the Q-Q relationship in a modern setting. Work by Rosenzweig & Wolpin (1980) represented the first empirical study that employed an IV strategy. In this, the authors argued that the occurrence of multiple births represented so-called 'exogenous' variation, and deemed this quasi-experimental methodology appropriate. Using a sample of Indian households, their findings supported the Q-Q mechanism. A multiple birth IV methodology was also performed by Black et al. (2005), who used a large sample of administrative records from Norway to estimate the effect of family size on school attainment. Without birth order controls, the Q-Q trade-off appears to hold, however controlling for birth order essentially wipes out the supposed Q-Q effect. The Black et al. analysis suggests that the standard Q-Q model needs to be rectified such that variation within families is accounted for. Furthermore, a study by Qian (2009) used variation across both time and region to estimate the impact of a softening of China's one child policy on child outcomes. The results from Qian's article contradicted the Q-Q model, as she found that increases in family size led to increases in child school enrollment.

3 The Historical Context

Effective comparisons of the demographic, economic and social historiography between Ireland and elsewhere are fraught with difficulty. In each of these categories, early 20th century Ireland is commonly viewed as an outlier within the European context. The island as a whole is widely regarded as not having participated in both the Industrial Revolution and the European demographic transition. However, a connection between demography and economic growth cannot be ruled out as Ireland's belated fertility transition was accompanied by a convergence in economic conditions (Ó Gráda & Walsh, 1995).

Nevertheless, simply looking at aggregated national statistics masks a substantial amount of regional Irish heterogeneity. Differences in the demographic setting have been highlighted by Ó Gráda (1991), who used the Henry-Coale measure, I_g, to explore inter-regional variation in Irish marital fertility.[3] Ó Gráda's results demonstrated a clear link between urbanization and marital fertility, as the two counties with the lowest levels of marital fertility in 1911 were Dublin (where most of the population live in Dublin city) and Antrim (which contains part of Belfast city). Using the I_g measure of marital fertility, Coale & Treadway (1986) estimated that England's fertility transition commenced in 1892, as between 1881 and 1911 the I_g measure of marital fertility dropped from 0.674 to 0.467. Considering the estimated time of this transition alongside the fact that the comparable 1911 levels of I_g in Dublin and Antrim were 0.582 and 0.598 respectively, presents us with the plausible proposition that both Belfast and Dublin city had also commenced a demographic transition by 1911.

Regional patterns of marital fertility are discussed in the following section, although from the above description it appears that the populations of both cities described above are suitable candidates to test the Q-Q hypothesis. However, despite a similarity in quantitative measures of marital fertility, the two conurbations differ across many other dimensions. Belfast's participation in the Industrial Revolution is unquestionable. By the beginning of the 20th century, Belfast had established itself as a world-leader in the textile, engineering and ship-building industries. Meanwhile, spin-off industries, such as rope-

[3]The I_g index is calculated by dividing a population's age-specific marital fertility by the same measure taken for married Hutterite women. The Hutterites are a religious group with the highest reliably recorded marital fertility.

making, provided further employment opportunities for the city's growing population. Indeed, in the second half of the 19th century, Belfast grew more rapidly than any other city within the British Isles (Clarkson, 1983, p.159).

The emergence of Belfast as a modern industrial center was not matched by Dublin, a city where economic conditions stagnated during the same period. This stagnation was illustrated by Daly (1982), who examined occupational statistics from the decennial census returns and found that the proportion of males engaged in manufacturing actually fell from 33% to 20% during the period 1841-1911. The emergence of mass-manufactured textiles marked the demise of Dublin's proto-industrial base, such as the silk garment industry. Thus, a substantial proportion of Dublin's working class drifted from secure, albeit unskilled, employment, to less secure general laboring. The economic disparity between the two cities was also apparent in average housing and accommodation conditions. In Dublin, large amounts of the working class were consigned to cramped and squalid multiple-occupancy tenement houses. This situation was largely avoided in Belfast, as it was typical for families to live in more comfortable individual dwellings.

[FIGURE 2 ABOUT HERE]

Early twentieth century Dublin was an administrative and commercial center, not a modern industrial one like Belfast. Figure two shows the occupational class distribution for those under twenty years of age, in both cities, stratified by gender, as reported in the 1911 Census of Ireland (BPP 1912-13). The cross-city occupational class divide is apparent. Belfast's young adults, both male and female, worked primarily in the city's labor intensive manufacturing sector. In total, there were 12,780 males and 14,946 females under twenty years of age employed in this sector. The comparable figures for Dublin – 1,996 and 939 respectively – underline Belfast's industrial superiority. In Dublin, young women were typically employed as domestic servants, whereas a large proportion of males were employed in commercial and administrative/professional classified occupations. A sizable proportion of the Dublin city male cohort was employed in industrial based occupations. However, the total size of this group only represents 15.6% of the equivalent group for Belfast. Figure 2 illustrates how differences in the economic composition of both cities affected occupational choices. Since Dublin did not have a labor intensive industrial base, it is sensible to deduce that there was a higher demand and more employment opportunities for teenagers in Edwardian Belfast.

A substantial amount of socioeconomic variation also existed within both cities, and a proportion of this variation can be explained by both ethnic and religious differences (although these were not mutually exclusive factors). In Ireland, Roman Catholics are descendants of the native Irish, while members of the Anglican Church of Ireland are primarily derived from either English settlers or a small group of the native Irish who converted from Catholicism – most likely as a vehicle for social advancement. However, the majority of Protestants in Belfast belonged to the Presbyterian faith – descendants of Scottish migrants. Dublin-based Protestants, largely composed of Episcopalian Church of Ireland followers, had a distinct socioeconomic advantage over the Roman Catholic cohort. This Protestant group was both over-represented in business, professional and senior civil service occupations and under-represented amongst the less well paid working-class occupational classes (Daly, 1982). Using the individual returns from the 1901 census, Hepburn (1996, p.82), was able to cross-tabulate occupational status and religion for Edwardian Belfast. Unlike Dublin, this distribution is more ambiguous, and, to understand it, Hepburn split the Protestant group between the Church of Ireland and Presbyterian faiths. All three religious groups were represented in similar enough proportions across the various occupational statuses, although the Presbyterians appear to have had a higher socioeconomic standing. The distributions for the Roman Catholic and Anglican Church of Ireland populations were almost identical, although this form of analysis ignores a substantial amount of within-occupation wage-discrimination which was purported to have been levied on Belfast Catholics (Hepburn, 1996).

Despite some socioeconomic disparity between the religious faiths, the above suggests that religion was not a substantial barrier to social advancement in either Belfast or Dublin. A key restraint, as Daly pointed out, was educational attainment. For example, in Dublin, entry into most clerical professions was contingent on schooling. Surprisingly, in contrast to the general backwardness which pervaded the island, Ireland's education system was comparatively quite advanced by 1911. Following the Irish Education Act of 1892, school fees for the majority of national/primary schools were abolished, while the same act also introduced compulsory school attendance for all children between the ages of six and fourteen (Ó Buachalla, 1988, p.26). However, it should be noted that mandatory attendance could be circumvented for children aged twelve and over, provided they had found a source of regular paid employment (Patterson, 1985, p.172).

The latter part of the 19th and early 20th century also saw a growth in intermediate/secondary school attendance in Ireland. The introduction of state-organized exams, in 1879, did much to foster this growth, since a successful exam certificate was a valuable currency for those wishing to join either the civil service or army (Coolahan, 1981, p.61). Typically, a combination of both monetary and opportunity costs restricted the poorest from graduating to secondary level, although there were many exceptions. Religious bodies, particularly the Catholic Christian Brothers order, built up a substantial network of secondary schools which enabled a large number of children from the poorest families to attain some form of secondary education. An example of this comes from Belfast, where the Christian Brothers and the Sisters of Mercy founded a number of schools in response to bishops Denvir and Dorrian's plea to break the cycle of poverty and low education amongst the city's Catholic population (Heatley, 1983, p.141).

From the brief review of the history of Irish education above, there appeared to be significant returns to education in both Belfast and Dublin. However, this additional schooling had relevant implications for a family's budget constraint. In the following sections, I evaluate the role which fertility played in educational attainment, by first demonstrating that, in 1911, the women of Belfast and Dublin restricted their levels of fertility, and secondly, that children from smaller families were more likely to achieve a higher level of education via the Q-Q trade-off.

4 Marital Fertility in Ireland

How prevalent was fertility control in 1911 Belfast and Dublin? To answer this, I use individual level data returns from the 1911 Census of Ireland. The National Archives of Ireland provide full and unrestricted access to the returns from both the 1901 and 1911 Irish censuses. In the context of this study, these returns differ in one important respect. In 1911, all married women were asked how many children they had given birth to and also how many of these children had survived.

[FIGURE 3 ABOUT HERE]

To measure regional marital fertility, I use the full sample of all married women in 1911 Ireland. In general, these data are reliable since the enumeration forms left "little room for mistake by person[sic] of ordinary intelligence," according to the National Archives.[4] However, using these data entails a few caveats. For example, research by Budd & Guinnane (1991) demonstrated how the Old Age Pensions Act of 1908 caused intentional age misreporting in 1911. Consequently, the age distribution for the latter age-groups is skewed. Figure 3 illustrates the 1911 age-distribution for Dublin, Belfast and the rest of Ireland. Clearly, a degree of age-heaping is present in all three series. Furthermore, figure 3 demonstrates how the old-age intentional age misreporting was less prevalent, although still present, in both Belfast and Dublin. To ensure accuracy, I restrict the sample of married woman to those aged between 18 and 49. Assuming that marriage represents the intention of a couple to start a family, I limit the sample to women married over two years to guarantee that couples have been given adequate time to commence family formation. The data are spatially categorized into 34 regional subsamples. One subsample contains all district electoral divisions in Belfast city, while another includes the relevant divisions for Dublin city.[5] The other 32 subsamples are drawn from the counties of the remaining electoral divisions.

Cross-regional variation in marital fertility is measured using estimates from a series of generalized-linear models. To facilitate meaningful regional comparisons using this cross-section of data, I control for cohort effects by estimating the number of children alive as a function of cubic trends in both age and marital duration. Similarly, the number of children who have died is also included in a factor variable to account for regional infant mortality differences. Given the nature of the dependent variable it is best to use a class of models specifically designed for count level data. Here, I follow

[4] `http://www.census.nationalarchives.ie/exhibition/dublin/census_day.html`

[5] The Belfast districts are: Belfast East, Clifton, Court, Cromac, Dock, Duncairn, Falls, Ormeau, Pottinger, Smithfield, St. Annes, St. Georges, Victoria, Windsor and Woodvale. The relevant divisions for Dublin are: Arran Quay, Clontarf, Donnybrook, Drumcondra, Fitzwilliam, Glasnevin, Inns Quay, Mansion House, Merchants Quay, Milltown, Mountjoy, Kilmainham, North City, North Dock, Pembroke, Rathmines-Rathgar, Rotunda, Royal Exchange, South City and South Dock.

previous studies which have addressed a similar question, such as Guinnane et al. (2006), and use the negative binomial distribution. The advantage of this model is that it allows parameter over-dispersion unlike the more restrictive Poisson distribution, which is also used in modeling count data but restricts the estimated variance. The parameter values from these generalized linear models are then used to predict the level of marital fertility in each region. These predictions, showing the average number of surviving children that a 40 year-old woman with a marital duration of 15 years and no child deaths had, alongside two standard error bars, are displayed in figure 4.

[FIGURE 4 ABOUT HERE]

The results displayed in figure 4 illustrate a remarkable difference in marital fertility between Ireland's two largest cities and the other regions. In Dublin and Belfast, the representative married woman of could be expected to have 2.2 children, while the average comparative figure for all other regions was 3.5. Additionally, figure 4 also demonstrates the high level of variation in this measure of marital fertility which existed across regions. Notably, married women in the counties which contain both cities (but lie outside the designated metropolitan area) have the closest comparative levels of marital fertility. Nevertheless, these estimates are still notably larger. For example, there was a 0.8 difference between Belfast and Antrim. Controlling for marital duration captures any age at marriage variation, so the estimates of marital fertility here are unbiased by age at marriage differences. Regardless, the age at marriage was actually lower in the two cities.[6] This within county variation shows that previous research, which used aggregated county-level data, substantially underestimates the prevalence of fertility control in Belfast and Dublin.

Figure 4 clearly shows the relatively high level of marital fertility control which was being exercised in both Belfast and Dublin. To understand how marital fertility varied within both cities, I repeat the same as above and include the following relevant covariates: whether a mother is fully literate (clearly indicated that she can read and write), whether the mother is in a household with a domestic servant present and religious group (with Roman Catholic as the base category). The coefficient estimates and their standard errors

[6]The average age at marriage was 23.2 in both Belfast and Dublin in this sample. The comparable mean for the rest of the country 25.0.

are displayed in table 1. The coefficients here represent the expected difference in the number of children alive on the logarithmic scale.

[TABLE 1 ABOUT HERE]

The results in table 1 reveal a number of similarities and also notable differences between the cities. Firstly, literacy is strongly associated with higher levels of fertility in both cities. Taking an exponentiation the coefficients allows us to treat them as multiplicative effects. In this instance, the literacy coefficient tells us that the predictive difference of literacy is around 11% in both cities. That literacy and fertility are positive related is surprising, and runs counter to the findings of Bleakley & Lange (2009). However, this is not necessarily a decisive objection. Firstly, literacy amongst mothers in both cities was very high (both above 90%). Secondly, these coefficients are not reporting the counter-factual impact of literacy and cannot be interpreted in an experimental or 'causal' sense. Assuming that illiteracy represented the poorest sector of society, a more plausible interpretation of these results suggests that the very poorest did not have sufficient means to have relatively large families. In other words, urban poverty imposed fertility restriction, not fertility control.

It is reasonable to assume that the presence of a domestic servant in a household indicates household wealth. This measure contrasts with literacy; as we expect the wealthiest households would have domestic servants. Furthermore, the discussion in section 3 revealed how religion can be used as a measure of socioeconomic standing in Dublin. The coefficients marking the presence of a domestic servant and religious group differ significantly across both the two cities. However, an interpretation of these coefficients draws essentially the same conclusion: after accounting for the very poorest (through the literacy measure) the wealthiest in both cities exercised a higher degree of fertility control. In Belfast, the presence of a domestic servant is associated with around a 5.7% decline in fertility, whereas the relationship is Dublin is a much weaker (1.5%). Nevertheless, in Dublin, fertility amongst the majority of the city's protestant groups (Anglican and Presbyterian) was far lower than for the comparatively less wealthy Roman Catholic cohort. No strong pattern emerges across religious groups in Belfast.

The above presents a strong case favoring Belfast and Dublin city's inclusion in the European demographic transition. Accordingly, I now consider the relationship between fertility restriction and schooling across both cities.

5 The Quantity-Quality Trade-Off

I assume that school attendance is partially explained by variation in parental resources. Therefore, the probability that child i attends school can be modeled as a simple linear function:

$$P(S_i \mid \tau_i, X_i) = a_0 + a_1\tau_i + a_2'X_i, \tag{1}$$

where τ measures child resources, endowed by the parents, and X_i is a vector of potentially confounding variables. The child Q-Q theory posits a negative relationship between sibship size and child resources. This relationship can be represented as follows:

$$\tau_i = c_1n_i + c_2w_i, \tag{2}$$

where n_i is the number of siblings in a household and w_i represents household wealth. Combining equations (1) and (2), and including household wealth (w_i) in the vector of potentially confounding variables yields the following reduced form linear equation:

$$P(S_i \mid n_i, X_i) = \beta_0 + \beta_1n_i + \beta_2'X_i, \tag{3}$$

which underpins the empirical strategy. The Q-Q trade-off is captured, in its reduced-form, using the $\beta_1(= a_1c_1)$ parameter. If the trade-off holds as theory suggests, I should find β_1 <0.

To empirically test this relationship, the complete population was drawn from the 1911 census returns for both Belfast and Dublin – where both cities are designated as before. Unfortunately, the census data are cross-sectional, without retrospective information on either completed family size or education. However, since these data contain the entire population, the number of observations is large enough to make accurate inferences using this cross-section. Specifically, I evaluate the presence of an extra sibling on the probability that a son or daughter between the age of 14 and 15 attends school. Here, school attendance is inferred from the census returns, as those attending school reported this as their occupational status. This indicator contains students at both primary and secondary level. For example, it was not uncommon for students to repeat their latter-primary years in order to obtain a more impressive school certificate (Parkes, 2010, p.50). However, while I cannot observe the standard or level of schooling, it is reasonable approximation to assume that school attendance in these age-groups is positively related to human capital formation.

The 14-15 age cohort is chosen for two reasons. Firstly, schooling was (almost) compulsory up until the age of 14 and consequently, reported school attendance was nearly universal up until this age. Whether or not school attendance was as high as reported in the census is certainly debatable. For example, school attendance amongst Dublin's poorest children, although registered in national school, was haphazard at best (Daly, 1982). The people's suspicions regarding the census enumerators, vividly shown by Guinanne & Budd (1991), allied to the fact that parents – almost exclusively in the cities (Akenson, 1970, p.345) – could be punished for truancy, suggests that the household head had an incentive to lie regarding school attendance below the age of 14. Secondly, as this age cohort is older, my sibship size variable is more likely to accurately reflect completed family size.

Estimating the effect of fertility on educational provision required these data to be trimmed across a number of dimensions. All single parent families are removed because variation in these observations is unrepresentative of purposeful fertility control. Similarly, I remove multiple family households, as deducing inter-family relationships for these households is problematic. The reliability of observations claiming to have been born outside wedlock is dubious, and therefore I restrict the sample to the children of mothers married at least 14 years. Similarly, I remove all women under the age of 30, as their records imply marriage before the age of 16. Furthermore, to ensure that the sample is not affected by age-cohort survival bias, I only include household heads and mothers under the age of 56. This action also removes any potential adverse effects of the aforementioned old-age pension misreporting. Finally, outliers with missing and evidently inaccurate data are also removed.[7]

[TABLE 2 ABOUT HERE]

These actions are necessary but have two potentially adverse effects. Firstly, the removed observations may be less likely to attend school, and secondly, the removal of observations will naturally decrease the precision of any statistical tests. Table 2 displays the numbers of observations for each age cohort alongside the proportion of those still in school. To illustrate the trends, I include both individuals aged 13 and 16. A comparison across these groups offers a good deal of comfort in relation to the above concerns. Firstly, we can see that the sample size for both cities does not decrease substantially. The decrease is more pronounced in Dublin. However, removing these data points appears to only have a slight effect on the probability of schooling for each age cohort. The extent of this effect varies only slightly by at most 1 percentage point for the 14-15 age cohort.

[FIGURE 5 ABOUT HERE]

Figure 5 plots school attendance across different age groups in both cities stratified by sibship size. Overall, the proportion in school is roughly twice as high in Dublin compared to Belfast. This observation is not unexpected since the 1911 census books report a similar disparity (1911 Census of Ireland, BPP 1912-13). This disparity is likely to stem from the better employment opportunities offered by Belfast's burgeoning industries, as shown in figure 2. However, the patterns of decline are roughly similar as the probability of school attendance is clearly a monotonically decreasing function of age. Figure 5 displays differences across the sibship size stratifications which are consistent with the Q-Q trade-off. Comparisons between the two diagrams show that the trade-off is much larger in Belfast.

[TABLE 3 ABOUT HERE]

There are a number of potential models which can be used to estimate the relationship between sibling size and the probability of schooling shown in equation (3). For

[7] These include observations where the number of children in the household exceeds the number which the mother reports to have given birth to.

robustness, I take the most flexible approach, and estimate a number of potentially appropriate models. To control for confounding variation, X_i contains the following relevant covariates: a dummy variable for those aged 15, gender, whether a domestic servant is present in the house, whether the household head can read and write, religious denomination (with Roman Catholics as the base category), the proportion of the siblings who have died in the household, father's age, mother's age and marital duration of the parents. I center the variable for marital duration and parent's age, to make the intercept term easier to interpret. Here, the intercept reflects the probability of schooling given that both parents were 45 years of age and married 20 years, conditional on the remaining variables. Additionally, I account for the criticisms of Black et al. (2005) and include observable sibling order as a factor variable in X_i.

Columns (1)-(3) of Table 3 present the coefficients of the baseline equation, estimated using the standard ordinary least squares linear model, for a pooled sample of both cities, Belfast and Dublin respectively. The influence of sibship size is relevant in both cities. In Belfast, the effect of an additional sibling appears to reduce the probability of school attendance by around 2.7%. The magnitude of the relationship is smaller, -1.0%, yet still statistically relevant in Dublin – as suggested in figure 5. Unsurprisingly, the pooled sample result in column (1) is a weighted average of the two cities (-2.0%). The other coefficient estimates yield additional information regarding the determinants of school attendance. Profound socioeconomic differences appear to have existed. A child from a household with a domestic servant is 35% and 24% more likely to attend school in Belfast and Dublin, whereas the comparable association between a father's literacy and school attendance was 7% and 11% in the two cities. A similar pattern emerges if we treat the proportion of children who the mother has given birth to but died as a measure of deprivation in Belfast but not Dublin.

As we have seen, variation across the different religious groups may have been indicative of both social class and ethnic differences. However, these socioeconomic differences were more evident in the Dublin sample. Consistent with this, the results for Belfast show how religious denomination accounted for little of the variation in school attendance. Clearly, ethnicity was not a barrier to school attendance in early 20th century Belfast. These results contrast strongly with the estimated Dublin sample coefficients –

where the being a member of the Protestant religions was strongly associated with school attendance. However, it is difficult to pin-point whether variation across these groups is caused by socioeconomic/income effects or ethnic effects. Regardless, the inclusion of religious group is controlling for potentially confounding effects which potentially bias the Q-Q parameter. Finally, father's age and marital duration, but not mother's age, are strongly correlated with school attendance. These variables are primarily included in order to control for potentially confounding sampling variation caused by the cross-sectional nature of these data. Therefore, an interpretation of their estimated coefficients, like the sibling order dummy variables, is not particularly illuminating.

Estimation using OLS entails three crucial assumptions. Firstly, I have assumed that the estimated Q-Q parameter (β_1) is unbiased. However, this critical assumption would be violated if the sibship size variable was correlated with the error term in any of the regression models. For example, it may be the case that the X_i vector does not adequately capture variation in household wealth, as in equation (3), leading to omitted variable bias. Secondly, it is reasonable to question whether the dependent variable can be modeled as a simple linear probability function. Finally, the estimated OLS regressions in columns (1)-(3) assume a parametric relationship between sibship size to the probability of school attendance. In other words, school attendance is assumed to fall monotonically and in a constant proportion as sibship size increases. This assumption would be violated if, for example, children with no siblings were less likely to have attended school compared to those with one sibling, and those with one sibling are more likely to attend than those with two, and so forth. Accordingly, I address these concerns in the following.

The validity of first assumption – that the relationship between sibling size and the probability of schooling is exogenous – can be tested by using instrumental variables (IV) with a two-stage least squares (2SLS) estimator. The use of this methodology in empirical economic research is not without criticism and I acknowledge recent contributions from both Deaton (2010) and Heckman & Urzúa (2010). However, for the purposes of the question at hand it serves as a useful check. The primary difficulty with this approach is to find a variable or variables which are correlated with the potentially endogenous variable but conditionally uncorrelated with the error term of the equation of interest. In the Irish census data, there is one variable which is a clear candidate as an instrument:

the presence of twin or multiple births. A similar strategy was pursued by Hatton & Martin (2010), who, to account for reverse causality, limit the designation of multiple birth families to those with a twin or multiple birth at last birth. The presence of twin births is rare ($\approx 0.9\%$), therefore, to maximize the number of twins and ensure that the instrument is sufficiently correlated with the family size, I perform the IV estimation on the pooled sample comprised of both cities.[8]

The first stage regression is displayed in column (4) of table 3. Clearly, the presence of a twin birth for the given level of parity has a strong positive effect on sibship size. Column (5) reports the second stage results. The magnitude of Q-Q trade-off increases from -2.0% to -2.8%, which indicates that if the OLS estimate in column (1) is biased, it is biased away from zero. However, the estimated precision of the trade-off decreases substantially. Consequently, it is not surprising that the Hausmann test, fails to reject the null hypothesis that the OLS estimate is consistent. The parameter estimates for all the other variables are essentially unchanged, and do not require further attention here.

The estimated variance of the 2SLS Q-Q coefficient warns against a decisive conclusion. Therefore, I address the issue of potential endogeneity using a different approach. Specifically, I estimate a multilevel model which captures spatial-level heterogeneity in both the intercept and sibship size slope terms. In essence, this methodology controls for endogeneity bias generated by spatially-varying omitted variables. For example, it is plausible that the variables included in X_i do not adequately control for confounding variation in income/wealth/social class. However, it is reasonable to assume that households with these shared characteristics reside in similar locations. In other words, it is more likely that a working class household lives on the same street as other working class households, and so forth. The census data contains information on the location of individuals for both district electoral division and street. Here, I use the much smaller street level district units and the model from equation (3) can be re-specified:

$$P(S_{ij}) = (\beta_0 + b_{i0})' + (\beta_1 + b_{i1})'n_j + \beta_2' X_{ij}, \tag{4}$$

where b_{i0} and b_{i1} represent the individual random effects for every i-th individual each

[8]Performing the analysis separating the cities does not alter the conclusions, although the estimates are too imprecise to warrant a detailed exposition here.

street j-th street in the dataset, so $(b'_{i0}, b'_{i1})' \sim N(0, \sigma^2 D)$ and β_0, β_1 and β_2 are the average, or fixed, effects which can be interpreted in the same manner as before.

The rare presence of multiple births necessitated a pooling of the cities in the 2SLS analysis. However, this necessity does not apply to the multilevel analysis, and I estimate two multilevel models, one for Belfast, and the other for Dublin. The results of the fixed effects for both multilevel models are displayed in columns (6) and (7) of table 3. The inclusion of street level random effects slightly changes the magnitude for some of these fixed effects. Nevertheless, the results here are very similar to those displayed elsewhere in table 3. Once again the estimated sibling size-scholar relationship is stronger for Belfast (-2.5%) compared to Dublin (-1.0%). These results indicate that the previous results, in columns (2) and (3), are not biased by any unobserved street-level heterogeneity.

The results so far have shown that the effect of sibling size on the probability of schooling is both statistically significant and not biased by endogeneity issues. However, while the least squares model produces unbiased linear estimates, there are additional classes of models which are better suited to modeling a binary variable. Here, I use a probit regression, which contains the linear model from equation (3) in the following generalized linear model:

$$P(S_i) = \Phi(\beta_0 + \beta_1 n_i + \beta'_2 X_i), \tag{5}$$

where Φ is the standard normal Cumulative Density Function and the other notation is as before. The most intuitive way to think about this is that there is some underlying continuous distribution defining human capital which is unobserved. We can observe whether or not an individual has remained in school and therefore this observation reflects the passing of some threshold in the unobserved continuous human capital variable. The coefficients in these models are not marginal effects – as in the least square models – and cannot be directly interpreted. However, computing the marginal effects of each variable is a reasonably straightforward procedure. Here, I implement the approach suggested by Kleiber & Zeileis (2008, p.126) and use the average of the sample marginal effects. These marginal effects and their associated standard errors are displayed in columns (8) and (9) of Table 3. These estimates are essentially identical to those produced by the linear model, and do not warrant further analysis.

[FIGURE 6 ABOUT HERE]

All the results above have assumed a parametric relationship between sibship size and school attendance. To test the validity of this assumption, I re-estimate the OLS models for both cities and include sibship size as a factor variable – with the no sibling group designated as the reference category. Figure 6 illustrates a summary of this analysis, showing the estimated coefficients for each sibship size cohort along with two standard error bars. The estimated coefficients can be interpreted as the probability of school attendance compared to the no sibling group – so a coefficient of -0.1 means that this cohort are 10% less likely to attend school than the group with no siblings.[9] To assess the extent of non-linearity, I include the Q-Q slope terms with the estimated trade-off slope coefficients of -0.027 and -0.01 for Belfast and Dublin respectively. Figure 6 indicates that a small degree of non-linearity is present. However, the line with the Q-Q slope rests comfortably inside the standard error region of each sibship size cohort in both cities as the parametric relationship assumption does not appear to be violated.

6 Cross-Cohort Variation

Table 3 demonstrates clear evidence in support of the child quantity-quality trade-off. Quite simply, children from larger families in both Belfast and Dublin were less likely to attend school after the age of 13. However, there was a striking difference in the magnitude of this trade-off between the cities, and this difference poses a very clear question – what determined the strength of the child Q-Q trade-off?

To answer this question, I split these data into a number of cohorts and examine how the relative strength of the sibling size school attendance relationship changes between groups. To this end, I perform the model-based recursive partitioning scheme proposed by Hothorn et al. (2008). This is a two-step procedure which first fits a parametric model to a data set and then tests for parameter instability across a number of specified partitioning variables. If parameter instabilities are found, the data set is split into various child nodes according to the partitioning variable with the highest instability. At

[9] Observations with 13 and 12 siblings in Belfast and Dublin respectively are not included in figure 6. The number of observations in either group is not large enough to warrant inclusion.

this point, split versions of the original parametric model are re-estimated and the tests for parameter instability are performed again. This algorithm works iteratively until no further partition nodes can be found.

Given the question at hand, the use of model-based recursive partitioning has a number of advantages. The model-based recursive partitioning algorithm allows the data to decide how the sample should be split. This process simplifies the analysis such that the various groups are ordered hierarchically. For example, the Q-Q trade-off might be stronger in Belfast because of differences between religions. In other words, the magnitude of the Q-Q trade-off may vary not according to location or social class, but religion and ethnicity. This methodology detects such splits, and avoids unnecessarily clustering data points, thus providing the most efficient platform to interpret the results. I proceed by estimating the effect of sibling size on school attendance, controlling for age, gender, parent's age and marital duration. These data are partitioned according to the following variables: city, religion, the presence of a domestic servant and father's literacy. The model is then estimated using a generalized linear model with a logistic distribution. Once again, there is no difference in the model if estimated using linear and generalized linear techniques. However, in this instance I use the generalized linear model as the graphical parameters (spinograms) can be easier to interpret. Instability in the parameters is detected using a Bonferroni-corrected significance level of $\alpha = 0.05$.

[FIGURE 7 ABOUT HERE]

[FIGURE 8 ABOUT HERE]

[FIGURE 9 ABOUT HERE]

The results of the recursive partitioning are displayed in figures 7 to 9. I initially estimated the model using the pooled sample. However, since first partition node was split by city, I produce two tree-based diagrams for both cities. Figure 7 shows Belfast, while figure 8 illustrates the Dublin tree. The plots in each leaf of figures 7 and 8 show the unconditional relationship between sibship size and school attendance. Therefore, I produce figure 9 to assist in the interpretation of these various diagrams, plotting the relevant conditional marginal effects for this relationship by dividing the logistic model coefficients by four.[10]

The difference between the both cities is apparent. We have already seen that school attendance was higher in Dublin. Despite this, some similarities between the cities existed. In particular, I find that the Q-Q trade-off was effectively absent amongst the wealthiest (those with a domestic servant) in both cities. Unsurprisingly, the spinogram for these nodes shows a very high proportion attending school. It appears that these families were wealthy to the point that an additional child could be absorbed without a significant dilution in the per-child resources. Only a portion of households in the sample (6%) had a domestic servant present, although no substantial differences existed in the Q-Q trade-off across either city or between religious groups. Religion, and by extension ethnicity, mattered little in terms of the Q-Q trade-off once this high-wealth threshold was achieved.

In Belfast, after splitting by domestic servant, the sample was recursively partitioned once again, by religious group. The results of this split show that Belfast's Roman Catholics were a lot less willing to engage in the Q-Q trade-off compared to those of other religious groups, as the Q-Q coefficient was -0.8% compared to -4.0%. Furthermore, no further partitions were found in the sample, indicating that socioeconomic differences only explain a portion of this disparity. For example, if profound economic differences were influencing the trade-off below the high-wealth threshold, we expect to see the poorest group partitioned – those with an illiterate household head. In other words, Belfast Catholics with literate household heads were no more likely to engage in the Q-Q trade-off than other religious groups with illiterates as household heads. The magnitude of the Q-Q effect in Dublin was much lower, as we have seen before. In contrast to Belfast, I find no evidence of vast differences, by religious cohort.

The results displayed in figures 7-9 are for a cross-section of data. However, there is substantial variation in this sample which helps us understand how the Q-Q trade-off mechanism evolved over time. Firstly, there is no relationship between sibling size and school attendance for the wealthiest, regardless of religion/ethnicity or location. This

[10]Dividing the coefficients by four gives a rough approximation of the marginal effect (Gelman & Hill, 2007, p.82). Linear models estimated using OLS produce almost identical coefficients to those displayed in figure 9.

result helps to explain why the findings of empirical studies based on contemporary data points have been so inconclusive. There appears to be some form of a wealth threshold, after which resource dilution amongst siblings has an insignificant impact on child outcomes. The results from modern studies in developed countries can be reconciled with this result. Firstly, it is reasonable to assume that continuous improvements in the standard of living have pushed the large majority of households above this wealth threshold. Secondly, the emergence of the modern welfare state attempts to provide a buffer which effectively stops families falling below this threshold.

Therefore, my estimates show how the Q-Q trade-off was solely practiced by the cohort of families below a high-wealth threshold. However, this cohort represented a sizable proportion of the populations in both cities, and there was sizable variation within this cohort worth consideration. The size of the trade-off was unambiguously larger in Belfast. Additionally, the coefficients from the various models understate the gulf between the two cities because school attendance was much higher in Dublin. Differences in the industrial nature, and hence teenage employment opportunities, provide the most sensible explanation for the aforementioned disparity. Belfast, as we have seen, had a modern industrial structure, whereas Dublin did not. Consequently, the decision to leave school for paid employment had greater relevance amongst Belfast teens than their Dublin counterparts. Following this line of thought, it is reasonable to assume that the nature of the Q-Q trade-off in Belfast, rather than Dublin, provides a more accurate representation of the trade-off elsewhere in late 19th and early 20th century Western Europe.

7 Conclusion

The emergence of unified growth theory has brought an imperative to understanding the historical relationship linking fertility and human capital. However, there is little evidence showing that the child Q-Q trade-off existed during the demographic transition. Constructing a new data set of Irish families from 1911, I found results supporting both Belfast and Dublin's participation in the European demographic transition. Furthermore, there is strong evidence for a Q-Q trade-off in both cities, as an extra sibling reduced the probability of school enrollment. However, the size of the trade-off was much larger in Belfast, a city with a modern industrial structure based on large-scale manufacturing. This suggests that the industrial revolution was an important element which stimulated the Q-Q mechanism. Additionally, the absence of the Q-Q effect amongst households with a domestic servant demonstrated how the trade-off vanished above a certain level of wealth. Hence, empirical results which use samples from post-demographic/sustained economic growth regimes are irrelevant tests of the micro-foundations of unified growth theory because of high living standards and the relatively low cost of schooling.

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Table 1: Number of Surviving Children Negative Binomial Regressions

	(1) Belfast	(2) Dublin
(Intercept)	1.106***	1.063***
	(0.015)	(0.016)
Age	0.284***	0.324***
	(0.039)	(0.046)
$Age^2/100$	-0.787***	-0.895***
	(0.108)	(0.128)
$Age^3/1000$	0.063***	0.073***
	(0.010)	(0.012)
Marital Duration	0.142***	0.143***
	(0.005)	(0.006)
Marital $Duration^2/100$	-0.232***	-0.273***
	(0.029)	(0.038)
Marital $Duration^3/1000$	0.005	0.014*
	(0.005)	(0.007)
Literate	0.097***	0.106***
	(0.013)	(0.014)
Servant in House	-0.059***	-0.015
	(0.014)	(0.013)
Church of Ireland	-0.023*	-0.101***
	(0.009)	(0.012)
Jewish	0.144**	0.158***
	(0.050)	(0.046)
Methodist	-0.020	-0.074
	(0.016)	(0.038)
Other Religion	0.014	-0.052
	(0.015)	(0.030)
Presbyterian	-0.004	-0.136***
	(0.009)	(0.029)
No. of Children Dead: 1	0.147***	0.190***
	(0.009)	(0.010)
No. of Children Dead: 2	0.159***	0.186***
	(0.011)	(0.013)
No. of Children Dead: 3	0.130***	0.153***
	(0.014)	(0.017)
No. of Children Dead: 4	0.095***	0.102***
	(0.019)	(0.022)
No. of Children Dead: 5	0.033	0.051
	(0.027)	(0.029)
No. of Children Dead: 6	0.003	0.062
	(0.038)	(0.039)
No. of Children Dead: 7	-0.080	-0.017
	(0.049)	(0.053)
No. of Children Dead: 8	-0.326***	-0.094
	(0.092)	(0.077)
No. of Children Dead: 9	-0.320**	-0.498***
	(0.104)	(0.119)
No. of Children Dead: 10	-0.375*	-0.452*
	(0.148)	(0.190)
No. of Children Dead: 11	-0.605*	-0.478
	(0.258)	(0.314)
No. of Children Dead: 12	-2.306*	-0.700
	(1.011)	(0.380)
No. of Children Dead: 13		-1.440
		(1.037)
N	31,655	25,220
R^2	0.291	0.263

Standard errors in parentheses. * p <0.05, ** p <0.01 and *** p <0.001.

Table 2: % in School Before and After Data Clean

	Belfast				Dublin			
Age	%	N	%	N	%	N	%	N
13	0.84	5,913	0.84	3,899	0.94	4,248	0.95	2,718
14	0.42	6,324	0.43	3,939	0.68	4,462	0.69	2,625
15	0.19	5,873	0.19	3,501	0.40	4,026	0.41	2,232
16	0.10	5,862	0.09	3,284	0.24	4,292	0.26	2,160

Table 3: Estimation Results – Linear, Multilevel and Generalized Linear Models

	(1) Scholar Both (OLS)	(2) Scholar Belfast (OLS)	(3) Scholar Dublin (OLS)	(4) No. Siblings Both (OLS)	(5) Scholar Both (2SLS)	(6) Scholar Belfast (Multilevel)	(7) Scholar Dublin (Multilevel)	(8) Scholar Belfast (Probit)	(9) Scholar Dublin (Probit)
(Intercept)	0.454***	0.511***	0.633***	3.827***	0.483**	0.519***	0.641***	0.021	0.127***
	(0.023)	(0.029)	(0.034)	(0.085)	(0.185)	(0.030)	(0.034)	(0.030)	(0.034)
Aged 15	-0.250***	-0.232***	-0.278***	0.126***	-0.249***	-0.235***	-0.277***	-0.229***	-0.264***
	(0.008)	(0.010)	(0.014)	(0.033)	(0.010)	(0.010)	(0.013)	(0.010)	(0.013)
Male	0.019*	0.056***	-0.037**	0.035	0.019*	0.057***	-0.037**	0.056***	-0.036**
	(0.008)	(0.010)	(0.013)	(0.032)	(0.008)	(0.010)	(0.013)	(0.010)	(0.013)
Father Age	0.010***	0.010***	0.008***	-0.030***	0.009***	0.009***	0.007***	0.010***	0.008***
	(0.001)	(0.002)	(0.002)	(0.005)	(0.002)	(0.002)	(0.002)	(0.001)	(0.002)
Mother Age	0.001	-0.000	0.002	-0.144***	0.000	0.000	0.002	-0.000	0.002
	(0.001)	(0.002)	(0.002)	(0.005)	(0.007)	(0.002)	(0.002)	(0.002)	(0.002)
Marital Duration	-0.007***	-0.005*	-0.010***	0.155***	-0.006	-0.004	-0.010***	-0.005*	-0.010***
	(0.002)	(0.002)	(0.003)	(0.006)	(0.008)	(0.002)	(0.003)	(0.002)	(0.003)
Literacy Father	0.085***	0.066**	0.106***	0.040	0.086***	0.056*	0.103***	0.070**	0.102***
	(0.016)	(0.022)	(0.025)	(0.065)	(0.017)	(0.022)	(0.025)	(0.023)	(0.025)
Church of Ireland	0.015	-0.024	0.096***	-0.402***	0.012	-0.027	0.086***	-0.025	0.099***
	(0.012)	(0.014)	(0.020)	(0.046)	(0.022)	(0.015)	(0.020)	(0.014)	(0.021)
Jewish	0.026	0.110	-0.044	-0.198	0.025	0.085	-0.048	0.112	-0.058
	(0.051)	(0.074)	(0.071)	(0.202)	(0.052)	(0.074)	(0.073)	(0.071)	(0.071)
Methodist	0.031	-0.001	0.135*	-0.353***	0.029	-0.008	0.110	-0.004	0.159*
	(0.021)	(0.023)	(0.064)	(0.084)	(0.027)	(0.023)	(0.064)	(0.023)	(0.071)
Other Religion	-0.031	-0.051*	-0.027	-0.348***	-0.034	-0.054*	-0.030	-0.052*	-0.028
	(0.020)	(0.022)	(0.049)	(0.080)	(0.026)	(0.023)	(0.049)	(0.022)	(0.048)
Presbyterian	0.009	-0.016	0.063	-0.402***	0.006	-0.021	0.055	-0.017	0.067
	(0.013)	(0.014)	(0.046)	(0.051)	(0.023)	(0.015)	(0.046)	(0.014)	(0.048)
Proportion Dead	-0.064**	-0.108***	0.008	-3.024***	-0.087	-0.087**	0.012	-0.100**	0.008
	(0.025)	(0.031)	(0.040)	(0.094)	(0.147)	(0.031)	(0.040)	(0.031)	(0.039)
Servant in House	0.296***	0.349***	0.237***	-0.238***	0.294***	0.301***	0.223***	0.314***	0.279***
	(0.018)	(0.025)	(0.025)	(0.070)	(0.021)	(0.026)	(0.026)	(0.025)	(0.029)
Dublin	0.235***			-0.140***	0.234***				
	(0.010)			(0.042)	(0.012)				
Order 2	-0.060***	-0.073***	-0.048**	0.711***	-0.055	-0.077***	-0.047**	-0.072***	-0.048**
	(0.011)	(0.014)	(0.018)	(0.044)	(0.036)	(0.014)	(0.018)	(0.014)	(0.018)
Order 3	-0.036*	-0.048**	-0.025	1.514***	-0.024	-0.049**	-0.026	-0.045*	-0.024
	(0.014)	(0.018)	(0.023)	(0.055)	(0.074)	(0.018)	(0.023)	(0.018)	(0.023)
Order 4	0.011	-0.021	0.052	2.287***	0.028	-0.027	0.051	-0.019	0.049
	(0.018)	(0.023)	(0.029)	(0.068)	(0.111)	(0.023)	(0.029)	(0.023)	(0.029)
Order 5	0.045	0.013	0.080*	2.967***	0.068	0.009	0.076*	0.015	0.077*
	(0.024)	(0.030)	(0.038)	(0.090)	(0.144)	(0.029)	(0.038)	(0.030)	(0.038)
Order 6	0.093**	0.090*	0.095	3.727***	0.121	0.081	0.088	0.086*	0.094
	(0.034)	(0.043)	(0.056)	(0.131)	(0.181)	(0.042)	(0.056)	(0.041)	(0.056)
Order 7	0.154**	0.157*	0.140	4.233***	0.186	0.142*	0.139	0.143*	0.150
	(0.058)	(0.070)	(0.100)	(0.226)	(0.210)	(0.069)	(0.099)	(0.067)	(0.102)
Order 8	0.073	-0.261	0.317	4.461***	0.107	-0.250	0.309	-1.276	0.361
	(0.131)	(0.196)	(0.179)	(0.519)	(0.250)	(0.193)	(0.177)	(12.442)	(0.215)
Order 9	0.437		0.433	5.704***	0.481		0.408		1.565
	(0.260)		(0.270)	(1.030)	(0.376)		(0.268)		(18.436)
Order 10	-0.362	-0.389		4.833**	-0.326	-0.401		-1.357	
	(0.448)	(0.436)		(1.781)	(0.505)	(0.426)		(29.168)	
Twin at Last Birth				0.889***					
				(0.169)					
No. Siblings	-0.020***	-0.027***	-0.010**		-0.028	-0.025***	-0.010**	-0.027***	-0.010**
	(0.002)	(0.003)	(0.004)		(0.048)	(0.003)	(0.004)	(0.003)	(0.004)
N	12,297	7,440	4,857	12,297	12,297	7,440	4,857	7,440	4,857
R-squared	0.176	0.128	0.135	0.369	0.175			0.108	0.107
First-Stage Partial F-Test					27.677				
Hausmann-Wu P-Value					0.874				
Number of Groups						1,555	1,372		

Standard errors in parentheses. * p <0.05, ** p <0.01 and *** p <0.001.

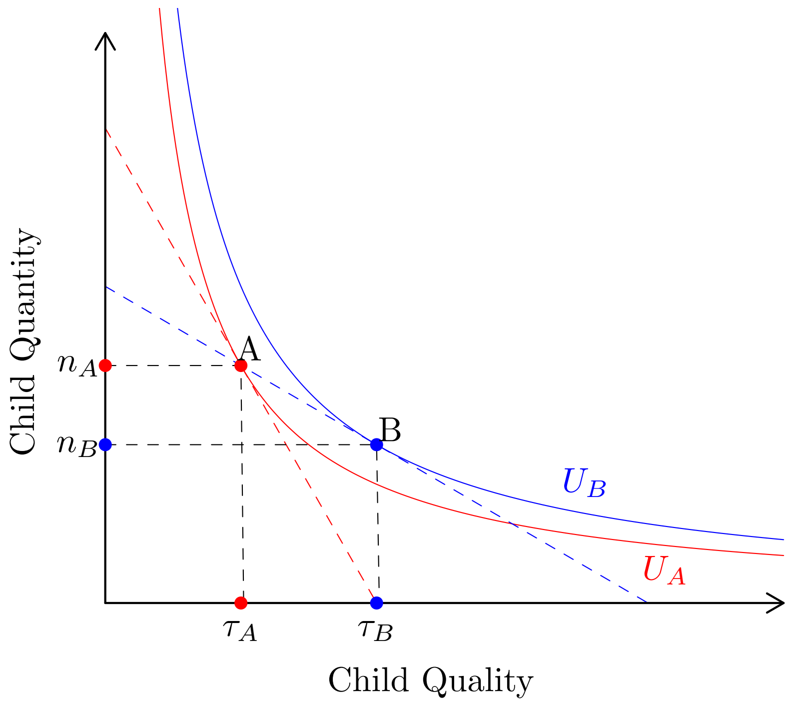
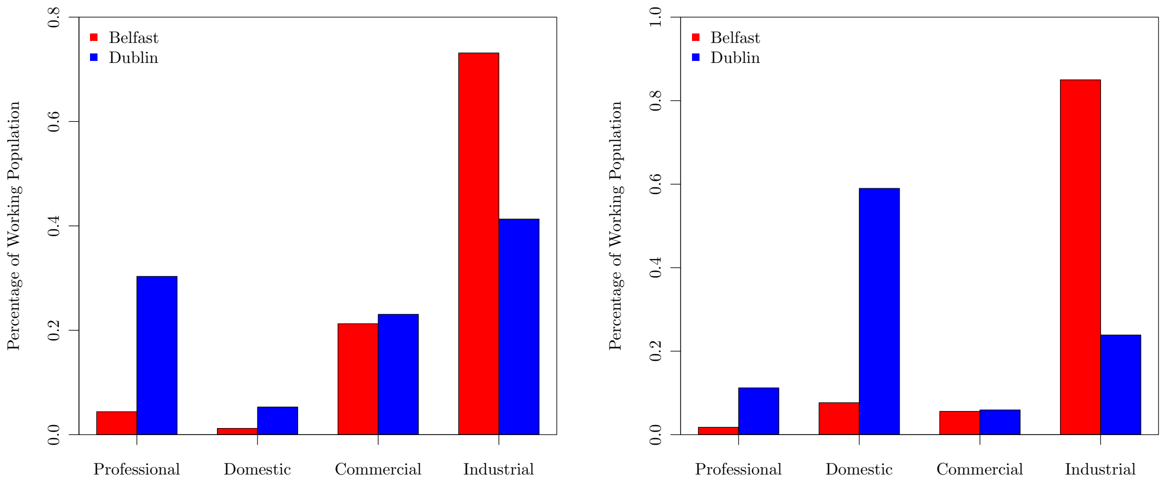


Figure 1: The quantity-quality trade-off



(a) Males (b) Females

Figure 2: Occupational distribution for those under twenty years of age in Belfast and Dublin, by gender. Source: 1911 Census of Ireland (BPP 1912-13)

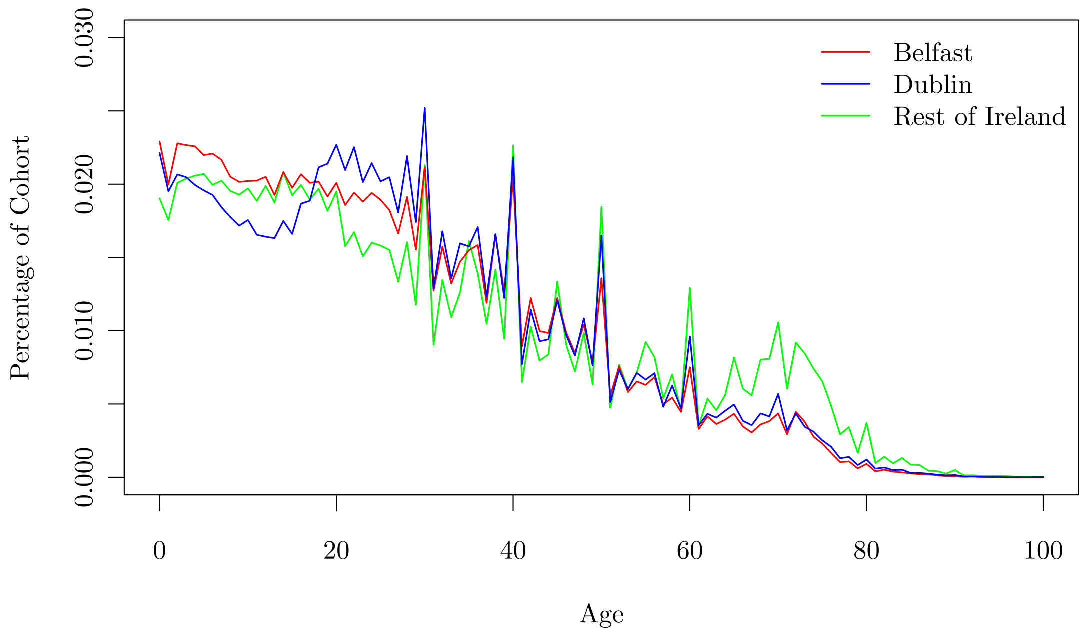


Figure 3: Population age distribution, Ireland 1911

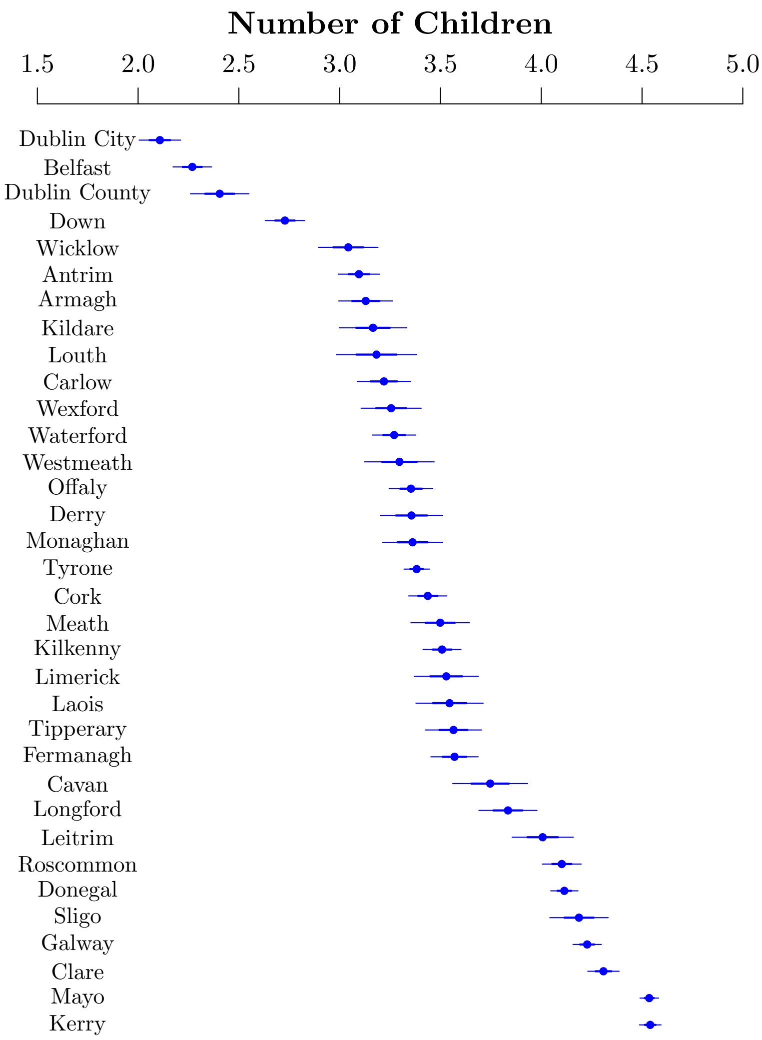
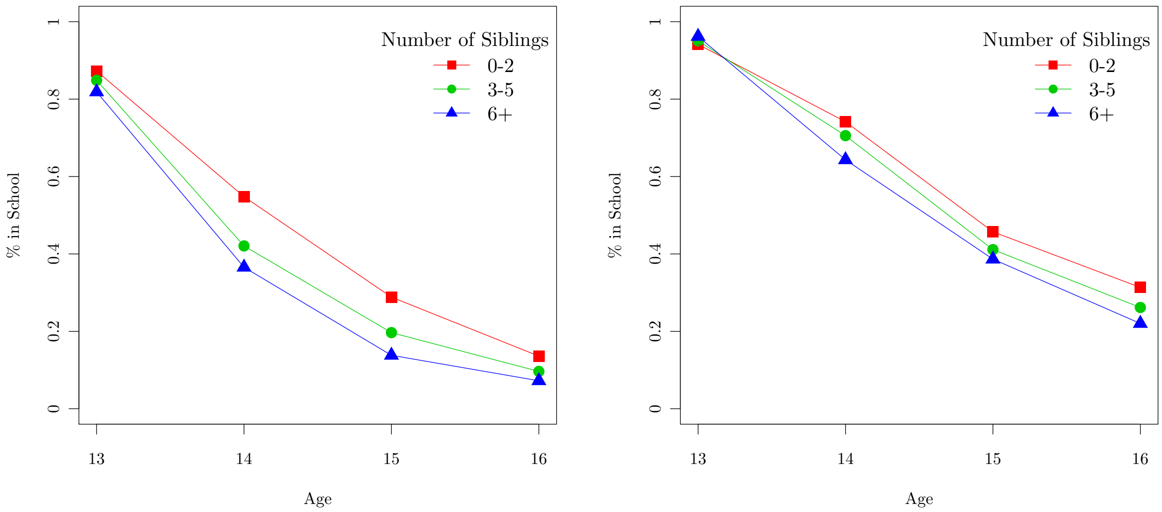


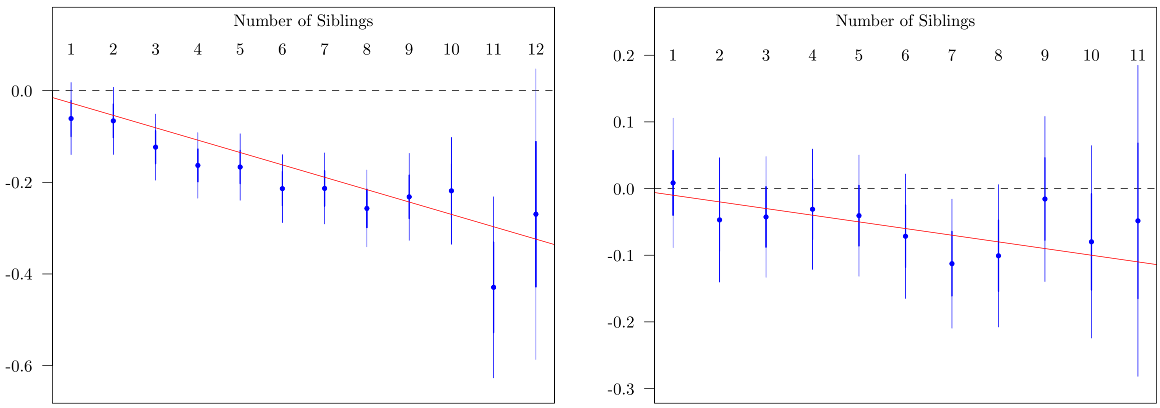
Figure 4: Predicted number of surviving children by region, Ireland 1911



(a) Belfast

(b) Dublin

Figure 5: School attendance by age cohort in Belfast and Dublin, 1911



(a) Belfast

(b) Dublin

Figure 6: Non-parametric Q-Q effects in Belfast and Dublin, 1911

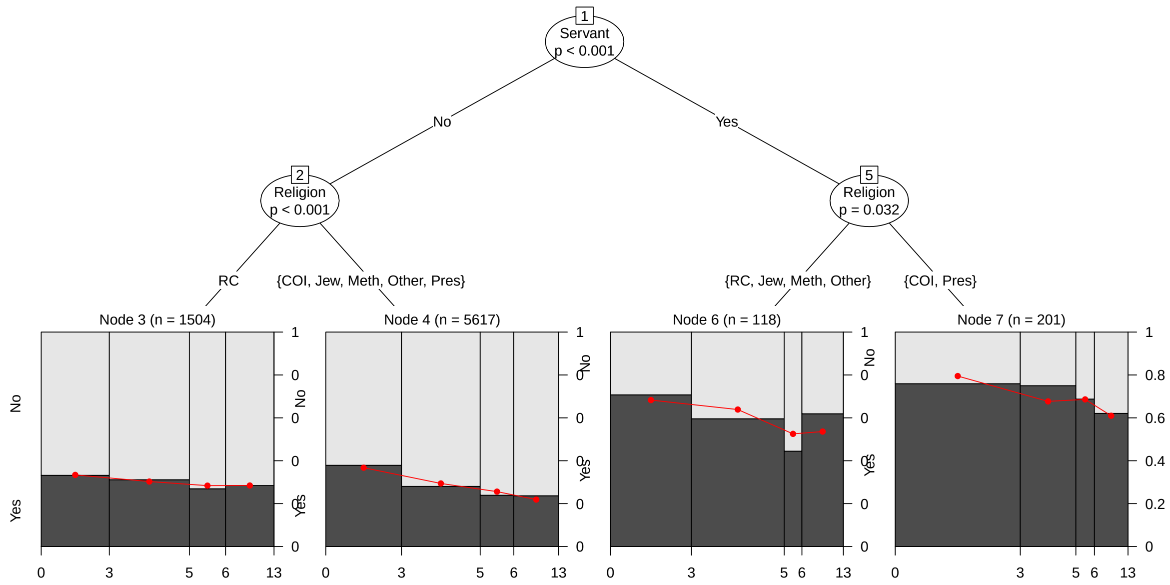


Figure 7: Q-Q Logistic regression tree for Belfast. The plots in the leaves give spinograms for school attendance versus sibling size.

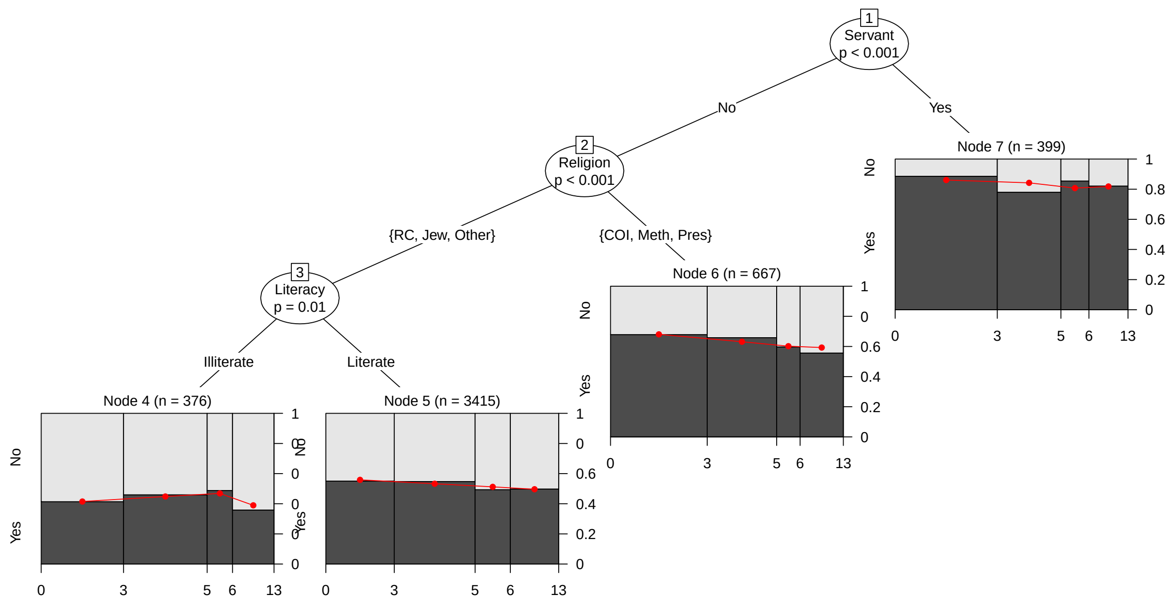


Figure 8: Q-Q Logistic regression tree for Dublin. The plots in the leaves give spinograms for school attendance versus sibling size.

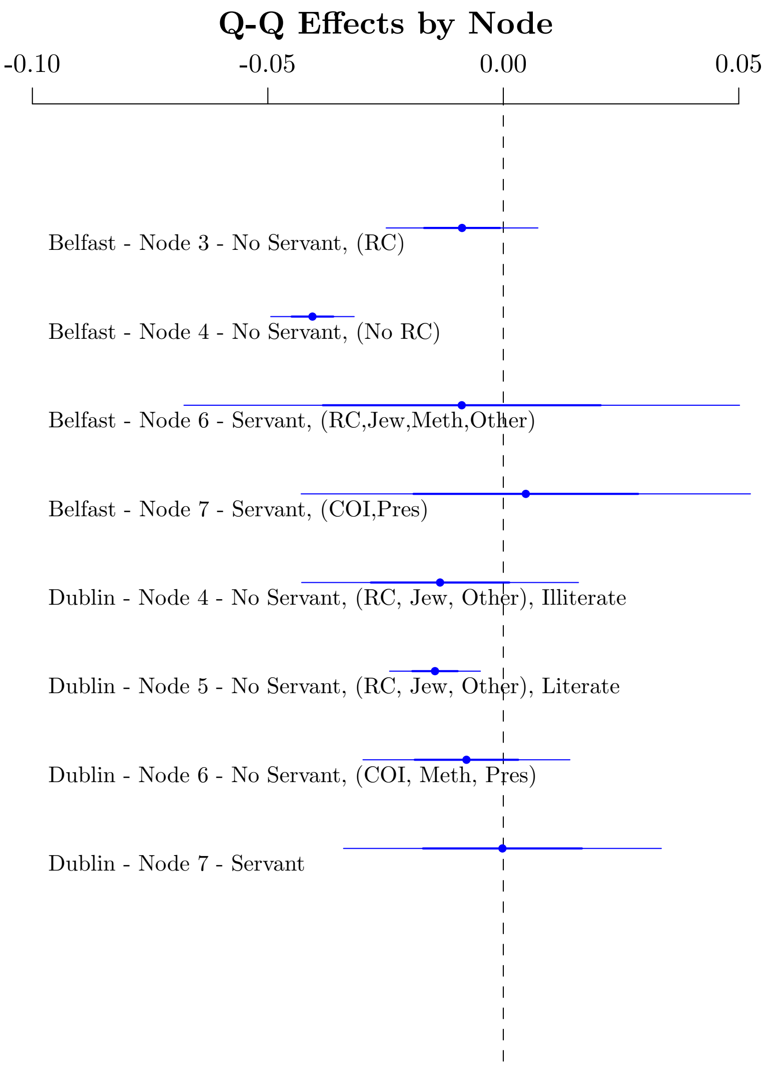


Figure 9: Predicted quantity-quality effect for each node.

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