Empirical Essays in Population Economics

Submitted by

Cevat Giray Aksoy

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Declaration of Authorship

I, Cevat Giray Aksoy, hereby declare that chapters 3 and 4 are entirely my own research work. Chapter 2 was conducted in collaboration with Professor Jeff Frank and Professor Christopher S. Carpenter.

Signed…………………………………………….. (Cevat Giray Aksoy)

Date:

Abstract

This thesis combines three empirical essays in population economics. The first chapter is a collaborative work with Professor Christopher Carpenter and Professor Jeff Frank. In this chapter we examine the earnings differences between sexual minorities and heterosexual individuals. Most prior work on labour market outcomes associated with a minority sexual orientation either lacks good information on labour earnings or only studies same-sex couples. We remedy this gap using a large individual level dataset from the United Kingdom that allows us to measure both constructs. We find a large lesbian earnings premium and gay male penalty in couples-based comparisons – similar to previous work – but no meaningful differential when we restrict attention to singles. The second chapter investigates the impact of unemployment on birth rates in England. It sheds light on the mixed results in the existing literature, particularly showing how the relationship between unemployment and fertility rates varies across demographic subgroups. This chapter also contributes to the existing literature by tackling the issue of endogeneity using a Bartik-style instrumental variable approach. The results of this study suggest that female unemployment tends to increase births, whereas male unemployment has the opposite effect. The third chapter explores the effects of house prices on fertility using a new instrumental variable strategy, exploiting exogenous variation in house prices induced by planning restrictions. Existing studies find a positive effect of house prices on fertility rate at the aggregate level. I show that evidence from a country with a highly regulated housing market suggests otherwise and the net effect is negative. I also find that home owners' birth rates respond positively to house price increases, whereas the opposite is true for renters. The estimates for those aged 20-29 imply that the negative effect of house prices on renters' birth rate is much larger than those implied by the older age group. In contrast, the results for those aged 30-44 show that the overall housing wealth effects are larger than those found in the 20-29 age band, and the home owner results are mainly driven by the older age group.

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Chapter 1

Introduction

This thesis consists of three empirical essays in population economics. The second chapter investigates the relationship between sexual orientation and labor market earnings. Most prior work on this topic has relied either on individual level surveys with small samples of sexual minorities or has used large samples of same-sex couples. This chapter uses a large individual level dataset from the United Kingdom and measures both constructs. It replicates the well-documented lesbian advantage and gay male penalty in couples-based comparisons but shows that these effects are absent in similarly specified models of non-partnered workers. This suggests both that couples-based samples overstate the true earnings differences attributable to a minority sexual orientation and that household specialization plays an important role in the lesbian earnings advantage. It also shows that there is no significant lesbian advantage or gay male penalty in London. Finally, there is a robust evidence that bisexual men earn significantly less than otherwise similar heterosexual men. The detailed discussion on how the effects reconcile with theories of specialization and discrimination has been provided. This chapter is forthcoming in the *Industrial and Labor Relations Review (ILR Review).*

The third chapter re-investigates the causal effects of local unemployment on fertility. It argues that contradicting results in the existing empirical research may have arisen due to a neglect of sub-demographic differences and failure to recognize endogeneity. It hypothesizes that male and female unemployment will have different impacts on fertility across subgroups of the population. Drawing on the UK Labor Force Survey and the Birth Statistics data from the Office for National Statistics, the results of this study suggest that female unemployment tends to increase births, whereas male unemployment has the opposite effect. More importantly, the reported results indicate the unemployment and fertility relation exhibits strong variation across demographic subgroups. Lastly, a persistent countercyclical fertility pattern is also documented at the county level. This chapter is published in the *B.E. Journal of Economic Analysis and Policy.*

Finally, the fourth chapter examines the effects of house prices on fertility rates using a new instrumental variable strategy, exploiting exogenous variation in house prices induced by planning restrictions. Existing studies find a positive effect of house prices on fertility rate at the aggregate level. I show that evidence from a country with a highly regulated housing market suggests otherwise and the net effect is negative. Using data from English counties, the instrumental variable estimates indicate that: a 10 per cent increase in house prices leads to a 2.8 per cent increase in births among owners and a 4.9 per cent decrease in births among renters. Once calculated at the mean home ownership rate the net effect is a 1.3 per cent fall in birth rates. In addition, I document that the positive home owners effect is primarily driven by the older cohort and the negative price effect among renters is mainly driven by those aged 20-29. A further assessment of house prices and fertility nexus reveals that these effects vary by regions and demographic subgroups. This chapter is released through the working paper series of the European Bank for Reconstruction and Development (EBRD), No. 192.

Chapter 2

Sexual Orientation and Earnings: New Evidence from the UK

2.1 Introduction

A growing literature in labor economics examines earnings differences between sexual minorities and heterosexuals using population representative datasets. To identify sexual minorities, researchers have used either: 1) *individual level data* with self-reports of a gay, lesbian, or bisexual orientation (Carpenter 2005, 2008a; Plug and Berkhout 2004, and others) or same-sex sexual behavior (Badgett 1995, Black et al. 2003, and others); or 2) *couples-based data* where sexual orientation is inferred through same-sex living arrangements and the identification of relationships between individual members of the household (Allegretto and Arthur 2001, Arabsheibani et al. 2004, Antecol et al. 2008, and others).[1] Two stylized facts have emerged from couples-based investigations: 1) men in cohabiting same-sex couples earn significantly less than men in different-sex relationships; and 2) women in cohabiting same-sex couples earn significantly more than women in different-sex relationships. In contrast, studies with individual level sexual orientation information generally (but not always) display smaller or insignificant earnings differences.

Because sexual minorities are only a small part of the overall population, the literature has struggled with a tradeoff between representativeness and sample size. Couples-based datasets such as population Censuses in Canada, the United States, and the United Kingdom yield very large samples of same-sex couples but do not identify the sexual identity of non-partnered individuals. In contrast, datasets with individual level information on sexual orientation or sexual behavior (e.g. the General Social Survey or the National Health and Nutrition Examination Survey) have generally been

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[1] The limitations of these alternative methods for identifying sexual orientation in large datasets have been discussed at length elsewhere (see, for example, Carpenter and Gates 2008).

much smaller in size, yielding very small numbers of sexual minorities. The few studies with individual level information on sexual orientation and reasonably large samples of sexual minorities have been limited to single states (e.g., Carpenter 2005), limited to young adults (e.g., Plug and Berkhout 2004, Sabia 2014), or lacked information on labor market earnings (Carpenter 2008a). As a result, it has been difficult to know whether differences in estimated earnings effects of a minority sexual orientation in different studies are due to differences in the samples, populations, or outcomes. Relatedly, it has been difficult to disentangle alternative theories underlying sexual orientation-based differences in labor market outcomes (e.g., specialization versus discrimination).[2]

We overcome these challenges by using confidential versions of the 2012-2014 UK Integrated Household Surveys (IHS) to which high quality labor market earnings data from the country's Annual Population Survey have been linked. These data allow us to identify large samples of sexual minority individuals – over 2,500 self-identified lesbians, gay men, and bisexuals (LGB) – through responses to a direct question about sexual orientation. Our sample is considerably larger than other studies using individual level sexual orientation information in the UK (Uhrig 2015 and Bryson 2014, described below), and indeed ours is the first population representative dataset with information on both sexual orientation and earnings for a large sample of adults from a single country. Moreover, our IHS data permit us to identify not only individual level sexual orientation but also same-sex partnerships. This means we can directly test for how measurement of sexual orientation (i.e., individual level self-reports versus same-sex partnerships) is

[2] Klawitter's (2015) meta-analysis of studies on this topic published between 1995 and 2012 showed that the sample size of sexual minorities and the measure of sexual orientation (couple-status versus sexual identity or sexual behavior) were both significantly related to the estimated earnings difference associated with a minority sexual orientation.

related to earnings differences between sexual minorities and heterosexuals. It also allows us to comment more directly on the possible explanations for earnings differentials. For example, the returns to specialization within the household should accrue to partnered rather than single individuals. In contrast, there is no clear prediction from economic theory on why partnered or non-partnered sexual minorities should suffer greater or lesser discrimination (though it could be that sexual orientation for partnered sexual minorities is more observable to employers, an issue we discuss below).

We show that having data on both partnered and non-partnered sexual minorities is substantively important. After controlling for observable determinants of earnings (such as education, location, and family structure), we find a positive and statistically significant earnings differential for partnered lesbians compared to partnered heterosexual women but no earnings differential for non-partnered lesbians compared with similarly situated non-partnered heterosexual women. We find a negative and marginally significant earnings penalty for partnered gay men compared to partnered heterosexual men but no earnings differential for non-partnered gay men compared with similarly situated non-partnered heterosexual men. Taking together the overall population of both partnered and non-partnered individuals, we find that the earnings difference associated with a gay sexual orientation for men is near zero, while the associated population-based earnings difference among women associated with a lesbian orientation is a premium of about 5.5 percent and is statistically significant.

The different results found for partnered and non-partnered sexual minorities, compared to heterosexuals of the same partnership status, are consistent with models of specialization within the household. Traditional heterosexual households specialize in market and non-market work, with disproportionate market activity done by the male partner. Even if same sex partnerships have the same degree of specialization, it will not be associated with gender. Everything else including degree of specialization being equal, the average partnered gay male (lesbian) will earn less (more) than the average partnered heterosexual male (female). This effect does not hold when comparing single gay men and lesbians to single heterosexuals, since there is no household specialization.

In addition to comparing partnered and non-partnered sexual minorities, we establish several other interesting facts about the sub-groups experiencing sexual orientation differences in earnings. The lesbian earnings advantage is driven by women without a university degree and by women who live outside of London, not by a metropolitan elite. There is a significant gay male earnings penalty in samples of older men (45-64 year olds), consistent with possible historical discrimination against gay men. We also find that bisexual men are estimated to earn significantly less than otherwise similar heterosexual men in the private sector but not in the public sector.

The paper is organized as follows: Section 2.2 reviews the relevant literature on sexual orientation-based differences in earnings. Section 2.3 describes the special license of the UK IHS data and the estimation framework. Section 2.4 presents the results, and Section 2.5 offers a discussion and concludes.

2.2 Literature Review

Our study contributes to the literature on sexual orientation and earnings among adults that uses population representative datasets to identify sexual minorities.[3] Badgett (1995) pioneered studies on this topic by identifying sexual minorities using information on reports of same-sex sexual behavior in the General Social Surveys (GSS), finding a significant gay male earnings penalty and a lesbian earnings advantage. Several follow-up studies found broadly similar results using behavior-based measures in the GSS and other data (Black et al. 2003; Blandford 2003; Carpenter 2007a).

There is less consistency in results across studies using datasets that identify sexual minorities through direct questions about sexual orientation identity (as opposed to same-sex sexual behavior). Carpenter (2008a) examined data from a large health survey in Canada and found that gay men had significantly lower personal incomes than otherwise similar heterosexual men while lesbians had significantly higher personal incomes than heterosexual women. Carpenter (2005) studied adults in California and found no evidence of significant earnings differentials for gay men or lesbians. Uhrig

[3] We do not review here correspondence studies that consistently show evidence of hiring discrimination against sexual minorities. For the US, Tilscik (2011) found strong evidence of discrimination against fictitious applicants who appeared to be gay. Similar experiments in other countries have also returned evidence of differential treatment against lesbians in Austria (Weichselbaumer 2003); against gay men in Greece (Drydakis 2009); and against gay men and lesbians in Sweden (Ahmed et al. 2013). For tractability, we also restrict attention here to studies of prime age adults similar to the sample we study here. A handful of studies have examined data on college students or young adults. Plug and Berkhout (2004) studied young people in the Netherlands and found very small earnings differences associated with a gay or lesbian orientation. Carpenter (2008b) examined young women in Australia and found that young lesbians had significantly lower personal incomes than similarly situated young heterosexual women. Sabia (2014, 2015) studied young adults in the US from the National Longitudinal Study of Adolescent Health (Add Health) and found that young gay men earn significantly less than young heterosexual men and that this difference cannot be explained by numerous controls for family and individual level heterogeneity. In contrast, young lesbians did not earn significantly different wages than otherwise similar young heterosexual women in the young adult sample. Finally, although we focus only on studies using large samples of data from representative surveys, it is worth noting that one study provides evidence on sexual orientation and salary among UK academics. Frank (2006) finds no evidence that gay or lesbian academics in the UK experience salary differences compared to otherwise similar heterosexual academics, although he does find differences in promotions.

(2015) used data from the UK Household Longitudinal Study (UKHLS) collected in 2011-2012 and found a statistically significant bisexual male earnings penalty of about 12 percent and a statistically significant lesbian earnings premium of about 12 percent, with no earnings differences experienced by gay men or bisexual women. Bryson (2014) used data from the UK's 2011 Workplace and Employment Relations Study (WERS) and found that bisexual men earn significantly less than similarly situated heterosexual men, while gay men and lesbians do not in general earn different wages than heterosexuals.

To gain larger sample sizes, much recent work on this topic has used data from population Censuses or administrative register data that identify sexual minorities through same-sex couples. This work was pioneered by a series of papers that used the 1990 Decennial Census to study same-sex unmarried partner couples in the United States (Black et al. 2000, Klawitter and Flatt 1998, Allegretto and Arthur 2001). Klawitter and Flatt (1998) and Allegretto and Arthur (2001) both found that men in same-sex couples in the 1990 Census earned significantly less than similarly qualified men in different-sex couples. Clain and Leppel (2001) used data from the 1990 US Census to show that women in same-sex partnerships earned significantly more than women in different-sex partnerships. Jepsen (2007) and Antecol et al. (2008) both used the 2000 US Census to further explore couples-based wage gaps. Jepsen (2007) found a significant lesbian premium with evidence that this premium was not driven by household specialization. Antecol et al. (2008) found a lesbian premium and gay male penalty with evidence that the premium might be due to human capital differences while the penalty might be due to discrimination. International studies have also examined sexual orientation-based differences in earnings using couples datasets. Arabsheibani et al. (2004, 2005) used the UK Labour Force Survey and found a couples-based gay male penalty and lesbian premium, while Ahmed and Hammarstedt (2010) used Swedish register data to identify couples who had formalized their relationship with the government and found a gay male earnings penalty.

Studies that infer sexuality from partnership are important since they tend to have much larger sample sizes than the existing individual-level samples in the literature. However, this work begs the question of whether or not partnered sexual minorities are representative of the overall sexual minority population. Carpenter (2008a) had a sufficiently large Canadian data set to provide a first answer to that question. That study found much larger differences in partner-based comparisons of total personal income versus population-based comparisons. However, total personal income may be misleading since it includes significant government transfer income. Transfers based on marital status or the presence of children in the household are likely to be correlated with sexual orientation. In the current study, we have data on labor market earnings and provide the first country-level study of sexual orientation and labor market earnings for a large population-representative sample of adults using large samples of sexual minority individuals.

2.3 Data Description and Empirical Approach

Our data come from a special license of confidential versions of the 2012-2014 UK Integrated Household Survey (IHS) with Annual Population Survey (APS) earnings variables linked to the individual records. The IHS is a large, representative household survey of UK residents similar to the March Current Population Survey in the United States. Approximately 350,000 individuals are sampled in each wave of the IHS. For our purposes, the key feature of these data is that the IHS asked respondents a direct question about their sexual orientation. Most studies in the literature on sexual orientation and earnings have relied on indirect methods for identifying sexual minorities, such as same-sex sexual behavior (as in some public health surveys) or, more commonly, the presence of a cohabiting same-sex partner (such as the UKLFS as used in Arabsheibani et al. 2005, 2004). Since people who do not have sex can still identify as sexual minorities, and since non-partnered sexual minorities may have different outcomes than cohabiting partnered sexual minorities, our individual level data on self-reported sexual orientation are preferred as a more comprehensive sample of the overall population of LGB individuals. Importantly, we also have information on which of the self-reported sexual minority individuals are in partnerships.

The IHS contains both a telephone and a face-to-face survey mode. In the telephone mode, respondents age 16 and older are asked "I will now read out a list of terms people sometimes use to describe how they think of themselves. (INTERVIEWER: read list to end without pausing. Note that 'Heterosexual or Straight' is one option; 'Gay or Lesbian' is one option.) 1. Heterosexual or Straight, 2. Gay or Lesbian, 3. Bisexual, 4. Other (Spontaneous DK/Refusal). As I read the list again please say 'yes' when you hear the option that best describes how you think of yourself. (INTERVIEWER: Pause briefly after each option during second reading)." In the face to face interviews, participants age 16 and older were shown a card that had the terms printed next to a number (such as "27. Heterosexual/Straight"). Individuals were then asked "Which of the options on this card best describes how you think of yourself? Please just read out the number next to the description." Notably, sexual minorities did not have to verbalize the words "gay", "lesbian", or "bisexual" to indicate their sexual orientation in either the telephone or face to face survey modes, which presumably reduced potential stigma.[4] Approximately 1.4-1.7 percent of individuals 16 and older self-identified as gay, lesbian, or bisexual in each wave of the IHS, which is similar to other large population-based surveys in the UK, US, and Canada (Joloza et al. 2010).

Individuals are asked about their employment status as well as their gross weekly pay before deductions.[5] In addition to the critical questions on sexual orientation and earnings, the IHS includes standard demographic characteristics such as sex, age, race/ethnicity, educational attainment, partnership/marital status, and the presence of children in the household. We restrict attention to individuals age 25 and older to focus on individuals most likely to have completed their education.[6]

We first estimate the relationship between sexual orientation and employment by estimating linear probability models separately by sex and partnership status.[7] These models take the form:

$$(1) \qquad \text{EMPLOYED}_i = \alpha + \beta_1 X_i + \beta_2(\text{GAY/LESBIAN})_i + \beta_3(\text{BISEXUAL})_i + \varepsilon_i$$

[4] In our empirical models below we include a dummy variable for interviews that were conducted face-to-face. The sexual orientation question was not asked in cases of 'proxy' interviews where a different member of the household provided the information. Forty-four percent of interviews were conducted either by proxy or for respondents under age 16. We exclude these observations without sexual orientation information.
[5] In results not reported but available upon request, our main results are robust to excluding small number of observations (less than a tenth of one percent of the full sample) with earnings less than £20 and more than £7500 per week.
[6] In results not reported but available upon request, we found that lowering our minimum age in the sample to 18 does not meaningfully change the results.
[7] Partnership is based on a dummy variable indicating the person is in any type of partnership (marriage, registered civil union, or a cohabiting partnership not officially recognized by the government).

where EMPLOYED is an indicator variable for being employed or being full-time employed, depending on the model. X is a vector of demographic and job variables that (depending on the model) include: age and its square; education dummies (degree levels, higher education qualification below degree level, A-levels, O-levels); race dummies (white, black, Asian, mixed race, other race); location dummies (London, England excluding London, Scotland, and Northern Ireland); and dummy variables for the presence of children in the household (any child <5, any child at least age 5). Note that in this model the relevant excluded category for sexual orientation is composed of individuals who report a heterosexual orientation. In all models we separately include dummy variables for people who reported 'other' to the sexual orientation question, who refused to provide a response, or who reported 'don't know' (although we do not report the coefficients in the results tables).[8] We also include in all models a dummy variable for interviews performed face-to-face. The error term ε is assumed to be well behaved, and we estimate standard errors robust to heteroscedasticity.

To assess the relationship between sexual orientation and earnings we estimate earnings models separately for males and females and for partnered and non-partnered individuals, among the sample of full-time workers. These models take the form:

$$(2) \qquad \text{LOG EARNINGS}_i = \alpha + \beta_1 X_i + \beta_2(\text{GAY/LESBIAN})_i + \beta_3(\text{BISEXUAL})_i + \varepsilon_i$$

where all variables are as described above.[9]

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[8] Appendix Table 2.1 reports demographic characteristics for individuals who did not provide a valid response to the sexual orientation question.

[9] We also estimated models where we included job characteristics (a private sector dummy, establishment size dummies, industry dummies, and occupation dummies), though a challenge in doing so is that these variables may be channels through which labor market discrimination operates, and thus it does not make sense to control for them in an attempt to distinguish discrimination from specialization. Prior work has demonstrated strong evidence of occupational sorting by sexual orientation (Antecol et al. 2008).

2.4 Results

Descriptive statistics

Table 2.1 presents descriptive statistics for demographic and employment characteristics from the IHS data broken down by self-reported sexual orientation and gender.[10] Self-identified gay men and bisexual men (compared to heterosexual men) are significantly more likely to have a university degree, less likely to be partnered, less likely to have children in the household, and more likely to live in London. Gay men (but not bisexual men) are less likely to belong to a racial/ethnic minority and more likely to live in England (rather than Wales, Scotland or Northern Ireland). In the raw data, gay men have significantly higher average weekly earnings than heterosexual men, while bisexual men have significantly lower average weekly earnings. There is no significant differential in full-time employment between heterosexual men, gay men and bisexual men.

Self-identified lesbians (compared to heterosexual women) are significantly more likely to have a university degree, less likely to belong to a racial/ethnic minority, less likely to have children in the household, and more likely to live in England and specifically in London. Notably, the partnership and presence of children differences between lesbians and heterosexual women are substantially smaller than those between gay men and heterosexual men. In the raw data, lesbians are significantly more likely to be full time workers and have higher average weekly earnings than heterosexual women. Bisexual women are significantly more likely than heterosexual women to

[10] We use the subsample of the IHS for which we have earnings information.

have a university degree, more likely to be partnered, and more likely to be full-time workers.[11]

Full-time employment

In Table 2.2 we examine the relationship between individual characteristics – including sexual orientation – and full-time employment (the likelihood of any employment is examined in Appendix Table 2.2) for men (columns 1 and 2) and women (columns 3 and 4).[12] We estimate models separately for the full sample in the top panel (combining partnered and non-partnered people and including a control for being in a partnership), for non-partnered individuals in the middle panel, and for partnered individuals in the bottom panel. Each column shows coefficients on the gay/lesbian and bisexual indicator variables; the odd numbered columns report estimates from models that only control for sexual orientation, while the even numbered columns add all the individual demographic characteristics (including residential location and presence of children). In column 2 of Table 2.2 we find that gay (bisexual) men are 4.5 (11.9) percentage points less likely to be working full-time than otherwise similar heterosexual men. Notably, this difference for gay men is driven by the partnered sample. Partnered

[11] It is worth noting that our estimates of the proportion of self-identified sexual minorities who report being partnered are independently interesting contributions to the literature since very few datasets have had information on sexual orientation at the individual level, particularly on a large national scale. Table 2.1 shows that a larger proportion of lesbians reports being partnered compared to gay men (69 percent of lesbians versus nearly 50 percent of gay men). These patterns – that the lesbian partnership rate is very similar to the partnership rate of heterosexual women and that the gay male partnership rate is substantially lower than the partnership rate of heterosexual men – were also found for adults in California (Carpenter and Gates 2008). Black et al. (2007) find a similar pattern using data from the GSS that identify sexual minorities from responses about same-sex sexual behavior. Our data also suggest that bisexual men have partnership rates (51.7 percent) that are more similar to those of gay men than to those of heterosexual men, while those of bisexual women (73.4 percent) are slightly higher than those of either lesbians or heterosexual women.

[12] Full-time workers are defined as employees working more than 30 paid hours per week (or 25 or more for the teaching professions).

gay men are 6.1 percentage points less likely to be working full time than otherwise similar partnered heterosexual men. In contrast, the difference for bisexual men is driven primarily in the non-partnered sample, where non-partnered bisexual men are 11.7 percentage points less likely to be working full time than otherwise similar non-partnered heterosexual men. These patterns are qualitatively identical for the analyses of the likelihood of any employment in Appendix Table 2.2.

The results for women in Column 4 of Table 2.2 for full-time employment show that lesbians are 8.2 percentage points more likely to be working full-time than otherwise similar heterosexual women, while bisexual women are 5.4 percentage points less likely to be working full-time. As with gay males, the lesbian difference in full-time employment (although of opposite sign to that for gay males) is predominantly driven by the partnered sample. Partnered lesbians are 15.4 percentage points more likely to be working full time than similar partnered heterosexual women. What differs for lesbians compared to gay males is that the differential reverses when we look at the likelihood that lesbians have any employment (as opposed to full-time employment) in the sample of partnered women after controlling for observables (see Appendix Table 2.2). This arises since heterosexual women in partnerships are more likely than lesbians to engage in part-time work.

These results are consistent with the model of specialization in traditional heterosexual partnerships. Partially, this may be the result of a substantially lower likelihood of children in the household for both gay men and lesbians, compared to their heterosexual counterparts (Table 2.1). This may reduce the need for partnered gay men to work full time in the same way as partnered heterosexual men. Conversely, lesbians on average have fewer childcare responsibilities than heterosexual women and can remain in full-time employment. (Black et al. 2007)

Earnings

Table 2.3 presents estimates of the association between minority sexual orientation and earnings among full-time workers. We focus on full-time workers to be consistent with most of the prior literature; we consider all workers in Table 2.4. The format of Table 2.3 follows Table 2.2 in that the top panel shows results for the full sample (combining partnered and non-partnered people and including a control for being in a partnership), the middle panel examines non-partnered individuals, and the bottom panel examines partnered individuals. Columns 1 and 2 present results for men, while columns 3 and 4 present results for women; the odd numbered columns include only the sexual orientation variables and year dummies, while the even numbered columns add all the demographic and family characteristics.

The results in Table 2.3 are striking. For all comparisons without controls for demographic characteristics in columns 1 and 3, we find that gay men and lesbians earn significantly more than heterosexual men and women, a finding that was previewed in Table 2.1. More importantly, once we control for education, age, and other characteristics in columns 2 and 4, we find important differences by partnership status. In the bottom panel of columns 2 and 4 comparing only partnered sexual minorities to otherwise similar partnered heterosexuals who are full-time workers – as is common in most of the prior literature – we find the usual pattern that partnered gay men earn significantly less than otherwise similar partnered heterosexual men, while partnered lesbians earn significantly more than otherwise similar partnered heterosexual women. In contrast, the middle panel of columns 2 and 4 for non-partnered individuals returns much smaller coefficients on the gay/lesbian indicator variables that are not statistically significant. The results for the full sample in the top panel of columns 2 and 4 confirm that the overall earnings effects of a gay or lesbian orientation are smaller than those implied by the partner-based comparisons, and only the estimate for lesbians is statistically significant in the combined sample. We also find a bisexual male earnings penalty relative to similarly situated heterosexual men that is approximately equal in partnered and non-partnered comparisons; in contrast, there is no earnings difference for bisexual women compared to otherwise similar heterosexual women, except for a marginally significant bisexual female earnings penalty among non-partnered individuals.[13]

In Tables 2.4a/4b and 2.5a/5b we report the sexual orientation coefficients in log earnings regressions for various subsamples, following the baseline specification in columns 2 and 4 of Table 2.3 for males and females, respectively. In addition to the estimates of the fully saturated model for full-time workers (reprinted in column 1 of Tables 2.4a/4b and 2.5a/5b for comparison purposes), we show results for samples that include all workers (including part-time workers) in column 2 of Tables 2.4a/4b. These models also include a control for being a full-time worker. For men, adding part-time

[13] Appendix Table 2.3 reports the values of all the coefficients in the fully saturated model (columns 2 and 4 in Table 2.3). Appendix Tables 2.4 and 2.5 (for men and women, respectively) show that these same basic patterns are robust to controlling additionally for sector of employment, establishment size, and industry of employment (either alone or in combination). Occupation controls do matter in one instance, however: the lesbian premium for non-partnered individuals only obtains after accounting for unrestricted occupation controls; including controls for establishment size, private sector, and industry dummies alone or in combination does *not* return a significant lesbian premium in the non-partnered sample. This confirms prior work that occupational sorting is important for understanding sexual orientation-based differences in labor market outcomes (Plug et al. 2014).

workers makes the negative gay male earnings effect become statistically significant, primarily due to an increase in the estimated negative earnings effect of a gay orientation for non-partnered men when part-time workers are added to the model. For women, the original patterns in column 4 of Table 2.3 remain, though the magnitudes on the lesbian coefficient are much larger in magnitude.

These earnings results shed light on the differential results in the literature when the sample is the full population of self-identified sexual minorities compared to samples that only identify those sexual minority individuals who are in partnerships. By having a large sample with both individual-level self-identification and partnership status, our results directly confirm that the significant earnings differentials are predominantly observed in partner-based samples.

Additional earnings effects

Columns 3 and 4 of Tables 2.4a/4b examine earnings differences when separating the sample by residence in London. Prior work has found this to be an important feature for understanding earnings of gay men (Arabsheibani et al. 2004). While there are many differences between London and the rest of the United Kingdom, one of the most salient in this context is that there is likely to be less discrimination on the basis of sexual orientation in London. Table 2.1 indicated London is a disproportionately popular residence choice for gay and bisexual men and lesbian and bisexual women. We find in column 3 of Table 2.4a that the estimated coefficient on the gay male indicator is positive in sign (suggesting a gay male premium in London), though it is not statistically significant. In contrast, we estimate in column 4 that gay men outside of London experience a significant wage penalty, and again this penalty is larger in partner-based samples. For bisexual men we estimate sizable earnings penalties both inside and outside of London, though the bisexual male coefficients are not statistically different from each other in the London versus non-London comparisons. These London-based patterns for men are interesting, as one could imagine that gay men would earn significantly less in London to compensate them for the city's more progressive attitudes. Instead, it appears that gay men with higher unobservable attributes may choose to move to London or alternatively there is less of a taste for discrimination in London.[14]

For women in Table 2.4b we also find an intriguing difference when we stratify by residence in London. Specifically, we find that the entire lesbian earnings advantage experienced by lesbians is found outside of London, with much smaller and statistically insignificant lesbian coefficients in London. Moreover, the London lesbian differential is the exact opposite of the finding for gay men (who are estimated to do systematically better than their heterosexual male counterparts only inside London). We find no strong difference in bisexual female earnings differences by residence inside or outside of London.

Columns 5 and 6 of Tables 2.4a/4b examine earnings effects of a minority sexual orientation for full-time workers separately by public vs. private sector of employment. This margin is potentially interesting since one might expect there to be stronger antidiscrimination protections in the public sector. In the UK, the public sector has a

[14] The disproportionately high representation of gay men and lesbians in London could also be related to their lower likelihood of having children in the household in a way that interacts strongly with differential returns to household specialization for sexual minorities compared to heterosexuals. Black et al. (2002) argue that the spatial distribution of gay and lesbian couples into disproportionately expensive, high-amenity locations reflects their differential consumption of non-child goods.

'positive duty' to address discrimination which goes beyond the relatively passive requirement upon the private sector not to discriminate.[15] Despite this, we find no meaningful differences in earnings effects of a gay sexual orientation for men in Table 2.4a by sector of employment. There is, however, some indirect evidence from columns 5 and 6 of Table 2.4a about possible discrimination against bisexual men: while bisexual men suffer an extremely large and statistically significant earnings penalty in the private sector, the estimated penalty in the public sector is small and insignificant. We acknowledge that the presence of a bisexual male earnings penalty and the absence of a gay male earnings penalty is puzzling and on its face is difficult to reconcile with simple theories of discrimination. For women, we do estimate a larger lesbian premium for public sector workers, though the estimate for the public sector sample is not statistically distinguishable from the insignificant lesbian coefficient in the private sector sample in column 5 of the top row of Table 2.4b.

Tables 2.5a and 2.5b present further results by demographic group, and the format of these tables follows that of Tables 2.4a/4b (including the fact that we reprint the full sample estimates in column 1). Columns 2 and 3 of Tables 2.5a/5b present results for 25-45 and 46-65 year old full-time workers, respectively; columns 4 and 5 present results for full-time workers with at least some university education versus those without any university education.

Results in Table 2.5a for men return evidence that older gay men experience an earnings penalty relative to similarly situated older heterosexual men. For women in Table 2.5b we document that the lesbian earnings premium is much larger and stronger

[15] Prior work on different UK data than we use here shows that lesbians earn relatively higher wages at employers with explicit antidiscrimination protections compared to those without (Bryson 2014).

in the sample of women without any university education. This is interesting since it has been hypothesized that highly educated sexual minorities might be more able to avoid some of the negative earnings effects of discrimination in the labor market; the fact that the lesbian advantage is observed in the relatively lower educated sample is less consistent with a simple taste-based discrimination explanation.

Effects by head of household status

Our argument on specialization and the lesbian (gay male) earnings premium (penalty) was based upon the gendered nature of heterosexual household specialization. It would hold if lesbian and gay male households specialized to the same degree as heterosexual households. However, if there are diminishing returns to market specialization, then if an average lesbian or gay male household specializes less (more) than an average heterosexual household, the premium would be increased (lessened) or the penalty would be decreased (increased). Of course, it is of interest in its own right as to whether specialization in lesbian or gay partnerships is less than in traditional heterosexual households.

We use information in the IHS to determine whether an individual in a partnership is a 'household head' or 'not a household head'.[16] If gay men and lesbians

[16] The IHS data include a measure for 'household reference person' (HRP). The Office of National Statistics defines the HRP as "the person who is the main owner, renter or in some other way responsible for the accommodation, and who has the highest income (and in some circumstances who has the highest income and is oldest). The rationale for this definition is that the main householder is the person who exerts the most influence on the household's living patterns and circumstances." This variable indicated that 44.4% of partnered heterosexual women were the HRP in their household. We are skeptical that this proportion accurately describes the conceptual construct we are interested in, and the HRP also has the problem that it defines household head status using earnings explicitly (and our outcome of interest is earnings). For these reasons, we chose to define an alternative version of 'household head' in the

in partnerships specialize less than similarly situated heterosexuals in partnerships, we would expect that the gay male penalty and the lesbian premium would be relatively evenly distributed between sexual minorities who are household heads and sexual minorities who are not household heads.

We test this hypothesis with the following model estimated separately by sex:

(3) $\text{LOG EARNINGS}_i = \alpha + \beta_1 X_i + \beta_2(\text{Gay/Lesbian and Household Head})_i + \beta_3(\text{Gay/Lesbian and Not Household Head})_i + \varepsilon_i$

where the sample consists of all partnered individuals in full-time work. Note that we also included dummies for the other sexual orientation categories (bisexual, 'other', and 'don't know'), but we do not report their coefficients. The main comparison group is *all* heterosexual individuals, as we did not want to compare primary earner lesbians to primary earner heterosexuals for the concern that partnered heterosexual women who are household heads are likely to be extremely positively selected, and thus the comparison between household head heterosexual women and household head lesbians would be difficult to interpret. Thus, we compare lesbian household heads and lesbians who are not household heads to all partnered heterosexual women, the large majority of whom are secondary earners (and recall the entire sample is conditioned on full-time work). Similarly, we compare gay male household heads and gay males who are not household heads to all partnered heterosexual men.

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following way: first, if one member of the partnership was a full-time worker and the other member was not a full-time worker, the full-time worker is coded as the household head. Second, if both members of the household are full-time workers, we coded as household head the person in the couple who was older. Third, if both members of the couple were full-time workers and were the same age, we used the 'first person listed in the record' as the household head. This approach returned 28.3% of partnered heterosexual women as 'household heads'.

The results are presented in Table 2.6 and provide some notable support in favor of household specialization underlying the lesbian premium relative to heterosexual women in the sample of partnered individuals.[17] To see this note that in column 2 of Table 2.6 we estimate that partnered lesbians who are household heads earn significantly more than similarly situated partnered heterosexual women by about 7 percent. Partnered lesbians who are not household heads do not earn significantly more than similarly situated partnered heterosexual women, though the point estimate also indicates a sizable premium. Importantly, we cannot reject that the coefficients on 'lesbian, household head' and 'lesbian, not household head' are equal. This is consistent with the idea that lesbian households specialize less than heterosexual households.

For gay men, we observe quite a different pattern than for lesbians. Specifically, the results in column 1 of Table 2.6 indicate that the earnings penalty experienced by partnered gay men compared to partnered heterosexual men accrues exclusively to the person in the partnership who is *not* the household head. Unlike the results for lesbians, we can reject equality of the coefficients between 'gay, household head' and 'gay, not household head'. This result provides support for the hypothesis that gay male households have significant levels of specialization. Since specialization in heterosexual households is often ascribed in large part to child-raising responsibilities, this is a surprising result given the small percentage of gay male households with young children.

[17] The sample sizes in Table 2.6 are smaller than those for the earnings analyses presented in the main paper by approximately 10,000 observations (approximately 6,000 men and 4,000 women). This is because there are many observations in the IHS where we observe earnings and work information for one member of the partnership but not the other member. Since our household head definition requires us to observe this information for both members of the couple, we necessarily drop these observations.

Oaxaca Blinder decompositions

Finally, we investigate Oaxaca Blinder decompositions following Arabsheibani et al. (2005).[18] Table 2.7 reports the mean predictions by group differences for the baseline specification. We estimate models separately by partnership status for comparisons of lesbian and gay men to their associated heterosexual counterparts, though for space considerations (and because partnership differences were not that important for the earlier results on the bisexual wage gap) we do not present results separately by partnership status for comparisons of bisexuals to heterosexuals. The top row shows differences between partnered gay men and partnered heterosexual men, the second row shows differences between non-partnered gay men and non-partnered heterosexual men, the third row shows differences between bisexual men and heterosexual men, the fourth row shows differences between partnered lesbians and partnered heterosexual women, the fifth row shows differences between non-partnered lesbians and non-partnered heterosexual women, and the bottom row shows differences between bisexual women and heterosexual women. Within each row we show the raw (unadjusted) wage gap between the two groups in column 1; the amount of the gap that can be accounted for by different endowments or characteristics in column 2; the amount of the gap that can be accounted for by different returns to characteristics or 'coefficients' in column 3; and the interaction in column 4.

For gay men compared to heterosexual men, recall that we did not find strong evidence of differences in average wages, with a limited negative effect comparing partnered men. In any case, the estimates in the top two rows of Table 2.7 indicate that the majority of any wage difference between gay men and heterosexual men – both in

[18] We use the method described in Jann (2008).

comparisons of partnered people and non-partnered people – can be attributed to different endowments, not different returns. Turning to comparisons between bisexual men and heterosexual men in the third row – where we found much larger earnings differences – the decomposition indicates that the vast majority of the earnings advantage experienced by heterosexual men can be attributed to their higher returns to characteristics, not their differential endowment of skills.[19]

Turning to the comparison of partnered lesbians to partnered heterosexual women in the fourth row of Table 2.7, we find that the lesbian earnings advantage documented in the sample of partnered people is approximately equally attributable to different endowments and different returns. The same is true but to a lesser extent for comparisons of non-partnered lesbians to non-partnered heterosexual women in the fifth row of Table 2.7 where we find a somewhat greater explanatory role for characteristics relative to returns.[20] Both of these cases contrast to the decomposition results for partnered and non-partnered gay men compared to partnered and non-partnered heterosexual men where we found that the mean gap was attributable much more to differential endowments compared with very little role for differential returns to

[19] Note that the wage gap in column 1 reports the mean predictions by groups without covariates, and a negative sign indicates that the non-heterosexual group experiences a premium in the raw unadjusted means compared to the heterosexual group. This is why there are negative signs on the wage gap in rows 1 and 2: partnered and non-partnered gay men earn *more* than partnered and non-partnered heterosexual men, respectively. Comparisons of their characteristics would predict that the partnered and non-partnered gay men would earn even more than the partnered and non-partnered heterosexual men based solely on characteristics since, for example, the gay men have higher education levels than the heterosexual men. This is why the coefficients on the characteristics in rows 1 and 2 are both even larger and negative than the raw wage gap. For the bisexual men the raw wage gap in column 1 is positive, meaning that the bisexual men earn much less than the heterosexual men in unadjusted comparisons.

[20] Again, note that for partnered and non-partnered lesbians compared to partnered and non-partnered heterosexual women, both raw wage gap estimates are negative, indicating that the lesbians earn more than the heterosexual women in unadjusted comparisons. As with the gay men, the coefficient on the characteristics is negative, suggesting that the partnered and non-partnered lesbians would earn even more than the partnered and non-partnered heterosexual women based solely on characteristics since, for example, lesbians have higher education levels than heterosexual women.

endowments. Finally, for the comparison of bisexual women and heterosexual women in the bottom row of Table 2.7, we find that all the earnings advantage for heterosexual women is due to differential returns to endowments as opposed to differential endowments, similar to the findings for bisexual men compared to heterosexual men in the third row.

2.5 Discussion and Conclusion

The main objective in this paper is to try to shed light on the somewhat contrasting results in the literature on sexual orientation and earnings for full population samples and those that only include partnered individuals. The latter studies have had the advantage of being much larger samples drawn from census data and other sources. But there has been a question as to how representative partnered individuals are over the whole lesbian and gay population. We have used what is to our knowledge the first countrywide dataset with both partnership status and self-identified sexual orientation combined with high-quality data on labor market earnings.

Over the population in full-time work (adopted as our main sample to be consistent with the bulk of the existing literature), we found a significant negative sexual orientation-based earnings coefficient for partnered gay men compared to partnered heterosexual men and a significant positive coefficient for partnered lesbians compared to partnered heterosexual women. There is no significant effect for non-partnered gay men or lesbians compared to non-partnered heterosexual men and women. The positive partnered lesbian effect is sufficiently strong that the lesbian coefficient on earnings remains significantly positive over the whole population sample. This does not hold for the negative earnings coefficient for gay men. Our results therefore are consistent with the literature: using partnered sexual minorities tends to show stronger effects than for those studies using the whole population of partnered and non-partnered individuals.

We have argued that these basic results are consistent with specialization. Traditional heterosexual partners typically involve gendered specialization, with the man more engaged in market activities than the woman, particularly given the higher prevalence of children among heterosexual couples. Other things equal, the average partnered heterosexual man will be more focused upon market activities than the average gay man. By the same argument, the average partnered lesbian will be more focused upon market activities than the average partnered heterosexual woman. And these differences should not accrue to non-partnered individuals. All of these specialization-based predictions hold in our data. Our findings that the lesbian premium among partnered individuals accrues approximately equally to lesbians who are household heads and lesbians who are not household heads also supports the idea that there is less specialization in a lesbian household.

An alternative hypothesis for why partnered lesbians may have an observed earnings premium not shared by non-partnered lesbians is that there is a high partnership selectivity effect. Partnered individuals may be selected as the more productive individuals, and the unobserved heterogeneity that facilitates forming a partnership may also be useful in the workplace. Moreover, this may vary by sexual orientation. We see from the descriptive data in Table 2.1, however, that lesbians are just as likely to be in a partnership as are heterosexual women. They are, however, more likely to be in full-time work. In this case, if the underlying selectivity effect is the same for both heterosexual and lesbian women, then the lesbian earnings differential among partnered women in full-time work should be less than the associated differential among non-partnered women in full-time work, since the average partnered lesbian in full-time work will have less favorable unobserved heterogeneity than the average partnered heterosexual woman. Since this is contrary to what we find, this casts doubt on the selectivity explanation for the lesbian premium for partnered women.[21]

While comparative specialization within the household is our preferred explanation for most of our results, there is some limited evidence for the presence of discrimination as an explanatory factor. Our results show that it is older gay men and partnered gay men that earn less than comparable heterosexual men. It is likely that the lack of a heterosexual marriage becomes more of a signal of sexual minority status as an individual gets older (Carpenter 2007, Frank 2007).[22] Partnered gay men may also be more observable as being gay than non-partnered counterparts. They may have photos of a same-sex partner or list their same-sex partner as a beneficiary, for example. If there is discrimination against gay men, these more observable individuals may bear a greater penalty. Further, the gay male penalty only occurs outside London, where there is likely to be a stronger taste for discrimination. Finally, the bisexual male penalty only

[21] In results not reported but available upon request we also found very little evidence of positive selection into partnership on the basis of education for gay men and lesbians, in contrast to prior results for gay men and lesbians in the United States (Carpenter and Gates 2008).

[22] Notably, there are other non-discrimination based explanations for the gay male earnings penalty accruing to older men. For example, there could be wealth effects for gay men associated with their much lower likelihood of raising children. In results not reported but available upon request we found that the significantly lower likelihood of full-time employment experienced by gay men in Table 2.2 is driven primarily by significantly lower full-time employment rates of older (i.e., 45-64 years old) gay men compared to similarly situated older heterosexual men. In contrast, employment rates for 25-44 year old gay men were not significantly different to those for similar heterosexual men. That the employment gap for gay men only is observed for the older sample suggests that wealth effects on earnings may be important even in the absence of any labor market discrimination.

occurs in the private sector and not the public sector where there are greater protections against discrimination.[23]

However, there is also evidence against the discrimination hypothesis in both the male and female comparisons. Among the full sample of partnered and non-partnered men, the presence of a large bisexual male penalty coupled with the absence of a gay male penalty is difficult to square with simple theories of taste-based discrimination. Similarly, among the full sample of women we observed that lesbians earn more than heterosexual women and that the premium occurs in samples of lower educated women and women outside of London – both places where we would normally expect greater discrimination if it existed. Also, the premium occurs among partnered and not non-partnered lesbians, and the same argument as with gay men suggests that these individuals will be more observable and therefore more subject to any discrimination.

Taken together, then, our unique samples of partnered and non-partnered sexual minorities and high quality data on earnings provide novel evidence supporting a role for specialization in explaining sexual orientation-based differences in labor market earnings, with less evidence for selectivity and at best limited and mixed support for discrimination. As more large-scale social science datasets add individual level information on sexual orientation, future work in other countries and contexts can continue to inform our understanding of how a minority sexual orientation shapes economic outcomes.

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[23] We also note that for both bisexual men and bisexual women, Table 2.7 indicated that the raw earnings penalty arises due to lower returns to bisexual individuals' characteristics rather than lower endowments for bisexual individuals. This pattern is quite consistent with discrimination against bisexual individuals, and indeed it is possible that there are different levels and patterns of discrimination against bisexual individuals compared to gay men and lesbians.

Table 2.1
Descriptive Characteristics – Demographics (among those with earnings information)
2012-2014 UK Integrated Household Surveys

Variables	Heterosexual men	Bisexual men	Gay men	Heterosexual women	Bisexual women	Lesbians
Age	44.91 (10.63)	43.63 (11.30)	41.95 (9.80)	44.23 (10.26)	41.45 (10.18)	40.78 (9.36)
Highest education level:						
University degree	0.308 (0.462)	0.409 (0.493) [A]	0.470 (0.499) [A]	0.307 (0.461)	0.427 (0.495) [B]	0.440 (0.497) [B]
Some higher education	0.118 (0.322)	0.119 (0.325)	0.121 (0.327)	0.131 (0.337)	0.103 (0.304)	0.139 (0.347)
A (‘advanced’) level	0.255 (0.436)	0.165 (0.372) [A]	0.202 (0.401) [A]	0.194 (0.395)	0.166 (0.372)	0.179 (0.383)
O (‘ordinary’) level	0.210 (0.407)	0.165 (0.372)	0.158 (0.365) [A]	0.278 (0.448)	0.252 (0.435)	0.194 (0.396) [B]
White	0.905 (0.293)	0.795 (0.405) [A]	0.952 (0.215) [A]	0.928 (0.258)	0.911 (0.284)	0.963 (0.189) [B]
Partnered	0.737 (0.440)	0.517 (0.501) [A]	0.497 (0.500) [A]	0.665 (0.472)	0.734 (0.442) [B]	0.690 (0.463)
Any Child <16	0.278 (0.448)	0.182 (0.387) [A]	0.012 (0.110) [A]	0.340 (0.474)	0.305 (0.461)	0.129 (0.335) [B]
England	0.744 (0.436)	0.744 (0.437)	0.829 (0.377) [A]	0.730 (0.444)	0.800 (0.401) [B]	0.789 (0.408) [B]
London	0.087 (0.282)	0.210 (0.409) [A]	0.226 (0.419) [A]	0.079 (0.270)	0.163 (0.370) [B]	0.113 (0.317) [B]
N. Ireland & Wales & Scotland	0.256 (0.436)	0.256 (0.437)	0.171 (0.377) [A]	0.270 (0.444)	0.200 (0.401) [B]	0.211 (0.408) [B]
Avg. Weekly Earnings	639.00 (515.30)	527.5 (316.30) [A]	677.10 (814.70) [A]	396.00 (411.80)	409.30 (278.40)	515.20 (310.10) [B]
Full-time worker	0.917 (0.275)	0.903 (0.296)	0.903 (0.296)	0.564 (0.496)	0.615 (0.487) [B]	0.807 (0.395) [B]
Sample Size	73318	176	1220	94810	429	839

Weighted means (standard deviations). Not reported (but included in the models) there are 7,020 men and 7,469 women who, when asked about sexual orientation, responded ‘other’, ‘don’t know’ or who refused a response. [A] The superscript letter A means statistically significant difference ($P < 0.05$) between the groups of gay men and bisexual men in contrast to the heterosexual men. [B] The superscript letter B means statistically significant difference ($P < 0.05$) between the groups of lesbians and bisexual women in contrast to the heterosexual women.

Table 2.2
Sexual Orientation and Full-time Employment
UK IHS 2012-2014, Adults age 25+

	Males		Females	
Controls for ➔	(1) Sexual orientation + year dummies	(2) + demographic characteristics (age, race, education, any kids, residence) + year dummies	(3) Sexual orientation + year dummies	(4) + demographic characteristics (age, race, education, any kids, residence) + year dummies
Full sample				
Gay/Lesbian	0.020	-0.045***	0.230***	0.082***
	(0.013)	(0.012)	(0.017)	(0.017)
Bisexual	-0.107***	-0.119***	-0.022	-0.054***
	(0.034)	(0.032)	(0.021)	(0.020)
R-squared	0.001	0.164	0.003	0.161
N	121206	121206	175285	175285
Non-partnered				
Gay/Lesbian	0.092***	-0.008	0.099***	-0.052*
	(0.019)	(0.018)	(0.032)	(0.031)
Bisexual	-0.069	-0.117***	-0.056	-0.146***
	(0.046)	(0.043)	(0.039)	(0.035)
R-squared	0.003	0.171	0.002	0.202
N	39508	39508	62650	62650
Partnered				
Gay/Lesbian	0.025	-0.061***	0.300***	0.154***
	(0.018)	(0.016)	(0.020)	(0.019)
Bisexual	-0.029	-0.029	-0.004	-0.014
	(0.047)	(0.047)	(0.025)	(0.023)
R-squared	0.001	0.178	0.004	0.149
N	81698	81698	112635	112635

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Specific controls in columns 2 and 4 include: a dummy variable for being interviewed face-to-face; age and its square; dummy variables for degree levels, higher education (HE qualification below degree level), A-levels, O-levels; race/ethnicity dummies (white, black, Asian, mixed race, other race); location dummies (London, England, Scotland, and Northern Ireland); the presence of children (any child <5 & any child ≥5) in the household; and year dummies. Columns 2 and 4 in the top panel also include a control for being in any kind of partnership.

Table 2.3
Sexual Orientation and Log Earnings, Full-time workers
UK IHS 2012-2014, Adults age 25+

	Males		**Females**	
Controls for ➔	(1) Sexual orientation + year dummies	(2) + demographic characteristics (age, race, education, any kids, residence) + year dummies	(3) Sexual orientation + year dummies	(4) + demographic characteristics (age, race, education, any kids, residence) + year dummies
All				
Gay/Lesbian	0.061***	-0.027	0.124***	0.054***
	(0.021)	(0.019)	(0.024)	(0.021)
Bisexual	-0.134**	-0.149***	0.004	-0.036
	(0.053)	(0.044)	(0.035)	(0.032)
R-squared	0.001	0.198	0.001	0.231
N	75017	75017	59221	59221
Non-partnered				
Gay/Lesbian	0.126***	-0.006	0.115**	0.029
	(0.027)	(0.025)	(0.047)	(0.037)
Bisexual	-0.038	-0.110	0.013	-0.097*
	(0.082)	(0.068)	(0.053)	(0.050)
R-squared	0.007	0.189	0.003	0.247
N	19905	19905	22385	22385
Partnered				
Gay/Lesbian	0.062*	-0.050*	0.124***	0.067***
	(0.032)	(0.028)	(0.028)	(0.025)
Bisexual	-0.164**	-0.189***	-0.002	-0.009
	(0.066)	(0.057)	(0.044)	(0.040)
R-squared	0.001	0.191	0.002	0.224
N	55112	55112	36836	36836

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Specific controls in columns 2 and 4 include: a dummy variable for being interviewed face-to-face; age and its square; dummy variables for degree levels, higher education (HE qualification below degree level), A-levels, O-levels; race/ethnicity dummies (white, black, Asian, mixed race, other race); location dummies (London, England, Scotland, and Northern Ireland); the presence of children (any child <5 & any child ≥5) in the household; and year dummies. Columns 2 and 4 in the top panel also include a control for being in any kind of partnership.

Table 2.4a
Sexual Orientation and Log Earnings, various subsamples, males
Specification is Table 2.3, Column 2
UK IHS 2012-2014, Adults age 25+

	(1) Baseline: full-time workers (Table 2.3, Column 2)	(2) All workers	(3) London residents, full-time workers	(4) Non-London residents, full-time workers	(5) Private sector full-time workers	(6) Public sector full-time workers
All males						
Gay	-0.027	-0.041**	0.051	-0.070***	-0.022	-0.030
	(0.019)	(0.021)	(0.039)	(0.020)	(0.024)	(0.027)
Bisexual	-0.149***	-0.185***	-0.190**	-0.126***	-0.174***	-0.017
	(0.044)	(0.062)	(0.076)	(0.054)	(0.051)	(0.079)
R-squared	0.198	0.182	0.179	0.191	0.199	0.199
N	75017	81734	6793	68224	58539	16459
Non-partnered males						
Gay	-0.006	-0.035	0.057	-0.041	0.011	-0.026
	(0.025)	(0.028)	(0.050)	(0.027)	(0.030)	(0.041)
Bisexual	-0.110	-0.134	-0.166	-0.095	-0.122*	-0.029
	(0.068)	(0.084)	(0.130)	(0.080)	(0.074)	(0.112)
R-squared	0.189	0.175	0.169	0.176	0.188	0.198
N	19905	21910	2237	17668	15804	4089
Partnered males						
Gay	-0.050*	-0.049	0.049	-0.094***	-0.059	-0.013
	(0.028)	(0.031)	(0.062)	(0.029)	(0.037)	(0.036)
Bisexual	-0.189***	-0.242***	-0.220**	-0.168**	-0.235***	-0.023
	(0.057)	(0.091)	(0.100)	(0.072)	(0.068)	(0.100)
R-squared	0.191	0.174	0.187	0.183	0.193	0.194
N	55112	59824	4556	50556	42735	12370

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 2.3. Models in the top panel also include a control for being in any kind of partnership.

Table 2.4b

Sexual Orientation and Log Earnings, various subsamples, females
Specification is Table 2.3, Column 4
UK IHS 2012-2014, Adults age 25+

	(1) Baseline: full-time workers (Table 2.3, Column 4)	(2) All workers	(3) London residents, full-time workers	(4) Non-London residents, full-time workers	(5) Private sector full-time workers	(6) Public sector full-time workers
All females						
Lesbian	0.054***	0.135***	0.014	0.063***	0.044	0.067**
	(0.021)	(0.028)	(0.071)	(0.020)	(0.028)	(0.030)
Bisexual	-0.036	-0.062	-0.059	-0.030	-0.048	-0.020
	(0.032)	(0.039)	(0.064)	(0.037)	(0.044)	(0.041)
R-squared	0.231	0.224	0.180	0.219	0.211	0.243
N	59221	103547	5753	53468	33695	25521
Non-partnered females						
Lesbian	0.029	-0.014	-0.000	0.036	0.008	0.063
	(0.037)	(0.064)	(0.092)	(0.040)	(0.051)	(0.055)
Bisexual	-0.097*	0.146**	-0.119	-0.094*	-0.087	-0.120
	(0.050)	(0.060)	(0.132)	(0.054)	(0.059)	(0.084)
R-squared	0.247	0.279	0.185	0.233	0.221	0.273
N	22385	34792	2749	19636	13259	9122
Partnered females						
Lesbian	0.067***	0.199***	0.036	0.075***	0.064*	0.072**
	(0.025)	(0.029)	(0.098)	(0.024)	(0.034)	(0.036)
Bisexual	-0.009	-0.040	-0.029	-0.003	-0.027	0.011
	(0.040)	(0.049)	(0.073)	(0.048)	(0.060)	(0.045)
R-squared	0.224	0.203	0.190	0.212	0.208	0.230
N	36836	68755	3004	33832	20436	16399

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 2.3. Models in the top panel also include a control for being in any kind of partnership.

Table 2.5a

Sexual Orientation and Log Earnings, by demographics, males

Specification is Table 2.3, Column 2

UK IHS 2012-2014, Adults age 25+

	(1) Baseline: full-time workers (Table 2.3, Column 2)	(2) 25-45 year olds, full-time workers	(3) 46-65 year olds, full-time workers	(4) Education greater than A-levels, full-time workers	(5) Education A-levels or less, full-time workers
All males					
Gay	-0.027	0.001	-0.092***	-0.029	-0.021
	(0.019)	(0.023)	(0.032)	(0.024)	(0.029)
Bisexual	-0.149***	-0.161***	-0.124*	-0.195***	-0.138*
	(0.044)	(0.055)	(0.070)	(0.057)	(0.074)
R-squared	0.198	0.214	0.177	0.111	0.087
N	75017	39069	35948	32640	38656
Non-partnered males					
Gay	-0.006	0.007	-0.041	-0.023	0.039
	(0.025)	(0.031)	(0.039)	(0.033)	(0.036)
Bisexual	-0.110	-0.094	-0.114	-0.159*	-0.126
	(0.068)	(0.089)	(0.088)	(0.089)	(0.115)
R-squared	0.189	0.206	0.173	0.108	0.073
N	19905	10814	9091	8219	10457
Partnered males					
Gay	-0.050*	-0.011	-0.134***	-0.035	-0.084*
	(0.028)	(0.033)	(0.050)	(0.036)	(0.047)
Bisexual	-0.189***	-0.221***	-0.121	-0.227***	-0.159*
	(0.057)	(0.066)	(0.105)	(0.075)	(0.093)
R-squared	0.191	0.210	0.167	0.101	0.080
N	55112	28255	26857	24421	28199

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 2.3. Models in the top panel also include a control for being in any kind of partnership.

Table 2.5b
Sexual Orientation and Log Earnings, by demographics, females
Specification is Table 2.3, Column 4
UK IHS 2012-2014, Adults age 25+

	(1) Baseline: full-time workers (Table 2.3, Column 4)	(2) 25-45 year olds, full-time workers	(3) 46-65 year olds, full-time workers	(4) Education greater than A-levels, full-time workers	(5) Education A-levels or less, full-time workers
All females					
Lesbian	0.054***	0.030	0.079*	0.028	0.100***
	(0.021)	(0.022)	(0.048)	(0.026)	(0.034)
Bisexual	-0.036	-0.030	-0.045	-0.075**	0.004
	(0.032)	(0.036)	(0.065)	(0.037)	(0.058)
R-squared	0.231	0.227	0.250	0.125	0.046
N	59221	31775	27446	29779	27158
Non-partnered females					
Lesbian	0.029	-0.024	0.136**	0.016	0.062
	(0.037)	(0.045)	(0.066)	(0.050)	(0.053)
Bisexual	-0.097*	-0.073	-0.142	-0.155**	0.049
	(0.050)	(0.050)	(0.128)	(0.066)	(0.062)
R-squared	0.247	0.249	0.260	0.131	0.073
N	22385	11160	11225	10663	10642
Partnered females					
Lesbian	0.067***	0.057**	0.056	0.037	0.119***
	(0.025)	(0.025)	(0.062)	(0.031)	(0.044)
Bisexual	-0.009	-0.013	0.000	-0.027	-0.004
	(0.040)	(0.048)	(0.073)	(0.042)	(0.072)
R-squared	0.224	0.217	0.248	0.125	0.035
N	36836	20615	16221	19116	16516

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 2.3. Models in the top panel also include a control for being in any kind of partnership.

Table 2.6
Sexual Orientation and Log Earnings, by Household Head Status for Sexual Minorities
Sample is partnered full-time workers
Specification is Table 2.3, Columns 2 and 4
UK IHS 2012-2014, Adults age 25+

	(1) Men	(2) Women
Gay/Lesbian & Household Head	0.002	0.071**
	(0.034)	(0.031)
Gay/Lesbian & Not Household Head	-0.141***	0.057
	(0.045)	(0.039)
R-squared	0.187	0.225
N	48688	32862

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Controls include: a dummy variable for being interviewed face-to-face; age and its square; dummy variables for degree levels, higher education (HE qualification below degree level), A-levels, O-levels; race/ethnicity dummies (white, black, Asian, mixed race, other race); location dummies (London, England, Scotland, and Northern Ireland); the presence of children (any child <5 & any child ≥5) in the household; and year dummies.

Table 2.7
Oaxaca Decompositions
Baseline Specification, with Demographic Controls
UK IHS 2012-2014, Adults age 25+

	(1) Wage Gap	(2) Characteristics	(3) Coefficients	(4) Interaction
Partnered gay men vs. partnered heterosexual men				
	-0.124	-0.254	0.010	0.120
	(0.061)	(0.104)	(0.053)	(0.100)
Non-partnered gay men vs. non- partnered heterosexual men				
	-0.064	-0.125	0.038	0.023
	(0.045)	(0.034)	(0.040)	(0.028)
Bisexual men vs. heterosexual men				
	0.181	0.048	0.232	-0.099
	(0.089)	(0.083)	(0.078)	(0.071)
Partnered lesbians vs. partnered heterosexual women				
	-0.128	-0.057	-0.048	-0.023
	(0.044)	(0.050)	(0.039)	(0.046)
Non-partnered lesbians vs. non-partnered heterosexual women				
	-0.085	-0.091	-0.043	0.049
	(0.081)	(0.074)	(0.061)	(0.053)
Bisexual women vs. heterosexual women				
	0.052	-0.006	0.081	-0.023
	(0.057)	(0.045)	(0.046)	(0.031)

Notes: For details on control variables, see notes to Table 2.3.

Chapter 3

The Effects of Unemployment on Fertility:

Evidence from England

3.1 Introduction

This paper examines how local unemployment affects household fertility outcomes. The standard economic models of fertility imply that unemployment has a potential offsetting impact on fertility, as it leads to a substantial fall in income. Assuming that children are a normal good, an increase in unemployment will have a negative impact on the demand for children in the current period, holding other factors constant. In societies with traditional gender roles, the income effect can be expected to be the main effect of male unemployment. Conversely, among females, unemployment decreases the opportunity cost of childrearing and may potentially increase birth rates. The final outcome will depend on individuals' expectations of the duration of joblessness and human capital depreciation as well as the strength of the net income effect.

There is a growing body of literature that investigates the impact of unemployment on household fertility decisions. However, the findings from the existing literature are mixed and occasionally contradictory.[1] In response, the current paper seeks to clarify the causal impact of unemployment on fertility and to demonstrate how previous mixed results may be due to differences in behavior among demographic subgroups. Ultimately, it concentrates on three key research questions:

1. Is the overall effect of unemployment on current fertility positive or negative?
2. Is the impact different for male and female unemployment and across different age groups?[2]
3. Are there further demographic characteristics that shape the fertility response to unemployment?[3]

The current paper builds on the recent contribution by Schaller (2015) who,

[1] See Sobotka, Vegard, and Philipov (2011) for a review of the earlier literature.
[2] 16–24 male, 25–34 male, 35–44 male, 16–24 female, 25–34 female, 35–44 female.
[3] Educational attainment, country of birth and marital status.

using US data, explicitly considers the role of gender-specific labor market conditions on fertility. It consequently reexamines the relationship between unemployment and fertility using English data with a particular focus on demographic subgroups. The main analysis starts with an estimation of a fixed effects model where current fertility by county and year is related to lagged unemployment while controlling for demographic characteristics and house prices. This is followed by a breakdown of the overall relationship by gender-specific unemployment and by age group. In order to address the potential endogeneity issue, an instrumental variables (IV) strategy is implemented based on the approach of Bartik (1991), in which labor demand shocks are used as an identifying source of variation. Although aggregate local unemployment leads to an increase in fertility, the analysis reveals important difference across demographic subgroups.[4] In particular, both the ordinary least squares (OLS) and IV estimations show that for the prime-aged cohort (25–34), male unemployment is negatively associated with the fertility, whereas female unemployment has a positive effect on birth rates. The results tend in a similar direction among the younger age group of 16–24, where male unemployment appears to have a negative impact on the current period fertility, while female unemployment has the opposite effect. These main findings are in line with the theoretical prediction outlined above. In addition, particular demographic factors differentially shape the fertility response across age groups. Specifically, education, country of birth and partnership status mediate the relationship between prime-aged individuals' labor market conditions and fertility more than that between youth labor market conditions and fertility.

[4] The countercyclical fertility behavior supports the predictions of Butz and Ward's (1979) model. They argue that as female employment rate increases in a country, negative labor demand shocks significantly reduce the cost of childbearing and women take advantage of this "joblessness" term.

The present paper thus sheds light on the mixed results in the existing literature, particularly showing how the relationship between unemployment and fertility varies across demographic subgroups. First and foremost, there are strong reasons to expect that the relationship between unemployment and fertility may vary across age groups. Second, in order to disentangle the effect of gender-specific unemployment, it is important to control for both male and female unemployment simultaneously. Finally, the current paper also contributes to the existing literature by tackling the issue of endogeneity using a Bartik-style IV approach.

The paper proceeds with an initial section introducing the conceptual framework and related literature. Section 3.3 subsequently focuses on data and methodology while Section 3.4 presents the results, and Section 3.5 concludes.

3.2 Conceptual Framework and Related Literature

In recent years, there have been a number of contributions exploring the cyclical nature of fertility. These studies show that the fall in fertility rates coincides with higher levels of female unemployment (Brewster and Rindfuss 2000; Esping-Andersen 2009; Engelhardt and Prskawetz 2004). The standard theoretical framework typically put forward builds on the work of Becker (1960) and depicts couples as utility-maximizing agents deciding on the number of children and on child-related expenditures. In this respect, two main approaches introduced children as a normal good into economic models of fertility. The quality and quantity approach (Becker 1960; Becker and Lewis 1973; Willis 1973) implies that an increase in income may have depressing effects on fertility. It relies on the fact that income elasticity for the number of children is substantially less than that for quality of children.[5] The timing of the fertility approach

[5] Becker (1960) altered his model and added a "quality" variable in order to explain the observed inverse

(Mincer 1963; Becker 1960) attributes the low opportunity cost of childrearing during recessions, which could imply a positive relationship between unemployment and fertility. Where "traditional" gender roles exist, the main effect of male unemployment can be expected to be the income effect. Conversely, female unemployment can be expected to have both a negative income effect and a potentially positive opportunity cost effect. In other words, an increase in male unemployment may lead to a fall in fertility.[6] While a rise in female unemployment may also cause a decline in fertility, it nevertheless has a potential substitution effect that goes in the direction of increasing births as opportunity cost becomes lower.[7]

In terms of the female income effect on fertility, previous studies appear to show that women choose to have fewer children as a result of an increase in the economic costs of childrearing (Easterlin 1973; Mincer 1963). This may occur because the "female time intensity" of children is still a key component in the price of fertility, and because the costs of having children for their careers and lifecycle income are substantial. In fact, it has been shown that the unemployment duration of a woman in the labor force is affected by her previous work experience and her lifetime allocation of time, which in turn is intimately related to her fertility decisions (Mincer 1963). An additional determinant affecting ideal timing of fertility is human capital depreciation. It may be argued that most women's human capital investment in schooling is completed before childrearing starts, and thus the cost of childrearing may also be related to women's educational attainment and to their employment

relationship between income and fertility. However, Jones, Schoonbroodt, and Tertilt (2008) show that without assuming a high elasticity of substitution between children and consumption, the quality–quantity approach does not sufficiently explain an inverse income–fertility relationship.

[6] Kravdal (2002) takes the discussion further and suggests that unemployed men are less attractive as a potential husband for the family formation, which may also reduce the possibility of fertility.

[7] Some papers in the literature focus on how fertility rates are affected by an exogenous change in household income. For example, Lindo (2010) shows that birth rates are negatively affected by lower household income due to job loss. Black et al. (2013) show that the 1970s coal boom in West Virginia caused an unexpected increase in income, which also led to an increase in fertility.

possibilities.[8]

Schaller (2015) provides a recent examination of this issue and investigates the differential impacts of male and female unemployment on fertility. Her results show that although birth rates follow a pro-cyclical pattern at the aggregate level in the United States, improvements in male labor market conditions are associated with increases in fertility, whereas improvements in female labor market conditions have the opposite effect. She also performs a particular examination of the demographic subgroups. Based on her analysis, the negative effect of unemployment becomes more pronounced for older people. Furthermore, single women and lower educated groups are highly impacted by the business cycles. The present paper shares with Schaller (2015) the emphasis on gender specific unemployment and demographic subgroup differences as a key determinant of the fertility outcome. An important difference is this paper additionally investigates how further demographic characteristics affect the unemployment–fertility relationship within age groups with different demographic characteristics, whereas Schaller (2015) does not make this distinction. Another contribution of my analysis comes from the fact that I can take house price changes into account. Recent evidence suggests that short-term increases in local house prices affect the fertility of home-owners positively but that of non-home-owners negatively in the United States (Dettling and Kearney 2014). Moreover, special attention should be devoted to England, because it exhibits different labor market properties from those of the United States, namely a considerable increase in female labor force participation (LFP) and a more rapid closing of the gender wage gap.[9]

[8] Heckman and Walker (1990) estimated semi-parametric reduced-form neoclassical models of life cycle fertility in Sweden and showed that rising female wages delay times to all conceptions and reduce total conceptions. Happel, Hill, and Low (1984) also argued that human capital accumulation is an important determinant of fertility timing.

[9] In England, the difference between the participation rate of men and women has shrunk remarkably from 14.5 percentage points in 1994 to 8.6 percentage points in the final quarter of 2011. Polachek and

The results in the literature so far are mixed and difficult to reconcile. Karaman Örsal, Dilan and Goldstein (2010) show a negative effect of both male and female unemployment on current fertility rates in 1976–2008 across 22 OECD (Organization for Economic Co-operation and Development) countries. A number of recent papers report findings that countries with higher female unemployment have lower number of births since the early 1990s (Adsera 2005; Ahn and Mira 2002; Brewster and Rindfuss 2000). In contrast, Ozcan, Mayer, and Luedicke (2010) show that the male unemployment delays the first birth, but female unemployment does not affect fertility in West Germany. Overall, these studies thus highlight the need for further systematic examination of the unemployment and fertility phenomenon.

3.3. Data and Methodology

The principal empirical methodology of this paper involves relating county-level fertility rates to lagged county-level unemployment rates and to control for time-varying county-level demographic characteristics. The following section briefly explains the main data sources and how the relevant variables are constructed.

3.3.1 Data

The fertility data used in this analysis come from the Office for National Statistics (ONS). It compiles counts of live births and stillbirths by age of mother and area of usual residence in England, years 1995–2011. Age-specific fertility rates (ASFRs) are constructed by dividing the number of births by the relevant female population using mid-year population estimates that are based on the censuses, in which female ages range between 16 and 44.[10] ASFRs provide an appropriate measure of varying fertility

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Xiang (2014) show that the gender wage gap is declining relatively more quickly in England – along with Canada and Korea – than in other countries.

[10] Age-gender cohorts are as follows: 16–24 male, 16–24 female, 25–34 male, 25–34 female, 35– 44 male, 35–44 female.

rates since they are unaffected by changes in population age distribution and are well suited for comparing fertility rates across age groups. In this analysis, ASFRs are based on age intervals of 16–24, 25–34 and 35–44.[11] As the best available measure of the labor market conditions prevailing at the time of the conception, births in calendar year, t, are matched with 1 year lagged, t – 1, data of Labor Force Survey (LFS) in the corresponding county.[12]

Access to the confidential LFS was vital to conduct this study, as it was used to construct county–year–age group-specific unemployment rates and county– year–age group-specific demographic characteristics. The sample is restricted in the age band 16–44 in order to study females who are of childbearing ages.[13] Table 3.1 presents the descriptive statistics for demographic characteristics and unemployment rates from the LFS, birth statistics from the ONS and house prices from the Department for Communities and Local Government.[14]

[11] The ONS's age grouping is utilized in this analysis.

[12] Both Birth Statistics and LFS data are available in a finer geography; however, due to small cell size in some areas, it was preferred to aggregate up to the ceremonial county level. There are 49 ceremonial counties in England. After the exclusion of the City of London and Rutland the remaining ceremonial counties are as follows: Bedfordshire, Berkshire, Bristol, Buckinghamshire including Milton Keynes, Cambridgeshire including Peter-borough, Cheshire consisting of Cheshire East, Cheshire West and Chester, Halton and Warrington, Cornwall including Isles of Scilly, Cumbria, Derbyshire including Derby, Devon including Plymouth and Torbay, Dorset including Bournemouth and Poole, County Durham including Darlington, Hartlepool and Stockton-on-Tees north of the River Tees, East Riding of Yorkshire, including Kingston-upon- Hull, East Sussex including Brighton and Hove, Essex including Southend-on-Sea and Thurrock, Gloucestershire including South Gloucestershire, Inner and Outer London, Greater Manchester, Hampshire including Portsmouth and Southampton, Herefordshire, Hertfordshire, Isle of Wight, Kent including Medway, Lancashire including Blackburn with Darwen and Blackpool, Leicestershire including Leicester, Lincolnshire including North Lincolnshire and North East Lincolnshire, Merseyside, Norfolk, North Yorkshire including Middlesbrough, Redcar and Cleveland, York and Stockton-on-Tees south of the River Tees, Northamptonshire, Northumberland, Nottinghamshire including Nottingham, Oxfordshire, Shropshire including Telford and Wrekin, Somerset including Bath and North East Somerset and North Somerset, South Yorkshire, Staffordshire including Stoke-on-Trent, Suffolk, Surrey, Tyne and Wear, Warwickshire, West Midlands, West Sussex, West Yorkshire, Wiltshire including Swindon, Worcestershire.

[13] Over the sample period, the median age difference between husband and wife was 2.1 years. Furthermore, only 6.4% of men and 3.4% of women who married since 1995 were more 10 years older than their spouse.

[14] All unemployment rates are based on the ILO definition (those who are out of work in the reference week, want a job, have actively sought work in the last 4 weeks and are available to start work within the next 2 weeks).

Figure 3.1 illustrates fertility and unemployment patterns among counties. Northumberland, Cheshire and Dorset have the lowest average fertility rates in England and the South East region along with the Greater London experience the highest fertility. Merseyside, Tyne and Wear and West Midlands are the areas with the highest unemployment rate over the sample period.

Figure 3. 2 shows trends in birth rates and unemployment rates by age groups at the national level.[15] Birth rates follow a decreasing trend for age group 1 after reaching a peak of 53.2 births per 1,000 women in 1995. With regard to the 25–34 band, birth rates rebounded after 2000 and reached its highest point of 112.2 in 2009. The oldest age group, 35–44, has experienced a tremendous rise in fertility rate, a steady increase from 21.3 to 35. Turning to the national time series data for unemployment rates, they notably differ in levels across age groups but follow a similar trend. The total unemployment rate reached its lowest point between 2002 and 2005, and gradually increased afterwards. Overall, substantial variation across counties in Figure 3.1 and considerable shift in birth trends across age groups in Figure 3.2 strongly suggest the inclusion of county-specific and age group-specific linear time trends. Additionally, from looking at the figure, birth rates appear to follow a countercyclical pattern over the analyzed time interval.

In all, a balanced panel is constructed for the 1994–2010 period, with 47 counties and three age groups. The final version of the dataset contains information on ASFRs, age- and gender-specific unemployment rates, educational attainment, marital status, ethnicity and country of birth.

[15] Trends in age-gender-specific unemployment rates do not distinctly differ from that of their age group counterparts.

3.3.2 Methodology

In order to obtain baseline estimates of the relationship between unemployment and fertility, the following fixed effect specification is employed:

$$ln(Y_{gct}) = \beta U_{gc(t-1)} + \psi X_{gc(t-1)} + \alpha_c + \theta_g + \gamma_t + \omega_c * T + \delta_g * T + \varepsilon_{gct}$$

The level of analysis is a county-year-age group cell. Y_{gct} is the birth rate in county c, age group g, in year t and $U_{gc(t-1)}$ is the lagged unemployment rate.[16] The county fixed effects, α_c, and age group fixed effects, θ_g, are included to control for differences in birth rates across counties and age groups owing to time invariant unobservable factors. The year fixed effects, γ_t, account for movements in fertility rates over time that are shared by all counties. The county-specific linear time trends, $\omega_c * T$, and age group-specific linear time trends, $\delta_g * T$, control for unobserved variables correlated with birth rates that change linearly over time within counties and age groups.[17] $X_{gc(t-1)}$ indicates lagged time-varying county-level demographic controls (country of birth, ethnicity, educational attainment and marital status) and house prices that account for changes in population composition and changes in the real estate market. The regressions are simultaneously carried out for both male and female unemployment in own age group.[18] All regressions are weighted by the relevant population of women in each cell.

This study uses the identification assumption that local unemployment rates are conditionally exogenous to household fertility outcomes. However, there are

[16] To be precise; t-1 refers to the year of conception.
[17] The inclusion of county-specific quadratic time trends was proved to be unimportant.
[18] The analysis for the age-gender-specific unemployment was performed with a similar specification in which both lagged male -MaleUnemp$_{gc(t-1)}$ – and female -FemaleUnemp$_{gc(t-1)}$ – unemployment rates are included in the same regression.

certain concerns associated with the use of unemployment rates as exogenous regressors. One of these is that local unemployment rates might be correlated with changes in other unobserved variables that may affect the fertility decision of individuals. Second, there may be a positive correlation between fertility and local labor supply. If birth rates increase due to the changes in local labor supply, then the unemployment measure may be picking up this relationship rather than the effect of local labor market demand. Lastly, the International Labour Organization (ILO) definition of unemployment may not be able to capture the full extent of the local labor market conditions, causing a measurement error. The estimation strategy to deal with these problems is to specify a variable that can account for demand-induced variation in unemployment, and can thus be used to obtain an unbiased estimate of the effect of unemployment on fertility. To this end, the IV approach was adopted to explore the robustness of the OLS results. The predicted unemployment rates were built based on the work of Bartik (1991), Blanchard and Katz (1992), Schaller (2015) and Anderberg et al. (2015) for the UK case, where the initial industry composition of employment is interacted with the corresponding national industry-specific trends in unemployment. In particular, the local industry composition by gender and age group at baseline, defined as the calendar year 1993, is combined with industry-specific unemployment rates by gender, age group and time at the national level over the sample period.[19] For each county, age group, gender and year industry predicted unemployment rates are constructed as follows:

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[19] Eight industries are used in the analysis based on a condensed version of the UK Standard Industrial Classification of Economic Activities, SIC (2007):"Agriculture, forestry, fishing, mining, energy and water supply", "Manufacturing", "Construction", "Wholesale, retail & repair of motor vehicles, accommodation and food services", "Transport and storage, Information and communication", "Financial and insurance activities, Real estate activities, Professional, scientific & technical activities, Administrative &support services", "Public admin and defense, social security, education, human health & social work activities", "Other services". The "industry unemployment rate" is defined as the unemployed by industry of last job as percentage of economically active by industry.

$$PredictedUnemp_{ghjt} = \sum_{k} \psi_{ghjk} UNEMP_{ghkt}$$

where ψ_{ghjk} is the share of industry k among employed individuals of age group g, gender h, county j at baseline, and where $UNEMP_{ghkt}$ is the unemployment rate, at the national level, in industry k for individuals of age group g, gender h and in time period t. Given that the predicted unemployment measure is a weighted average of the national industry-specific unemployment rates, these weights reflect the baseline local industry composition in the relevant gender and age group.

The IV approach has a number of attractive features. Most importantly, the estimates cannot be affected by contemporaneous omitted variables since the only local input into the predicted unemployment rates is the industry structure at baseline and these rates cannot be related to any contemporaneous (during the sample period) omitted variables. Furthermore, as mentioned earlier, birth rates in a county are a function of both local labor supply and labor demand. It is for this reason that the use of observed changes in local labor market confounds the results. Instead, the IV uses labor demand shocks as an identifying source of variation and act as an exogenous change in local labor demand. Additionally, the predicted unemployment rate during the next period relies only on initial local industry composition and national-level industry-specific unemployment rates that influence the gender composition of employment opportunities. However, one might be concerned for the earlier time periods of the panel. Later in this paper, this issue is investigated by dropping some of the years at the beginning of the sample period from the estimated model. This estimation does not send any warning signals that the main results are substantially affected by underlying serial correlation in county-specific circumstances.

3.4 Results

3.4.1 OLS Specifications

Table 3.2 presents the results from the OLS estimation. Column 1 reports the estimation with all fixed effects included; column 2 adds basic demographic characteristics (education, ethnicity, country of birth and partnership status) and an additional control variable (house prices); column 3 adds county-specific linear time trends; and column 4 adds age group-county-specific linear time trends.

The specification in the first column yields a positive and statistically insignificant coefficient of 0.008. After adding more controls and time trends, the results consistently show that the overall unemployment rate is positively associated with fertility rates. To assess whether different age groups are more likely to move in response to an economic shock, I include age group-county-specific linear time trends in column 4. The coefficient of the main interest remains similar to the ones in columns 1–3, implying that age groups do not systematically move into different counties in response to adverse labor market conditions. According to the estimate from the fully saturated model in column 4, a 1 percentage point increase in the unemployment rate is associated with a 1.3% increase in birth rates, which is significant at the 1 % level. This finding suggests that fertility moves countercyclically, and, accordingly the substitution effect dominates any negative income effect over the sample period. The main reasons for this finding are twofold: On the LFP side, the difference between the LFP of men and women has shrunk considerably since the mid-1990s. On the earnings side, the gender pay gap has been following a downward trend.[20]

[20] More specifically, the LFP difference between men and women fell from 14.5% age points in 1994 to 8.6 percentage points in the final quarter of 2011. In terms of the gender pay gap, based on median hourly earnings excluding overtime, it has narrowed for full-time employees, to 9.1% compared with 17.4% in 1997. The gap for all employees has also followed a downward trend to 19.5 %, down from 27.5% in 1997. In addition, Polachek and Xiang (2014) find that the pay gap in England is declining relatively more quickly than in other countries.

the rest of the analysis is carried out based on the specification in column 3, in which I control for observable demographic characteristics and house prices. I also exploit the panel aspects of the data by including county and year fixed effects as well as the county-specific linear time trends.

The analysis continues with a stratification of the regressions with age-specific characteristics so as to gain further insight into the demographic basis of this result. The upper panel of Table 3.3 suggests that the fertility is the most responsive to unemployment rates between the ages of 25 and 34. In this age group, a 1 percentage point increase in the unemployment rate leads to a 1.32% increase in fertility. The incidence of youth unemployment also has a positive and marginally significant impact on fertility. The older group, 35–44, by contrast, shows no significant impact of unemployment on birth rates. Altogether, the results in this table suggest that age groups react differently when they are exposed to local unemployment shocks. The next table, therefore, proceeds to further examine the age group–gender characteristics.

The differences between age- and gender-specific unemployment rates are highlighted in Table 3.4 in which I expect to find that female unemployment will be positively associated with fertility, whereas male unemployment will have the opposite impact. Overall, the results are in line with this notion. For example, a 1 percentage point increase in female unemployment leads to a 1.29% increase in births for the prime age group. For men, the negative and significant effects are concentrated among the younger cohorts, 16–24 and 25–34. However, for the age group 35–44, there are insignificant coefficients on unemployment for both males and females. This finding could be because there is a weak relationship between older cohorts' fertility decision and their labor market status. The following section of the paper is concerned with endogeneity of unemployment, and IV estimation results are presented.

3.4.2 IV Estimation

As discussed in Section 3.3.2, the potential sources of bias in the OLS results are the possibility of reverse causality and/or some unobservables that are affecting fertility rates other than unemployment. I start by considering the full sample and estimate the effect of overall unemployment (aged 16–44) on fertility. Confirming the OLS findings in Table 3.2, columns 1–4, I find that the IV coefficients have the same sign; however, they are larger in magnitude. In the first stage, the predicted unemployment rates are significantly correlated with the endogenous variable and in the expected direction. Since the IV estimates can be interpreted as the impact of a fall in local labor demand on fertility, finding higher coefficients suggest that these estimates reflect the local spillovers in unemployment. In other words, they capture: (1) the main effect due to being unemployed, (2) the risk of being in unemployed in the near future and (3) expectations about future wage growth.

In order to quantify the different effects of the predicted unemployment rate, I proceed to separately estimate the relationship across age groups and age group–gender cohorts. In Table 3.3, the point estimates on age groups are all positive and statistically significant at conventional levels. Column 2, in bottom panel of Table 3.4, strongly confirms previous findings that male and female unemployment have different impacts on fertility and reveals the importance of sub-group characteristics. In parallel to the main hypothesis of this paper, the results imply that a one percentage point increase in male unemployment leads to a 2.18% decrease in birth rates whereas same amount of increase in female unemployment leads to a 6.26% increase in birth rates for the age group 25–34. This shows that the positive effect of female unemployment on fertility is much larger for prime aged women. The results for men again show that unemployment has a significant negative effect on fertility, with the effect being stronger at later ages. Taken together, these results suggest that unemployment is an important determinant of fertility behavior. The following part of the results proceeds to explore whether the responses are homogenous across demographic characteristics.

3.4.3 Analysis by Demographic Characteristics

Having detailed information in UK LFS enables a filtering out of the effects of unemployment on fertility by education, country of birth and marital status. In order to allow the point estimates on county and year fixed effects to vary across subgroups, I construct covariates and estimate the relationship for each age group. In Table 3.5, columns 1–3 present estimates for the OLS specifications and columns 4–6 present estimates for the IV specifications. Because the cell size in industry employment composition by demographic characteristics is small at baseline, some IV estimates have lower number of observations. However, this does not cause excluded variables to show a weak partial correlation with unemployment. Although not reported due to space restrictions, the values for first-stage F-statistics are consistently higher than 10. Nonetheless, one caveat is that the measurement error of group-specific unemployment rates may lead to biased estimates, so these results should be interpreted with caution.

Looking at the estimates presented by educational attainment in Panel (A) of Table 3.5, I find that the coefficients on unemployment are negative and significant for the 16–24 age group, but they are positive for older cohorts who are highly educated. In the further education category, unemployment also seems to be positively associated with fertility. The fact that the effect is more pronounced on the degree level may be attributable to the fact that women take advantage of the low opportunity cost of childbearing to prevent future career interruption which occurs due to the transition to motherhood. The IV estimates again are larger in magnitude and tend in the same direction.

Panel B shows results separately for UK-born and non-UK-born cohorts. Coefficients on unemployment for non-UK borns in both OLS and IV estimations are negative and mostly larger in magnitude. This result may be explained by the fact that immigrants are more sensitive to cyclical increases in unemployment than those of natives. For UK borns, the effects of unemployment on fertility tend to be positive.

Turning now to the evidence on partnership in panel (C), I detect clear differences by marital status. Contrary to my expectations, I find that unemployment has an only negligible influence on fertility at younger ages, but it exhibits stronger association among older cohorts. Indeed, the coefficient of unemployment is negative and significantly different from zero for singles at more advanced ages. Looking at the results for married cohorts, unemployment is positively associated with fertility and the effect is mostly concentrated at the prime age cohort.

3.4.4 Robustness Checks

Additional robustness checks are conducted in order to detect whether the main findings remain stable to different specifications. The first column in Table 3.6 presents results for teenage fertility. Results in the upper panel of column 1 indicate that unemployment has a positive and marginally significant impact on birth rates. As can be seen from the bottom panel, both male and female unemployment maintain the expected signs of direction. Although the effects are small in magnitude, the results are still in line with the main findings of this paper.

Between columns 2 and 4, I further discuss the important issue of the validity of my instrument, which is introduced in Section 3.4.2. The strong correlation of my instrument with the endogenous variable is apparent from all the first-stage results presented in Appendix Table 3.1. Although the IV results seem robust to a number of alternative specifications, one may argue that unemployment is not necessarily a purely demand-driven measure and is also affected by the changes in labor supply. Finally, I build on my instrument and use employment growth index as an alternative measure to assess the robustness of my results. The measure reflects exogenous labor demand for females and males in each age group and is constructed as follows: for each county, age group and year, I start with a variable measuring the proportion of employment based on local industry structure at baseline. Next, similar to Aizer (2010), I construct an annual employment growth index for each gender and age group, by interacting baseline measure with national trends in employment growth rates in industries dominant in the county, then collapsing over industries within each county– year–age group cell:

$$EmpGrowth_{ghjt} = \sum_{k} \psi_{ghjk} EMPGROWTH_{ghkt}$$

where ψ_{ghjk} is the share of industry k among employed individuals of age group g, gender h, county j at baseline, and where $EMPGROWTH_{ghkt}$ is the industry employment growth rate, at the national level, in industry k for individuals of age group g, gender h and in time period t. The first-stage results show that all measures of unemployment are well correlated with the instruments. On the whole, the IV estimates presented in Table 3.6 alleviate concerns about the biased estimates induced by the changes in labor supply and are consistent with the baseline results.

Reassuringly, I find that: (1) fertility moves counter-cyclically over the business cycle; (2) the prime age group is more responsive to changes in unemployment; (3) male unemployment is negatively and female unemployment is positively associated with births, with these effects often being smaller and mostly significant. Finally, the results introduced in columns 5–7 show that the outcome is not driven by the serial correlation and present estimates after dropping the first 4 years of the sample. Across all age groups, the coefficients consistently remain qualitatively unchanged. Taken together, I obtain results that are very similar to the baseline estimates and provide meaningful insights on the countercyclical nature of births along with group-specific differences across social strata in England.

3.5 Conclusion

The aim of this study was to assess the causal effect of unemployment on fertility in England and to evaluate how this effect varies across sub-demographic groups. Its main finding is that the substitution effect dominates the income effect at the aggregate level, implying a countercyclical fertility pattern that may be attributable to changes in female labor market outcomes over the sample period. Additionally, the relevance of gender-specific unemployment is clearly supported by the current findings and the results confirm the main predictions of dynamic fertility models. The findings indicate that female unemployment tends to increase fertility, as women take advantage of the low opportunity cost of childbearing in the form of mothers' time. Male unemployment goes in the opposite direction, which implies an income effect. Returning to the questions posed at the beginning of this study, it is now possible to state that a comparison of age groups reveals that unemployment is more likely to affect the fertility of younger cohorts, rather than older ones. A speculative reason for this is the possibility that the former are more able to postpone their fertility until economic conditions recover, while labor market conditions may play a less important role for couples whose “fertile” lifetime is nearing its end. The expected variation in the unemployment and fertility relation by educational attainment, marital status and country of birth is also documented. Although the study has successfully demonstrated the aforementioned findings, it is however limited by the use of total birth rates, and the findings cannot be transferable to birth orders. Future research should therefore concentrate on the investigation of birth orders while considering demographic subgroup characteristics.

All in all, the present study confirms previous findings and provides additional evidence suggesting the existence of strong variation across sub-demographic groups, while showing that different age groups and genders react differently to local unemployment shocks.

Figure 3.1: Mean Unemployment Rates and Birth Rates Across Counties

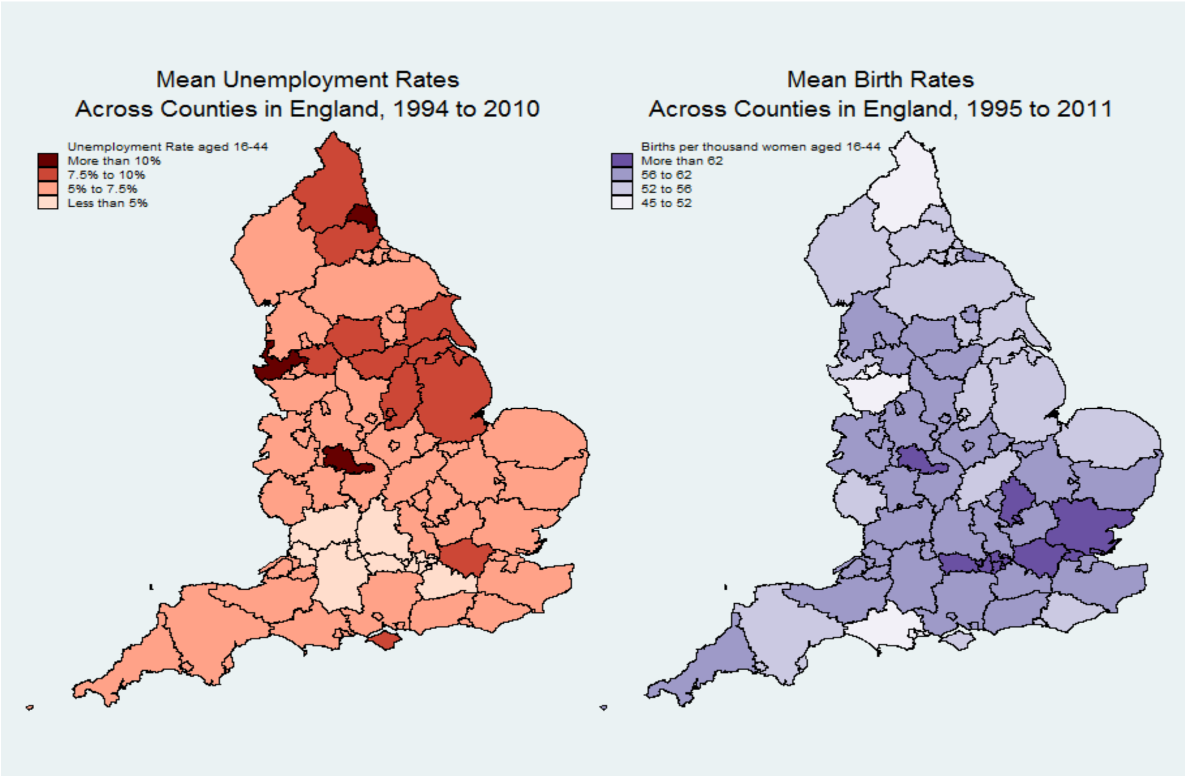

Figure 3.2: Age-Specific Unemployment Rates and Fertility Rates

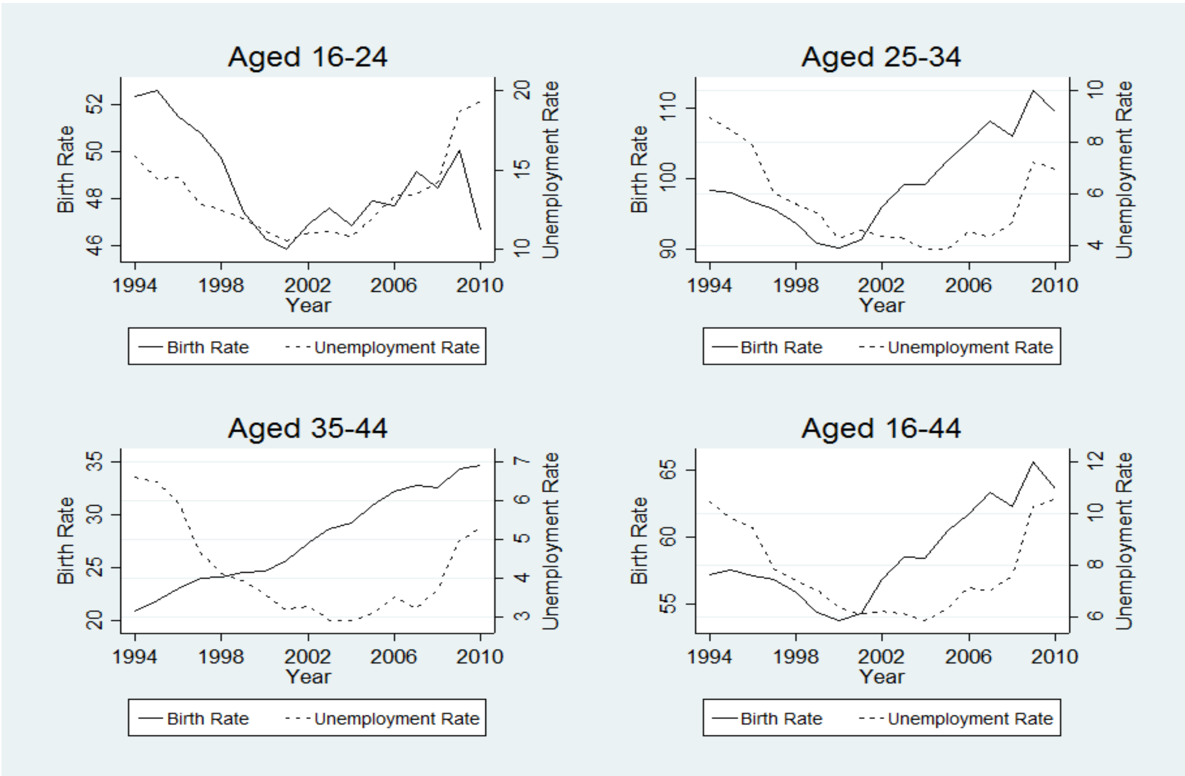

Table 3.1: Summary Statistics

Variables	Mean	Standard Deviation
Birth Rates by Age Group		
Aged 16-24	48.69	8.705
Aged 25-34	99.61	11.23
Aged 35-44	27.74	7.689
Unemployment by Age Group		
Aged 16-24	13.40	4.512
Aged 25-34	5.772	5.314
Aged 35-44	4.197	2.064
Unemployment by Age Group & Gender		
Female Aged 16-24	11.46	4.450
Female Aged 25-34	5.207	2.476
Female Aged 35-44	3.961	1.861
Male Aged 16-24	15.09	5.604
Male Aged 25-34	5.922	3.582
Male Aged 35-44	4.384	2.790
Single	0.500	0.060
Married	0.418	0.058
Divorced/Widowed	0.082	0.014
UK Born	0.910	0.071
Non UK Born	0.090	0.043
White	0.931	0.077
Other Ethnicities	0.090	0.058
Higher Education	0.238	0.061
Further Education	0.245	0.028
Compulsory Education or less	0.517	0.060
House Prices	£145,482	£112,045
N	2,397	

Notes: The table provides within cell means for 47 counties used in the baseline specification. House prices are CPI adjusted to 2005 pounds.

Table 3.2: Effect of Unemployment on Fertility - Alternative OLS & IV Specifications

	(1) OLS	(2) OLS	(3) OLS	(4) OLS	(5) IV	(6) IV	(7) IV	(8) IV
Unemployment Rate	0.0085	0.0131***	0.0141***	0.0130***	0.0110	0.0252***	0.0266***	0.0263***
	(0.0060)	(0.0032)	(0.0035)	(0.0036)	(0.0074)	(0.0065)	(0.0067)	(0.0071)
Higher Education		0.0025	0.0026	0.0055		0.0002	0.0021	0.0045
		(0.0039)	(0.0038)	(0.0038)		(0.0037)	(0.0034)	(0.0035)
Further Education		-0.0142***	-0.0153***	-0.0028		-0.0159***	-0.0160***	-0.0034*
		(0.0019)	(0.0019)	(0.0019)		(0.0021)	(0.0019)	(0.0019)
Single		0.0053	0.0033	0.0029		0.0073**	0.0055	0.0036
		(0.0044)	(0.0046)	(0.0051)		(0.0034)	(0.0041)	(0.0047)
Non UK Born		0.0058	0.0101**	0.0202***		0.0032	0.0091**	0.0182***
		(0.0043)	(0.0041)	(0.0047)		(0.0040)	(0.0041)	(0.0046)
Other Ethnicities		0.0436***	0.0517***	0.0372***		0.0372***	0.0475***	0.0349***
		(0.0123)	(0.0146)	(0.0127)		(0.0123)	(0.0144)	(0.0133)
House Prices		-0.0003	0.0005*	0.0006**		0.0003*	0.0011***	0.0012***
		(0.0005)	(0.0003)	(0.0003)		(0.0002)	(0.0002)	(0.0002)
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
County Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age Group Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
County Specific Trends	No	No	Yes	Yes	No	No	Yes	Yes
County-AgeGroup Trends	No	No	No	Yes	No	No	No	Yes
N	2397	2397	2397	2397	2397	2397	2397	2397

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parentheses and clustered at age group-county level. Dependent variable: log fertility rate by year, age group and county. Fertility rates are constructed by dividing the number of births by the relevant female population using mid-year population that are based on censuses. House prices (10,000s) are CPI adjusted to 2005 pounds. Unemployment rates are calculated based on the ILO definition. All specifications are weighted by the total number of women in each cell.

Table 3.3: Effect of Unemployment on Fertility - Age Group Specifications

	(1)	(2)	(3)
	Aged 16-24	Aged 25-34	Aged 35-44
	OLS	OLS	OLS
Unemployment rate	0.0044***	0.0132***	0.0011
	(0.0008)	(0.0011)	(0.0019)
	IV	IV	IV
Unemployment rate	0.0087***	0.0274***	0.0068*
	(0.0018)	(0.0028)	(0.0031)
Year Fixed Effects	Yes	Yes	Yes
County Fixed Effects	Yes	Yes	Yes
County Specific Trends	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes
House Prices	Yes	Yes	Yes
N	799	799	799

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 3.2.

Table 3.4: Effect of Unemployment on Fertility – Age Group-Gender Specifications

	(1)	(2)	(3)
	Aged 16-24	Aged 25-34	Aged 35-44
	OLS	OLS	OLS
Male unemployment rate	-0.0011***	-0.0089***	-0.0036
	(0.0003)	(0.0029)	(0.0054)
Female unemployment rate	0.0044***	0.0129***	0.0038
	(0.0008)	(0.0010)	(0.0022)
	IV	IV	IV
Male unemployment rate	-0.0032***	-0.0218***	-0.0189
	(0.0011)	(0.0067)	(0.0116)
Female unemployment rate	0.0099***	0.0626***	0.0483***
	(0.0032)	(0.0099)	(0.0063)
Year Fixed Effects	Yes	Yes	Yes
County Fixed Effects	Yes	Yes	Yes
County Specific Trends	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes
House Prices	Yes	Yes	Yes
N	799	799	799

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 3.2.

Table 3.5: Effect of Unemployment on Fertility - Demographic Group Specifications

		(1) Aged 16-24	(2) Aged 25-34	(3) Aged 35-44	(4) Aged 16-24	(5) Aged 25-34	(6) Aged 35-44
(A) Education		OLS	OLS	OLS	IV	IV	IV
Degree level or above	Unemp. Rate	-0.0272***	0.0111**	0.0027***	-0.0962***	0.0438***	0.0066**
		(0.0090)	(0.0046)	(0.0008)	(0.0373)	(0.0123)	(0.0025)
Further Education	Unemp. Rate	-0.0080**	0.0053**	-0.0008	-0.0275**	0.0192*	0.0024
		(0.0029)	(0.0023)	(0.0028)	(0.0108)	(0.0086)	(0.0081)
Comp. Edu. or below	Unemp. Rate	-0.0052*	0.0056	0.0038*	-0.0204**	0.0178*	0.0254
		(0.0025)	(0.0036)	(0.0018)	(0.0086)	(0.0086)	(0.0161)
	N	750	750	750	750	750	750
(B) Country of Birth		OLS	OLS	OLS	IV	IV	IV
UK Born	Unemp. Rate	0.0075***	0.0111***	-0.0012	0.0177***	0.0277***	0.0107
		(0.0008)	(0.0019)	(0.0080)	(0.0016)	(0.0058)	(0.0131)
Non UK Born	Unemp. Rate	-0.0328***	-0.0088***	-0.0014**	-0.0896***	-0.0249***	-0.0023**
		(0.0059)	(0.0023)	(0.0006)	(0.0079)	(0.0068)	(0.0010)
	N	764	764	764	764	764	764
(C) Marital Status		OLS	OLS	OLS	IV	IV	IV
Single	Unemp. Rate	-0.0002	-0.0026***	-0.0031**	0.0006	-0.0037***	-0.0063***
		(0.0002)	(0.0006)	(0.0012)	(0.0011)	(0.0013)	(0.0021)
Married	Unemp. Rate	0.0001*	0.0054**	0.0003**	0.0003	0.0079***	0.0017**
		(0.0000)	(0.0020)	(0.0001)	(0.0006)	(0.0029)	(0.0006)
	N	799	799	799	799	799	799

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 3.2.

Table 3.6: Effect of Unemployment on Fertility – Robustness Checks

	(1) Pred. Unemp. Rates	(2) Employment Growth Index	(3) Employment Growth Index	(4) Employment Growth Index	(5) Exclusion of Years 95-98	(6) Exclusion of Years 95-98	(7) Exclusion of Years 95-98
	IV	IV	IV	IV	IV	IV	IV
	Aged 16-19	Aged 16-24	Aged 25-34	Aged 35-44	Aged 16-24	Aged 25-34	Aged 35-44
Unemployment Rate	0.0016*	0.0039***	0.0109***	-0.0012	0.0038**	0.0225***	0.0042*
	(0.0007)	(0.0012)	(0.0033)	(0.0008)	(0.0017)	(0.0068)	(0.0020)
Male Unemp. Rate	-0.0002**	-0.0040***	-0.0335***	-0.0008***	0.0003**	-0.0255***	-0.0006*
	(0.0001)	(0.0009)	(0.0106)	(0.0002)	(0.0001)	(0.0093)	(0.0003)
Female Unemp. Rate	0.0011**	0.0066***	0.0521***	-0.0012	0.0034***	0.0499***	0.0138**
	(0.0004)	(0.0012)	(0.0125)	(0.0012)	(0.0007)	(0.0044)	(0.0052)
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
County Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
County Specific Trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Demographic Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
House Prices	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	799	799	799	799	611	611	611

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 3.2.

Chapter 4

The Effects of House Prices on Birth Rates:

New Evidence from England

4.1 Introduction

The price changes in the housing market have been a much-debated topic in the United Kingdom over the past years. According to the Land Registry data, average house prices have increased by nearly 290 per cent between 1995 and 2013, from £67,000 to £234,000, in England with some counties experiencing a more than 900 per cent rise in house prices. Subsequently, these strong movements in the property market generated inevitable effects on households' wealth and disposable income. For example, the British newspaper *The Guardian* reported in a recent (2015) article that:

"Based on Land Registry and HMRC data, a homebuyer earning the median salary for their region in 1995 would have had to spend between 3.2 times and 4.4 times their salary on a house, depending on where they lived. In 2012-13, the median house price had risen to between 6.1 times and 12.2 times median regional incomes."

In 2013, *The Telegraph*, another British newspaper, argued that household consumption was sensitive to changes in housing wealth:

"Rising house prices in many areas of the country are creating an opportunity for homeowners who are asset rich but cash poor to release equity from their homes."

In 2011, the BBC quoted the chief executive of the Shelter, a housing charity, who said:

"We have become depressingly familiar with first-time buyers being priced out of the housing market, but the impact of unaffordable rents is more dramatic. With no cheaper alternative, ordinary people are forced to cut their spending on essentials like schools and families."

Lastly, in a brief article in 2007, the BBC cited a reader's comment to reflect the importance of the price and availability of housing on families:

"House prices are currently the main obstacle to stable family life and stable parenting."

Indeed, this unique housing market structure, and its close link to household spending decisions, suggest a plausible connection between housing tenure and having children in England. In an attempt to unravel this relationship, I investigate how birth rates are affected by changes in house prices separately for home owners and renters through income and price effects. There are two ways of looking at this interaction. On the one hand, given that housing wealth is a major component of the household's assets for home owners and empirically housing wealth and consumption tend to move in the same direction (for example, Campbell and Cocco, 2007; Case, Quigley and Shiller, 2005 & 2011), an increase in house prices should lead to an increase in demand for children, if children are a normal good.[1] For tenants, on the other hand, higher rent payments may force them to reduce their consumption in line with the increase in cost of housing and should generate a negative price effect on the demand for children.

Despite the logical appeal of these arguments, to date there have been few studies that set out to present an analysis of this relationship. The current available evidence mainly comes from the United States (for example, Lovenheim and Mumford, 2013; Dettling and Kearney, 2014) and from Hong Kong (Yi and Zhang, 2010).[2] However, a further analysis is necessary as the housing market in England stands out in international

[1] Becker's (1960) quality and quantity approach suggests that a rise in income may have depressing effects on birth rates if income elasticity for the number of children is considerably lower than that for quality of children.

[2] I do not review the fertility literature in detail here. For studies using US data, see Butz and Ward (1979), Currie and Schwandt (2014) and Schaller (2016). Studies using data from other countries include, among others: Adsera (2005), Ahn and Mira (2002), Aksoy (2016), Ozcan, Mayer and Luedicke (2010). See Sobotka, Vegard, and Philipov (2011) for a wider review of the earlier literature.

context. It is mainly because: (i) houses in England are overvalued, and the prices are among the highest in the world (Kuenzel and Bjørnbak, 2008); and (ii) prices exhibit extreme volatility – real house prices in England as a whole are substantially more volatile than in the most volatile metro areas in the United States (Hilber and Vermeulen, 2016); (iii) in terms of size, houses are substantially smaller than in the United States and any other European countries (Morgan and Cruickshank, 2014); (iv) house prices have risen faster than in any other OECD country and have far outstripped earnings growth (Hilber, 2015); and (v) private sector rents are the highest in Europe, and renters spend nearly 40 per cent of their income on paying their rent in comparison to the European average of 28 per cent (National Housing Federation, 2015). Therefore, special attention should be devoted to England as it exhibits different housing market characteristics than other countries.

Another reason for concern is that some previous studies are subject to a possible endogeneity bias. For example, assuming that the housing supply is fixed in the short term, if people who plan to have a child and expand their families demand larger houses, then the direction of causation runs from birth rates to house prices. This implies a case of reverse causation. In addition, there are strong reasons to suspect that the existing studies suffer from omitted variable bias. Since house prices for dwelling units are determined by a number of factors, the failure to control for demand shifters that play a part in affecting house prices as well as birth rates may bias estimates of the house price-fertility rate nexus. For instance, holding all else equal, a rise in household liquid assets (that is, non-housing wealth) or one-person family formation, may potentially affect both constructs. In order to establish a causal relationship and be able to assess the effect of house price on birth rates, I propose a clear identification strategy that takes the aforementioned factors into account.

This paper extends the emerging literature on house prices and birth rates by using novel source of variation from England. More specifically, I exploit the regulatory constraints on the housing market by refusal rates for major development projects in English counties, which is generated by outdated planning regulations. The current planning system maintains many of the mechanisms introduced in the Town and Country Planning Act of 1947 which has been characterized as quite complex and inefficient, preventing new housing developments coming forward. This constrained housing supply can be attributed to the lack of incentives of local authorities to respond to the planning permission applications in a positive and timely manner. There are several reasons put forward to explain how regulatory constraints can affect house prices. For example, Barker (2004) suggests that the housing supply in England does not respond to price signals, mostly because of the constraints embedded in the planning system. Rouwendal and Vermeulen (2007), and Saiz (2010) show that restrictions on land supply have negatively affected new housing construction despite the increases in house prices. Similarly, Hilber and Vermeulen (2016) argue that planning rules are particularly important in driving up house prices in England and have a substantive positive impact on the house price-earnings elasticity.

In this paper, I use rich data sets of English counties spanning from 1996 to 2014 to investigate the effects of house prices on birth rates. I first present OLS results of how fertility rates respond to increases in house prices, estimated separately for home owners and renters. To address omitted variables bias I (i) extensively control for demographic characteristics, and (ii) take full advantage of the panel data structure by including county, year and age group fixed effects as well as county-specific linear and quadratic time trends. While the net effect of house prices on birth rates is negative, the OLS estimation establishes a strong positive relationship between house prices and fertility rates among home owners and a negative relationship among renters. The association between house prices and birth rates is stronger in supply constrained areas, especially in south-east England, including London. Moreover, results from the demographic subgroups and alternative house price measures analyses also confirm the main empirical findings. Lastly, the IV specifications yield coefficients that are larger in magnitude in which the fully saturated specification indicates that a 10 per cent increase in house prices is associated with 4.9 per cent fall in births among renters and 2.8 per cent increase in births among home owners. At the mean ownership rate the net effect is negative and leads to a 1.3 per cent fall in births.

The contribution of this paper to the literature is three-fold. First, I provide evidence on the impact of house prices on fertility rates from a country with a highly restricted housing market, which has not been studied previously. These findings provide new evidence on how country-specific factors influence birth rates by showing the importance of housing market regulations. Second, I deal with the endogeneity of house prices in a robust manner and I exploit a novel source of variation, that is refusal rates for major development projects. This direct measure of planning restrictiveness makes a distinctive contribution to the literature and distinguishes the current study from previous papers which rely on survey data. Third, I carefully disentangle the impact of house prices from other factors and tackle the possible endogeneity of house prices with respect to internal migration (that is, mobility) by accounting for net population changes for each county-year cell. The focus on England provides a distinct addition to the literature and this research will further our knowledge about an increasingly important and understudied socio-economic problem.

The remaining part of the paper proceeds as follows: the next section discusses the planning system and process in England. Section 4.3 introduces the data description and sources. The fourth section presents the empirical strategy and the fifth section presents the results. Lastly, section 4.6 offers a discussion and concludes.

4.2 The Planning System and Process in England

In 1947, the UK parliament legislated a bill called The Town and Country Planning Act which nationalized the right to develop land and established a new system, indicating that individual planning permission was needed from a local authority for all land development and housing projects. The legislation has continued to evolve on a number of occasions with the last major changes made in 2011, namely by the Localism Act (2011), in an attempt to move towards decentralized planning.[3]

The planning system delegates regional planning bodies (RPBs) to oversee land use and number of houses being built in England. In this centralized system, the RPBs decide the allocation of land and impose housing targets on local (planning) authorities through the five year plans (to avoid confusion, I will henceforth replace the

[3] Over the years, the planning system has undergone a number of major changes, the titles of which are as follows: Town and Country Planning Act (1990) introduced the concept of "plan-led development", Planning and Compulsory Purchase Act (2004) initiated compulsory acquisition of land for development and other planning purposes, Planning Act (2008) aimed to accelerate the process for granting permissions for major new infrastructure projects such airports, roads, dams, energy facilities and so on.

term local (planning) authority by county).[4] The local planning committees then allocate land use based on planning applications received. As outlined by the regulations, the standard process for obtaining planning permission is as follows. Upon receiving applications, the counties publicize the projects to start the formal consultation phase which lasts for at least three weeks. After this period, a county needs to decide within 8 weeks for minor projects, for example household projects cases, and within 13 weeks for major development projects such as large housing sites.[5]

However, granting planning permission is not as straightforward as it may seem. Despite the various alterations made over time, the regulatory framework is still identified as one of the main barriers to new housing supply (see Cheshire and Sheppard, 2002; Barker, 2004 and 2006; Evans and Hartwich, 2005, Hilber and Vermeulen, 2016) and characterized by three main properties: (i) centralized land use approach and complications in determining the level of housing supply in counties; (ii) complex application process and under-resourced planning departments; (iii) the lack of financial incentives for counties to deliver essential infrastructure and amenities.

Although the primary goal of the centralized system is to encourage the housing supply, this top-down system fails to take annual housing demand changes and associated price movements into account. Indeed, the Department of Communities and Local Government's (DCLG) 2011 report suggests that regional strategies have antagonized local residents, setting them against development plans and housing targets. Such opposition is mainly driven by concerns that adequate public services (road,

[4] In England the local planning authorities are 32 London borough councils, 36 metropolitan borough councils, 201 non-metropolitan district councils, 55 unitary authority councils, the City of London Corporation and the Council of the Isles of Scilly.
[5] This time frame initiated by the Labour government in 2002 in order to shorten the processing time.

public transportation, hospital, schools and so on) will not be provided and house prices will be gradually eroded. Consequently, the planning authorities mainly focus on the interests of the local residents and tend to preserve the land from development, which distorts the local housing market equilibrium.

Ineffectiveness of the current housing planning practices also arises from the fact that authorities do not have sufficient resources and planners (such as technical staff and legal services) to respond to the complex application process in a timely manner.[6] For example, household projects and major development projects are subject to the same application procedure, one that often leaves too few resources available to properly evaluate the large-scale development applications. This, in turn, results in delays or refusal of submissions in order to meet targets set by the central system.[7] Political concerns are also widely embedded in local level planning. Cambridge Centre for Housing and Planning Research (CCHPR) (2014) – commissioned by the DCLG – reports that elected members (that is councilors – representatives elected to a county) may refuse applications that planning officers have recommended for approval based on non-technical grounds.[8]

Turning to the lack of financial incentives, counties are unwilling to support housing growth due to lack of funding for infrastructure provision. In the current system, the government uses different resources (such as census data, school

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[6] In addition to the public consultation, local planning authorities have to consult a number of departments (for example the local highway authority, environment agency, English Heritage, health and safety executive and so on) who may be influenced by a proposed housing project.

[7] According to the Home Builders Federation, as of January 2015, more than 150,000 applications were waiting to advance from the "outline" planning stage to the "detailed" planning stage.

[8] Local councilors have a conclusive say in planning decisions made in their area. Each planning committee consists of local councilors who come from different political parties. When a committee makes a decision, councilors usually consider (though they are not lawfully obliged to do so) a recommendation report written by planning staff. Councilors may accept the recommendation directly or discuss the application further before deciding to accept or refuse the application.

enrollment and so on) to determine the amount of funding.[9] However, these data sources are mainly backward-looking and are not updated annually. Consequently, the funding stream is not based on the actual needs of the local communities, and disadvantages counties that experience rapid housing growth (Barker, 2004). In other words, new housing developments impose a formidable fiscal burden on counties and the internalization of the associated cost by local residents.

As briefly introduced above, the planning structure is not able to cope effectively with the house prices in response to changes in demand, causing a mismatch between the goals of the central system and those of local communities. I therefore propose to use refusal rates for major development projects to capture the effects of regulatory constraints on house prices. This serves as a "catch-all" variable and combines land scarcity, policy restrictiveness, and local land use policy measures. The advantages and disadvantages of this measure will be further discussed in the following section.

4.3 Data

The data used in this paper come from

- the UK Land Registry
- the Department for Communities and Local Government (DCLG)
- the confidential version of the UK Labour Force Survey (LFS)
- the Office for National Statistics (ONS)
- the Family Resources Survey (FRS).

[9] The UK government uses the "four-block" model (the relative needs, the relative resource, the central allocation and the floor damping) to allocate funds to counties since 2006. The calculations mainly rely on demographic, economic and social data to define the share of revenue spending for each county given a fixed national total. This model also recognizes the differences in the amount of local income which individual counties have the potential to raise. The full details of the calculations can be found in the *Local Government Finance Report* which is published annually by parliament.

The level of analysis is age group-county-year cell and the details on how the balanced panel data was constructed are provided below.

4.3.1 Data on House Prices and Birth Rates

I obtain house price index (HPI) data from the UK Land Registry over the period 1995-2013. This index is based on actual residential sales transactions and includes information on more than 24 million sales and uses over 7 million identified pairs to compile the index. More specifically, it takes the difference in prices between January 1995 and the last time sale price, then averages these changes at county level. In addition, it accounts for seasonality as well as depreciation or appreciation of properties over time and is designed for "like for like" evaluation between houses.

It uses a sample size that is larger than all other data available and also has a number of other advantages.[10] First the repeat sales method minimizes the concerns about unobserved heterogeneity in house prices through following the same property over time. Second the index includes all cash and mortgage transactions and third it is based on actual property transactions, in other words, it eliminates the bias that arises from mortgage valuation approvals (see Chandler and Disney, 2014).

I obtained the decisions on major (residential) development projects data from the DCLG and constructed the refusal rates based on the following definition: the proportion of development projects with 10 or more dwellings that was refused by a planning authority. To construct the panel data, I matched counties and local planning

[10] The sample size in ONS, Nationwide and Halifax house price indices are considerably smaller. It is because they are solely based on mortgage lending and exclude cash transactions. Further information about the index can be found at: https://www.gov.uk/guidance/about-the-price-paid-data.

authorities based on 2001 borderlines.

4.3.2 Time Variant County Characteristics, Homeownership Rates, and Birth Rates

I control for a number of time-variant determinants of birth rates in my main specification. More specifically, for each county-age group-year cell, I control for gross weekly wages, unemployment rate, gross household assets excluding housing wealth, share of college graduates, share of foreign born population, share of other ethnicities, share of one-person family, share of non-married individuals and net internal migration (that is, net change in population by a county in a given year).[11]

I use the LFS to construct demographic characteristics, labour market controls and information on housing tenure. The LFS is a nationally representative survey of the whole population of the United Kingdom and the current sample size is about 41,000 households for every quarter. The main advantage of the LFS in the current framework is that the low tier geographical classification allows for aggregation and matching the corresponding data at the local planning authority level. Mean home ownership rates are constructed for each age group-county cell and kept constant at baseline to minimize endogeneity concerns and to account for compositional changes over time. It is important to note that home ownership rates have remained largely unchanged over the sample period.

Data on internal migration come from the ONS. It refers to residential moves

[11] The unemployment measure is based on the International Labor Organization (ILO) definition which outlines unemployment as all people who are without work but are nevertheless available for and seeking employment. Gross weekly wages are CPI adjusted and calculated for both males and females while excluding the weekly earnings less than £25 and more than £10,000 a week.

between different geographic areas within England and provides the most complete components of change for the mid-year population series which are updated annually. In particular, I use net migration estimates that reflect the number of movements across counties and serve as a satisfactory explanation for sorting patterns in response to changes in house prices.

The data on household (liquid) assets, including fixed term investments, are obtained from the Family Resources Survey (FRS). The FRS is a continuous survey of households. For each financial year (April to March) more than 20,000 private households are interviewed and the survey provides detailed information on the income resources and conditions of families in England. I use this information to disentangle any possible confounding effects of other types of wealth.

The birth count data used in this analysis are from the ONS. These statistics include counts of live births and stillbirths by mother's area of usual residence and age group. Age-specific fertility rates (ASFRs) are constructed for age bands 20-29 and 30-44 by dividing the number of births by the corresponding female population in a given county-age group-year cell. Because the data do not provide information to determine the date of conception, I matched the ASFRs with one year lagged, *t-1*, county level characteristics and house prices. Using all previously mentioned data, I have 5,624 age group-county-year observations.

4.3.3 Descriptive Statistics and Trends

In Table 4.A, I present descriptive statistics by age groups, 20-29 (column 1) and 30-44 (column 2). The age groups differ from each other in a number of ways. For instance, the fertility rate is higher for the age band 20-29, at around 90 births per 1,000 women. The same cohort also earns and saves less on average, is more likely to be un-employed and renting, and is less likely to live alone, be married and hold a university degree. Age band 30-44 on average earns and saves more, is less likely to be unemployed or renting and is more likely to live alone, be highly educated and married. With regards to the country of birth and ethnicity, there are almost no substantial differences between the two groups.

Figures 4.1 and 4.2 show trends in age-specific fertility rates and house prices at the national level. The birth rate for those aged 20-29 appears to fluctuate strongly: it fell to 84 births per 1,000 women in 2002, and then reached their peak of 96 births in 2007 before sharply declining again. With regard to the 30-44 age band, birth rates follow a pro-cyclical trend and reached their highest point of 64 in 2008, sharply dropping until 2010 and rebounding afterwards. Overall, differences in demographic and labour market characteristics as well as trends highlight the importance of accounting for other factors that may confound the association between house prices and birth rates.

Table 4.B presents the means and standard deviations for house prices and refusal rates. In terms of house prices, a first look at the data shows that house prices vary greatly across regions. Over the sample period, the mean house price was £172,001 with a standard deviation of £103,024 and considerably higher in the south, including Greater London, than in the rest of England. Table 4.B also reports the average refusal rates. Similar to house prices, the refusal rates exhibit a significant amount of variation across regions and are the highest in the south.

Figure 4.3 plots histograms of the share of refusal rates for each region over the sample period in which the charts have been put on a common x-scale. Altogether, it suggests that refusal rates vary across regions independently of the price patterns. Figure 4.1 in the appendix plots the variation in refusal rate trends across regions. It shows that the trends are non-linear due to the onset of the financial crisis and some regions have experienced larger increases in refusal rates. Moreover, the regions that are large and exhibit substantial variation in shares appear to stand out sharply in terms of house prices. Lastly, Figure 4.4 shows that weekly rents closely track movements in house prices.

4.4 Estimation Methodology

4.4.1 Empirical Strategy

In order to identify the causal effects of house prices on current birth rates, it is vital to have a variable that is exogenous to birth rates and strongly correlated with county-level house prices. As introduced above, I use refusal rates for major development projects as a source of exogenous variation in house prices in England that successfully captures the regulatory and supply side constraints on the housing market. Similar to Dettling and Kearney (2014), the fully saturated model specification that I estimate is given by the equation:

$$\begin{aligned} log(Birth_{cgt}) \;=\; & \beta_0+\beta_1 HP_{c(t-1)}+\beta_2 HP_{c(t-1)}*Own_{cg}+\beta_3 Own_{cg} \\ & +X_{cg(t-1)}+\theta_c+\delta_g+\gamma_t+\gamma_{c(t-1)}+\varepsilon_{cgt} \end{aligned} \tag{1}$$

where c, g, and t index counties, age groups and years respectively. The data set consists of a balanced panel of 148 counties, two age groups (20-29 and 30-44) and each observed over from 1995 to 2013. Birth$_{cgt}$ is the log of the age-specific fertility rates for a given county, age group and year. $HP_{c(t-1)}$ is the log of house price index and shows how an increase in house prices affects the relationship between house prices and fertility rates among renters and existing owners who might plan to move to a larger property with an additional child. $HP_{c(t-1)}*Own_{cg}$, house price index is interacted with a baseline measure of ownership rates (1995) and captures the relationship between house prices and fertility rates among home owners.

$X_{cg(t-1)}$ is a vector of control variables and has three main components: (i) labour market controls, (ii) demographic characteristics, (iii) non-housing wealth and net population change. First, to account for pro-cyclical variation in labour market outcomes, I include the unemployment rate and the natural log of average weekly gross household income. Second, to adjust for the effect of demographic structure on fertility rates I directly control for time varying observable demographic characteristics. These variables are as follows: share of non-UK born population, share of college graduates, share of non-white individuals, share of one-person family, share of non-married individuals and share of individuals who came to England less than a year ago. Third, I include net population change to account for potential sorting (that is, mobility) patterns. More specifically, migration can be expected to affect house prices from the demand side. Hence, it is included to improve precision and, if correlated with house prices, also to address potential omitted variable bias. The log of non-housing wealth disentangles the effect of the housing wealth from other possible confounding wealth effects.

To account for other unobservable characteristics, I exploit the panel aspects of the data set. In particular, county fixed effect, θ_c , and age group fixed effect, δ_g, are included to minimize all variation in birth rates caused by factors that vary across counties as well as age groups and are constant over time. Year fixed effects, γ_t , are included to eliminate the time variant macro-economic shocks that lead changes in fertility rates shared by all counties and age groups over time. County-specific linear and quadratic time trends, $\gamma_{c(t-1)}$, remove variation in intra-county fertility rates caused by factors that are county specific over time. ε_{cgt} is the unobserved determinants of fertility. Standard errors are clustered at age group-county level and all regressions are weighted by the relevant female population in each cell. In fully saturated models, the birth rate-house price estimates are identified exploiting within-county variation in house prices in which county-specific time trends are expected to minimize the unobservable effects that may be correlated with other explanatory variables.

4.4.2 Instrumentation Strategy

The primary aim of this paper is to provide causal effects of house prices on fertility rates. However, there are inherent reverse causality and omitted variables bias issues if the above relation was estimated using OLS. For instance, if people who plan to have a child demand larger houses, this may eventually lead the direction of causation to run from birth rates to house prices. They may also be jointly affected by some omitted factors. Furthermore, measurement error in house prices may cause attenuation bias. An alternative strategy that addresses these issues would be to use an instrumental variable that affects county level house prices yet are unrelated to fertility rates and to re-estimate equation (1) using 2SLS method.[12] This method would be useful

[12] I perform 2SLS estimation following Balli and Sorensen (2013) to avoid a "forbidden regression" (see, Wooldridge 2002). To be specific, I estimate 2SLS structural-equation model with two instruments: refusalrate and refusalrate*ownership and use vce(robust) to account for heteroskedastic errors.

in establishing causal links.

To find a valid instrument, I focus on one of the main factors that drives the house prices upwards in England: a restrictive planning system (Barker, 2006, DCLG 2011, CCHPR 2014, Hilber and Vermeulen, 2016). The regime is restrictive and mostly characterized by high refusal rates for (new) major development projects. It is defined as the fraction of housing projects with 10 or more dwellings that were rejected by a local planning authority and a commonly used measure to capture regulatory restrictiveness (Cheshire and Sheppard, 1989; Preston et al., 1996; Hilber and Vermeulen, 2016).

Since there is a large time gap between planning approval and a dwelling being built, and to account for pro-cyclical changes in decisions owing to high housing demand, I use three-year moving average refusal rates (that is the three years leading up to the current period). This also allows me to use a more informative measure rather than concentrating on year-to-year variations which may not be able to fully capture the scope and functions of the planning system. I assume that the link between three-year moving average refusal rates for major development projects and county level house price trends would not have been systematically correlated with county level birth rates. There are three justifications for this identifying assumption. First, using moving averages helps to reduce potential endogeneity concerns arising from changes in refusal rates that are driven by the local demographic structure. It is mainly because these estimates cannot be affected by contemporaneous omitted variables. Second, it seems that the decisions on applications are more likely to be shaped by the central government. As discussed in section 2, the housing targets are exogenously determined and given to the counties by the RPBs leaving them little room for maneuver to cope successfully with the house prices in response to changes in local housing demand. Third, targets are set by using backward-looking data, and hence should not be influenced by the current birth rates. Within this context, identification is achieved by an exclusion restriction that refusal rates for major development projects should affect birth rates only through supply induced variation in house prices.

Empirically, the refusal rates are highly positively correlated with house prices over the sample period at the aggregate level. However, the main concern of whether refusal rates exogenously determine birth rates requires further investigation. One potential concern is that a more profitable real estate market may encourage developers to apply for more projects, leading to artificially high refusal rates in some counties. In this case, my instrument would be correlated with the second-stage residuals. Although including county and year fixed effects in the estimating equation should address this issue, I additionally examine the validity of my results by using alternative instruments: (i) change in project approval delay rate before and after the 2002 policy reform, (ii) baseline refusal rates in 1994, (iii) average refusal rates between 1995 and 2013, and (iv) a different measure: the number of accepted dwellings over baseline housing stock, as an instrumental variable. This measure attempts to isolate the effects from the number of permissions asked. That is, if counties of the same size systematically grant the similar number of permissions with the different number of applications received (that is, high versus low), then refusal rates would be contaminated by the number of applications. In such a case, estimates would be both qualitatively and quantitatively different. It is worth noting that the results of these alternative measures are similar to the presented IV results, thus confirming the validity of the IV estimates.

4.5 Empirical Results

This section presents three sets of results. I first show Ordinary Least Squares (OLS) estimates. I then focus on IV/2SLS results outlined in section 4.2. Lastly, I present findings for demographic subgroups and with different house price measures.

4.5.1 OLS Estimates

Table 4.1 presents the results from the OLS estimation where the dependent variable is the log of the age-specific fertility rate. I estimate models separately for the full sample in the top panel, for those aged 20-29 in the middle panel, and for those aged 30-44 in the bottom panel. Column 1 reports the estimation with all fixed effects included (county, year and age group); column 2 adds demographic characteristics and net change in population; column 3 adds labour market characteristics and gross household assets excluding housing wealth; column 4 adds county-specific linear time trends; column 5 adds county-specific quadratic time trends.

The HousePrice*OwnershipRate interaction coefficient in the first column yields a positive and statistically significant estimate of .002 and the House Price coefficient yields a negative and statistically significant estimate of -.011. More specifically, the coefficient on the former measure the estimated effect in the hypothetical case of only homeowners (that is, age group-year-county cell with a home ownership rate of 1) and the latter quantify the estimated effect in the hypothetical case of zero homeowners (that is, age group-year-county cell with a home ownership rate of 0). After adding more controls and time trends (columns 2 to 5), the results consistently and significantly show that for home owners an increase in house price is positively associated with birth rates whereas the opposite is true for renters. In the fully saturated model (column 5), the magnitude of the relationship indicates that a 10 per cent increase in house prices leads to a 0.8 per cent rise in fertility rates among home owners. For renters, the same amount of increase in the house prices causes a 1.7 per cent fall in the fertility rate. These results confirm Dettling and Kearney's (2014) finding that house prices significantly and differently affect birth rates among home owners and renters. The net effect at the mean home ownership rate is negative, which suggests that birth rates move counter-cyclically and the negative price effect dominates the income effect in England. This finding is different than those found in the United States (Dettling and Kearney, 2014) and there are a number of possible explanations: first, it may be because of the adverse effects of a restrictive housing market on renters who spend around 50 per cent (as high as 70 per cent in London) of their gross disposable income on rent.[13] Second, house prices have risen much faster than earnings which has made getting onto the property ladder even harder for the average first-time buyer. Third, unlike the United States, the rental cost has closely tracked movements in house prices while earnings have remained relatively stable, subsequently leading to a fall in disposable income. Fourth, young families, who are mostly renters, have been priced out of the rental markets and rental accommodation has become unaffordable. Altogether, these factors can potentially create barriers to family formation and make having children financially "prohibitive" for renters, leading to a fall in birth rates at the aggregate level over the sample period.

In the middle and bottom panel of Table 4.1, I show that the negative price effect

[13] Author's own calculation from the English Housing Survey, 2013-2014.

among renters is mainly driven by those aged 20-29. In contrast, the positive effect of house prices on birth rates is driven primarily by the older cohort (those aged 30-44), in which people are significantly more likely to be home owners and less likely to postpone their childbearing. These findings support the notion that housing costs exert downward pressure on the fertility outcome of young adults and that there is a connection between getting on the property ladder and building a family.

Regarding the other variables listed in Table 4.2 in the Appendix, unemployment rate and non-housing wealth positively and significantly affect the birth rates. The coefficient on gross weekly wages is also positive but insignificant in all models. Recall that I included net population change in an attempt to account for mobility and, hence, expected to see positive coefficients. Indeed, the positive and significant point estimates suggest that the variable does a good job in addressing the sorting issue in response to changes in house prices. Inclusion of county-specific time trends makes the coefficients larger both for home owners and for renters. This implies that the county-specific fertility trends driven by the omitted fertility determinants tend to move in the opposite direction of the trends in house prices over the sample period.[14]

The analysis continues with a stratification of the regressions with county level housing supply characteristics so as to gain further insights into the housing market basis of this result. More specifically, I split the sample by the observations in the upper quartile of home ownership distribution and versus the ones in the lower quartile. The first two columns of Table 4.3 suggest that the fertility rate is the most responsive to house price changes in more supply constrained counties where ownership rate is

[14] Inclusion of ownership-county specific and age group-county specific linear time trends made little difference to the values in column 5.

above the 75th percentile. In these counties, a 10 per cent increase in house prices leads to a 1.2 per cent increase in birth rates among home owners and a 2.3 per cent decrease in birth rates among renters. It is mainly because the supply shortage leads to higher house prices, generating a larger income effect for home owners and a larger negative price effect for renters. In areas where home ownership rate sits at the bottom quartile, coefficients are smaller in magnitude and mostly significant at the conventional levels. Altogether, the results in this table suggest that the effects are more pronounced and seem to be driven by supply constrained counties.

The differences across demographic groups are highlighted in Table 4.4 in which I expect to find that non-native, less educated and non-white groups to be “more” affected by increases in house prices. Overall, the estimates are parallel to this notion. For example, a 10 per cent increase in house prices leads to a 2.3 per cent decrease among the foreign born population. For home owners, the same amount of increase is associated with 1.6 per cent in birth rates. Overall, similar to the findings in Table 4.1, renters between the ages of 20 and 29 seem to be more negatively affected than the older cohort. On the one hand, may be because this cohort witnessed substantially higher rises in rents, they earn less on average and are less likely to borrow to finance their child-related expenses. On the other hand, the significantly higher coefficient on house prices and birth rates among the older cohort may suggest that this cohort benefits from the long-standing real house price growth. In other groups, the results maintain the expected sign of direction while being significant in most cases. The following section of this paper is concerned with endogeneity of house prices and instrumental variable estimation results are presented.

4.5.2 IV Estimates

In this section I present the IV estimates of the house price-fertility relationship using county level refusal rates for major development projects as an instrument for county-level house prices. As previously discussed in section 4.2, if the estimates suffer from omitted variables that are not picked up by the linear trends and the fixed effects, the OLS specifications will yield biased point estimates. In addition, if high fertility rates lead to a rise in housing demand and, in turn, increase house prices, the OLS results will provide wrong statistical inference.

Before discussing the instrumental variable estimates, a discussion on the validity and power of the instrument is needed. Table 4.1 in the appendix presents the first stage estimates of the instrument and successively adds more controls in the models similar to Table 4.1. In all specifications, the first-stage relationship between refusal rate and house prices is strongly positive: three-year moving average refusal rates are significantly associated with house price growth at the 1 per cent level (column 1 in Appendix Table 4.1), and this relationship is robust to the inclusion of demographic and labour market controls (columns 2 and 3) as well as county-specific time trends (columns 4 and 5). Overall, the instrument seems to exert stronger positive effects on house prices as more controls are included and have predictive power. Recall that refusal rates are used to capture the restrictiveness of the planning system and higher refusal rates typically lead to increases in house prices. It is an expected outcome, since the planning system in England tends to be protective. The results for the first stage F-test also show that the first-stage relationships are fairly strong.

In addition, I employ alternative measures for the instrumental variable, including the change in the approval delay rate (before and after the 2002 reform) of major residential projects (that is, the number of decisions that are delayed over 13 weeks in any given county-year cell relative to all decisions made in that county-year cell), baseline refusal rates in 1994, average refusal rates between 1995 and 2013 and the number of accepted dwellings over baseline housing stock. The change in the delay rate measure uses the exogenous variation generated by the policy reform introduced in 2002 to shorten the length of the planning application process.[15] The point estimates at the first stage are positive as expected and mostly statistically significant (these models not shown). However, the first-stage F-tests in some cases are less than 10, so results with alternative measures should be interpreted with caution.[16]

Table 4.2 presents the IV estimates in which I replicate the OLS specifications from Table 4.1. Again, I only report the coefficients of main interest. The IV-2SLS fixed-effects framework results show that: (i) the impact of house prices on birth rates among home owners is positive and significant at 1 per cent confidence, with a point estimate of 0.028 (standard error 0.008 in column 5 of Table 4.2), (ii) the impact of house prices on birth rates among renters is negative and significant at 1 per cent confidence, with a point estimate of 0.049 (standard error 0.016 in column 5 of Table 4.2). Overall, in each model, the IV coefficients are larger than the OLS estimates. The OLS point estimates range from 0.2 per cent to 0.8 per cent for home owners and from 1.1 per cent to 1.7 per cent for renters. The corresponding interval for the IV results range from 0.6 per cent to 2.8 per cent for home owners and 2 per cent to 4.9 per cent for renters.

[15] I thank Christian Hilber for suggesting I use the delay rate as an alternative instrument and his generosity in sharing these data. The merits of this measure have been discussed at length by Hilber and Vermeulen (2016).

[16] See Table 4.3 in the Appendix.

In the middle and bottom panels of Table 4.2, I report the estimates by age groups, following the baseline specification. I document important differences in these models. Both in the middle and bottom panels, I find the usual pattern that home owners' birth rates respond positively to house price increases, whereas the opposite is true for renters. However, the middle panel columns for those aged 20-29 show that the negative effect of house prices on renters' birth rate is much larger than those implied by the older age group. In contrast, the results for those aged 30-44 show that the overall housing wealth effects are larger than those found in the 20-29 age band, and the home owner results in the top panel are mainly driven by the older age group. Altogether, these findings suggest that the causal relationship between house prices and birth rates holds for both home owners/renters and among those aged 20-29 and 30-44.

Table 4.3 presents results based on the housing supply constraints. In general, the pattern remains similar to the OLS findings. Column 3 presents results for the counties with "low ownership rate" in which the IV results are almost three times larger than the OLS estimates and significant at conventional levels. For the counties with high ownership rate, in column 4, the IV results are also substantially larger and also significant at 1 per cent.

However, the IV results for the demographic groups are not so strong. First stage values of the F test for triple interactions are less 10 than in below degree level estimates. Therefore, the OLS estimates may be preferable to IV estimates for these categories. Other estimates are significant at conventional levels and the effects of house prices on birth rates are even stronger than those from the OLS regressions for both home owners and renters. In particular, the coefficients on the foreign born and other ethnicity interactions exceed the OLS estimate considerably, indicating that these groups are the most affected. The following part of the results explore whether the responses remain robust to alternative housing price measures.

4.5.3 Robustness Checks

The results presented thus far demonstrate the fact that house price is an important determinant of the fertility rate outcome. Even though the effects vary with estimation features of the model, an increase in house price clearly has a positive impact on birth rates among home owners and has a negative impact on birth rates among renters. Nevertheless, I conduct additional analyses in order to detect whether the main findings remain stable to the different measures of the house prices, namely, median house prices, lower quartile house prices and lagged house prices – that is (t-2) and (t-3) –. In fact, the results mostly hold through columns 1 to 3 and 5 to 7 of Table 4.5. For the median and lower quartile house prices, the magnitude of the point estimates is similar to those found in the main OLS and IV specifications. When I use the lagged house prices, the effects get smaller and become insignificant at higher lags. Nevertheless, the results are still in line with the main findings of this paper.

Table 4.6 investigates birth rate differences when splitting the sample by region (that is, based on mother's area of usual residence) in which I observe quite a distinctive pattern for birth rates. Prior work has reported that local scarcity of open developable land is the greatest in south east England and this region-based analysis may contribute to the understanding of the house prices and fertility relationship (Hilber & Vermeulen, 2016). While there are many differences between (Greater) London and the rest of England, one of the most noticeable in this context is that house prices are likely to exert stronger effects on birth rates in London due to high variation in property prices and rents. I find in column 7 of Table 4.6 that the estimated coefficient on the interaction term is positive in sign, considerably larger in magnitude (compared with the main specification) and statistically significant at the 1 per cent level. In contrast, in column 7, estimates show that renters in London experience a considerably low fertility rate, and again this effect is much larger than the baseline results. For southern England (excluding Greater London), I estimate sizable effects for both home owners and renters, though the coefficients are smaller than those documented for London. Results for the east and west Midlands return evidence that house prices do not significantly affect birth rates in these regions. There is, however, evidence from columns 1 to 3 about significant effects of house prices on birth rates in the northern part of the country. These region-based findings are particularly noteworthy, because they provide supportive evidence for the validity of the main argument of this paper and confirm the causal link between house prices and birth rates. In addition, the region estimates somewhat indicate that the main results are mostly driven by the regions in southern England including Greater London.[17]

Table 4.7 presents further results by housing boom and bust periods from the fully saturated OLS and IV specifications. I find that the impact of house prices on fertility rates is larger during "boom" years. Table 4.8 examines house price effects with additional controls. In column 1, I reprint the full sample IV estimates, columns 2-5 show the results of including alternative housing market controls (share of first time buyers at the regional level, share of two or more bedroom apartments at the regional

[17] I do not report region results by age groups for the sake of space. Overall, these estimates are mostly in line with those reported in Tables 4.1 and 4.2.

level, share of new dwellings at the regional level, and number of transactions (sales) at the county level) individually. Table 4.8 indicates that the coefficients on the birth rates are largely invariant to the inclusion of additional housing market characteristics, with no exception. I also estimated models in combination with each other and house prices interacted with the regional level controls, and these did not fundamentally change the main findings.[18]

4.6 Conclusion

Using the refusal rates for major development projects as an instrumental variable for house prices, the results presented in this paper show that house prices significantly affect birth rates in England. I found a significant positive birth rate coefficient for home owners and a significant negative birth rate coefficient for renters. There are also significant effects for younger (aged 20-29) and older (30-44) home owners and renters. The positive "home owner" effect is mainly driven by those aged 30-44 and the negative "renter" effect is driven by those aged 20-29. At the aggregate level the net effect is negative, in other words the negative price effect dominates the income effect. The expected variation in house prices and fertility by demographic characteristics is also documented.

The stronger effects of house prices found in the models where the sample was separated according to the residency in London and counties exhibit high ownership rate. The results for foreign-born population and those with a less than degree level education are also particularly large in magnitude. Moreover, the findings do not depend on the estimation methodology used, even though I do find relatively larger coefficients

[18] These results are not presented in the current tables but are available upon request.

when I instrument for county-level house prices.

The stronger effects of house prices found in the models where the sample separated by the residency in London and counties exhibit high ownership rate. The results for foreign-born population and those with less than degree level education are also particularly large in magnitude. Moreover, the findings do not depend on the estimation methodology used, even though I do find relatively larger coefficients when I instrument for county level house prices.

I also find similar results when I use different measures of house prices. In both OLS and IV results I find evidence that the median and lower quartile house prices significantly affect birth rates. This finding, again, is reproduced when I instrument for house prices. For the lagged house prices, the results indicate that house prices have a significant impact on fertility rates. However, the point estimates attenuate towards zero at higher lags. These findings for the different length of lags cast an important behavioral interpretation of the fertility responses to housing market trends. That is, people seem to take house price changes in previous years into account.

The findings of this study are potentially important from a public policy standpoint: if the negative effect of house prices on fertility rates is mainly driven by the younger cohort (that is, couples put off having children because of not being able to afford suitable accommodation), it may be possible to reverse this trend through the design of better housing and/or child benefit schemes. One example would be to scrap the Help-to-Buy ISA’s maximum purchase cap of £250,000 and £450,000 for London. This cap limits couples to buy two or more-bedroom family homes and creates a barrier for young potential first-time buyers. If such government schemes help people to not only get a foot on the housing ladder but also afford buying a family home rather than a flat, they could decrease the number of households with children renting privately and reduce the average age of the first-time mother. I offer these suggestions with a cautionary reminder that further individual level analysis and careful case studies are necessary to explain the causal mechanism and to design more effective policy responses.

Lastly, although the study has successfully demonstrated the aforementioned findings, it is however limited by the use of total birth rates, and the findings cannot be transferable to birth orders and comment on the quality-quantity trade-off. Future research should, therefore, concentrate on the investigation of birth orders and the potential relationship between house prices and the quality of children.

Figure 4.1: Age-Specific Fertility Rates (aged 20-29) and House Prices

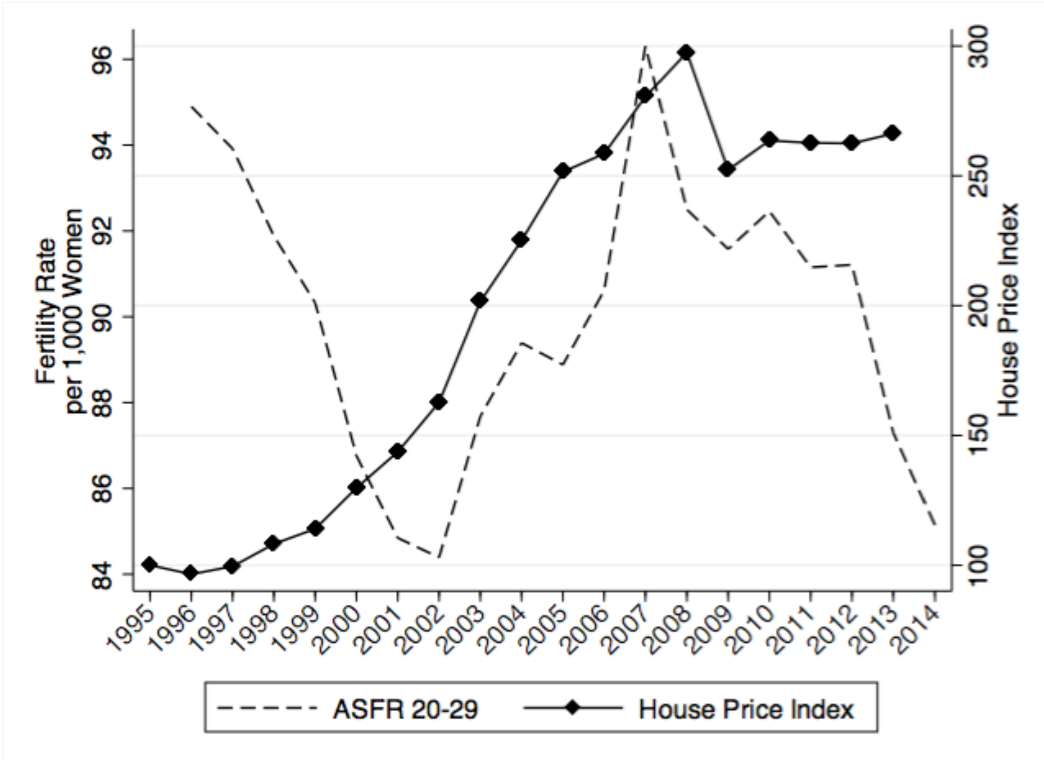


Figure 4.2: Age-Specific Fertility Rates (aged 30-44) and House Prices

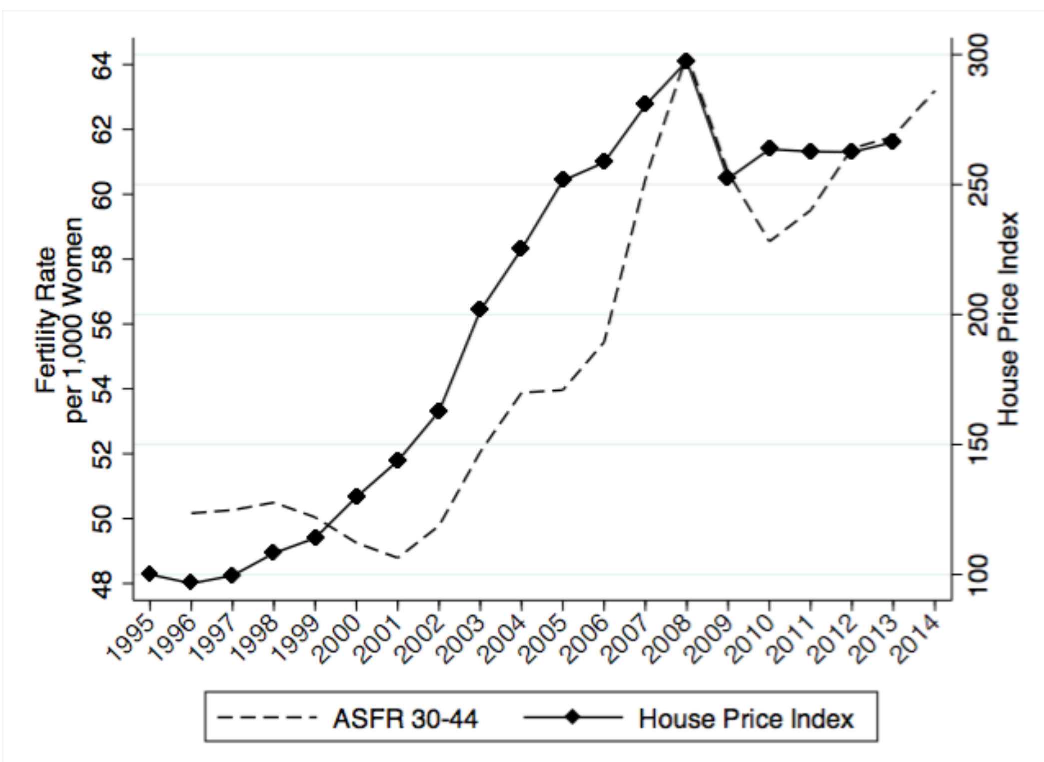


Figure 4.3: Shares of Refusal Rates across Regions

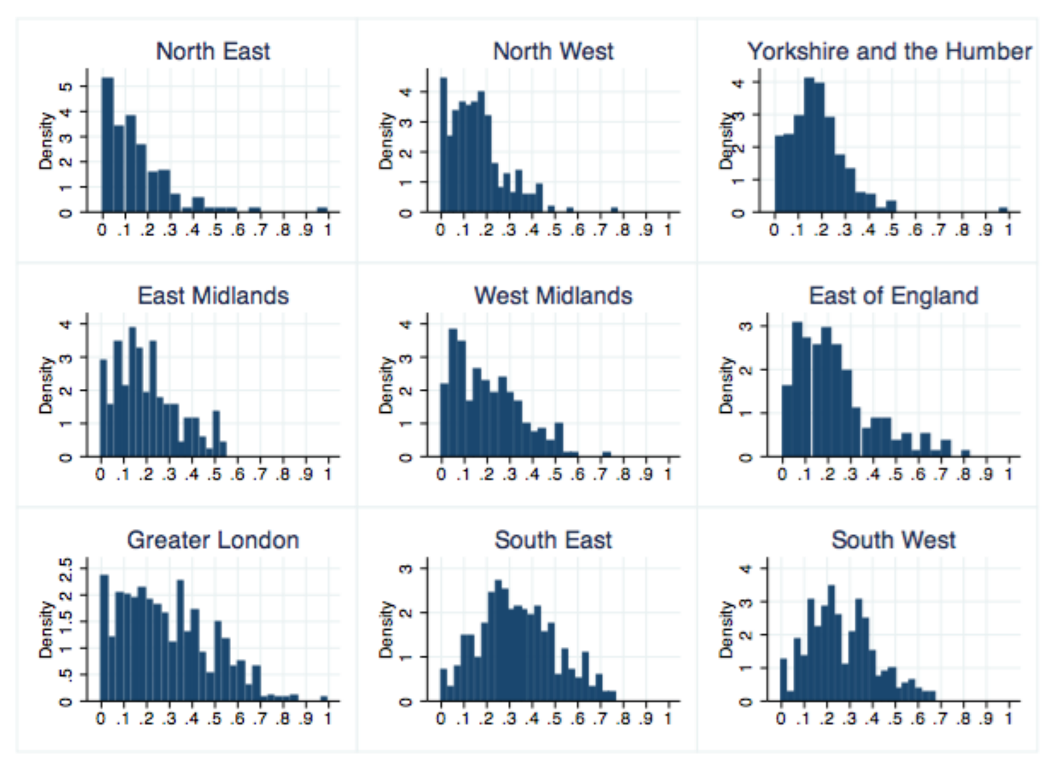


Figure 4.4: Rents and House Prices

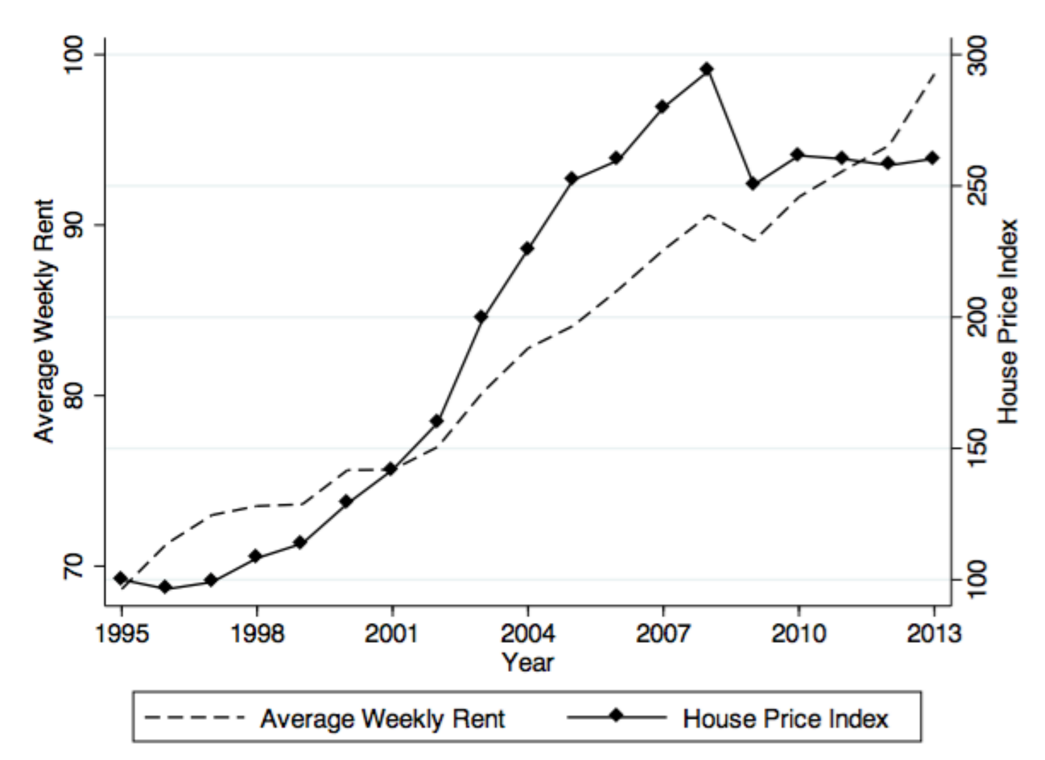


Table 4.A: Descriptive Characteristics

Variables	Aged 20-29	Aged 30-44
Fertility Rate (1000s)	90.06 (22.83)	55.47 (21.26)
Home Ownership Rate	54.49 (15.81)	69.88 (12.46)
Home Ownership Rate (1995)	60.21 (14.18)	73.87 (11.34)
Gross Weekly Wages	333.67 (122.66)	518.47 (203.66)
Unemployment Rate	9.48 (5.14)	5.32 (3.21)
Gross Household Liquid Assets	3,463 (2,998)	5,687 (4,781)
Degree Level	27.62 (11.47)	31.05 (12.10
A Level	27.72 (7.35)	20.25 (5.49)
Below A Level	44.66 (11.31)	48.70 (10.33)
Foreign Born	14.64 (14.02)	16.28 (15.70)
Other Ethnicities	13.65 (15.19)	12.95 (14.77)
One Person Family	7.94 (4.39)	10.41 (4.29)
Single	77.97 (8.45)	28.11 (9.41)
N	2812	2812

Notes: Within cell means (standard deviations). The table provides within cell means for the age-specific demographic characteristics for the 148 counties used in the analysis. Fertility rates are constructed by dividing the number of births by the corresponding female population using mid-year population estimates based on censuses, in which female ages range between 20 and 44. Source for county level birth data is UK Office for National Statistics. The UK Labour Force Survey is being used to construct county-year-age group specific unemployment rates, gross weekly wages and demographic characteristics. The gross weekly wages are calculated by dividing self-reported gross annual pay by the number of weeks worked in the same calendar year and are CPI adjusted to 2005 pounds. Unemployment rates refer to the per centage of economically active people who are unemployed by ILO standards. Household assets are self-reported and exclude housing wealth

Table 4.B: Descriptive Characteristics

Regions	House Prices	House Price Index	Refusal Rate	Net Population Change
North East	82,593 (30,750)	150.68 (57.48)	17.35 (12.63)	178 (190)
North West	86,628 (35,749)	161.02 (60.44)	17.70 (10.83)	502 (1898)
Yorkshire and the Humber	88,102 (37,969)	157.75 (61.29)	19.15 (11.14)	1454 (2169)
East Midlands	101,340 (43,747)	180.30 (66.53)	21.75 (12.92)	3075 (2565)
West Midlands	103,408 (43,567)	171.93 (61.13)	21.64 (14.53)	1361 (2310)
East of England	122,050 (51,059)	205.57 (76.21)	24.32 (16.54)	4176 (3486)
Greater London	237,185 (128,183)	259.37 (110.70)	31.60 (20.63)	2404 (2263)
South East	154,706 (66,743)	216.42 (80.01)	34.34 (15.98)	2280 (3140)
South West	131,603 (52,734)	213.40 (82.16)	25.57 (13.91)	2242 (1915)

Notes: Within cell means (standard deviations). This table provides aggregate level variables averaged across 9 regions (148 counties) and 19 years (1995-2013) used in the analysis. Land Registry data on House Prices and House Price Index are based on reports from the individual house price records of all residential property sales in England. Data on Refusal Rates for Major Development Projects are obtained from the Department of Communities and Local Government. It is defined as the proportion of housing projects consisting of at least 10 dwellings that was refused by a local planning authority in one calendar year. Net Population Change data come from the Office for National Statistics – Population Estimates Unit and provide detailed information on the components of population change for counties, London boroughs and districts in England. According to the ONS (2015), the estimated resident population of an area includes all those people who usually live there, regardless of nationality. Arriving international migrants are included in the usually resident population if they remain in England for at least a year. Emigrants are excluded if they remain outside England for at least a year, which is consistent with the United Nations definition of a long-term migrant. Armed forces stationed outside of England are excluded. Students are taken to be usually resident at their term time address. Internal migration flows presented in the table reflect the number of movements that cross local authority boundaries.

Table 4.1: House Prices and Birth Rates – OLS Estimations

Controls for ➔	(1) OLS Fixed effects	(2) OLS + demographic characteristics	(3) OLS + labor market characteristics	(4) OLS + county linear trends	(5) OLS + county quadratic trends
Full sample					
HPI*Ownership	0.002***	0.004*	0.007**	0.009**	0.008***
	(0.000)	(0.002)	(0.003)	(0.004)	(0.002)
House Price(HPI)	-0.011***	-0.017***	-0.013***	-0.019**	-0.017***
	(0.002)	(0.004)	(0.003)	(0.007)	(0.004)
R-squared	0.303	0.489	0.578	0.704	0.816
N	5624	5624	5624	5624	5624
Aged 20-29					
HPI*Ownership	0.002*	0.002*	0.003*	0.004***	0.004***
	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)
House Price(HPI)	-0.014***	-0.019***	-0.018***	-0.024***	-0.020***
	(0.004)	(0.006)	(0.005)	(0.007)	(0.005)
R-squared	0.298	0.454	0.608	0.704	0.803
N	2812	2812	2812	2812	2812
Aged 30-44					
HPI*Ownership	0.004***	0.007***	0.012***	0.014***	0.013***
	(0.001)	(0.002)	(0.003)	(0.004)	(0.003)
House Price(HPI)	-0.002	-0.004*	-0.003**	-0.006*	-0.004*
	(0.002)	(0.002)	(0.001)	(0.003)	(0.002)
R-squared	0.343	0.466	0.661	0.713	0.835
N	2812	2812	2812	2812	2812

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors, clustered at the age group-county level, are in parentheses. Column 1 includes: age, year and county fixed effects, column 2 adds: share of non-UK born population, share of college graduates, share of non-white individuals, share of one-person family, share of non-married individuals, share of households with at least one child and share of individuals who came to England less than a year ago and net population change, column 3 adds: unemployment rate, the average weekly gross household income and non-housing wealth, column 4 adds: county linear time trends and column 5 adds: county quadratic time trends.

Table 4.2: House Prices and Birth Rates – IV/2SLS Estimations

Controls for ➔	(1) IV Fixed effects	(2) IV + demographic characteristics	(3) IV + labor market characteristics	(4) IV + county linear trends	(5) IV + county quadratic trends
Full sample					
HPI*Ownership	0.006^{**}	0.014^{**}	0.023^{**}	0.026^{**}	0.028^{***}
	(0.002)	(0.006)	(0.008)	(0.010)	(0.008)
House Price(HPI)	-0.020^{***}	-0.028^{***}	-0.035^{***}	-0.043^{***}	-0.049^{***}
	(0.006)	(0.009)	(0.011)	(0.014)	(0.016)
R-squared	0.298	0.439	0.517	0.710	0.802
N	5624	5624	5624	5624	5624
Aged 20-29					
HPI*Ownership	0.009^{**}	0.008^{**}	0.016^{**}	0.011^{***}	0.012^{**}
	(0.004)	(0.003)	(0.007)	(0.003)	(0.005)
House Price(HPI)	-0.030^{***}	-0.037^{***}	-0.044^{***}	-0.051^{***}	-0.062^{***}
	(0.009)	(0.012)	(0.013)	(0.016)	(0.019)
R-squared	0.363	0.496	0.562	0.704	0.819
N	2812	2812	2812	2812	2812
Aged 30-44					
HPI*Ownership	0.012^{***}	0.020^{***}	0.030^{**}	0.034^{***}	0.036^{***}
	(0.004)	(0.006)	(0.012)	(0.010)	(0.011)
House Price(HPI)	-0.008	-0.013^{***}	-0.014^{***}	-0.019^{**}	-0.021^{***}
	(0.005)	(0.004)	(0.004)	(0.008)	(0.007)
R-squared	0.383	0.488	0.601	0.744	0.825
N	2812	2812	2812	2812	2812

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 4.1.

Table 4.3: House Prices and Birth Rates – High vs. Low Ownership Rates

	(1) OLS Low Ownership (25th Per centile)	(2) OLS High Ownership (75th Per centile)	(3) IV Low Ownership (25th Per centile)	(4) IV High Ownership (75th Per centile)
Full sample				
HPI*Ownership	0.006*	0.012***	0.015**	0.025***
	(0.003)	(0.003)	(0.006)	(0.007)
House Price(HPI)	-0.008**	-0.023***	-0.017**	-0.064***
	(0.003)	(0.006)	(0.007)	(0.021)
R-squared	0.758	0.777	0.813	0.828
N	1406	1406	1406	1406
Aged 20-29				
HPI*Ownership	0.007	0.009***	0.029**	0.052***
	(0.012)	(0.002)	(0.012)	(0.014)
House Price(HPI)	-0.011**	-0.031**	-0.036**	-0.091***
	(0.004)	(0.011)	(0.015)	(0.028)
R-squared	0.764	0.782	0.833	0.835
N	1406	1406	1406	1406
Aged 30-44				
HPI*Ownership	0.004**	0.021***	0.015***	0.069***
	(0.001)	(0.005)	(0.004)	(0.020)
House Price(HPI)	-0.004	-0.015***	-0.013*	-0.040***
	(0.004)	(0.004)	(0.007)	(0.012)
R-squared	0.742	0.798	0.799	0.823
N	1406	1406	1406	1406

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 4.1.

Table 4.4: House Prices and Birth Rates – Various Interactions

	(1) Full Sample OLS	(2) Full Sample IV	(3) Aged 20-29 OLS	(4) Aged 20-29 IV	(5) Aged 30-44 OLS	(6) Aged 30-44 IV
HPI*Ownership*Foreign-born	0.016*	0.038*	0.008*	0.044***	0.010**	0.048*
	(0.007)	(0.020)	(0.003)	(0.017)	(0.004)	(0.020)
House Price(HPI)*Foreign-born	-0.023**	-0.072**	-0.030***	-0.085***	-0.018**	-0.032**
	(0.010)	(0.031)	(0.007)	(0.018)	(0.003)	(0.012)
R-squared	0.713	0.748	0.799	0.801	0.756	0.813
N	5624	5624	2812	2812	2812	2812
HPI*Alevelorbelow	0.005*	0.023	0.003	0.014	0.006	0.051*
	(0.002)	(0.013)	(0.008)	(0.022)	(0.010)	(0.023)
House Price(HPI)*Alevelorbelow	-0.018**	-0.051**	-0.011**	-0.036**	-0.021**	-0.056*
	(0.007)	(0.018)	(0.004)	(0.013)	(0.004)	(0.026)
R-squared	0.778	0.743	0.796	0.825	0.788	0.802
N	5624	5624	2812	2812	2812	2812
HPI*Ownership*OtherEthnicities	0.004	0.012*	0.009**	0.018**	0.004**	0.019***
	(0.004)	(0.007)	(0.003)	(0.008)	(0.001)	(0.005)
House Price(HPI)*OtherEthnicities	-0.005*	-0.019***	-0.016*	-0.049**	-0.013*	-0.012*
	(0.002)	(0.006)	(0.008)	(0.020)	(0.005)	(0.006)
R-squared	0.746	0.788	0.754	0.810	0.780	0.804
N	5624	5624	2812	2812	2812	2812

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 4.1.

Table 4.5: House Prices and Birth Rates – Alternative House Price Measures

	(1) OLS Median House Prices	(2) OLS Lower Quartile House Prices	(3) OLS Average House Price, *t-2*	(4) OLS Average House Prices, *t-3*	(5) IV Median House Prices	(6) IV Lower Quartile House Prices	(7) IV Average House Price, *t-2*	(8) IV Average House Prices, *t-3*
Full Sample								
HPI*Ownership	0.011*	0.003*	0.008*	0.003	0.023*	0.014**	0.012*	0.006
	(0.005)	(0.001)	(0.004)	(0.017)	(0.012)	(0.005)	(0.006)	(0.025)
House Price(HPI)	-0.017***	-0.011***	-0.008*	-0.006	-0.040***	-0.041***	-0.022**	-0.016
	(0.005)	(0.003)	(0.004)	(0.011)	(0.013)	(0.013)	(0.008)	(0.018)
R-squared	0.802	0.788	0.765	0.741	0.813	0.788	0.772	0.766
N	5624	5624	5328	5032	5624	5624	5328	5032
Aged 20-29								
HPI*Ownership	0.006*	0.011**	0.005	0.000	0.013**	0.019*	0.006	0.002
	(0.003)	(0.004)	(0.006)	(0.010)	(0.005)	(0.010)	(0.011)	(0.013)
House Price(HPI)	-0.025**	-0.014**	-0.010**	-0.008	-0.070***	-0.033**	-0.013	-0.010
	(0.010)	(0.005)	(0.004)	(0.008)	(0.019)	(0.013)	(0.010)	(0.015)
R-squared	0.798	0.777	0.801	0.765	0.757	0.794	0.786	0.731
N	2812	2812	2664	2516	2812	2812	2664	2516
Aged 30-44								
HPI*Ownership	0.018**	0.016***	0.013**	0.006	0.038**	0.030***	0.029**	0.011
	(0.007)	(0.005)	(0.005)	(0.006)	(0.014)	(0.009)	(0.012)	(0.012)
House Price(HPI)	-0.008**	-0.010**	-0.005*	-0.003*	-0.020**	-0.018*	-0.019***	-0.007
	(0.003)	(0.004)	(0.003)	(0.002)	(0.008)	(0.009)	(0.006)	(0.007)
R-squared	0.813	0.801	0.801	0.743	0.813	0.743	0.779	0.745
N	2812	2812	2664	2516	2812	2812	2664	2516

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 4.1.

Table 4.6: House Prices and Birth Rates – by Regions

Regions ➔	(1) North East	(2) North West	(3) Yorkshire and the Humber	(4) East Midlands	(5) West Midlands	(6) East of England	(7) Greater London	(8) South East	(9) South West
OLS									
HPI*Ownership	0.007*	0.011***	0.005	0.007**	0.016	0.011***	0.017***	0.012***	0.006**
	(0.004)	(0.003)	(0.011)	(0.002)	(0.018)	(0.002)	(0.005)	(0.003)	(0.002)
House Price(HPI)	-0.004**	-0.016**	0.002	-0.009	-0.007	-0.019***	-0.031***	-0.022***	-0.010***
	(0.001)	(0.006)	(0.009)	(0.006)	(0.015)	(0.005)	(0.009)	(0.005)	(0.003)
R-squared	0.786	0.752	0.744	0.756	0.778	0.772	0.789	0.769	0.790
N	456	836	570	342	532	380	1216	722	570
IV									
HPI*Ownership	0.018	0.034***	0.018*	0.013	0.022	0.037***	0.040***	0.025***	0.018**
	(0.013)	(0.011)	(0.009)	(0.010)	(0.027)	(0.010)	(0.013)	(0.008)	(0.007)
House Price(HPI)	-0.012**	-0.054**	-0.010	-0.014	-0.021	-0.059**	-0.071***	-0.049***	-0.028***
	(0.005)	(0.021)	(0.032)	(0.012)	(0.036)	(0.018)	(0.022)	(0.015)	(0.008)
R-squared	0.802	0.767	0.765	0.771	0.745	0.761	0.802	0.804	0.810
N	456	836	570	342	532	380	1216	722	570

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 4.1.

Table 4.7: House Prices and Birth Rates –Boom and Bust Periods

	(1) IV Baseline Estimate (Table 4.2, Column 5)	(2) OLS Housing Boom 1995-2008 & 2012-2013 (inc.)	(3) OLS Housing Bust 2009-2011 (inc.)	(4) IV Housing Boom 1995-2008 & 2012-2013 (inc.)	(5) IV Housing Bust 2009-2011 (inc.)
Full sample					
HPI*Ownership	0.028***	0.013**	0.004*	0.029***	0.013*
	(0.008)	(0.005)	(0.002)	(0.009)	(0.006)
House Price(HPI)	-0.049***	-0.022***	-0.008**	-0.056***	-0.015**
	(0.016)	(0.007)	(0.003)	(0.018)	(0.006)
R-squared	0.802	0.758	0.747	0.766	0.734
N	5624	4736	888	4736	888
Aged 20-29					
HPI*Ownership	0.012**	0.008***	0.003	0.015**	0.005
	(0.005)	(0.002)	(0.004)	(0.006)	(0.011)
House Price(HPI)	-0.062***	-0.021***	-0.008**	-0.050***	-0.021**
	(0.019)	(0.006)	(0.003)	(0.015)	(0.008)
R-squared	0.819	0.763	0.735	0.774	0.724
N	2812	2368	444	2368	444
Aged 30-44					
HPI*Ownership	0.036***	0.011***	0.008*	0.035***	0.016**
	(0.011)	(0.003)	(0.004)	(0.009)	(0.006)
House Price(HPI)	-0.021***	-0.007*	-0.003	-0.024***	-0.007*
	(0.007)	(0.003)	(0.002)	(0.007)	(0.004)
R-squared	0.825	0.772	0.730	0.780	0.733
N	2812	2368	444	2368	444

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 4.1.

Table 4.8: House Prices on Birth Rates – Alternative Controls

	(1) IV Baseline Estimate (Table 4.2, Column 5)	(2) IV Baseline Estimate + Share of First Time Buyers (Regional Level)	(3) IV Baseline Estimate + Share of Two or More Bedroom Apartments (Regional Level)	(4) IV Baseline Estimate + Share of New Dwellings (Regional Level)	(5) IV Baseline Estimate + Number of Transactions (County Level)
Full Sample					
HPI*Ownership	0.028***	0.025***	0.018**	0.024**	0.023**
	(0.008)	(0.007)	(0.007)	(0.010)	(0.008)
House Price(HPI)	-0.049***	-0.051***	-0.043***	-0.055***	-0.042***
	(0.016)	(0.017)	(0.014)	(0.016)	(0.013)
R-squared	0.802	0.805	0.811	0.809	0.813
N	5624	5624	5624	5624	5624
Aged 20-29					
HPI*Ownership	0.012**	0.018*	0.015*	0.015**	0.013**
	(0.005)	(0.008)	(0.007)	(0.006)	(0.005)
House Price(HPI)	-0.062***	-0.054***	-0.050***	-0.060***	-0.058***
	(0.019)	(0.016)	(0.015)	(0.018)	(0.019)
R-squared	0.819	0.827	0.830	0.829	0.835
N	2812	2812	2812	2812	2812
Aged 30-44					
HPI*Ownership	0.036***	0.029***	0.033***	0.039***	0.030**
	(0.011)	(0.009)	(0.009)	(0.013)	(0.012)
House Price(HPI)	-0.021***	-0.018***	-0.025***	-0.024**	-0.019**
	(0.007)	(0.005)	(0.008)	(0.010)	(0.008)
R-squared	0.825	0.815	0.820	0.818	0.816
N	2812	2812	2812	2812	2812

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 4.1.

Appendix Table 2.1
Descriptive Characteristics – Demographics (among those with earnings information)
2012-2014 UK Integrated Household Surveys

Variables	DK/Refuse/Other response to sexual orientation question, males	DK/Refuse/Other response to sexual orientation question, females
Age	44.05 (10.40)	43.70 (10.36)
Degree level	0.319 (0.466)	0.317 (0.465)
Higher ed.	0.114 (0.318)	0.127 (0.333)
A level	0.231 (0.421)	0.181 (0.385)
O level	0.215 (0.411)	0.271 (0.444)
White	0.892 (0.310)	0.904 (0.294)
Partnered	0.721 (0.449)	0.647 (0.478)
Any Child <16	0.296 (0.456)	0.305 (0.461)
England	0.749 (0.434)	0.747 (0.435)
London	0.111 (0.314)	0.111 (0.314)
N. Ireland & Wales & Scotland	0.251 (0.434)	0.253 (0.435)
Avg. Weekly Earnings	662.70 (806.20)	422.20 (312.70)
Full-time worker	0.925 (0.264)	0.640 (0.480)
Sample Size	7,020	7,469

Weighted means (standard deviations).

Appendix Table 2.2
Sexual Orientation and Any Employment
UK IHS 2012-2014, Adults age 25+

	Males		Females	
	(1)	(2)	(3)	(4)
Controls for ➔	Sexual orientation + year dummies	+ demographic characteristics (age, race, education, any kids, residence) + year dummies	Sexual orientation + year dummies	+ demographic characteristics (age, race, education, any kids, residence) + year dummies
All				
Gay/Lesbian	0.026***	-0.026**	0.085***	-0.028*
	(0.012)	(0.011)	(0.015)	(0.015)
Bisexual	-0.103***	-0.114***	-0.071***	-0.078***
	(0.032)	(0.030)	(0.022)	(0.020)
R-squared	0.001	0.153	0.001	0.155
N	121206	121206	175285	175285
Non-partnered				
Gay/Lesbian	0.107***	0.014	0.013	-0.098***
	(0.017)	(0.016)	(0.031)	(0.030)
Bisexual	-0.078*	-0.124***	-0.125***	-0.180***
	(0.045)	(0.042)	(0.040)	(0.035)
R-squared	0.003	0.171	0.002	0.169
N	39508	39508	62650	62650
Partnered				
Gay/Lesbian	0.022	-0.043***	0.122***	0.005
	(0.016)	(0.015)	(0.016)	(0.015)
Bisexual	-0.003	-0.010	-0.047*	-0.032
	(0.040)	(0.037)	(0.026)	(0.024)
R-squared	0.001	0.157	0.001	0.156
N	81698	81698	112635	112635

See notes to Table 3.

Appendix Table 2.3

Expanded set of Coefficient Estimates, Fully Saturated Model

(i.e., Columns 2 and 4 of Table 2.3) 2012-2014 UK Integrated Household Surveys

	(1)	(2)	(3)	(4)
	Non-partnered males	Partnered males	Non-partnered females	Partnered females
Gay	-0.006	-0.050*	0.029	0.067***
	(0.025)	(0.028)	(0.037)	(0.025)
Bisexual	-0.110	-0.189***	-0.097*	-0.009
	(0.068)	(0.057)	(0.050)	(0.040)
Other	-0.049	-0.015	-0.053	0.015
	(0.063)	(0.049)	(0.063)	(0.060)
Refused	0.043**	0.033**	0.026	0.017
	(0.020)	(0.016)	(0.017)	(0.018)
S.O. Nonresponse	-0.155***	-0.010	-0.090***	0.006
	(0.045)	(0.012)	(0.031)	(0.018)
Age	0.069***	0.068***	0.051***	0.059***
	(0.004)	(0.002)	(0.003)	(0.003)
Age-squared	-0.001***	-0.001***	-0.001***	-0.001***
	(0.000)	(0.000)	(0.000)	(0.000)
Degree Level	0.561***	0.606***	0.708***	0.665***
	(0.016)	(0.010)	(0.020)	(0.017)
Higher Ed.	0.391***	0.403***	0.454***	0.366***
	(0.019)	(0.012)	(0.021)	(0.018)
A Level	0.250***	0.283***	0.291***	0.209***
	(0.016)	(0.010)	(0.021)	(0.017)
O Level	0.099***	0.144***	0.170***	0.114***
	(0.016)	(0.011)	(0.020)	(0.017)
Face to Face	-0.010	-0.035***	-0.009	-0.028***
	(0.010)	(0.006)	(0.009)	(0.007)
Ethnicity	Yes	Yes	Yes	Yes
Location	Yes	Yes	Yes	Yes
Family	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes
Observations	19905	55112	22385	36836
R^2	0.189	0.191	0.247	0.224

See notes to Table 2.3.

Appendix Table 2.4
Sexual Orientation and Log Earnings, Sensitivity to Job Controls
UK IHS 2012-2014, Adults age 25+ with full-time employment, males

Controls for ➔	(1) Sexual orientation only + basic + family (i.e., baseline specification)	(2) Sexual orientation only + basic + family + only private sector cont.	(3) Sexual orientation only + basic + family + only establishment size cont.	(4) Sexual orientation only + basic + family + only industry cont.	(5) Sexual orientation only + basic + family + only occupation cont.	(6) (all but occ. cont.) Sexual orientation only + basic + family + private sector + establishment size + industry controls
All males						
Gay	-0.027	-0.024	-0.028	-0.009	-0.017	-0.008
	(0.019)	(0.019)	(0.018)	(0.018)	(0.018)	(0.018)
Bisexual	-0.149***	-0.150***	-0.145***	-0.141***	-0.147***	-0.138***
	(0.032)	(0.044)	(0.044)	(0.043)	(0.041)	(0.043)
R-squared	0.198	0.199	0.217	0.220	0.257	0.238
N	75017	75017	75017	75017	75017	75017
Non-partnered males						
Gay	-0.006	-0.004	-0.010	0.011	-0.000	0.009
	(0.025)	(0.025)	(0.025)	(0.024)	(0.023)	(0.024)
Bisexual	-0.110	-0.112*	-0.099	-0.096	-0.117*	-0.089
	(0.068)	(0.068)	(0.067)	(0.065)	(0.063)	(0.065)
R-squared	0.189	0.189	0.207	0.215	0.254	0.230
N	19905	19905	19905	19905	19905	19905
Partnered males						
Gay	-0.050*	-0.047*	-0.047*	-0.031	-0.038	-0.026
	(0.028)	(0.028)	(0.027)	(0.028)	(0.027)	(0.027)
Bisexual	-0.189***	-0.186***	-0.191***	0.184***	-0.176***	-0.184***
	(0.057)	(0.058)	(0.057)	(0.057)	(0.053)	(0.057)
R-squared	0.191	0.192	0.212	0.212	0.249	0.232
N	55112	55112	55112	55112	55112	55112

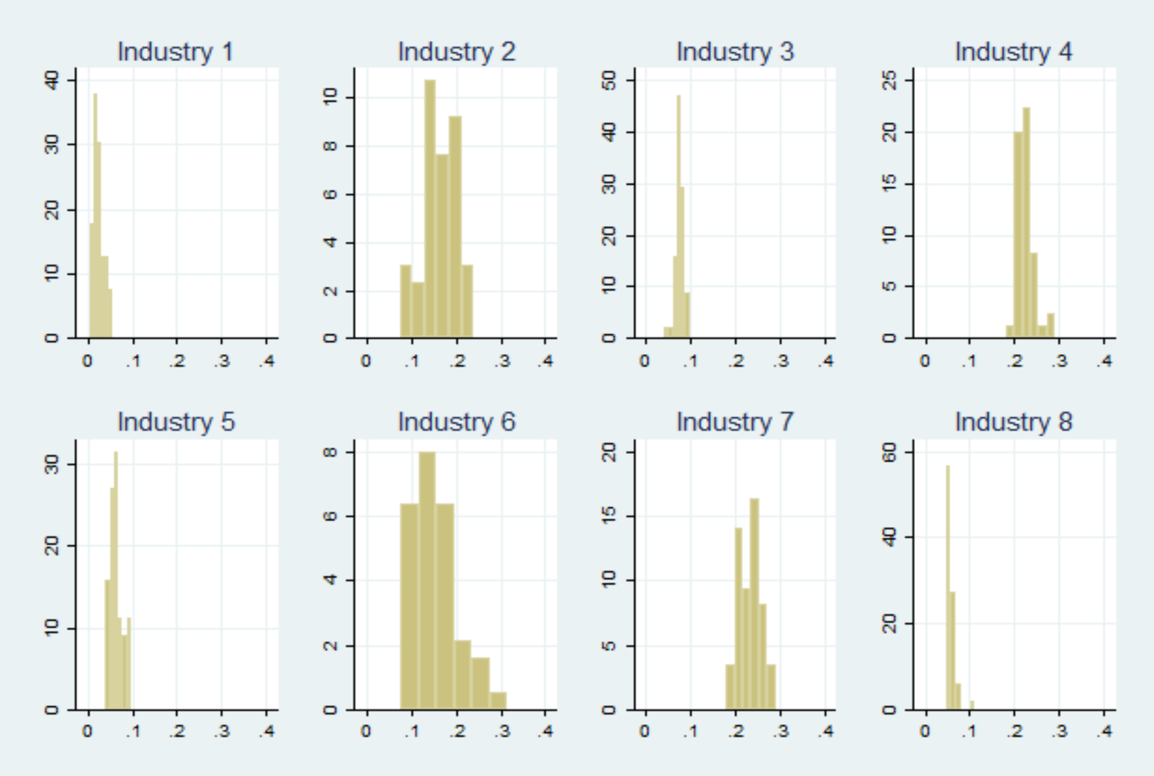
See notes to Table 2.3.

Appendix Table 2.5
Sexual Orientation and Log Earnings, Sensitivity to Job Controls
UK IHS 2012-2014, Adults age 25+ with full-time employment, females

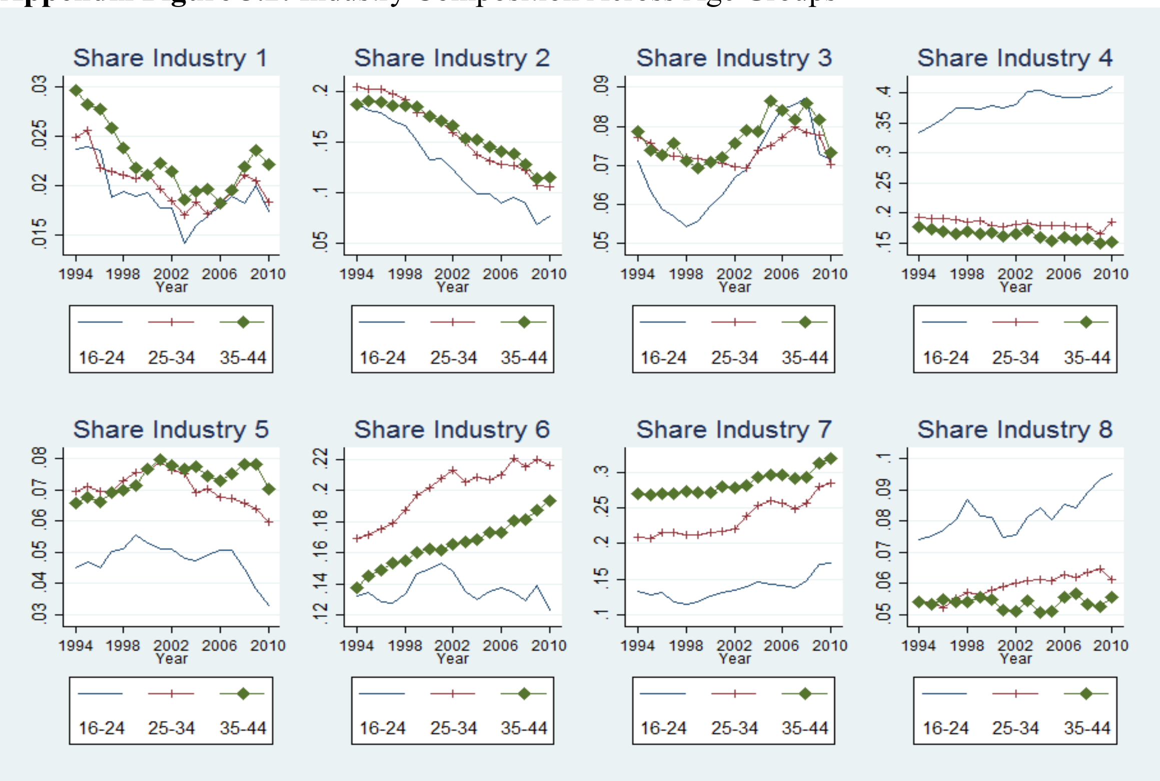
Controls for ➔	(1) Sexual orientation only + basic + family (i.e., baseline specification)	(2) Sexual orientation only + basic + family + only private sector cont.	(3) Sexual orientation only + basic + family + only establishment size cont.	(4) Sexual orientation only + basic + family + only industry cont.	(5) Sexual orientation only + basic + family + only occupation cont.	(6) (all but occ. cont.) Sexual orientation only + basic + family + private sector + establishment size + industry controls
All females						
Lesbian	0.054***	0.053**	0.048**	0.056***	0.079***	0.050**
	(0.021)	(0.021)	(0.020)	(0.020)	(0.019)	(0.020)
Bisexual	-0.036	-0.035	-0.030	-0.033	-0.028	-0.026
	(0.032)	(0.032)	(0.032)	(0.032)	(0.029)	(0.032)
R-squared	0.231	0.232	0.259	0.259	0.369	0.285
N	59221	59221	59221	59221	59221	59221
Non-partnered females						
Lesbian	0.028	0.028	0.029	0.038	0.074**	0.038
	(0.037)	(0.037)	(0.036)	(0.036)	(0.033)	(0.035)
Bisexual	-0.096*	-0.092*	-0.077	-0.112**	-0.078*	-0.086*
	(0.050)	(0.050)	(0.051)	(0.047)	(0.045)	(0.049)
R-squared	0.247	0.248	0.275	0.280	0.383	0.306
N	22385	22385	22385	22385	22385	22385
Partnered females						
Lesbian	0.068***	0.066***	0.058**	0.068***	0.083***	0.058**
	(0.025)	(0.025)	(0.025)	(0.025)	(0.023)	(0.025)
Bisexual	-0.009	-0.010	-0.010	0.000	-0.008	-0.001
	(0.041)	(0.041)	(0.041)	(0.040)	(0.037)	(0.040)
R-squared	0.223	0.224	0.251	0.248	0.362	0.274
N	36836	36836	36836	36836	36836	36836

See notes to Table 2.3.

Appendix Figure 3.1: Share of Each Industry Across Counties



Appendix Figure 3.2: Industry Composition Across Age Groups



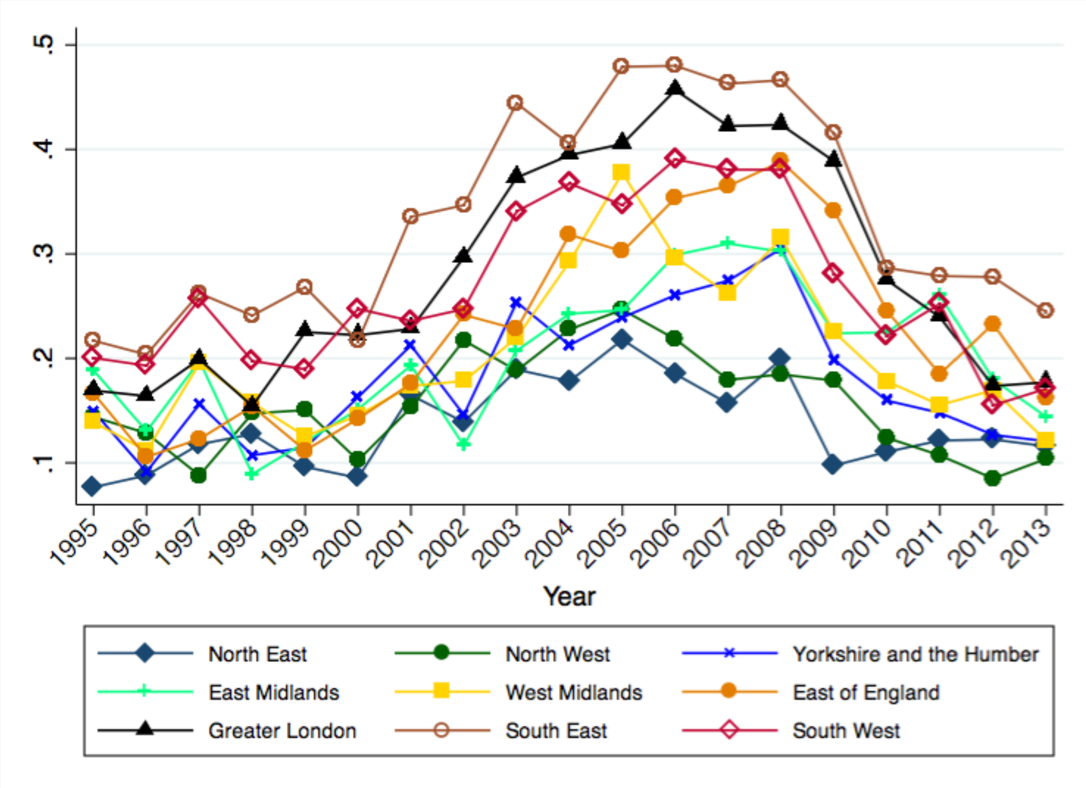
Notes: Industry 1: Agriculture Industry 2: Manufacturing Industry 3: Construction Industry 4: Distribution/Hotel/ Restaurant Industry 5: Transport and Communications Industry 6: Banking/Finance/Insurance Industry 7: Public Administration/Education/Health Industry 8: Other Services

Appendix Table 3.1: First-stage Estimations

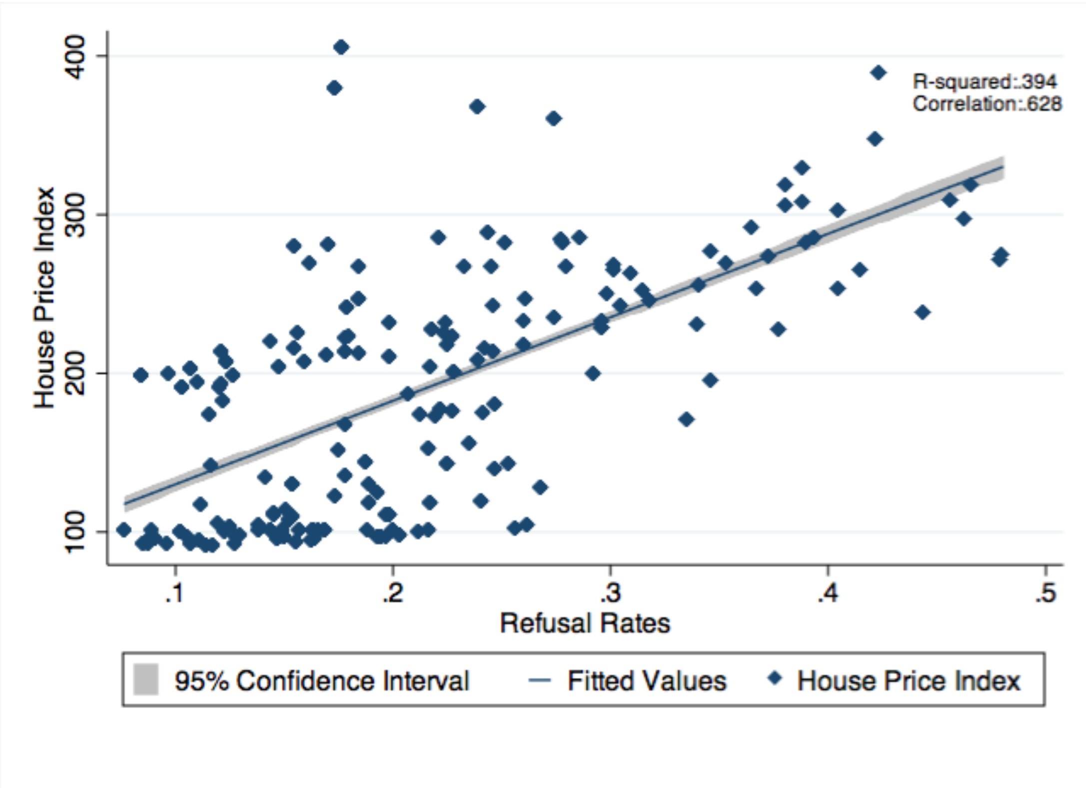
	(1)	(2)	(3)	(4)
	Aged 16-44	Aged 16-24	Aged 25-34	Aged 35-44
	First Stage	First Stage	First Stage	First Stage
Predicted Unemployment Rate	1.094***	1.107***	0.0995***	1.042***
	(0.0333)	(0.0388)	(0.0226)	(0.0229)
Year Fixed Effects	Yes	Yes	Yes	Yes
County Fixed Effects	Yes	Yes	Yes	Yes
Age Group Fixed Effects	Yes	No	No	No
County-Age Group Trends	Yes	No	No	No
County Specific Trends	Yes	Yes	Yes	Yes
N	2397	799	799	799
1^{st} Stage R^2	0.901	0.764	0.836	0.809
1^{st} Stage F Statistic	29.60	24.52	51.32	54.91

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 3.2.

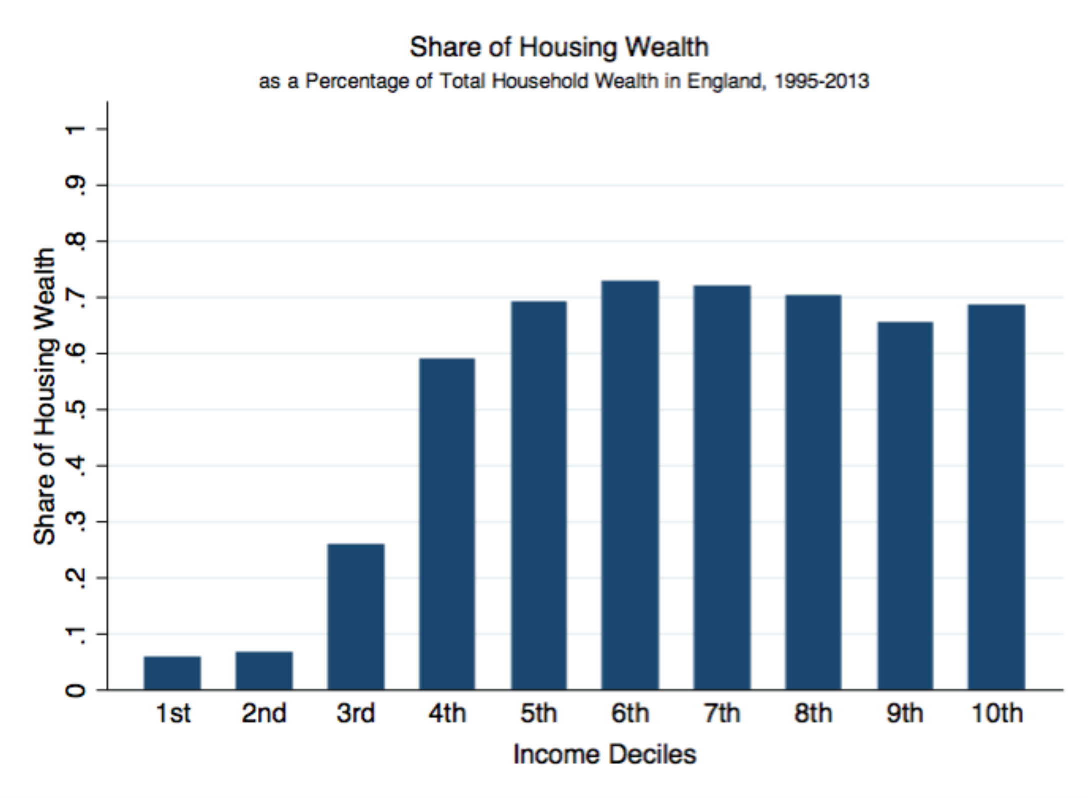
Appendix Figure 4.1: Variation in Refusal Rates across Regions



Appendix Figure 4.2: Refusal Rates and House Prices by Regions



Appendix Figure 4.3: Housing Wealth and Household Wealth



Appendix Table 4.1: First-Stage Regressions of Refusal Rates on County Level House Prices

Controls for ➔	(1) FS Fixed effects	(2) FS + demographic characteristics	(3) FS + labor market characteristics	(4) FS + county linear trends	(5) FS + county quadratic trends
3 Year MA Refusal Rates	0.038***	0.046**	0.044***	0.053***	0.050***
	(0.008)	(0.018)	(0.013)	(0.017)	(0.015)
First Stage R-squared	0.402	0.455	0.466	0.703	0.826
F Statistic	12.02	11.70	13.83	15.65	18.46
N	5624	5624	5624	5624	5624

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 1.

Appendix Table 4.2: Expanded set of Coefficient Estimates, (i.e., Column 5 of Table 4.1)

	(1) OLS	(2) IV
HPI*Ownership	0.008**	0.028***
	(0.002)	(0.008)
House Price(HPI)	-0.017***	-0.049***
	(0.004)	(0.016)
Gross Weekly Wages $_{ct-1}$	0.001	0.004
	(0.003)	(0.010)
Unemployment Rate $_{ct-1}$	0.012**	0.020*
	(0.004)	(0.009)
Gross Household Assets $_{ct-1}$	0.023***	0.039**
	(0.007)	(0.016)
Degree Level $_{ct-1}$	-0.009*	-0.023*
	(0.004)	(0.012)
Foreign Born $_{ct-1}$	0.001	0.004
	(0.003)	(0.005)
Other Ethnicities $_{ct-1}$	0.006	0.008
	(0.006)	(0.010)
One Person Family $_{ct-1}$	-0.008	-0.012
	(0.008)	(0.019)
Single $_{ct-1}$	-0.001	-0.003
	(0.004)	(0.007)
Net Population Change $_{ct-1}$	0.012**	0.032**
	(0.004)	(0.013)
Fixed Effects	Yes	Yes
County Specific Time Trends	Yes	Yes
R-squared	0.816	0.802
N	5624	5624

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%.

Appendix Table 4.3: House Prices and Birth Rates – Alternative IV/2SLS Estimations

Instruments ➔	(1) Change in Delay Rate	(2) Average Refusal Rates between 1995-2013	(3) Refusal Rates in 1994	(4) Change in Delay Rate*Refusal Rates	(5) The number of accepted dwellings over baseline housing stock
Full sample					
HPI*Ownership	0.035**	0.024*	0.019	0.032***	0.017*
	(0.013)	(0.015)	(0.010)	(0.010)	(0.008)
House Price(HPI)	-0.053**	-0.045**	-0.039*	-0.049***	-0.036**
	(0.020)	(0.018)	(0.019)	(0.015)	(0.014)
R-squared	0.822	0.777	0.781	0.801	0.768
N	5624	5624	5624	5624	5624
First stage F Statistic	12.05	9.78	7.82	13.56	8.55

Notes: * significant at 10%; ** significant at 5%; *** significant at 1%. For details on control variables, see notes to Table 1.

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