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# Game of zones: The political economy of conservation areas[*]

**Abstract:** We develop and test a simple theory of the conservation area designation process in which we postulate that the level of designation is chosen to comply with interests of local homeowners. Conservation areas provide benefits to local homeowners by reducing uncertainty regarding the future of their area. At the same time, the restrictions impose a cost by limiting the degree to which properties can be altered. In line with our model predictions we find that an increase in preferences for historic character by the local population increases the likelihood of a designation, and that new designations at the margin are not associated with significant house price capitalisation effects.

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## 1 Introduction

One of the key motivations for a variety of spatial planning policies is to solve coordination problems inherent to free markets. Among such policies historic preservation occupies a leading position in terms of the rigidity of the related regulations as well as the complexity of related social and private costs and benefits. These policies restrict individual property

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rights in order to protect buildings with a particular aesthetic, cultural or historic value. In doing so the policy may overcome a coordination problem by ensuring that owners can no longer "freeride" on the character of nearby buildings while making inappropriate changes to their own properties. In other words it may help to solve a so-called prisoner's dilemma (Holman and Ahlfeldt, 2015). A welfare-maximising preservation policy must take into account the social costs and benefits of preservation incurred by the wider society and even future generations. It is therefore unlikely that designation decisions that are considered socially optimal are also in the interest of local homeowners. In this paper we analyse whether owners are able to 'game the system' to their advantage i.e. whether the designation status of each zone in a neighbourhood is determined by the preferences of the homeowners residing there. We answer this question by deriving a model of the designation process, in which a planner acts as an agent of local homeowners, and then empirically testing its predictions.

Our theory distinguishes between a *heritage effect*, which can be internal or external, i.e. the effect of the appearance of a historic building on the perceived value of the house itself (internal) or nearby houses (external), and a *policy effect*, which results from the legal treatment of the designation policy. We argue that with positive heritage effects, the policy benefits the owners by removing uncertainty regarding the future of the neighbourhood, i.e. the presence of the heritage effect. These benefits are opposed by the costs of regulation (in the form of development restrictions and maintenance obligations) so that the net effect of the policy is ambiguous. Our theoretical framework predicts positive but diminishing returns to the share of designated land within a neighbourhood. Taking on the assumption that the planner acts in the interests of local homeowners we can derive a condition for the (political) equilibrium level of designation. This condition generates two empirically testable hypotheses. Firstly, new designations will result from increases in the local preferences for heritage. Secondly, in equilibrium, the marginal costs and benefits of designation will offset each other, resulting in a zero impact of new designations on house prices. At all other locations in a neighbourhood the effect will be positive.

We test these implications using two different empirical approaches. Firstly, we estimate the effect of changes in neighbourhood composition—what we define as gentrification—on the likelihood of designations. We address endogeneity concerns using an instrumental variable approach. Secondly, we use a variant of the difference-in-differences (DD) method to estimate the effect of new designations on the market value of properties. Our method is well suited to accommodate the endogenous nature of the treatment as it allows for heterogeneous pre-existing trends and a spatiotemporal structure of the treatment effect. Our analysis is based on the whole of England, making use of 1 million property transactions from 1995 to 2010 and approximately 8,000 designated conservation areas (CAs), of which 915 were designated in the same observation period. We also make use of ward-level education data from the UK census for 1991, 2001 and 2011 in order to analyse the effect of changing neighbourhood characteristics on the designation status. Previewing our results, we find that an increase in the local share of residents holding a university or college degree leads to an expansion of the designated area. The property price effect inside newly designated CAs turns out not to be statistically distinguishable from zero. We find evidence that the effect just outside the CA boundary is positive and significant. These results are in line with the political equilibrium policy level, suggesting that the planner adheres to local homeowner interests.

Our analysis of the CA designation process adds to a growing body of literature on the political economy of housing markets, which implicitly or explicitly assumes that property owners are able to influence political outcomes in their own interest (e.g. Ahlfeldt and Maennig, 2015; Brunner and Sonstelie, 2003; Brunner et al., 2001; Cellini et al., 2010; Dehring et al., 2008; Fischel, 2001a, 2001b; Hilber and Mayer, 2009; Oates, 1969). We also contribute to a literature that investigates policies related to spatial externalities (Hansen and Libecap, 2004; Koster and Rouwendal, 2014; Libecap and Lueck, 2011; Rossi-Hansberg et al., 2010), and a literature that investigates the costs and benefits of restrictive planning regimes (e.g. Cheshire and Hilber, 2008; Cheshire et al., 2011; Hilber and Vermeulen, 2015). Our results are also relevant to research that has looked into the value amenities add to neighbourhoods and cities more generally (e.g. Bayer et al., 2007; Brueckner et al., 1999; Chay and Greenstone, 2005; Cheshire and Sheppard, 1995; Glaeser et al., 2001; Sheppard, 2014). Notably, there is also a growing body of literature that investigates the property price effects of designation policies, mostly focussed on the U.S. (e.g. Asabere et al., 1989; Asabere and Huffman, 1994; Asabere et al., 1994; Coulson and Lahr, 2005; Coulson and Leichenko, 2001; Glaeser, 2011; Leichenko et al., 2001; Noonan, 2007; Noonan and Krupka, 2011; Schaeffer and Millerick, 1991).

The key contribution of this study is to provide insights into the political economy of CA designation and to examine whether the outcome follows local homeowners' interests. We also make a number of more specific, though still important, contributions. Firstly, the theoretical framework we develop lends a structure to the designation process that helps to interpret the existing evidence that has typically been derived from ad hoc empirical models. Secondly, our analysis of CA effects on property prices is one of the few rigorous analysis of this kind available for Europe (e.g. Ahlfeldt and Maennig, 2010; Koster et al., 2014; Lazrak et al., 2013) and the first to analyse England. It is unique in terms of the size and spatial detail of the data set and special in its focus on the spatial modelling of heritage externalities. Thirdly, our difference-in-differences analysis of the designation effects on property prices is one of the few studies that uses a quasi-experimental research design to separate the policy effect of designation from correlated location effects (Been et al., 2015; Koster et al., 2014; Noonan and Krupka, 2011). Fourthly, we make use of novel variations of the DD approach to identify temporal and spatial trends and discontinuities in policy effects, which could be applied more generally to programme evaluations. Fifthly, we provide one of the few empirical analyses of the determinants of heritage designation (Maskey et al., 2009; Noonan and Krupka, 2010, 2011). More generally, we establish a novel connection between the spatial outcome of a political bargaining process and one of the most striking contemporary urban phenomena: gentrification.

The structure of the paper is as follows. The next section introduces our theoretical model of heritage designations and the institutional setting. Section 3 presents our empirical strategy. A presentation and discussion of our empirical results is in section 4. The last section concludes.

# 2 Theory and context

## 2.1 Theoretical framework

We assume that a linear neighbourhood exists along a spatial dimension $x$ on the interval [0,1]. At each point along $x$ there exists a small zone of housing which may be designated as a CA as a whole or not.[1] Specific to each zone, housing units are endowed with distinct levels of *internal* heritage, which encompasses interior and exterior features that provide utility to the occupant. This internal heritage level monotonically declines in distance from the neighbourhood centre: $h(x) = \bar{h}g(x)$, where $g(x) \geq 0$ is a heritage density function with a strictly negative first derivative $g_x < 0$ and $\bar{h} \geq 0$ is a scale parameter that reflects the overall neighbourhood endowment with heritage.[2] Each unit of internal heritage at location $x$ also provides utility to occupants at all other locations through a spatial externality as it contributes to the neighbourhood character. In the absence of regulation owners have no control over redevelopments at other locations that are potentially detrimental to the character of the neighbourhood. The long-run expected heritage character, thus, depends on a preservation probability $0 \leq \pi < 1$. Abstracting from a spatial decay in the heritage externality we define the *external* heritage $H$ as the aggregate of the expected long-run preserved internal heritage across all locations in the neighbourhood. A social planner may designate zones as CAs from the centre out to a point $D \leq 1$ in the neighbourhood. We assume the effect of the policy is to raise the preservation probability to full certainty.[3] The level of external heritage is therefore equal to the integral of the internal heritage distribution up to $D$ plus the integral of the preservation probability $\pi$ times the distribution thereafter:

$$H(D) = \int_0^D h(x)\,dx + \pi \int_D^1 h(x)\,dx \tag{1}$$

In keeping with intuition the partial derivatives reveal that designation increases external heritage at a decreasing rate: $H_D = (1-\pi)h(D) > 0$; $H_{DD} = (1-\pi)h_D < 0$. To owners, the designation of their zone comes at the cost of restricted development rights $c(x \leq D)$. Utility increases in locational amenities $a(x)$, the consumption of a composite numeraire good $X$ and housing space $L$:

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1 The planner can either designate the whole zone or none of the zone, consistent with the idea of CAs as ensembles of buildings that work together to produce a desirable local character. Protection of single buildings is covered by listed building status. Designating a zone is assumed to approximate a marginal increase in the level of designation for the whole neighbourhood. Essentially the zone represents an infinitely small part of the whole neighbourhood.

2 This downward slope could be rationalised if a neighbourhood grew outwards from the centre and if heritage is related to housing age.

3 Our argument does not depend on the assumption of full preservation probability, only that preservation is *more likely* inside CAs.

$$U(x) = U(h(x), H(D), c(x \leq D), a(x), X, L) \quad (2)$$

, with $U_h > 0, U_H > 0, U_c < 0, U_a > 0; U_X > 0; U_L > 0; U_{XX} < 0; U_{LL} < 0; U_{DD} < 0$.[4] At any zone not being newly designated (i.e. where $x \neq D$) the marginal utility from designating an additional zone is strictly positive: $U_D(x \neq D) = U_H H_D > 0$. For zones that are themselves designated the marginal utility is $U_D(x \neq D) = U_H H_D + U_C$. We posit that a social planner acts on behalf of the owners and we thus designate all zones where the effect of being designated is non-negative.

*Hypothesis 1: The political equilibrium designation share $D^*$ increases with the preferences for heritage $U_H$*

We can solve for the political equilibrium level of designation $D^*$ by setting the marginal utility from designation to zero ($U_D(x \neq D) = U_H H_D + U_C = 0$), i.e. we find the zone where the costs and benefits of designation cancel out:[5]

$$D^* = g^{-1}\left(\frac{-U_C}{U_H(1-\pi)\bar{h}}\right) \quad (3)$$

In keeping with intuition, this political equilibrium designation share increases with preference for heritage and the heritage endowment of the neighbourhood and decreases in the cost of designation and the preservation probability ($D^*_{U_H} > 0, D^*_{h_0} > 0$, $D^*_{U_C} < 0$, $D^*_{\pi} < 0$). While all predictions are amenable to empirical tests, in principle, data limitations dictate that we concentrate on the first prediction ($D^*_{U_H} > 0$) when taking equation (3) to the data.

To develop a testable hypothesis on whether the equilibrium condition is fulfilled, i.e. the planner sets $D = D^*$, we incorporate capitalisation effects in the next step. We first assume that individuals maximise their utility defined above subject to a budget constraint $W = X + \theta(x)L$, where $\theta(x)$ is a housing bid rent. The utility function is quasi-concave in $X$ and

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[4] The diminishing marginal utility of designation $U_{DD} < 0$ does not depend on diminishing marginal utility of external heritage $U_{HH}$. In fact, $U_{HH}$ may be positive as long as the effect does not outweigh the decreasing returns to designation given by the downward slope in the internal heritage distribution: $U_{HH} < -\frac{U_H H_{DD}}{H_D H_D} H_{DD}$ so that $U_{DD} < 0$.

[5] Under the assumptions made there will be no designation if the marginal costs of designation exceed the marginal benefits even at the centre ($U_H H_D(D = 0) < -U_C$) and full designation if the marginal benefits of designation at least equalise marginal costs at the outer margin ($U_H H_D(D = 1) \geq -U_C$). For the remaining cases ($U_H H_D(D = 0) \leq -U_C > U_H H_D(D = 1)$) the solution in the text applies.

$L$, competition is perfect, thus maximisation results in the Marshallian demand functions $X^d(W, \theta(x))$, and $L^d(W, \theta(x))$ and the following indirect utility function:

$$V(x) = V\big(h(x), H(D), c(x \leq D), a(x), X^d(W, \theta(x)), L^d(W, \theta(x))\big) \tag{4}$$

, with $V_W > 0$ and $V_\theta < 0$ (given by the Envelope Theorem). Assuming perfect mobility, a central condition of the spatial equilibrium is that ceteris paribus any utility effect from a change in designation must be compensated for by an adjustment in rent to keep utility constant at an exogenous reservation level, i.e. $V_D dD = -V_\theta d\theta$ and, thus, $\theta_D = -V_D/V_\theta$.

*Hypothesis 2: In the political equilibrium D, designation of a zone leads to a zero capitalisation effect inside the zone but a positive effect in the rest of the neighbourhood.*

The marginal capitalisation effect of designation for the designated zone ($x = D$) and zones in the rest of the neighbourhood ($x \neq D$) is:

$$\theta_D(x = D) = -\frac{U_H H_D + U_C}{U_\theta};\ \theta_D(x \neq D) = -\frac{U_H H_D}{U_\theta} \tag{5}$$

If the planner sets the level of designation to the political equilibrium described above (where $U_H H_D = -U_C$) there will be a zero designation effect inside new designations $\theta_D(x = D^*) = 0$ and a positive effect outside $\theta_D(x \neq D^*) > 0$. This provides a direct test of whether planners in practice design designation policies in the interest of local homeowners. Note that if the designation share was greater than our defined political equilibrium ($D' > D^*$) then the marginal costs would exceed the benefits for the marginal zone (i.e. $U_H H_D < -U_C$) and there would be a negative capitalisation effect inside marginal designations: $\theta_D(x = D') < 0$. This is a result that we would also be able to detect with our empirical tests.

A final note is due on the nature of the political equilibrium we have defined. If the planner were able to compensate owners for the cost of designation incurred in the form of development restrictions it would be efficient and equitable to extend the CA further out. The intuition is that the marginal benefits of designation spread to all zones in the neighbourhood while the marginal costs are only incurred by the marginal zone.[6] The political equilibrium $D^*$ therefore is a second-best solution for the neighbourhood, which avoids distributional conflicts in a world where transfers are not available (such as in England). However, the neighbourhood optimum with transfers is a first-best solution only for the neighbourhood, but is not necessarily a social optimum as it does not take into account either the benefits of preservation to residents in other neighbourhoods or the impact of supply restrictions on rents at the housing market level.

## 2.2 Institutional context

In England, the designation of CAs started in 1967 and continues today under provisions 69 and 70 of the Planning Act 1990 (Listed Buildings and Conservation Areas).[7] Conservation areas are those that have been identified as having "special architectural or historic interest, the character or appearance of which is desirable to preserve or to enhance" (Section 69). The Planning Policy Guidance Note 15 (PPG15) states that a CA "may form groups of buildings, open spaces, trees, historic street patterns, village greens or features of historic or archaeological interest. It is the character of the areas rather than individual buildings that CAs seek to enhance." Conservation areas are designated on the grounds of local and regional criteria. Following designation, the Local Authority has more control over minor developments and the demolition of buildings (Botrill, 2005). However, the protection an area receives when it is designated a CA is determined at the national level to reflect the wider interests of society.

In 2011 there were around 9,800 CAs in England. Conservation areas vary in character and size. Many have strong historical links, for example, an architectural style associated with a certain period. Besides these characteristics, designation is made based on softer benefits said to have emanated from CA designation, including the creation of a unique sense of place-based identity, encouraging community cohesion and promoting regeneration (HM

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6 If there are $N$ equally populated zones in along $x$ the neighbourhood optimum ($NO$) will be $D^{NO} = g^{-1}\big(U_C/\big[N \times U_H(1-\pi)\bar{h}\big]\big)$, which increases with the number of zones $D_N > 0$.

7 However, the first legislation to protect the historic environment was enacted in 1882 when the Ancient Monuments Protection Act was passed to protect a small number of designated ancient monuments. More statutory measures came into force in the ensuing years, but it was the passage of the Ancient Monuments Consolidation and Amendment Act in 1913 that set out a more comprehensive legislative framework for the protection of ancient monuments.

Government, 2010).[8] This ‘instrumentalisation’ of conservation policy, which seeks to encompass heritage values, economic values and public policy outcomes, has been identified as a key shift in the English policy context (Pendlebury, 2009; Strange, 2003). This is reflective of the notion of heritage not as a single definable entity, but as a political, social, cultural and economic “bundle of processes” (Avrami, 2000 cited in Pendlebury, 2009: 7).

In combination with bottom-up schemes leading to designation (e.g. community-led designation), the complex heritage preservation agenda, which pursues a multitude of objectives, and the institutional setting with responsibilities shared across several institutional layers creates significant scope for organised interest groups like property owners to influence the outcome of a political bargaining process.

# 3 Empirical strategy

## 3.1 Designation process

To develop an empirical test of Hypothesis 1, that an increase in the heritage preferences $U_H$ leads to increases in the designation share $D^*$, we specify some parametric versions of the more general functions used in section 2. In particular, we assume a linear distribution of internal heritage: $h(x) = \bar{h}(1 - x)$ and a Cobb-Douglas utility function: $U(x) = a(x)e^{\varphi h(x)}e^{\gamma H(D)}e^{c(x \leq D)}X^{\delta}L^{1-\delta}$. The political equilibrium designation share is then

$$D^* = 1 - \frac{c}{(1-\pi)\gamma\bar{h}} \tag{6}$$

We adopt the common assertion that the demand for urban consumption amenities increases in education and income (Brueckner et al., 1999; Carlino and Saiz, 2008; Falck et al., 2014; Glaeser and Gottlieb, 2006; Shapiro, 2006; van Duijn and Rouwendal, 2013, 2014). It has been shown that rich and well-educated people have a higher willingness to pay for living in CAs (Koster et al., 2014) and especially so for living in more distinctive areas (Ahlfeldt and Holman, 2014). Building on this literature, we assume that the preference for

[8] For details see HM Government (2010): *The Government’s Statement on the Historic Environment for England*. London: DCMS.

heritage $\gamma_n$ in a neighbourhood $n$ at time $t$ is related to the share of people in the neighbourhood who hold a higher education certificate $(DEG_{nt})$[9] with the following functional form:

$$\gamma_{nt} = {DEG_{nt}}^{\vartheta} e^{-\varepsilon_{nt}} \tag{7}$$

where $\vartheta > 0$ such that the relationship is positive. The selection of DEG, rather than years of schooling, as an educational proxy is driven by data availability. Typically, those with (without) degrees have 16 (12) years of schooling. Therefore, in a neighbourhood with a 25% degree share, a doubling of the degree share is roughly equivalent to one year of schooling. Since the purpose of our empirical exercise is to evaluate the causal impact of changes in heritage preferences on designation status – and not the causal impact of education on heritage preference – it is sufficient to assume that $\vartheta$ captures a correlation between education and heritage preferences. $\varepsilon_{nt}$ is a random disturbance term capturing determinants of heritage preferences that are not correlated with education. Rearranging the equilibrium designation share equation (6), substituting the education degree proxy relationship in (7) and taking logs we arrive at the following empirical specification:

$$\log(1 - D_{nt}) = \alpha - \vartheta \log(DEG_{nt}) - \omega_n + \varepsilon_{nt} \tag{8}$$

$$\text{where } \alpha = \log(1 - \pi) - \log(c) \text{ and } \omega_n = \log(\bar{h}_n) + l_n. \tag{9}$$

The $n$ subscripts correspond to the individual 'neighbourhoods' of our theoretical model and we choose to represent these empirically as UK census wards. Wards are the smallest geographical areas that are comparable between 1991 and 2011 censuses. Because of data availability we use the census years of 1991 and 2011 as our time periods. All idiosyncratic time-invariant location components $l_n$ (location-specific determinants of designation not modelled in our theory) and the unobserved heritage endowment $\bar{h}_n$ of a neighbourhood $n$ as captured by $\omega_n$ as well as the preservation probability $\pi$ and the costs to owners of conservation policies are removed by taking first-differences:

$$\Delta\log(1 - D_n) = \Delta\alpha - \vartheta\, \Delta\log(DEG_n) + \Delta\varepsilon_n \tag{10}$$

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9 We also use income as a proxy for a subsample of our data set – results are reported in the appendix (section 5.1).

Our estimation equation now depicts that a neighbourhood change reflected in a positive change in (logged) educational degree share causes the (logged) share of non-designated land on the left-hand side to decrease. This is simply another way of saying that a positive change in educational degree leads to a higher designation share, although the transformation is non-linear. Note that we implicitly assume that we are in equilibrium in the sense that all areas that should be designated at $t$ are in fact designated. This is in line with Ahlfeldt and Holman (2014) who show that the distinctiveness of CAs tends to decline in their year of designation. Moreover, we estimate our model using a long difference between 1991 and 2011, which is more than two decades after the start of the policy and the initial wave of designations. By 1991 more than 78% of the CAs existing in 2011 had already been designated. Results for the smaller differences between 1991–2001 and 2001–2011, respectively, are reported in the appendix.

Equation (10) evidently follows from a stylised model world. In the empirical implementation we add a number of covariates to control for alternative determinants of designation. The ongoing designation is then only determined by the local changes in preferences, the steady ageing of buildings and the effects on heritage, which are differentiated out. To control for contagion effects in designation we add the initial (1991) designation share which we instrument with the share in 1981 to avoid a mechanical relationship with the dependent variable. A number of variables are added to account for heterogeneity in the net benefits of designation and abilities to express (collective) opinions in a political bargaining that may influence the designation decision. These include the initial (1991) degree share, the homeownership rate, the household size, the average population age and the share of foreigners (both in initial shares and changes). We alter the baseline model in a number of robustness checks to account for institutional heterogeneity at the Travel to Work Area (TTWA) level, neighbourhood appreciation trends and, to the extent possible, the historic and physical quality of the housing stock.

In practice, however, it is difficult to control for all determinants of designation that are external to our model. One particular concern is that areas can be designated if the heritage is threatened by poor maintenance in a declining neighbourhood. Such derelict will likely be negatively correlated with our explanatory variable and is unlikely to be fully captured by the control variables we have at hand. At the same time, the policy itself could make it more likely that educated people are attracted to designated areas due to a different valuation of uncertainty (reverse causality). Since an OLS estimation of equation (10) can result in a significant bias in either direction we make use of instrumental variables $z_n$, which predict changes in education, $\rho(z_n, \Delta\log DEG_n) \neq 0$, but must be conditionally uncorrelated with the differenced error term, $\rho(z_n, \Delta\varepsilon_{nt}) = 0$. We argue that rail station density (in London, additionally Underground station) and also effective employment accessibility (both time-invariant in levels) are good predictors of neighbourhood gentrification (Florida, 2002; Glaeser et al., 2001).[10] We also argue that it is unlikely that these level variables directly impact on the likelihood of designation conditional on the unobserved heritage endowment in the fixed effects $\omega_n$.

Another empirical concern is that, theoretically, a decrease in preferences for heritage must provoke a reduction of the designated area. The abolishment of CAs, however, is extremely rare in England so our data is left-censored (we do not observe increases in the share of non-designated land). Since we are interested in testing whether the mechanisms emphasized by the model are at work, and not simply the causal effect of changes in degree share on designation share, we take the model to the data using a Tobit approach:

$$Y_n^* = \Delta\alpha - \vartheta\,\Delta\log(DEG_n) + \Delta\varepsilon_n, \qquad \Delta\varepsilon_n \sim N(0, \sigma^2) \tag{11}$$

, where $Y_n^* = \Delta\log(1 - D_n)$ is a latent variable and the observed variable is defined as follows

$$Y_n = \begin{cases} Y_n^*, & if\ Y_n^* = \Delta\log(1 - D_n)\ < 0 \\ 0, & if\ Y_n^* \quad\quad\quad\quad\quad\quad \geq 0 \end{cases} \tag{12}$$

## 3.2 Equilibrium designation

To test Hypothesis 2, that there will be a zero capitalisation effect inside and a positive effect outside newly designation CAs, we employ a combination of hedonic (Rosen, 1974) and difference-in-differences (DD) (Ashenfelter and Card, 1985) methods. We begin with a DD model that estimates the average designation effect within CAs and within areas near to conservations areas from a before-and-after comparison to various control groups. We then

[10] Our measure of effective employment accessibility aggregates employment in surrounding regions weighted by distance, as in Ahlfeldt et al. (2015). Transport infrastructure is captured by a kernel density measure (Silverman, 1986) with a radius of 2km, which is considered to be the maximum distance people are willing to walk (Gibbons and Machin, 2005).

estimate a DD specification that allows for different price trends before and after the designation date in the treatment and control groups. One objective of this 'time trend DD' model is to deal with potential endogeneity driven by gentrification. Finally, we estimate a 'distance gradient DD' model that measures the treatment effect at different distances from the boundary both inside and outside the newly designated CAs.

*Difference-in-differences (DD)*

We define a group of 912 'treated' CAs as those that were designated between 1996 and 2010 to ensure we observe property transactions both before and after the designation date. Our counterfactuals are established via various control groups of housing units that are similar to the treated units but are themselves not treated. These control groups are discussed in more detail in the results section and in the appendix (section A3.2).

Our baseline DD model takes the following form:

$$p_{it} = \beta^I I_i + \beta^E E_i + \beta^{IPost}(I_i \times Post_{it}) + \beta^{EPost}(E_i \times Post_{it}) + X_i'\mu + f_n + Y_t + \epsilon_{it} \qquad (13)$$

where $p_{it}$ is the natural logarithm of the transaction price for property $i$ in time period $t$. $I_i$ is an internal treatment indicator, defined as a dummy variable equal to one if the observation is internal to a treated CA. $E_i$ is an external treatment indicator, similarly indicating observations just outside a treated CA within an area that is potentially exposed to spillovers. While our standard models use a spillover area of 500m we also experiment with various alternative spatial specifications. $Post_{it}$ is a dummy variable indicating whether the transaction year $t$ is equal to or greater than the designation year, $X_i$ is a vector of controls for property, neighbourhood and environmental characteristics, $f_n$ is a set of $n$ location fixed effects and $Y_t$ are year effects. The $\beta^{IPost}$ and $\beta^{EPost}$ parameters give the difference-in-differences estimates of the designation effect on the properties within and just outside a CA. We show in appendix 3.2 that $\beta^{IPost}$ is equal to the net marginal policy (designation costs and benefits) effect while $\beta^{EPost}$ reflects the pure (albeit spatially discounted) policy benefit.

*Time trend difference-in-differences*

The standard DD specification (13) identifies the policy treatment effect under some arguably restrictive assumptions. Firstly, the treatment and control groups follow the same trend before and after the treatment. Secondly, the treatment occurs at a singular and *a priori* known date and affects the level (and not the trend) of the outcome variable. These assumptions are evidently violated if the outcome variable does not respond immediately to the treatment, for example, because of costly arbitrage or in anticipation of the treatment, e.g. because of an investment motive by buyers. In our case, a positive pre-trend can also be associated with the gentrification that causes designation according to our theoretical model, a reverse causality problem. A positive pre-trend can also be associated with development pressure, which could affect designation by reducing the preservation probability.

To address these limitations of the standard DD we refine the model to accommodate differences in trends across the treatment and the control group. Our linear time trend DD model takes the following form:

$$\begin{aligned} p_{it} = {} & \beta^{I} I_i + \beta^{IYD}(I_i \times YD_{it}) + \beta^{E} E_i + \beta^{EYD}(E_i \times YD_{it}) + \beta^{IPost}(I_i \times Post_{it}) \\ & + \beta^{IPostYD}(I_i \times Post_{it} \times YD_{it}) + \beta^{EPost}(E_i \times Post_{it}) \\ & + \beta^{EPostYD}(E_i \times Post_{it} \times YD_{it}) + X_i'\mu + f_n + Y_t + \epsilon_{it} \end{aligned} \qquad (14)$$

where $YD_{it}$ is the number of years since the designation date, with the pre-designation years having negative values. We interact $YD_{it}$ with $Post_{it}$ to allow the trend in the treatment group relative to the control groups to change after the designation date. In other specifications we use quadratic rather than linear trends and evaluate the fit of the parametric polynomial function using a semi-parametric version of (14) that replaces the $YD_{it}$ variables with full sets of years-since-designation effects (details in appendix 3.2).

Controlling for unobserved trends makes this functional form similar to a regression discontinuity design (RDD) with a time running variable (Anderson, 2014). Gentrification trends that may be endogenous to the treatment are expected to behave smoothly around the designation date. Only the designation itself changes abruptly at the designation date. Therefore, a significant 'dis-in-diff' parameter ($\beta^{IPost}$ or $\beta^{EPost}$) can be entirely attributed to the treatment even where the usual DD assumption of homogeneous trends is violated. Furthermore, if the assumption of homogeneous trends does hold then significant trend differences ($\beta^{IYD}$,$\beta^{EYD}$, $\beta^{IPostYD}$ or $\beta^{EPostYD}$) describe designation effects that occur in anticipation or accumulate over time.

*Distance gradient difference-in-differences*

Our last empirical model estimates the treatment effect of designation at different distances from the CA boundaries as well as the discontinuity in the effect at the border itself. Estimating the distance trend in treatment effects relaxes two of the abstractions from our theory. Firstly, it relaxes the assumption that there is no spatial decay in heritage externalities.[11] We are able to detect the decline in external effects over distance and whether there are external effects that are spatially confined outside newly designated CAs. Similarly, we are able to detect whether there are positive effects towards the centre of a CA where the internal heritage density is greater. Secondly, it relaxes the assumption that marginal designations are infinitely small. In reality, there may be cases where a new CA corresponds to multiple zones in in our stylised model. In this case it would be the last zone, or the outer edge of the newly designated CA, where we would expect a zero effect.

Estimating the discontinuity in the designation effect at the border allows us to test for the existence of the policy cost, in the spirit of an RDD with a boundary distance running variable (Gibbons et al., 2013). The policy benefits from designation are expected to decay smoothly across the CA boundary since these are based on the preservation of a visual amenity. Whilst the internal heritage benefit is expected to end suddenly at the border, it is time-invariant in the short-run and will be captured by a fixed effect. The policy costs from designation (property rights restrictions) are the only time-varying heritage-related effect that ends abruptly at the boundary. These will therefore be captured by the discontinuity parameter. The DD with spatial trends we estimate takes the following form:[12]

$$\begin{aligned} p_{it} = {} & \beta^{I} T_i + \beta^{ID}(T_i \times D_i) + \beta^{IPost}(T_i \times Post_{it}) + \beta^{IDPost}(T_i \times D_i \times Post_{it}) \\ & + \beta^{O} O_i + \beta^{OD}(O_i \times D_i) + \beta^{OPost}(O_i \times Post_{it}) \\ & + \beta^{ODPost}(O_i \times D_i \times Post_{it}) + X_i'\mu + f_n + Y_t + \epsilon_{it} \end{aligned} \tag{15}$$

[11] In reality heritage externalities are likely to decline quite steeply in distance (Ahlfeldt & Maennig, 2010; Holman & Ahlfeldt, 2015; Rossi-Hansberg et al., 2010).

[12] In models with historical CAs as control groups the following terms are also included $\beta^{CD}(C_i \times D_i) + \beta^{EC} EC_i + \beta^{ECD}(EC_i \times D_i)$, where $C_i$ indicates internal to control CA and $EC_i$ external to control CA. This ensures that spatial effects are estimated conditional on the spatial trends in control CA.

where $D_i$ is the distance from the property to the CA boundary (internal distances are negative values), $O_i$ indicates properties outside a treated CA and $T_i$ indicates the CA that is nearest to a property that was treated at any point of the study period. In order to fully explore the extent of spatial externalities $O_i$ indicates a larger area outside CAs rather than just within 500m as indicated by $E_i$ in previous models.[13] As with the time trend DD specification we also estimate an expanded model specification in which we allow for quadratic distance trends and semi-nonparametric specifications replacing the distance variable with some distance bin effects. In this specification, the coefficient $\beta^{IPost}$ gives the intercept of the internal effect (i.e. the internal effect at the boundary) and $\beta^{IDPost}$ estimates how this changes with respect to internal distance. Jointly, these terms capture the net policy costs and benefits of designation inside designated areas. A zero $\beta^{IPost}$ coefficient would reflect a zero effect at the boundary and would be in line with the equilibrium condition derived in the theory section. A negative $\beta^{IDPost}$ would be in line with the existence of policy benefits (due to increased preservation probability) that spill over with decay. The parameters $\beta^{OPost}$ and $\beta^{ODPost}$ allow for a spatial discontinuity in the treatment effect at the boundary and heterogeneity in spatial trends inside and outside the treated areas. As with $\beta^{IDPost}$, a jointly negative $\beta^{IDPost} + \beta^{ODPost}$ would be in line with the decaying policy benefits external to the CA. The discontinuity at the border is measured by the external intercept term $\beta^{OPost}$. A statistically significant positive estimate would indicate a cost of the policy (inside CAs). A jointly positive effect of $\beta^{IPost} + \beta^{OPost}$ would in turn indicate the existence of policy benefits (outside CAs).

# 4 Data

We have compiled two distinct data sets for the two stages of the empirical analysis. Both data sets make use of data provided by English Heritage. These include a precise Geographical Information System (GIS) map of 8,167 CAs in England, the Conservation Areas Survey containing information on community support and risk status (average condition, vulnerability and trajectory of a CA) and a complete register of listed buildings.

For the analysis of the determinants of designation we use UK census wards as a unit of analysis. Shares of designated land within each census ward are computed in GIS. Various

[13] Specifically, the empirical analysis uses properties within 1,400m of the treated CA.

ward level data on educational level, age, ethnical background, average household size and homeownership status and vacancy rate were obtained from the UK census. Any changes in ward boundaries between the years were corrected for using the online conversion tool GeoConvert.[14] For robustness tests we also collected a measure of the ward's average income (Experian). The instrumental variables station density and employment potential are created using data from nomis (workplace employment) and the Ordinance Survey (rail stations). The average turnover in housing is approximated as the number of properties transacted per year in a ward as recorded in the Nationwide Building Society data set (see below).

For the analysis of the capitalisation effects of designation we use transactions data related to mortgages granted by the Nationwide Building Society (NBS) between 1995 and 2010. The data for England comprise 1,088,446 observations and include the price paid for individual housing units along with detailed property characteristics. These characteristics include floor space ($m^2$), the type of property (detached, semi-detached, flat, bungalow or terraced), the date of construction, the number of bedrooms and bathrooms, garage or parking facilities and the type of heating. There is also some buyer information including the type of mortgage (freehold or leasehold) and whether they are a first-time buyer. Importantly, the transaction data includes the full UK postcode of the property sold, allowing it to be assigned to grid-reference coordinates.

With this information it is possible within GIS to calculate distances to CA borders and to determine whether the property lies inside or outside these borders. Furthermore, it is possible to calculate distances and other spatial measures (e.g. densities) for the amenities and environmental characteristics such as National Parks, as well as natural features like lakes, rivers and coastline. The postcode reference also allows us to observe a merger of transactions and various household characteristics (median income and ethnic composition) from the UK census, natural land cover and land use, various amenities such as access to employment opportunities, cultural and entertainment establishments and school quality. A more detailed description of all the data used is in the appendix (section 4.1).

[14] Available at: http://geoconvert.mimas.ac.uk/

# 5 Empirical results

## 5.1 Designation process

Table 1 reports the results of our Tobit model of the designation process defined in equations (10–12). The non-instrumented baseline model is in column (1). As predicted by our theory, increases in educational levels that are presumably correlated with heritage preferences are associated with reductions in the share of non-designated land. More precisely, an increase in the degree share by 1% is associated with a 0.12% reduction in the share of non-designated land. This decrease corresponds to a $0.12\% \times (1 - \overline{D}_{t-1})/\overline{D}_{t-1} = 2.61\%$ increase in the share of designated land for a ward with the mean of the positive initial designation share $\overline{D}_{t-1} = 4.4\%$. The effect substantially increases once we instrument the change in degree share using rail station density and employment potential (column 2). This increase is in line with unobserved (positive) deterioration trends that a) increase the likelihood of designation and b) are negatively correlated with changes in degree share. Introducing the instruments, the effect of a 1% increase in degree share on the share of non-designated land increases to 0.88%, which for a ward with the mean initial designation share $\overline{D}_{t-1}$ corresponds to an increase in the designated land share of approximately 19%. While we have argued that our estimates are supposed to reflect a causal estimate of gentrification (proxied by degree shares) on designation probabilities and not necessarily a causal effect of degree share on heritage preferences, a parameter estimate of $\hat{\vartheta} = 0.88$ is at least indicative of heritage preferences increasing relatively steeply in education. It is notable that increases in the share of designated land are also positively correlated with high initial levels of degree shares.

The remaining columns in Table 1 provide variations of the benchmark model (2). We add TTWA effects to control for unobserved institutional heterogeneity in column (3). Column (4) adds several CA characteristics that capture historic quality (e.g. number of listed buildings), risk (e.g. various measures capturing vulnerability and trajectory, which may affect the preservation probability) and development pressure (e.g. vacancy rate). The latter includes a measure of property price appreciation, which we obtain from ward-level regressions of log property prices on a time trend (and property controls, see appendix section 3.1 for details). With this variable we control for a potentially positive correlation between owners' risk aversion and the value of their properties – typically their largest assets. This is a potentially important control since a larger risk aversion increases the benefit from a policy that increases certainty regarding the future of the neighbourhood and, thus, potentially increases the optimal designation share. It is a demanding control since positive price trends are potentially endogenous to changes in neighbourhood composition and may thus absorb some of the gentrification effect on designation. Model (5) replicates the benchmark model on a reduced sample of predominantly residential wards to ensure that the results are not driven by commercial agents, which we do not model in our theory.[15] None of these model alterations changes the education effect substantially. Model (6) tests for an interaction effect between homeownership rate and degree share. We find that the (positive) impact of neighbourhood change on designation shares (interaction term) is particularly large in high homeownership areas (see column 6). This is in line with a political economy literature that suggests that homeowners tend to form well-organised interest groups (e.g. Ahlfeldt and Maennig, 2015; Brunner and Sonstelie, 2003; Dehring et al., 2008; Fischel, 2001a).

The results in Table 1 offer some further interesting insights on the potential determinants of designation. We do not find evidence in support of contagion effects in designation, i.e. designated land shares do not tend to increase where shares were initially high. The likelihood of designation rises with ward population age, which could be related to a higher appreciation of heritage by the elderly. The likelihood declines in the share of foreigners, which could likewise reflect a lower appreciation among people with different cultural backgrounds. An alternative and potentially complementary explanation may be a lack of familiarity with the institutional context and, thus, a difficulty with `gaming the system'.

*Further robustness*

While our IVs comfortably pass the typical statistical tests, we experiment with four alternative sets of IVs. We also split up the 1991–2011 long difference into two shorter differences (1991–2001 and 2001–2011), use the change in income as a proxy for heritage preferences (for 2001–2011) and run the baseline model in OLS, keeping only observations with positive changes in shares of designated land. The results are presented in the appendix (section 5.2) and support those discussed here.

[15] In the results reported we drop wards with more workers than inhabitants, which amount to about 7.4% of the total sample. The results do not change qualitatively even if we drop the top quintile according to the same metric.

**Tab. 1. Designation process**

| | (1) Tobit | (2) IV Tobit | (3) IV Tobit | (4) IV Tobit | (5) IV Tobit | (6) IV Tobit |
|---|---|---|---|---|---|---|
| | Δ log share non designated land (t) | | | | | |
| Δ log degree share (t) (ϑ) | -0.112*** | -0.875*** | -0.754*** | -0.794*** | -0.874*** | -0.871*** |
| | (0.022) | (0.105) | (0.136) | (0.100) | (0.100) | (0.103) |
| log degree share (t-1) | -0.116*** | -0.426*** | -0.401*** | -0.394*** | -0.438*** | -0.403*** |
| | (0.012) | (0.043) | (0.060) | (0.042) | (0.042) | (0.041) |
| log designation share (t-1) | -0.005*** | 0.003* | 0.005** | 0.004** | 0.004** | 0.003* |
| | (0.001) | (0.002) | (0.002) | (0.002) | (0.002) | (0.002) |
| Δ log homeownership (t) | 0.207*** | 0.618*** | 0.563*** | 0.582*** | 0.658*** | 0.530*** |
| | (0.034) | (0.067) | (0.082) | (0.073) | (0.070) | (0.061) |
| log homeownership (t-1) | 0.134*** | 0.195*** | 0.208*** | 0.220*** | 0.238*** | 0.588*** |
| | (0.020) | (0.023) | (0.026) | (0.029) | (0.027) | (0.065) |
| Δ log average household size (t) | 0.037 | -0.336*** | -0.205** | -0.346*** | -0.454*** | -0.121 |
| | (0.050) | (0.074) | (0.082) | (0.076) | (0.086) | (0.074) |
| log average household size (t-1) | -0.027 | -0.304*** | -0.289*** | -0.376*** | -0.229*** | -0.353*** |
| | (0.058) | (0.074) | (0.082) | (0.077) | (0.078) | (0.076) |
| Δ log pop age (t) | -0.014 | -0.277*** | -0.214** | -0.332*** | -0.477*** | -0.078 |
| | (0.068) | (0.081) | (0.084) | (0.091) | (0.100) | (0.084) |
| log pop age (t-1) | -0.109*** | -0.252*** | -0.275*** | -0.288*** | -0.232*** | -0.263*** |
| | (0.055) | (0.062) | (0.068) | (0.074) | (0.066) | (0.063) |
| Δ log share of foreigners (t) | 0.004 | 0.075*** | 0.066*** | 0.074*** | 0.071*** | 0.051*** |
| | (0.011) | (0.015) | (0.017) | (0.015) | (0.015) | (0.014) |
| log of share of fore (t-1) | -0.003 | 0.079*** | 0.051*** | 0.079*** | 0.083*** | 0.071*** |
| | (0.007) | (0.013) | (0.016) | (0.013) | (0.013) | (0.012) |
| log price trend | | | | 0.017 | | |
| | | | | (0.022) | | |
| Δ log vacancy rate (t) | | | | -0.003 | | |
| | | | | (0.010) | | |
| log vacancy rate (t-1) | | | | -0.009 | | |
| | | | | (0.015) | | |
| log turnover in housing | | | | -0.007 | | |
| | | | | (0.006) | | |
| log listed buildings density | | | | -0.003 | | |
| | | | | (0.004) | | |
| log of share of building from pre1945 | | | | -0.021*** | | |
| | | | | (0.006) | | |
| average condition (1 best, 4 worst) | | | | -0.069*** | | |
| | | | | (0.020) | | |
| average vulnerability score (1 low, 8 high) | | | | -0.052*** | | |
| | | | | (0.019) | | |
| average trajectory score (-2 improving, +2 deteriorating) | | | | 0.037 | | |
| | | | | (0.038) | | |
| Δ log degree share (t) x homeownership (t-1) | | | | | | -0.953*** |
| | | | | | | (0.138) |
| Constant | 0.490** | 1.470*** | 1.565*** | 1.801*** | 1.351*** | 1.724*** |
| | (0.235) | (0.286) | (0.323) | (0.360) | (0.300) | (0.299) |
| *TTWA Effects* | NO | NO | YES | NO | NO | NO |
| *Residential wards only* | NO | NO | NO | NO | YES | NO |
| *CHI2* | - | 328.334 | 617.186 | 491.909 | 312.116 | 332.841 |
| *EXOG_P* | - | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| *OVERID* | - | 0.001 | - | 0.435 | 5.805 | 0.242 |
| *OVERIDP* | - | 0.981 | - | 0.509 | 0.016 | 0.623 |
| *Observations* | 7965 | 7965 | 7965 | 7965 | 7379 | 7965 |

Notes: See the data section for a description of control variables. First-stage results are in Table A5. IVs are station density, employment potential and the degree share in 1981 in all models except model (1). Model (4) includes a dummy variable indicating 60 wards for which no price trend could be computed due to insufficient transactions (trends are set to zero). Standard errors in parentheses. $^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$.

## 5.2 Equilibrium designation

*Difference-in-differences*

Table 2 shows the results from an estimation of the standard DD equation (13) for different selections of control groups and fixed effects. Each model includes controls for property, location and neighbourhood characteristics, year effects and location fixed effects to hold unobserved time-invariant effects constant. Column (1) is a naïve DD using the mean price trend of all properties located beyond 500m of a treated CA as a counterfactual. Columns (2) to (7) provide more credible counterfactuals by restricting the control group to properties that are presumably similar to the treated properties. Column (2), with ward fixed effects, and (3), with nearest CA fixed effects, provide a spatial matching by restricting the sample to properties within 2km of a treated CA, where many unobserved *location* characteristics are likely to be similar. In column (4) we impose the additional restriction that properties in the control group must fall within 500m of the boundaries of a historically designated CA (before 1996), which increases the likelihood of unobserved *property* characteristics being similar. While areas that are designated at any point in time are likely to share many similarities, the diminishing returns to designation in our theoretical framework also imply that heritage-richer areas should generally be designated first. To evaluate whether the designation date of the treated CAs, relative to those in the control groups, influences the DD estimate, we define CAs designated 1996–2002 as a treatment group and form control groups based on CAs designated just before (1987–1994) or right after (2003–2010) in columns (5) and (6). Finally, in column (7) we use environmental, property and neighbourhood characteristics to estimate the propensity of being in a treated (1996–2010) CA over a historical (<1996) CA. Then the treated CAs are matched to their 'nearest-neighbour', i.e. the most similar non-treated CA, based on the estimated propensity score (Rosenbaum and Rubin, 1983). A fixed effect is defined for each treated CA and its nearest-neighbour control CA such that the treatment effect is estimated by the direct comparison between the treated CA and its nearest neighbour.

We anticipate that the strength of the counterfactual increases as we match the treatment and control group based on proximity (2 & 3), proximity and qualifying for designation (4, 5 & 6) and qualifying for designation and a combination of various observable characteristics (7). As the credibility of the counterfactual increases, the statistical significance of the treatment effect tends to decrease. Benchmarked against the nationwide property price trend both the internal effect (Inside × Post) and the external effect (Within 500m × Post) are significant at the 5% level. The magnitudes of these effects are of similar size, implying a 2.8% premium for houses inside newly designated CAs and a 2.3% premium outside. The spatial matching (2 & 3) renders the internal treatment effect insignificant (2 & 3). With further refinements in the matching procedure the external effect also becomes insignificant. Table 2 results, thus, suggest that designation does not lead to significant property price adjustments. Evidence is weak for positive (policy) spillovers to nearby areas.

**Tab. 2. Conservation area premium – designation effect**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| | log property transaction price | | | | | | |
| Inside treated CA × Post designation | 0.028*** (0.009) | 0.014 (0.009) | 0.014 (0.010) | 0.003 (0.012) | -0.024 (0.070) | -0.077 (0.111) | -0.003 (0.013) |
| Within 500m buffer of treated CA × Post des. | 0.023*** (0.004) | 0.013*** (0.004) | 0.012*** (0.005) | 0.004 (0.006) | 0.012 (0.027) | -0.005 (0.022) | -0.005 (0.010) |
| Inside treated CA | -0.043*** (0.009) | -0.038*** (0.009) | -0.048*** (0.010) | -0.037*** (0.012) | -0.062 (0.057) | 0.029 (0.108) | -0.024 (0.021) |
| Within 500m buffer of treated CA | -0.010** (0.004) | -0.004 (0.004) | -0.011** (0.005) | 0.005 (0.005) | 0.003 (0.030) | 0.006 (0.023) | -0.002 (0.013) |
| Hedonic controls | YES | YES | YES | YES | YES | YES | YES |
| Location controls | YES | YES | YES | YES | YES | YES | YES |
| Neighborhood cont. | YES | YES | YES | YES | YES | YES | YES |
| Year effects | YES | YES | YES | YES | YES | YES | YES |
| Ward effects | YES | YES | - | - | - | - | - |
| Nearest treat. CA effects | - | - | YES | YES | YES | YES | - |
| Matched CA effects | - | - | - | - | - | - | YES |
| Treatment group: CAs designated | 1996-2010 | 1996-2010 | 1996-2010 | 1996-2010 | 1996-2002 | 1996-2002 | 1996-2010 |
| Control group | Full England sample | Within 2km of treated CA | Within 2km of treated CA | Within 500m of CA designated before 1996 & within 2km of treated CA | Within 500m of CA designated 1987-1995 & within 4km of treated CA | Within 500m of CA designated 2003-2010 & within 4km of treated CA | Within 500m of pre-1996 CA matched on propensity score |
| $R^2$ | 0.921 | 0.922 | 0.915 | 0.915 | 0.861 | 0.864 | 0.909 |
| AIC | -587,375 | -156,426 | -130,469 | -67,046 | -5,408 | -8,475 | -41,184 |
| Observation | 1,088k | 302k | 302k | 178k | 21k | 32k | 133k |

Notes: Standard errors in parentheses are clustered on location fixed effects. Conservation area control groups in columns (4)-(7) have separate fixed effects for the areas inside and outside a CA. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

*Time trend DD*

In total, we estimate 10 versions of the time trend DD specification outlined in equation (14). These include models that feature linear (1–5) and quadratic (6–10) trends and several of the control groups utilised in Table 2. Below, we provide a discussion of the general theme emerging from these models, but relegate the full set of parametric estimates to Table A11 in the appendix to save space.

Figure 1 provides a graphical illustration of the predicted effect of being in the treatment group over the control group against years-since-designation. Figure 2 provides an analogical illustration of the external treatment effect, i.e. the spillovers onto areas adjacent to the designated CAs. A horizontal red line is drawn at the mean of the pre-treatment effects in order to illustrate the differences between the time trend DD results and those of the standard DD.

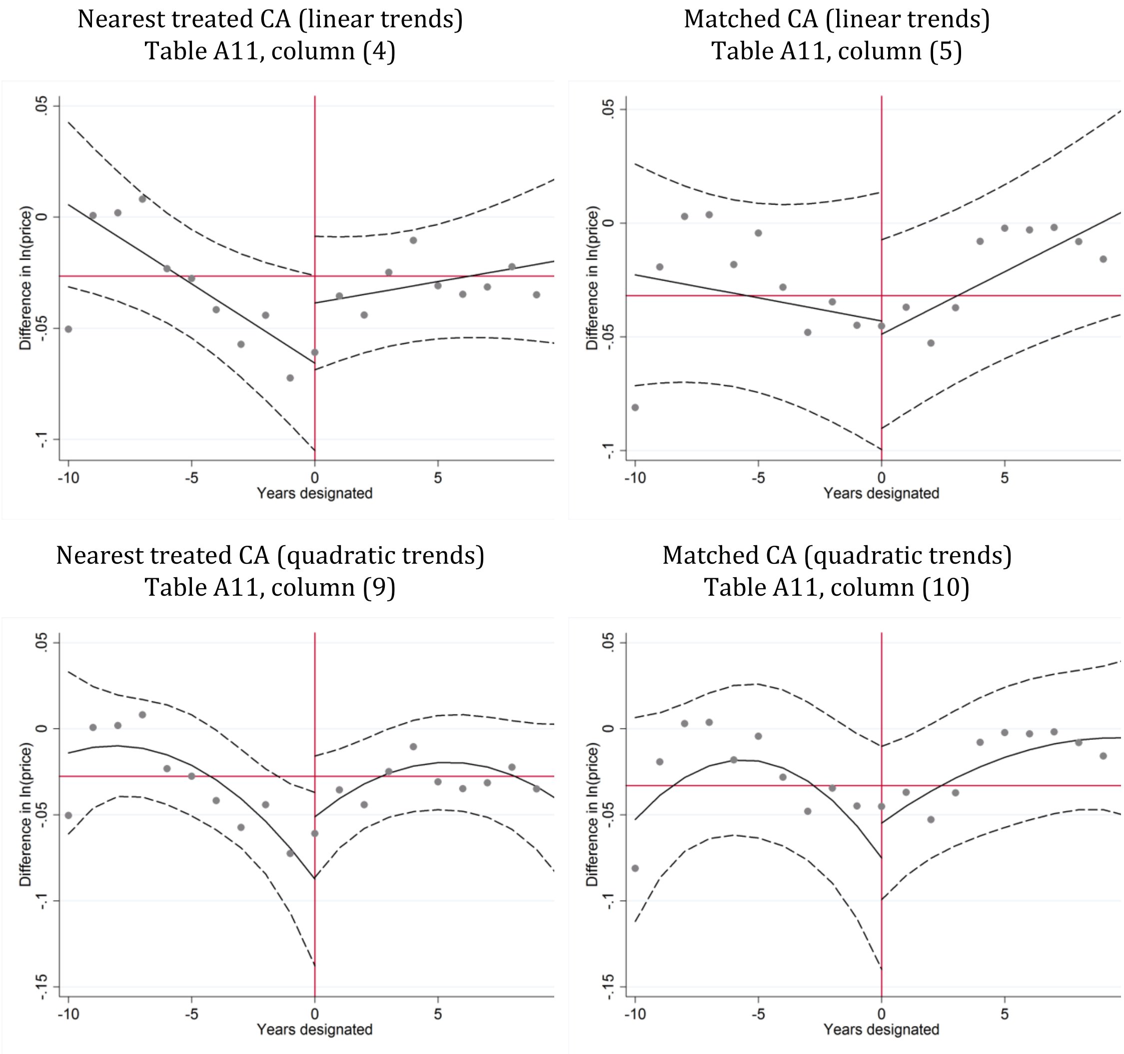
With respect to our theory, there are two key insights that emerge from the results reported in those figures. Evidence of a discontinuity in the price trend at the time of designation is weak, in particular within CAs. However, there seems to be a positive shift in price trends after designation within CAs as well as areas near to CAs. With respect to functional form those figures suggest that the linear and quadratic forms used in the parametric DD estimates (black lines) are generally supported by the more flexible semi-parametric estimates for the 'years-since-designation bins' (grey dots).[16]

These results are generally in line with the broader evidence emerging from a comparison of the parametric estimates across all specifications. Across the 10 specifications we estimate (see Table A11 in the appendix) the external (Within 500m × Post) 'dis-in-diff' parameter estimate is significant in four of 10 specifications at the 5% level and in one half of the specifications at the 10% level. The internal (Inside × Post) parameter is only significant in one specification at the 10% level (column 8). Despite the flexibility in pre- and post-trends the 'dis-in-diff' parameter is near to zero in virtually all models. This suggests primarily that there exists a significant treatment effect exactly at the treatment date only for the external area. This interpretation is in line with the predictions of our theoretical model. The positive change in the internal price trend after a CA has been designated (Inside treated CA × Post designation × Years designated) is significant at the 5% level in seven of the 10 models. This may be regarded as evidence of a cumulative internal effect of the designation policy. There is also a faster appreciation in the external area post-designation that is significant in four of the 10 models.

[16] Confidence bands for the semi-parametric 'bins' model are presented in appendix Figure A3.

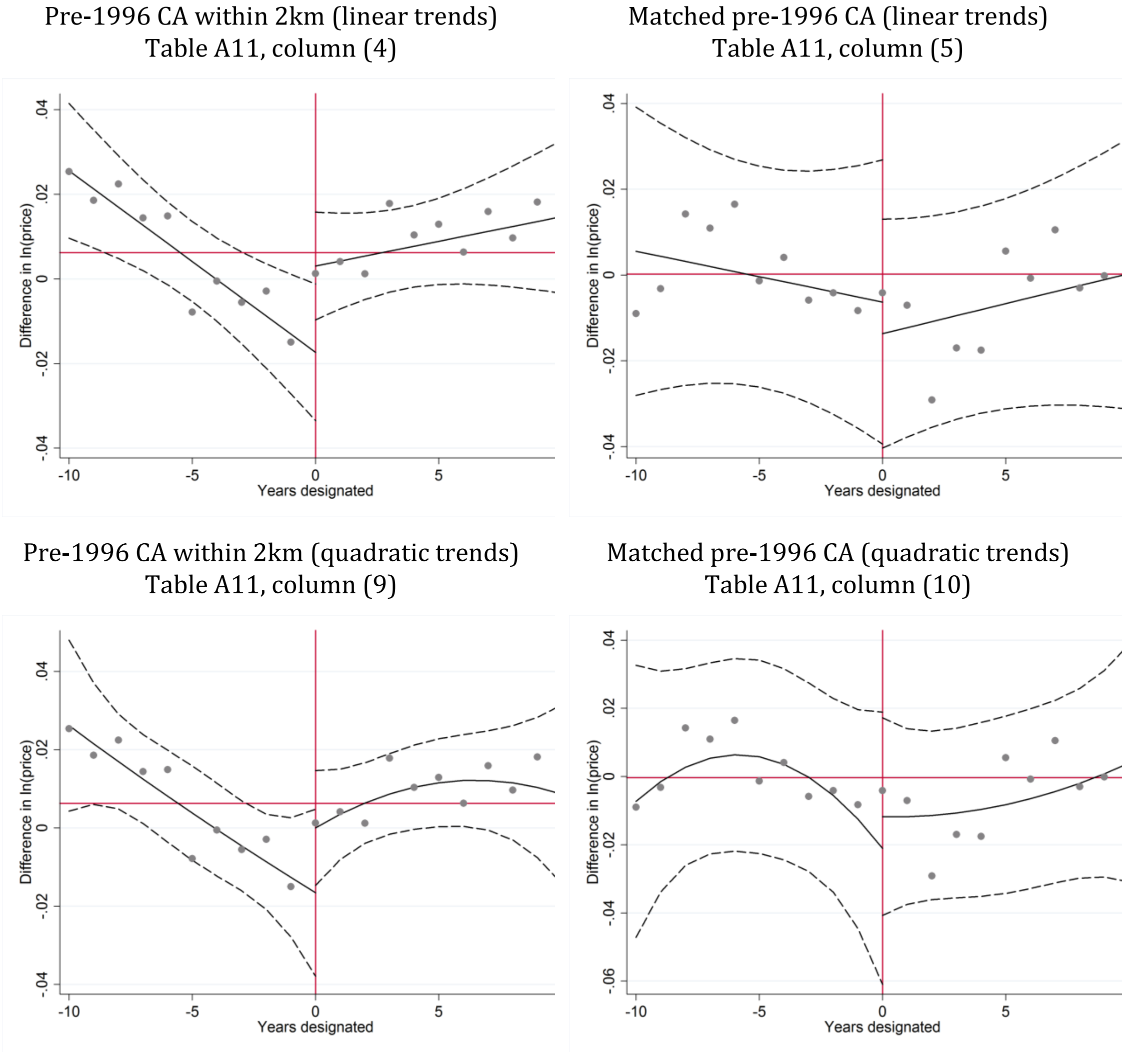
In short, the time trend DD confirms that the designation policy causes no immediate effect inside the CA but shows instead that it increases the speed of price appreciation over time. The time trend DD also uncovers that areas external to the CA receive an immediate shift in prices at the designation date in line with our theoretical hypothesis.

**Fig. 1. DD effects with time trends: Internal estimates**




Note: The solid lines are graphical illustrations of the parametric estimates presented in Table A11 and estimated using equation (14). The dashed lines indicate the 95% CI which are calculated using standard errors of multiplicative interaction terms presented by Aiken and West (1991). The grey dots plot the point estimates of 'years-since-designation bins' effects obtained from separate regression described and presented in more detail in the appendix. The horizontal red line illustrates the mean of the pre-treatment estimates.

**Fig. 2. DD effects with time trends: External estimates**



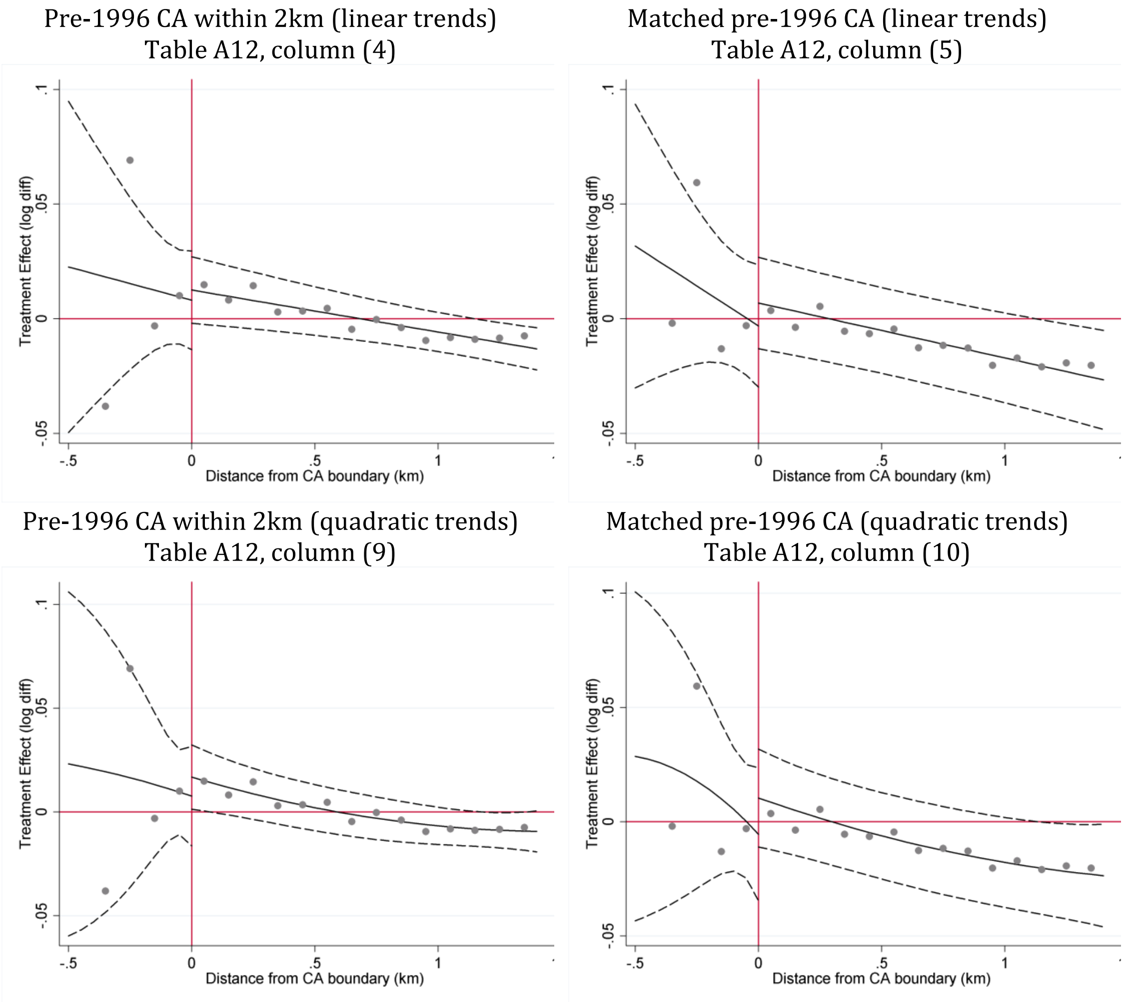

Notes: The solid lines are graphical illustrations of the parametric estimates presented in Table A11 and estimated using equation (14). The dashed lines indicate the 95% CI which are calculated using standard errors of multiplicative interaction terms presented Aiken and West (1991). The grey dots plot the point estimates of 'years-since-designation bins' effects obtained from separate regression described and presented in more detail in the appendix. The horizontal red line illustrates the mean of the pre-treatment estimates.

*Distance gradient DD*

As with the time trend DD, we estimate 10 versions of the distance gradient DD model (specification 15) that feature linear (1–5) and quadratic (6–10) trends and several of the control groups utilised in Table 2. Again, we summarise the evidence here and report all parametric estimates in the appendix (Table A12).

Figure 3 illustrates the designation effect at varying (internal and external) distances from the CA boundary. One interesting feature of Figure 3 is a positive shift in the designation effect that arises as one crosses the boundary from the inside to the outside of CAs. This positive shift is suggestive of a policy cost that exists inside CAs but not outside. We find a positive discontinuity coefficient (Outside × Post) in all 10 models we estimate, although the parameter is always statistically insignificant. In none of the models is the designation effect just inside the boundary significantly positive, which is in line with our predictions for capitalisation effects under equilibrium designation policy and a spatial decay in heritage externalities. With the strongest control groups (matched pre-1996 CA) the effect is not only statistically but also economically insignificant. In contrast, there is at least one estimate that points to a positive and significant (at 5% level) 1.6% effect just outside the CA as predicted by our theory (bottom-left panel in Figure 3). While the effect is only significant within 100m of the CA boundary, this is precisely where we expect a positive effect in a world with spatial decay in heritage (housing) externalities.

**Fig. 3. DD effects with spatial trends**




Notes: The solid lines are graphical illustrations of the parametric estimates presented in Table A12 and estimated using equation (15). The dashed lines indicate the 95% CI which are calculated using standard errors of multiplicative interaction terms presented by Aiken and West (1991).

Figure 3, more generally, is suggestive of a localised heritage character as the treatment effect increases towards the centre for the CA and decreases in external distance to the boundary. The designation effect becomes zero after about 700m at the latest, which is in line with existing evidence of a relatively steep decay in heritage and housing externalities (Ahlfeldt and Maennig, 2010; Lazrak et al., 2013; Rossi-Hansberg et al., 2010).

Briefly summarised, the results of the distance gradient DD yields estimates that are qualitatively consistent with our theory, even though the statistical significance, overall, is marginal at best. Yet it seems fair to conclude that across the treated CAs owners – at least on average – are not harmed by designation. There is some evidence that owners just outside a CA receive some benefit.

# 6 Conclusion

Historic preservation policies are among the most restrictive planning policies used to overcome coordination problems in the housing market internationally. These policies aim at increasing social welfare at the cost of constraining individual property rights. From the perspective of owners of properties in conservation areas (CA), the policy may help to solve a collective action problem, preventing owners from freeriding on the heritage character of nearby buildings while inappropriately altering their own property. If property owners value the heritage character of nearby buildings and can influence the designation process they will seek out a (local) level of designation where the marginal costs of designation equate the marginal benefits. An increase in the marginal benefit of designation will lead to an increase in designation activity. If the planner acts on behalf of the local owners, additional designations in a neighbourhood will not lead to an adverse impact on those being designated.

We provide evidence that is supportive of this scenario using two empirical approaches that follow from a simple model of equilibrium CA designation. First, we present a neighbourhood level IV Tobit analysis that reveals a positive impact of an increase in degree share, which is presumably (positively) correlated with heritage preferences, on the share of designated land. Gentrification, by increasing the value of neighbourhood stability to local owners, can cause designation. Second, we use variations of the difference-in-differences (DD) method to estimate temporally and spatially variant capitalisation effects of designation on newly designated areas as well as spillovers to adjacent areas. These methods qualify more generally as a useful tool for programme evaluations where a treatment is suspected to lead to an impact on (spatial or temporal) trends and discontinuities. Within newly designated CAs we find no significant short-run effects of designation and some evidence of positive capitalisation effects in the long run. There is some evidence of positive spillovers onto properties just outside.

These results suggest that the policy is either deliberately adhering to the interests of local owners or, as suggested in the literature on the political economy of housing markets, homeowners are able to successfully influence the outcome of local policies in their interest. It is therefore unlikely that the policy is welfare-maximising on a wider geographic scale. Depending on the general restrictiveness of the planning system, historic preservation may constrain housing supply and generate welfare losses. The net welfare effect to a wider housing market area is an interesting and important question that we leave to future research.

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# Technical appendix to Game of zones: The political economy of conservation areas

## 1 Introduction

This appendix complements the main paper and is not designed to stand alone or as a replacement. We order the sections in this appendix analogically to the sections in the main paper to facilitate cross-referencing. Section 2 provides a less technical discussion of how a planner determines the designation share and adds to the theory section of the main paper. Section 3 complements the empirical strategy section of the main paper by providing a more detailed discussion of the control variables in the Tobit designation process models. The section also links the reduced form difference-in-differences parameters to the marginal policy effect in the theoretical model. Section 4 provides a detailed overview of the data we use, its sources and how they are processed. Finally, section 5 complements the empirical results section of the main paper by showing the extended estimation outputs and results of a variety of robustness tests and model alterations not reported in the main paper for brevity.

## 2 Theory and context

### 2.1 Theoretical framework

This section provides an intuitive discussion of how a planner determines the designation share. Figure A1 illustrates a linear neighbourhood in which the heritage endowment of each zone declines in the distance from the neighbourhood centre, the nucleus of the development of the neighbourhood. For simplicity, we assume a linear decay, but the exact functional form is not crucial for the model implications: $h(x) = \bar{h}(1 - x)$, where $h_0$ is the heritage level at the centre. At any loca-

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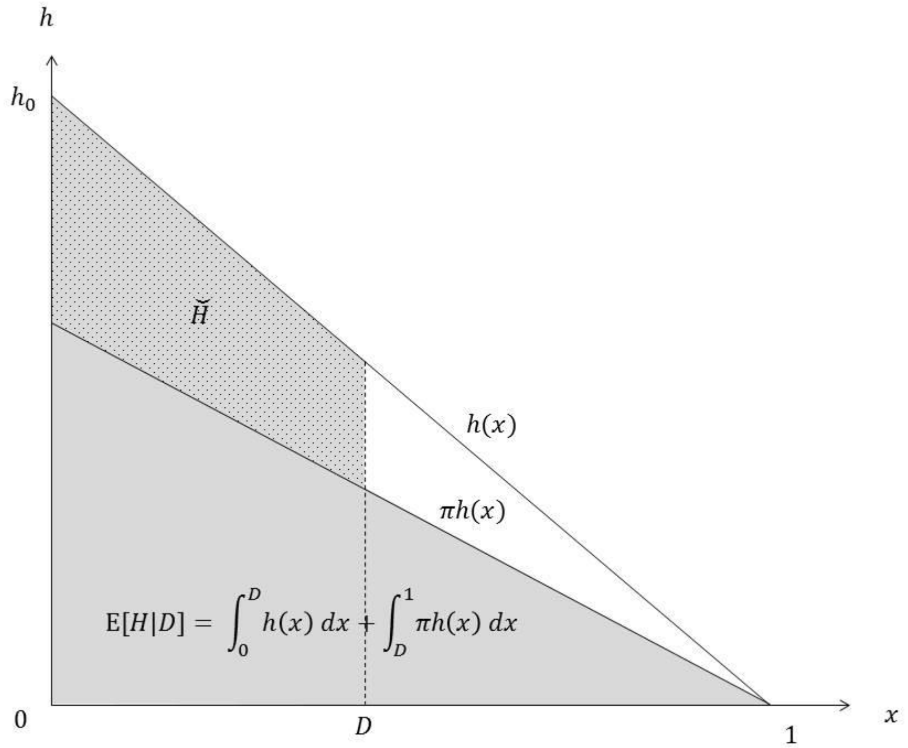
tion $x$ in the neighbourhood the expected preserved heritage is $\bar{h}(1-x)\,\pi$. The preservation probability is $\pi < 1$ if a zone is not designated because detrimental redevelopments over time affect the expected heritage character of the neighbourhood. The effect of the preservation policy is simply to increase the expected heritage to $\bar{h}(1-x)$ at all locations $0 \leq x \leq D$, because no detrimental redevelopments can take place within the protected area. The expected external heritage $E[H|D]$ (defined as $H$ in the main paper to save space) without preservation is indicated by the grey-shaded area in Figure A1. The expected amount of external heritage saved by the preservation policy is illustrated as the black-dotted area $\breve{H}$ which denotes the difference in (expected) external heritage between a scenario with no designation and a scenario with a designation share $D$. Because the heritage distribution is downward sloping it is immediately obvious that additional designations will increase the preserved external heritage at a decreasing rate. Figure A1, thus, intuitively establishes one of the stylised facts of our theory, that the social marginal benefits of designation are downward sloping.

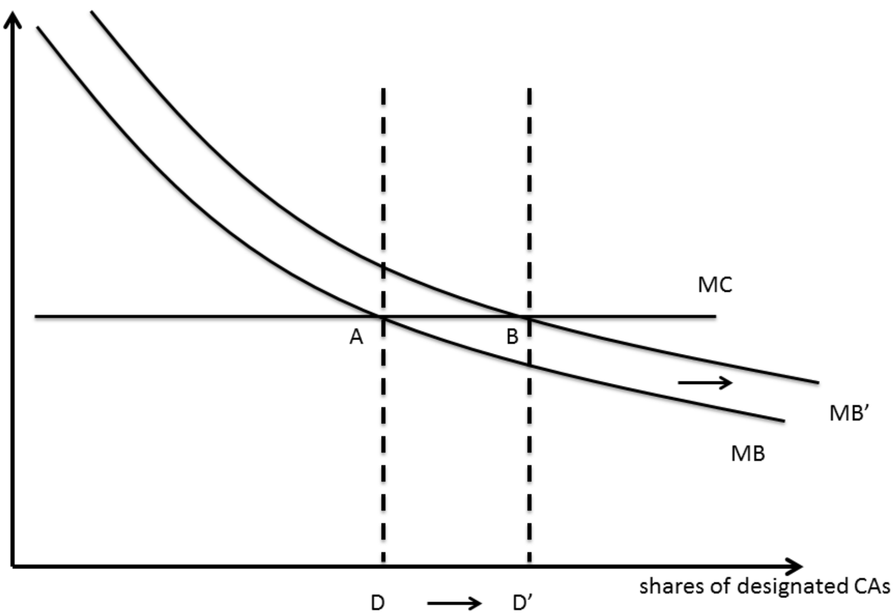
**Fig A1. Expected heritage distribution with partial designation**




Notes: The function $h(x)$ gives the internal heritage at each zone in the neighbourhood. The expected external heritage is equal to the grey-shaded area and is the integral of $h(x)$ up to the designation share plus the integral of $\pi$ times this $h(x)$ from the designation share until the neighbourhood limit at $x = 1$. The stippled area marked $\breve{H}$ is the amount of expected external heritage preserved by the policy.

The downward sloping social marginal benefits (MB) function is depicted in Figure A2. For simplicity, we assume that the marginal costs (MC) of designation are constant, i.e. the cost of not being allowed to redevelop the own house according to changing needs is the same for everyone. At point *A*, the designation share *D* adheres to local homeowner interests. The representative homeowner in each zone along *x* is happy with the designation status of their zone. However, this is not a (local) welfare-maximising equilibrium since further extension to *B* would benefit all owners in zones to the left of *A* and to the right of *B* as they would profit from increasing the expected heritage in the neighbourhood without experiencing a change in marginal cost. In zones between A and *B*, however, the social marginal benefit would also increase, but the increase would not compensate for the private marginal costs associated with a change in the designation status from undesignated to designated. Thus, if the planner cannot compensate owners between *A* and *B* because such transfer policies do not exist (such as in England), the planner chooses *A* as a second-best policy for the neighbourhood (this solution ignores demand for heritage from outside the neighbourhood and supply effects to the wider housing market area).

**Fig A2. Designation equilibrium**




If there is, however, a change in preferences and residents develop a greater taste for external heritage $\gamma$, e.g. because of gentrification, the marginal benefits curve shifts to the right. The planner adapts to this situation and raises the designation share to set marginal benefits equal to marginal costs again. This new equilibrium is illustrated by point *B* where the designation share increases to D'.

# 3 Empirical strategy

## 3.1 Designation process – control variables

This section provides a detailed description and motivation of the control variables we use to account for the determinants of conservation area (CA) designation that are unrelated to the mechanisms modelled in our theory. In particular, we try to control for composition effects, neighbourhood sorting, heterogeneity in terms of homeownership and whether the heritage in a neighbourhood is at particular risk.

We add the initial period (1991) degree share for two reasons. First, we assume that the highly educated derive higher (net-)benefits from neighbourhood heritage. To the extent that this group is capable of more efficiently articulating their will in a political bargaining a higher degree share will make the designation more likely. It is important to control for the initial degree share since levels and changes may be correlated in either direction. On the one hand there may be catch-up growth in the degree share of less educated regions, i.e. mean reversion. On the other hand, people with degrees may be more likely to move to areas with an already high share of people with degrees, which would imply a self-reinforcing process leading to spatial segregation.

We also include a control for the extent of designation in the initial period (1991). The share of designated land area in the total ward area would be (positively) correlated with the change in the designation share if designations spark further designations as in a contagion model. Initial designation also helps to control for the possibility that the skilled may be attracted to areas with a lot of designated land. To avoid a mechanic relationship between the dependent variable and the lagged designation share we instrument designation in the initial period (1991) by its lagged value, i.e. the designation share in 1981.

Another set of controls is driven by the interest in homeowners within the designation process. Homeowners experience extra benefits/costs from designation since, unlike renters, they are not compensated for changes in neighbourhood quality by increases in rents. Homeowners thus have additional incentives to engage in political bargaining. Similar to the other controls, homeownership status enters in lagged levels and differences. In a final specification we also add an interaction of the logged change in degree with homeownership (rescaled to a zero mean to make coefficients comparable). We use average household size (both in differences and lagged levels) because larger households are more likely to lobby against designation and the resulting constraint on available floor space. We control for further neighbourhood characteristics by including average population age and the share of foreigners inside a ward (also both in differences and lagged levels). We expect older residents to have a stronger appreciation of heritage, making it more likely that they would lobby for designation. Conversely, a high share of foreigners is expected to be negatively correlated with designation. Foreigners, on average, might not know the planning system that well and may perhaps find it more difficult to form interest groups. Moreover, they might value English heritage differently due to their cultural background.

A larger risk aversion increases the benefit from a policy that increases certainty regarding the future of the neighbourhood and, thus, potentially increases the optimal designation share. To control for a potentially positive correlation between owners' risk aversion and the value of their properties – typically their largest assets – we add a measure of neighbourhood appreciation. We generate ward-level property price trends in $n$ separate auxiliary regressions of the following type:

$$\log(P_{itn}) = \alpha_n + X_{ni}b_n + \beta_n T_t + \varepsilon_{itn} \qquad \text{(A1)}$$

where $X$ is a vector of property and neighbourhood characteristics and $T$ is a linear time trend. To avoid a reverse effect of designation on the property price trend we only consider transactions that occur outside CAs.

A second set of controls deals with potential development risk. Areas that experience development pressure or are in poor and/or declining condition may be more likely to be designated in order to protect against the threats to the heritage character of the neighbourhood. We use the vacancy rate, a density measure of listed buildings, housing turnover, the share of buildings pre-1945 as well as score measures for a CA's condition, vulnerability and trajectory provided by English Heritage to capture development pressure. We expect that neighbourhoods with few vacancies will be put under higher development pressure. Vacancies enter the specification both in differences and lagged levels. The reason for the differenced term is that a change in development pressure is likely to lead to a change in designation status as a result. We argue that the lagged level may also capture changes (not just levels) in development pressure. As an example, it seems likely that general population growth would put greater development pressure on neighbourhoods with lower vacancy rates. By using the total number of houses sold between 1995 and 2010 we introduce an alternative measure of development pressure. The share of houses built before

1945 serves as an indicator of potential heritage. If we are not in a steady state, building age could affect the change in designation share. The score measures reflect the development risk inside a CA and come from a survey provided by English Heritage. The higher the condition score, the worse the heritage conditions. A higher vulnerability and a higher trajectory are also indicated by higher scores. Except for the score variables, all control variables enter our empirical specification in logs.

While taking first-differences of the empirical specification will remove all time-invariant ward-specific effects that might impact on the level of designation (including the heritage itself), it will not help if there are location-specific effects that impact on the *changes* in designation status. For example, if there is heterogeneity across Local Authorities (LAs) about how difficult or easy it is to designate, arising from different bureaucratic practices, then this would affect changes in designation for all wards within a particular LA. We therefore estimate a fixed effects specification for the 166 English Travel to Work Areas (TTWAs). The TTWAs are designed to approximate city regions, which can be described as somehow self-contained economic areas, from a job market perspective. By applying a TTWA fixed effect model we are therefore able to control for unobserved socio-economic heterogeneity across TTWAs.

## 3.2 Difference-in-differences

This section motivates the difference-in-differences approach for the estimation of the marginal policy effect. Firstly, in a simple cross-sectional hedonic estimation we illustrate how the policy and heritage effects are difficult to disentangle. Secondly, we lay out how the difference-in-differences treatment effect is used to estimate the marginal policy effect laid out in terms of the structural parameters of our model.

*Cross-sectional hedonics*

Let us assume a parametric Cobb-Douglas utility function as in section 3.1 in the main paper: $U(x) = a(x)e^{\varphi h(x)}e^{\gamma E[H|D]}e^{-c\tilde{D}(x)}X^{\delta}L^{1-\delta}$, where $h(x)$ is the internal heritage endowment (i.e. heritage character of the specific housing unit), $\varphi$ is the internal heritage preference parameter, $E[H|D]$ is the external heritage (i.e. expected heritage of surrounding units, which depends on the designation policy, defined as $H$ in the main paper to save space) and is conditional on the designation share as defined above. $\gamma$ is the external heritage preference parameter and $c$ represents the costs of designation policies which arise from the development restrictions imposed inside CAs. The cost to a representative individual is $e^{-c\tilde{D}(x)}$ and depends on their zone's designation status $\widetilde{D}(x)$, a binary function of $x$ which takes the value of one if $x \leq D$ and zero otherwise (more generally defined as $c(x \leq D)$ in the main paper). Utility maximisation and a reservation utility level of $\overline{U} = 1$ due to perfect spatial competition imply an equilibrium bid rent for a representative homeowner $\theta(x) = (1-\delta)\left[\delta^{\delta} W a(x) e^{\varphi h(x)} e^{\gamma E[H|D]} e^{-c\widetilde{D}(x)}\right]^{\frac{1}{1-\delta}}$, or, in logs:[1]

$$\ln\theta(x) = \tau + \frac{1}{1-\delta}\ln a(x) + \frac{\varphi h(x)}{1-\delta} + \frac{\gamma E[H|D]}{1-\delta} - \frac{c\widetilde{D}(x)}{1-\delta} \qquad \text{(A2)}$$

The following heritage and policy effects determine the bid rent:

$$Policy\ cost = \frac{c\widetilde{D}(x)}{1-\delta} \qquad \text{(A3)}$$

$$External\ heritage\ effect\ (conditional\ on\ designation) = \frac{\gamma E[H|D]}{1-\delta} \qquad \text{(A4)}$$

$$Internal\ heritage\ effect = \frac{\varphi h(x)}{1-\delta} \qquad \text{(A5)}$$

Consider the cross-sectional reduced form equation:

$$p_{it} = \aleph I_i + \mathrm{X}_i'\mu + f_n + Y_t + \epsilon_{it} \qquad \text{(A6)}$$

where $p_{it}$ is the natural logarithm of the transaction price for property $i$ in time period $t$, $I_i$ is a dummy variable equal to one if the observation is internal to a treated CA, $\mathrm{X}_i$ is a vector of controls for property, neighbourhood and environmental characteristics, $f_n$ is a set of $n$ location fixed effects and $Y_t$ are year effects. The coefficient $\aleph$ on the $CA_i$ dummy identifies the policy cost associated with the location of a property inside a CA $\widetilde{D}(x) = 1$. The policy cost should have a negative effect on logged house prices. The coefficient also partly identifies the internal heritage effect. Specifically, it identifies the value of the difference between the mean internal heritage inside CAs and the mean internal heritage outside CAs (i.e. $\varphi/(1-\delta)\left(\overline{h_{CA_\iota=1}} - \overline{h_{CA_\iota=0}}\right)$). This should be positive because the policymaker would normally designate areas that have the most heritage. Finally, under the existence of some spatial decay in externalities, it also identifies the value of the difference inside and outside CAs in the external heritage effect (i.e. $\gamma(1-\delta)\left(\overline{E[H|D]_{CA_\iota=1}} - \overline{E[H|D]_{CA_\iota=0}}\right)$). This is a function of internal heritage and will therefore also be positive.

[1] Where $\tau$ is a constant and equal to: $\ln(1-\delta) + \frac{\delta}{1-\delta}\ln\delta + \frac{1}{1-\delta}\ln W$.

The coefficient ℵ thus reflects a composite effect of policy costs, policy benefits and correlated internal heritage effect. Furthermore, in reality the actual distribution of internal heritage is unknown and there is likely a spatial decay to externalities, further complicating the estimate.[2] In practice, ℵ will also be affected by unobserved neighbourhood characteristics that are correlated with the distance to the CA. A positive ℵ parameter, at best, tells us only that the overall higher levels of heritage (internal and external) combined with the policy benefits of conservation outweigh the policy costs. This does not provide a comprehensive evaluation of the policy effect itself. To try and disentangle these effects we implement a different empirical approach.

*Difference-in-differences*

Using the difference-in-differences (DD) approach to estimate the marginal effect of a change in designation status offers an improved identification.

Our empirical difference-in-differences specification is equation (13) from the main paper:

$$p_{it} = \beta^I I_i + \beta^E E_i + \beta^{IPost}(I_i \times Post_{it}) + \beta^{EPost}(E_i \times Post_{it}) + X_i'\mu + f_n + Y_t + \epsilon_{it} \qquad \text{(A7)}$$

Table A1 illustrates the conditional mean prices (after controlling for time effects) for the treatment and control group in the pre- and post-treatment periods. It is important to note that the year fixed effects $Y_t$ capture the general development of prices over time. Without this feature it would be necessary to control for the overall growth in price between the pre- and post-treatment periods via the inclusion of a non-interacted version of $Post_{it}$.

**Tab. A1. Interpretation of DD parameters**

| Conditional mean of prices | Pre | Post |
|---|---|---|
| Treated (Internal) | $\bar{p}_{Pre}^{Treat} = \beta^I$ | $\bar{p}_{Post}^{Treat} = \beta^I + \beta^{IPost}$ |
| Control | $\bar{p}_{Pre}^{Con} = 0$ | $\bar{p}_{Post}^{Con} = 0$ |
| $Treatment\ Effect = (\bar{p}_{Post}^{Treat} - \bar{p}_{Pre}^{Treat}) - (\bar{p}_{Pre}^{Con} - \bar{p}_{Pre}^{Con})$ | | |
| $Treatment\ Effect = ([\beta^I + \beta^{IPost}] - [\beta^I]) - ([0] - [0])$ | | |
| $Treatment\ Effect = \beta^{Post}$ | | |

Notes: The conditional mean of prices in the treatment group in the pre-period is denoted $\bar{p}_{Pre}^{Treat}$. This represents the log of prices conditional on fixed and year effects $(f_n + Y_t)$ and controls $X_i$. The same notation is used for the other groups.

[2] In a general case the estimate would be equal to:
$$\aleph = \frac{\varphi}{1-\delta}\left(\overline{h_{CA_i=1}} - \overline{h_{CA_i=0}}\right) + \frac{\gamma}{1-\delta}\left(\overline{E[H|D]_{CA_i=1}} - \overline{E[H|D]_{CA_i=0}}\right) + \frac{c}{1-\delta}$$

Our treatment coefficient $\beta^{IPost}$ essentially differentiates across the treatment and control groups before and after designation and is thus defined as follows:

$$\beta^{IPost} = (\bar{p}_{Post}^{Treat} - \bar{p}_{Pre}^{Treat}) - (\bar{p}_{Post}^{Con} - \bar{p}_{Pre}^{Con}) \qquad \text{(A8)}$$

Let us assume that the relationship between the observed conditional mean and the theoretical bid rent is given by:

$$\bar{p}_{Post}^{Treat} = \theta_{Post}^{Treat} + u_{Post}^{Treat} \qquad \text{(A9)}$$

where $u_{Pre}^{Treat}$ are partially unobservable factors specific to properties in the Treated-Post cell. The same relationship applies to the other cells (Treated-Pre, Control-Post and Control-Pre). At the heart of our identification strategy we assume that the price trends unrelated to the policy are the same within the treatment and the control group. The typical identifying assumption on which the difference-in-differences identification strategy relies can be expressed as follows:

$$(u_{Post}^{Treat} - u_{Pre}^{Treat}) = (u_{Post}^{Con} - u_{Pre}^{Con}) \qquad \text{(A10)}$$

The credibility of the counterfactual rests on the likelihood that the treatment group, in the absence of the intervention, would have followed a trend that is similar to that of the control group. An appropriate definition of the control group is therefore a critical element of the identification strategy. We therefore consider a number of different control groups in which we try to reduce the potential heterogeneity between properties in the treatment and control group.

The first treatment group is a spatial match where we choose the observations that fall within a 2km buffer surrounding CAs that changed designation status during the observation period (1995–2010). As an alternative, we consider a number of matching procedures that rest on the idea that properties inside CAs generally share similarities. Properties in CAs that did not change designation status therefore potentially qualify as a control group. To make the areas in the treatment and control group more similar, we select CAs based on similarities with those in our treatment group (Rosenbaum and Ruben, 1983). For the matching procedure we only make use of variables that turn out to have significant impact in the auxiliary propensity score matching regression.[3] We use a nearest neighbour matching procedure, which produces a broader and a narrower group.

Under the assumptions made it is straightforward to demonstrate that the DD treatment coefficient gives the pure policy effect we are interested in. Combining the theoretical bid rent of equation (A2) with the definition of $\bar{p}_{Post}^{Treat}$ in equation (A9) gives the conditional mean price of (treated) properties inside newly designated CAs before (pre) and after (post) designation:[4]

$$\bar{p}_{Pre}^{Treat} = \tau + \frac{1}{1-\delta}\ln a_i + \frac{\varphi h_i}{1-\delta} + \frac{\gamma E[H|D]}{1-\delta} + u_{Pre}^{Treat} \tag{A11}$$

$$\bar{p}_{Post}^{Treat} = \tau + \frac{1}{1-\delta}\ln a_i + \frac{\varphi h_i}{1-\delta} + \frac{\gamma}{1-\delta}\left(E[H|D] + \frac{dE[H|D]}{dD}\right) - \frac{c\widetilde{D}_i}{1-\delta} + u_{Post}^{Treat} \tag{A12}$$

where a new designation is represented as an increase in designation share $D$. For a control group sufficiently far away to not be exposed to the heritage externality we similarly get:

$$\bar{p}_{Pre}^{Con} = \tau + \frac{1}{1-\delta}\ln a_i + \frac{\gamma E[H|D]}{1-\delta} + u_{Pre}^{Con} \tag{A13}$$

$$\bar{p}_{Post}^{Con} = \tau + \frac{1}{1-\delta}\ln a_i + \frac{\gamma E[H|D]}{1-\delta} + u_{Post}^{Con} \tag{A14}$$

where there is (by definition) no new designation. Given the common trend assumption of equation (A10), $\beta^{IPost}$ identifies the pure net policy effect of designation:

$$\beta^{IPost} = \frac{\gamma}{1-\delta}\frac{dE[H|D]}{dD} - \frac{c\widetilde{D}(x)}{1-\delta} \tag{A15}$$

In the empirical implementation of the DD strategy we also consider alternative treatment groups that consist of properties just outside CAs, which are potentially exposed to spillovers, but not to the cost of designation. The interpretation of the external treatment coefficient can be derived analogically where designation leads to benefits but without the associated costs:

[3] A list of significant controls in propensity score matching regressions is included in the next subsection.

[4] Where the theoretical locations $x$ have been replaced by observed housing transactions $i$.

$$\bar{p}_{Pre}^{Treat} = \tau + \frac{1}{1-\delta}\ln a_i + \frac{\gamma E[H|D]}{1-\delta} + u_{Pre}^{Treat} \quad \text{(A16)}$$

$$\bar{p}_{Post}^{Treat} = \tau + \frac{1}{1-\delta}\ln a_i + \frac{\gamma}{1-\delta}\left(E[H|D] + \frac{dE[H|D]}{dD}\right) + u_{Post}^{Treat} \quad \text{(A17)}$$

Under the common trends assumption the treatment coefficient reflects the pure policy benefit associated with the reduction in uncertainty as predicted by the stylised theory:

$$\beta^{EPost} = \frac{\gamma}{1-\delta}\frac{dE[H|D]}{dD} \quad \text{(A18)}$$

*Propensity score matching regression*

In order to determine the control group for the difference-in-differences specification, a propensity score matching approach was employed. We used a stepwise elimination approach in order to determine which variables have a significant impact on the propensity score. With a significance level criterion of 10% the following variables remained in the final CA propensity score estimation:

**CA characteristics**: Urban, commercial, residential, industrial, world heritage site, at risk and Article 4 status.

**Environmental characteristics**: Land Cover Type 9 (Inland bare ground), Land Cover Type 3 (mountains, moors and heathland), distance to nearest national nature reserve, distance to nearest national park, national park (kernel density) and area of outstanding natural beauty (kernel density).

**Neighbourhood characteristics**: Median income and ethnicity Herfindahl index

**Amenities**: Distance to nearest bar, distance to nearest Underground station, distance to nearest hospital, distance to nearest motorway and distance to nearest TTWA centroid.

*Semi-parametric temporal and spatial estimations of treatment effects*

We estimate a semi-parametric version of the time trend DD specification (14) that replaces the $YD_{it}$ variables with a full set of years-since-designation bins. We group transactions into bins depending on the number of years that have passed since the CA they fall into or are near to had been designated. Negative values indicate years prior to designation. These bins ($b$) are captured by a set of dummy variables $PT_b$:

$$p_{it} = \sum_b \beta_b^I (PT_i^b \times I_i) + \sum_b \beta_b^E (PT_i^b \times E_i) + \sum_b \beta_b PT_i^b + X_i' \mu + f_n + Y_t + \epsilon_{it} \qquad \text{(A19)}$$

The parameters $\beta_b^I$ and $\beta_b^E$ give the difference in prices between treatment and control groups in each years-since-designation bin $b$. The results of this semi-parametric estimation are plotted in Figure A3 in appendix 5.2. In order to allow for a casual inspection of the fit of the parametric models the semi-parametric point-estimates are also plotted in Figure 1 (internal) and Figure 2 (external) of the main paper.

As with the time trend DD, we relax the parametric constraints of the distance gradient DD by replacing the distance variable in equation (15) with distance bins:

$$p_{it} = \sum_d \beta_d \left(DB_i^d \times T_i\right) + \sum_d \beta_d^{Post} \left(DB_i^d \times T_i \times Post_{it}\right) + X_i' \mu + f_n + Y_t + \epsilon_{it} \qquad \text{(A20)}$$

where $DB_i^d$ are positive (external) and negative (internal) distance bins from the designation area boundary and $\beta_d^{Post}$ are $d$ treatment effect parameters at different distances inside and outside the CA. If the planner designates according to local homeowner interests then the bin that corresponds to the locations just inside the treated CA should indicate a zero treatment effect. This may or may not be associated with a positive effect for the bins that are deepest inside the CA. Furthermore, if there are significant externalities associated with the designation (and heritage in general) then the bins just outside the boundary should indicate a positive effect. A lower effect for further out bins would indicate a spatial decay to this externality. The results from this specification are presented Fig A4. in appendix 5.2 and in Figure 3 of the main paper.

# 4 Data

## 4.1 Data sources

*Housing transactions*

The transactions data relates to mortgages for properties granted by the Nationwide Building Society (NBS) between 1995 and 2010. The data for England comprise 1,088,446 observations and include the price paid for individual housing units along with detailed property characteristics. These characteristics include floor space ($m^2$), the type of property (detached, semi-detached, flat, bungalow or terraced), the date of construction, the number of bedrooms and bathrooms, garage or parking facilities and the type of heating. There is also some buyer information, including the type of mortgage (freehold or leasehold) and whether they are a first-time buyer.

Importantly, the transaction data includes the full UK postcode of the property sold allowing it to be assigned to grid-reference coordinates. With this information it is possible within a Geographical Information Systems (GIS) environment to calculate distances to CA borders and to determine whether the property lies inside or outside these borders. Furthermore, it is possible to calculate distances and other spatial measures (e.g. densities) for the amenities and environmental characteristics that will be used as control variables. Since the data set refers to postcodes rather than individual properties, it is not possible, however, to analyse repeated sales of the same property. This is a limitation shared with most property transaction data sets available in England, including the land registry data.

*Neighbourhood characteristics*

The main variables used for estimating the capitalisation effects of neighbourhood characteristics are median income and ethnic composition. The income data is a model-based estimate of median household income produced by Experian for Super Output Areas of the lower level (LSOA). This is assigned to the transaction data based on postcodes. The data on ethnicity was made available by the 2001 UK census at the level of Output Area (OA). Shares of each of the 16 ethnic groups and a Herfindahl index[5] were computed to capture the ethnic composition of neighbourhoods.

*Environmental variables*

The environmental variables capture the amenity value of environmental designations, features of the natural environment, different types of land cover and different types of land use.

Geographical data (in the form of ESRI shapefiles) for UK national parks, areas of outstanding natural beauty, and national nature reserves are available from Natural England. National parks and areas of outstanding natural beauty are protected areas of countryside designated because of their significant landscape value. National nature reserves are "established to protect sensitive features and to provide 'outdoor laboratories' for research" (National England website). Straight line distances to these designations were computed for the housing units as geographically located by

[5] The Herfindahl index ($HI$) is calculated according to the following relation: $HI = \sum_{i=1}^{N} s_i^2$, where $s_i$ is the share of ethnicity $i$ in the LSOA, and $N$ is the total number of ethnicities.

their postcodes. Furthermore, density measures that take into account both the distance to and the size of the features were created. We apply a kernel density measure (Silverman, 1986) with a radius of 2km which is considered to be the maximum distance people are willing to walk (Gibbons and Machin, 2005).

The location of lakes, rivers and coastlines are available from the GB Ordinance Survey. The distance to these features is also computed for the housing units from the transaction data. The UK Land Cover Map, produced by the Centre for Ecology and Hydrology, describes land coverage by 26 categories as identified by satellite images. We follow Mourato et al. (2010) who construct nine broad land cover types from the 26 categories. Shares of each of these nine categories in 1km grid squares are calculated and the housing units take on the value of the grid square in which they reside.

The generalised Land Use Database (GLUD) available from the Department for Communities and Local Government gives area shares of nine different types of land use within Super Output Areas, lower level (LSOA). These nine land use types are domestic buildings, non-domestic buildings, roads, paths, rail, domestic gardens, green space, water and other land use. These shares are assigned to the housing units based on the LSOA in which they are located.

*Amenities*

The locational amenities variables capture the benefits a location offers in terms of accessibility, employment opportunities, quality of schools, and the proximity of cultural and entertainment establishments.

Employment accessibility is captured both by the distance to the Travel to Work Area (TTWA) centroid and a measure of employment potentiality. TTWAs are defined such that 75% of employees who work in the area also live within that area. Thus, they represent independent employment zones and the distance to the centre of these zones is a proxy for accessibility to employment locations. A more complex measure of accessibility is the employment potentiality index (Ahlfeldt, 2011).[6] This is computed at the Super Output Area, lower level (LSOA) and represents an average of employment in neighbouring LSOAs weighted by their distance.

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[6] Further detail on the construction of the employment potentiality measure is provided in section 4.2.

Key Stage 2 (ages 7–11) assessment scores are available from the Department for Education at the Super Output Area, middle layer (MSOA). School quality is thus captured at the housing unit level by computing a distance-weighted average of the KS2 scores of nearby MSOA centroids.[7]

Geographical data on the locations of motorways, roads, airports, rail stations and rail tracks are available from the GB Ordinance Survey. Distances were computed from housing units to motorways, A-roads, B-roads and rail stations to capture accessibility. Buffer zones were created around the motorways and roads along with distance calculations to rail tracks and airports in order to capture the disamenity noise effects of transport infrastructure.

Further data on local amenities were taken from the Ordinance Survey (police stations, places of worship, hospitals, leisure/sports centres) and OpenStreetMap (cafés, restaurants/fast food outlets, museums, nightclubs, bars/pubs, theatres/cinemas, kindergartens and monuments, memorials, monuments, castles, attractions, artwork). The number of listed buildings was provided by English Heritage. Kernel densities for these amenities were computed for housing units using a kernel radius of 2km and a quadratic kernel function (Silverman, 1986). The radius of 2km is consistent with amenities having a significant effect on property prices only when they are within walking distance.

[7] This is calculated as an Inverse Distance Weighting (IDW) with a threshold distance of 5km and a power of 2.

**Tab. A2. Variable description**

| | |
|---|---|
| Dependent Variable | |
| Price | Per square metre transaction price in British pounds of the corresponding floor space (expressed as a natural logarithm). Transaction data from the Nationwide Building Society (NBS). |
| Independent Variables | |
| CA Effects | Dummy variables denoting property transactions taking place within the boundaries of a currently existing CA, in a CA at the time when designated or where the designation date is unknown as well as various buffer areas surrounding current or treated CAs. |
| Fixed Effect Control | Travel to Work Areas, nearest CA catchment areas and interactives with year effects. |
| Housing information | Set of property variables from the NBS including: Number of bedrooms, number of bathrooms, floor size (in square metres), new property (dummy), building age (years), tenure (leasehold/freehold), central heating (full: gas, electric, oil, solid fuel), central heating (partial: gas, electric, oil, solid fuel), garage (single or double), parking space, property type (detached, semi-detached, terraced, bungalow, flat/maisonette). |
| Neighbourhood information | Set of neighbourhood variables including: media income (2005, LSOA level), share of white population at total population (2001 census, output area level), share of mixed population at total population (2001 census, output area level), share of black population at total population (2001 census, output area level), share of Asian population at total population (2001 census, output area level), share of Chinese population at total population (2001 census, output area level), Herfindahl of ethnic segregation (including population shares of White British, White Irish, White other, Mixed Caribbean, Mixed Asian, Mixed Black, Mixed other, Asian Indian, Asian Pakistani, Asian others, Black Caribbean, Black African, Black other, Chinese, Chinese other population, 2001 census output area). |
| Conservation area Characteristics | Set of characteristic variables for CAs from English Heritage including: Conservation area land use (dummy variables for residential, commercial, industrial or mixed land use), CA type (dummy variable for urban, suburban or rural type), CA size (dummy for areas larger than a mean of 128,432.04 square meters), CA (square meter), CA has an Article 4 Direction implemented (dummy), oldness of CA (dummy for areas older than a mean of 1981), CA at risk (dummy), CA with community support (dummy), CA is World Heritage Site (dummy). |
| Environment Characteristics and Amenities | Set of locational variables processed in GIS including: National parks (distance to, density), areas of outstanding beauty (distance to, density), natural nature reserves (distance to, density), distance to nearest lake, distance to nearest river, distance to nearest coastline, land in 1km square: Marine and coastal margins; freshwater, wetland and flood plains; mountains, moors and heathland; semi-natural grassland; enclosed farmland; coniferous woodland; broad-leaved/mixed woodland; urban; inland bare ground. |
| Other amenities | Set of locational variables created in GIS including: Average key stage 2 test score (MSOA averages as well as interpolated in GIS), distance to electricity transmission lines, A-roads (distance to, buffer dummy variables within 170m), B-roads (distance to, buffer dummy variable within 85m), motorway (distance to, buffer dummy variable within 315m; buffer distances refer to the distance were noise of maximum speed drops down to 50 decibel), distance to all railway stations, distance to London Underground stations, distance to railway tracks, distance to bus stations, distance to airports, densities of cafés, restaurants/fast food places, museums, nightclubs, bars/pubs, |

| | |
|---|---|
| | theatres/cinemas, kindergartens, monuments (memorial, monument, castles, attraction, artwork), hospitals, sports/leisure centres, police stations and worship locations, distance to Travel to Work Areas, employment potentiality (based on TTWAs with a time decay parameter of 0.073). |
| Neighbourhood Distance Controls | Set of neighbourhood distance dummy variables created in GIS including: Distances outside CA border (up to 50m, 100m, 150m, 200m, 250m, 300m, 350m, 400m, 1km, 2km and 3km), distances inside CA border (up to 50m, 100m, 150m, 200m). |

## 4.2 Further notes on data methods

*Employment potentiality*

The employment potentiality index is computed at the Super Output Area, lower level (LSOA) and represents an average of employment in neighbouring LSOAs weighted by their distances. Employment potentiality is calculated for each Lower Layer Super Output Area $i$ (LSOA) based on employment in all other LSOAs $j$ using the following equation:

$$\mathrm{E}P_i = \sum_j E_j e^{-a\, d_{ij}} \tag{A21}$$

where $d$ measures the straight line distance converted into travel time assuming an overall average speed of 25km/h (Department for Transport, 2009) and $E$ is the absolute number of workers in the respective LSOA. We use the spatial decay parameter of $a = -0.073$ estimated by Ahlfeldt (2013). Internal distances are calculated as:

$$d_{ij=i} = \frac{1}{3}\sqrt{\frac{Area_i}{\pi}} \tag{A22}$$

*Kernel densities for national parks, areas of outstanding natural beauty and national nature reserves*

The kernel density is a measure that takes into account both the proximity and the size of NPs, AONBs and NNRs. Every 100x100m piece of designated area is assigned a point and the density of these resulting points calculated for 10km kernels and a quadratic kernel function (Silverman, 1986, p.76, equation 4.5) around each housing unit using a kernel density method. The result is similar to calculating a share of, for example, NP area within a circle, the one difference being that the points are additionally weighted by distance to the housing units according to a normal distribution.

*Buffers for motorways and roads*

The buffer sizes for the different roads are as follows: B-road (85m), A-road (170m) and motorway (315m). These distances are calculated based on how far it is expected that the noise from traffic travelling at the speed limit of the respective roads (Steven, 2005) would decline to an assumed disamenity threshold level of noise of 50db (Nelson, 2008).

*Land cover map broad categories*

**Tab. A3. Land cover broad categories as defined by Mourato et al. (2010)**

| | |
|---|---|
| 1 | Marine and coastal margins |
| 2 | Freshwater, wetlands and flood plains |
| 3 | Mountains, moors and heathland |
| 4 | Semi-natural grasslands |
| 5 | Enclosed farmland |
| 6 | Coniferous woodland |
| 7 | Broad-leaved/mixed woodland |
| 8 | Urban |
| 9 | Inland bare ground |

# 5 Estimation results

## 5.1 Designation process

In order to test our theoretical implication that changes in heritage preferences lead to changes in designation we estimate the regression model as outlined in section 3.1 in the main paper. The prediction of the model is that positive changes in heritage preferences should lead to negative changes in the share of non-designated land in a neighbourhood. OLS regression results are reported in Table A4. We drop all zeros and identify the effect based on the sample of observations with observable changes in CA shares. The standard OLS estimates without (1) and with a basic set of composition controls (2) are insignificant. Due to the potential sources of bias in OLS discussed in the main paper (section 3.1) we re-estimate the two models using our instrumental variables. The 2SLS estimates (3) and (4) are in line with the Tobit results reported in the main paper and support the theory that a positive change in degree share leads to higher designation.

**Tab. A4. Designation regressions: OLS/2SLS models**

| | (1)<br>OLS<br>Δ log designa-<br>tion share (t) | (2)<br>OLS<br>Δ log designa-<br>tion share (t) | (3)<br>2SLS<br>Δ log designa-<br>tion share (t) | (4)<br>2SLS<br>Δ log designa-<br>tion share (t) |
|---|---|---|---|---|
| Δ log degree share (t) | -0.016 | 0.002 | -0.602*** | -0.871*** |
| | (0.013) | (0.014) | (0.096) | (0.247) |
| log degree share (t-1) | | -0.015 | | -0.379*** |
| | | (0.013) | | (0.105) |
| log designation share (t-1) | | 0.001 | | 0.006* |
| | | (0.001) | | (0.004) |
| Δ log homeownership (t) | | 0.041 | | 0.492*** |
| | | (0.032) | | (0.140) |
| log homeownership (t-1) | | 0.011 | | 0.056 |
| | | (0.023) | | (0.036) |
| Δ log average household size (t) | | 0.140 | | -0.483** |
| | | (0.107) | | (0.193) |
| log average household size (t-1) | | 0.209*** | | -0.107 |
| | | (0.032) | | (0.125) |
| log pop age (t-1) | | 0.126*** | | -0.025 |
| | | (0.041) | | (0.103) |
| Δ pop age (t) | | 0.183*** | | -0.222 |
| | | (0.047) | | (0.164) |
| log foreigner share (t-1) | | -0.019*** | | 0.083*** |
| | | (0.007) | | (0.031) |
| Δ foreigner share (t) | | 0.004 | | 0.068*** |
| | | (0.007) | | (0.026) |
| Constant | -0.040*** | -0.782*** | 0.361*** | 0.299 |
| | (0.011) | (0.169) | (0.066) | (0.497) |
| *IV* | - | - | YES | YES |
| *Controls* | - | YES | - | YES |
| $R^2$ | 0.001 | 0.047 | -0.733 | -0.445 |
| *F* | 1.516 | 15.628 | 38.934 | 5.724 |
| *AIC* | -871.268 | -925.893 | -1.359 | -268.685 |
| *OVERID* | . | . | 2.936 | 2.103 |
| *OVERIDP* | . | . | 0.087 | 0.147 |
| *Observations* | 1580 | 1580 | 1580 | 1580 |

Notes: Sample includes wards with changes in designation share. See the data section for a description of control variables. IVs are station density, employment potential and the degree share in 1981. Standard errors in parentheses and clustered on fixed effects. *$p < 0.05$, **$p < 0.01$, ***$p < 0.001$.

Table A5 reports the first-stage results to the second-stage results reported in Table 1 in the main paper. IVs are (conditionally) positively correlated with the change in degree share, and initial designation share, respectively.

**Tab. A5. Standard IV models – First-stage regressions**

| | (1) Δ log degree share (t) | (2) Δ log degree share (t) | (3) Δ log degree share (t) | (4) Δ log degree share (t) | (5) Δ log degree share(t) x home-ownership (t-1) | (6) log degree share (t) |
|---|---|---|---|---|---|---|
| rail station density | 0.098*** | 0.100*** | 0.070*** | 0.102*** | 0.021*** | -0.033 |
| | (0.026) | (0.024) | (0.019) | (0.020) | (0.006) | (0.208) |
| employment potentiality | 2.14E-8*** | 2.08E-8*** | 2.85E-8*** | 2.97E-8*** | 1.46E-9 | 7.54E-8 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| pred. Δ log degree sh. (t) x homeowners. (t-1) | | | | | 0.481*** | |
| | | | | | (0.024) | |
| log degree share (t-1) | 0.005*** | 0.006*** | 0.003** | -0.415*** | | 0.828*** |
| | (0.001) | (0.001) | (0.001) | (0.011) | | (0.019) |
| log designation share (t-2) | -0.021 | -0.020 | -0.021*** | 0.005*** | -0.025*** | -0.005*** |
| | (0.017) | (0.015) | (0.008) | (0.001) | (0.008) | (0.001) |
| Δ log homeownership (t) | 0.527*** | 0.540*** | 0.636*** | 0.596*** | -0.007 | -0.707*** |
| | (0.063) | (0.062) | (0.074) | (0.078) | (0.030) | (0.181) |
| log homeownership (t-1) | 0.145*** | 0.174*** | 0.228*** | 0.183*** | 0.213*** | -0.536*** |
| | (0.030) | (0.033) | (0.045) | (0.041) | (0.019) | (0.131) |
| Δ log average hh. size (t) | -0.445*** | -0.400*** | -0.495*** | -0.529*** | 0.162* | -0.153 |
| | (0.076) | (0.067) | (0.079) | (0.089) | (0.068) | (0.286) |
| log average hh. size (t-1) | -0.235*** | -0.277*** | -0.250** | -0.091 | -0.006 | -1.318** |
| | (0.070) | (0.069) | (0.086) | (0.095) | (0.045) | (0.442) |
| log pop age (t-1) | -0.087 | -0.040 | -0.289*** | 0.001 | 0.008 | 0.584 |
| | (0.052) | (0.055) | (0.072) | (0.059) | (0.033) | (0.335) |
| Δ pop age (t) | -0.321*** | -0.256*** | -0.490*** | -0.552*** | 0.155*** | 0.216 |
| | (0.086) | (0.068) | (0.095) | (0.079) | (0.042) | (0.356) |
| log foreigner share (t-1) | 0.080*** | 0.083*** | 0.079*** | 0.076*** | -0.005 | 0.053 |
| | (0.008) | (0.009) | (0.009) | (0.007) | (0.003) | (0.045) |
| Δ foreigner share (t) | 0.091*** | 0.087*** | 0.093*** | 0.077*** | -0.003 | 0.009 |
| | (0.019) | (0.016) | (0.020) | (0.016) | (0.003) | (0.068) |
| Log price trend | | | 0.001 | | | |
| | | | (0.028) | | | |
| Δ log vacancy rate (t) | | | 0.037** | | | |
| | | | (0.012) | | | |
| log vacancy rate (t-1) | | | 0.070*** | | | |
| | | | (0.013) | | | |
| Log listed buildings | | | 0.008 | | | |
| | | | (0.004) | | | |
| log turnover in housing transactions (t) | | | -0.016** | | | |
| | | | (0.006) | | | |
| log of share of building from pre1945 | | | 0.016*** | | | |
| | | | (0.004) | | | |
| Constant | 0.687** | 0.537* | 1.457*** | 0.242 | 0.052 | -0.739 |
| | (0.233) | (0.219) | (0.342) | (0.309) | (0.171) | (1.446) |
| *TTWA FE* | - | YES | - | - | - | - |
| *Housing cond.* | - | - | YES | - | - | - |
| *Residential wards* | - | - | - | YES | - | - |
| *Observations* | 7965 | 7965 | 7965 | 7379 | 7965 | 7965 |
| *F* | 592.006 | . | 339.162 | 508.799 | . | 1852.756 |
| $R^2$ | 0.708 | 0.742 | 0.719 | 0.709 | 0.960 | 0.717 |

Notes: Columns (1–4) are first stages of columns (2–5) in Table 1 in the main paper. Columns (5–6) are first stages of model (6) in the main paper. See the data section for a description of control variables. IVs are station density, employment potential and the degree share in t-2 all models. Model (3) includes a dummy variable indicating 60 wards for which no price trend could be computed due to insufficient transactions. We derive the instrument (predicted Δ log degree share (t) x homeownership (t-1)) for the interaction term in model (5) by interacting homeownership (t-1) with the predicted values of an auxiliary regression where we regress Δ log degree share on the exogenous variables, i.e. on the standard IVs and controls. Standard errors in parentheses and clustered on fixed effects. $^{*}p< 0.05$, $^{**}p< 0.01$, $^{***}p< 0.001$.

We have tried four sets of instrumental variables with our benchmark model (Table 1, column 2 in the main paper). The coefficient estimates reported in Table A6 remain qualitatively similar and quantitatively close to the main model. First-stage results are reported in Table A7. The alternative instruments, again, pass the validity tests. Only the overidentification test is failed by specification (1) using employment potentiality and museum density as instruments.

**Tab. A6. Alternative IV models**

| | (1) *Δ* log designation share (t) | (2) *Δ* log designation share (t) | (3) *Δ* log designation share (t) | (4) *Δ* log designation share (t) |
|---|---|---|---|---|
| *Δ* log degree share (t) | -0.828*** | -0.860*** | -0.845*** | -0.875*** |
| | (0.113) | (0.115) | (0.111) | (0.117) |
| log degree share (t-1) | -0.408*** | -0.421*** | -0.415*** | -0.427*** |
| | (0.047) | (0.047) | (0.046) | (0.048) |
| log designation share (t-1) | 0.003 | 0.003 | 0.003 | 0.003 |
| | (0.002) | (0.002) | (0.002) | (0.002) |
| *Δ* log homeownership (t) | 0.594*** | 0.612*** | 0.604*** | 0.610*** |
| | (0.070) | (0.071) | (0.070) | (0.071) |
| log homeownership (t-1) | 0.194*** | 0.196*** | 0.194*** | 0.197*** |
| | (0.023) | (0.023) | (0.023) | (0.023) |
| *Δ* log average household size (t) | -0.313*** | -0.329*** | -0.324*** | -0.334*** |
| | (0.077) | (0.078) | (0.077) | (0.078) |
| log average household size (t-1) | -0.281*** | -0.295*** | -0.289*** | -0.299*** |
| | (0.075) | (0.076) | (0.075) | (0.076) |
| log pop age (t-1) | -0.240*** | -0.246*** | -0.243*** | -0.246*** |
| | (0.062) | (0.062) | (0.062) | (0.062) |
| Δ pop age (t) | -0.270*** | -0.280*** | -0.277*** | -0.273*** |
| | (0.083) | (0.083) | (0.082) | (0.082) |
| log foreigner share (t-1) | 0.074*** | 0.077*** | 0.075*** | 0.078*** |
| | (0.014) | (0.014) | (0.014) | (0.014) |
| Δ foreigner share (t) | 0.070*** | 0.073*** | 0.072*** | 0.075*** |
| | (0.016) | (0.016) | (0.016) | (0.016) |
| Constant | 1.394*** | 1.436*** | 1.419*** | 1.438*** |
| | (0.289) | (0.291) | (0.289) | (0.291) |
| *Controls* | YES | YES | YES | YES |
| *IV* | YES | YES | YES | YES |
| *Observations* | 7965 | 7965 | 7965 | 7968 |
| *CHI2* | 319.851 | 318.289 | 321.092 | 316.186 |
| *EXOG_P* | 0.000 | 0.000 | 0.000 | 0.000 |
| *OVERID* | 2.289 | 0.084 | 0.500 | 0.233 |
| *OVERIDP* | 0.130 | 0.772 | 0.479 | 0.629 |
| *Instruments (as densities except employment pot.)* | Employment potentiality<br>Museum | Employment potentiality<br>Coffee place | Employment potentiality<br>Bar | Rail station<br>Coffee place |

Notes: Baseline model is in column (2) in Table 1 in the main paper. First-stage results are in Table A7. See the data section for a description of control variables. Standard errors in parentheses and clustered on fixed effects. $^{*}p< 0.05$, $^{**}p< 0.01$, $^{***}p< 0.001$.

Furthermore, we split the long difference between 1991 and 2011 into two shorter differences of 1991 to 2001 and 2001 to 2011. For the latter short difference, we moreover use the change in income instead of change in degree as a proxy for heritage preferences. The coefficient estimates remain qualitatively similar to the main model and are reported with their first stages in tables

A8 and A9. The coefficient of the key variable is slightly smaller than in the benchmark specification of the short difference between 1991 and 2001 (column 4) and considerably larger for the period between 2001 and 2011 (column 8). In columns (9–12) we use income as a proxy of heritage preference. Focussing on the benchmark specification in the final column, doubling income more than quadruples the designation share. The respective instruments are valid and sufficiently strong. Overall, the results are in line with our theory; increases in heritage preferences, proxied by changes in degree share or change in income, lead to increases in designation shares.

**Tab. A7. Alternative IV models – First-stage regressions**

| | (1) *Δ* log degree share (t) | (2) *Δ* log degree share (t) | (3) *Δ* log degree share (t) | (4) *Δ* log degree share (t) |
|---|---|---|---|---|
| employment potentiality | 3.07E-8*** | 2.95E-8*** | 2.85E-8*** | |
| | (0.000) | (0.000) | (0.000) | |
| museum density | 0.086 | | | |
| | (0.053) | | | |
| coffee place density | | 0.004 | | -0.007 |
| | | (0.004) | | (0.005) |
| bar density | | | 0.004 | |
| | | | (0.003) | |
| rail station density | | | | 0.196*** |
| | | | | (0.018) |
| log degree share (t-1) | -0.409*** | -0.410*** | -0.411*** | -0.409*** |
| | (0.010) | (0.010) | (0.010) | (0.009) |
| log designation share (t-2) | 0.005*** | 0.005*** | 0.005*** | 0.005*** |
| | (0.001) | (0.001) | (0.001) | (0.001) |
| *Δ* log homeownership (t) | 0.521*** | 0.516*** | 0.521*** | 0.534*** |
| | (0.064) | (0.063) | (0.067) | (0.061) |
| log homeownership (t-1) | 0.137*** | 0.135*** | 0.141*** | 0.128** |
| | (0.032) | (0.034) | (0.034) | (0.039) |
| *Δ* log average household size (t) | -0.465*** | -0.463*** | -0.455*** | -0.441*** |
| | (0.070) | (0.070) | (0.070) | (0.077) |
| log average household size (t-1) | -0.272*** | -0.276*** | -0.257*** | -0.240*** |
| | (0.067) | (0.066) | (0.061) | (0.064) |
| log pop age (t-1) | -0.099 | -0.099 | -0.088 | -0.101 |
| | (0.051) | (0.052) | (0.053) | (0.052) |
| Δ pop age (t) | -0.314*** | -0.316*** | -0.312*** | -0.345*** |
| | (0.086) | (0.090) | (0.085) | (0.086) |
| log foreigner share (t-1) | 0.081*** | 0.082*** | 0.081*** | 0.087*** |
| | (0.009) | (0.009) | (0.009) | (0.010) |
| Δ foreigner share (t) | 0.090*** | 0.091*** | 0.091*** | 0.091*** |
| | (0.019) | (0.019) | (0.019) | (0.018) |
| Constant | 0.039 | 0.051 | 0.035 | -0.015 |
| | (0.092) | (0.094) | (0.091) | (0.091) |
| *Controls* | YES | YES | YES | YES |
| *Observations* | 7965 | 7965 | 7965 | 7968 |
| *F* | 568.539 | 566.433 | 573.506 | 525.781 |
| $R^2$ | 0.706 | 0.706 | 0.707 | 0.705 |

Notes: Second-stage results are in Table A6. See the data section for a description of control variables. Standard errors in parentheses and clustered on fixed effects. $^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$.

## Tab. A8. Short differences and income model

| | (1)<br>1991-2001 | (2)<br>1991-2001 | (3)<br>1991-2001 | (4)<br>1991-2001 | (5)<br>2001-2011 | (6)<br>2001-2011 | (7)<br>2001-2011 | (8)<br>2001-2011 | (9)<br>2001-2011 | (10)<br>2001-2011 | (11)<br>2001-2011 | (12)<br>2001-2011 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | Δ log designation share (t) | | | | | | |
| *Δ* log degree share (t) | -0.017** | -0.216*** | -0.066*** | -0.483*** | 0.477*** | 1.653*** | -0.010 | -2.129** | | | | |
| | (0.009) | (0.021) | (0.014) | (0.079) | (0.052) | (0.126) | (0.080) | (0.919) | | | | |
| log degree share (t-1) | | | -0.056*** | -0.185*** | | | -0.117*** | -0.535*** | | | | |
| | | | (0.007) | (0.024) | | | (0.027) | (0.182) | | | | |
| log designation share (t-1) | | | -0.003*** | 0.004*** | | | -0.012*** | -0.009*** | | | -0.014*** | 0.004 |
| | | | (0.001) | (0.001) | | | (0.003) | (0.003) | | | (0.002) | (0.007) |
| *Δ* log homeowner-ship (t) | | | -0.056*** | 0.386*** | | | -0.117*** | 0.732* | | | -0.027 | 1.194*** |
| | | | (0.007) | (0.056) | | | (0.027) | (0.385) | | | (0.116) | (0.434) |
| log homeowner-ship (t-1) | | | 0.129*** | 0.077*** | | | -0.122 | 0.340*** | | | 0.098** | 0.777*** |
| | | | (0.028) | (0.014) | | | (0.115) | (0.127) | | | (0.042) | (0.237) |
| *Δ* log average household size (t) | | | 0.068*** | -0.245*** | | | 0.057 | -0.727 | | | 0.190 | 0.074 |
| | | | (0.013) | (0.062) | | | (0.037) | (0.450) | | | (0.181) | (0.272) |
| log average house-hold size (t-1) | | | 0.004 | -0.162*** | | | 0.219 | -0.099 | | | 0.278*** | 0.129 |
| | | | (0.037) | (0.049) | | | (0.185) | (0.177) | | | (0.095) | (0.149) |
| log pop age (t-1) | | | -0.027 | -0.158*** | | | 0.241** | 0.041 | | | 0.285** | -1.364** |
| | | | (0.037) | (0.036) | | | (0.095) | (0.185) | | | (0.112) | (0.559) |
| Δ pop age (t) | | | -0.109*** | -0.188*** | | | 0.389*** | -0.107 | | | 0.519** | -2.009** |
| | | | (0.033) | (0.056) | | | (0.112) | (0.362) | | | (0.217) | (0.899) |
| log foreigner share (t-1) | | | -0.044 | 0.057*** | | | 0.557*** | -0.004 | | | -0.025* | 0.101** |
| | | | (0.048) | (0.011) | | | (0.211) | (0.016) | | | (0.015) | (0.046) |
| Δ foreigner share (t) | | | 0.001 | 0.121*** | | | -0.017 | -0.001 | | | -0.026 | -0.104** |
| | | | (0.004) | (0.025) | | | (0.014) | (0.038) | | | (0.028) | (0.048) |
| *Δ* log income | | | | | | | | | -0.218*** | -9.330*** | -0.142** | -7.305*** |
| | | | | | | | | | (0.069) | (2.024) | (0.070) | (2.364) |
| log income (t-1) | | | | | | | | | | | -0.144*** | -0.909*** |
| | | | | | | | | | | | (0.037) | (0.261) |
| Constant | 0.159*** | 0.224*** | 0.489*** | 0.864*** | 0.317*** | -0.126*** | -1.436*** | 0.367 | 0.549*** | 2.881*** | 0.007 | 13.647*** |
| | (0.005) | (0.009) | (0.143) | (0.167) | (0.022) | (0.043) | (0.472) | (0.900) | (0.027) | (0.524) | (0.556) | (4.552) |
| *IV* | - | YES | - | YES | - | YES | - | YES | - | YES | - | YES |
| *Observations* | 7965 | 7965 | 7965 | 7965 | 7966 | 7966 | 7966 | 7966 | 7966 | 7966 | 7966 | 7966 |
| *CHI2* | | 103.847 | | 202.519 | | 170.741 | | 203.917 | | 21.242 | | 88.061 |
| *EXOG_P* | | 0.000 | | 0.000 | | 0.000 | | 0.012 | | 0.000 | | 0.000 |
| *OVERID* | | 7.555 | | 1.413 | | 1.385 | | 19.198 | | 13.526 | | 0.741 |
| *OVERIDP* | | 0.006 | | 0.235 | | 0.239 | | 0.000 | | 0.000 | | 0.389 |

Notes: See the data section for a description of control variables. First-stage results are in Table A9. Standard errors in parentheses. * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.

**Tab. A9. Short differences and income model – First-stage regressions**

| | (1) 1991-2001 *Δ* log degree share (t) | (2) 1991-2001 *Δ* log degree share (t) | (3) 1991-2001 log designation share (t-1) | (4) 2001-2011 *Δ* log degree share (t) | (5) 2001-2011 *Δ* log degree share (t) | (6) 2001-2011 log designation share (t-1) | (7) 2001-2011 *Δ* log income (t) | (8) 2001-2011 *Δ* log income (t) | (9) 2001-2011 log designation share (t-1) |
|---|---|---|---|---|---|---|---|---|---|
| rail station density | 0.055 | 0.053$^{*}$ | -0.003 | -0.062$^{***}$ | 0.038$^{***}$ | 0.059 | -0.012 | 0.018 | 0.066 |
| | (0.049) | (0.021) | (0.208) | (0.010) | (0.008) | (0.151) | (0.037) | (0.029) | (0.159) |
| employment potentiality | 0.000$^{***}$ | 0.000$^{***}$ | 0.000 | -0.000$^{***}$ | -0.000 | 0.000 | 0.000$^{*}$ | 0.000 | 0.000 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| log degree share (t-1) | | 0.055 | 0.053$^{*}$ | -0.003 | -0.062$^{***}$ | 0.038$^{***}$ | | | |
| | | (0.049) | (0.021) | (0.208) | (0.010) | (0.008) | | | |
| log designation share (t-2) | | 0.007$^{***}$ | 0.828$^{***}$ | | 0.000 | 0.922$^{***}$ | | 0.002$^{**}$ | 0.927$^{***}$ |
| | | (0.001) | (0.019) | | (0.000) | (0.010) | | (0.001) | (0.009) |
| *Δ* log homeownership (t) | | 0.586$^{***}$ | -0.613$^{**}$ | | 0.408$^{***}$ | -1.137$^{***}$ | | 0.172 | -1.232$^{***}$ |
| | | (0.067) | (0.232) | | (0.048) | (0.328) | | (0.101) | (0.343) |
| log homeownership (t-1) | | 0.061$^{**}$ | -0.431$^{***}$ | | 0.143$^{***}$ | -0.114 | | 0.110$^{***}$ | -0.141 |
| | | (0.022) | (0.118) | | (0.018) | (0.102) | | (0.028) | (0.126) |
| *Δ* log average household size (t) | | -0.534$^{***}$ | -0.161 | | -0.424$^{***}$ | 0.664 | | -0.009 | 0.733 |
| | | (0.044) | (0.325) | | (0.093) | (0.725) | | (0.089) | (0.733) |
| log average household size (t-1) | | -0.253$^{***}$ | -1.519$^{***}$ | | -0.139$^{*}$ | 0.273 | | -0.004 | 0.227 |
| | | (0.041) | (0.436) | | (0.059) | (0.258) | | (0.067) | (0.249) |
| log pop age (t-1) | | 0.004 | 0.555 | | -0.154$^{***}$ | 0.744$^{*}$ | | -0.217$^{***}$ | 0.896$^{*}$ |
| | | (0.048) | (0.337) | | (0.045) | (0.350) | | (0.058) | (0.345) |
| Δ pop age (t) | | -0.231$^{***}$ | 0.311 | | -0.325$^{***}$ | 0.152 | | -0.362$^{**}$ | 0.245 |
| | | (0.051) | (0.370) | | (0.077) | (0.455) | | (0.118) | (0.443) |
| log foreigner share (t-1) | | 0.110$^{***}$ | 0.085$^{*}$ | | 0.004 | -0.064 | | 0.015$^{**}$ | -0.035 |
| | | (0.009) | (0.042) | | (0.005) | (0.045) | | (0.005) | (0.046) |
| Δ foreigner share (t) | | 0.267$^{***}$ | 0.061 | | 0.026$^{*}$ | 0.023 | | -0.010 | -0.021 |
| | | (0.017) | (0.084) | | (0.012) | (0.065) | | (0.007) | (0.066) |
| Log income (t-1) | | | | | | | | -0.114$^{***}$ | 0.191 |
| | | | | | | | | (0.020) | (0.101) |
| Constant | 0.297$^{***}$ | 0.278 | -0.336 | 0.389$^{***}$ | 0.790$^{***}$ | -3.160$^{*}$ | 0.255$^{***}$ | 1.880$^{***}$ | -5.076$^{**}$ |
| | (0.008) | (0.209) | (1.438) | (0.005) | (0.221) | (1.479) | (0.004) | (0.239) | (1.602) |
| *Controls* | - | YES | YES | - | YES | YES | - | YES | YES |
| *Observations* | 7965 | 7965 | 7965 | 7966 | 7966 | 7966 | 7966 | 7966 | 7966 |
| *F* | 134.968 | 557.956 | 1891.124 | 73.689 | 464.362 | 3091.590 | 8.301 | 17.028 | 2640.502 |
| $R^2$ | 0.124 | 0.590 | 0.717 | 0.095 | 0.614 | 0.856 | 0.004 | 0.103 | 0.856 |

Notes: Second-stage results in Table A8 (instrumented models) See the data section for a description of control variables. Standard errors in parentheses and clustered on fixed effects. $^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$.

## 5.2 Equilibrium designation

*Baseline DD results: Extended results*

Table A10 below reports the CA effects as well as a set of estimated hedonic implicit prices (housing characteristics in particular) for the difference-in-differences estimation given by equation (13) in the main paper. In keeping with intuition, housing units with more bathrooms and bedrooms fetch higher prices, as do detached, semi-detached and bungalows (over the omitted category flats/maisonettes). The sales price of terraced housing is significantly larger than sales prices of otherwise comparable flats/maisonettes in some, but not all, models. Larger floor spaces are associated with higher prices but with diminishing effects. There is a premium for new properties. Leased properties are of less value than those owned. Properties with parking spaces, single garages and double garages sell for higher prices than those without any parking facilities. There is a house price premium for properties with central heating over other types of heating. In order to control for a potentially non-linear relationship between housing age and house prices we included a series of house age bins. In order to separate the effects of pure building age (which may be associated with deterioration) from the build date (which may strongly determine the architectural style) we allow for age cohort and building data cohort effects. Since the 'New property' variable identifies all properties where the build age is zero years, the omitted category from the age variables is 1–9 years. All of the bins for properties older than this indicate significant negative premiums. The negative premium increases with age, mostly quickly over the first few categories and then more slowly until the penultimate category and finally decreases for buildings over 100 years. The effect of the build date is also non-linear. The general tendency is for buildings built in earlier periods to have higher prices than buildings built in the omitted period 2000–2010. However, this effect becomes insignificant in the 60s and 70s; periods associated with the architectural styles of the post-war reconstruction phase that are today less appreciated than other styles. The greatest premium is attached to houses built pre-1900, the earliest category.

**Tab. A10. Conservation area premium – designation effect**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| | Log property transaction price | | | | | | |
| Inside treated CA × Post designation | 0.028*** | 0.014 | 0.014 | 0.003 | -0.024 | -0.077 | -0.003 |
| | (0.009) | (0.009) | (0.010) | (0.012) | (0.070) | (0.111) | (0.013) |
| Within 500m buffer of treated CA × Post des. | 0.023*** | 0.013*** | 0.012*** | 0.004 | 0.012 | -0.005 | -0.005 |
| | (0.004) | (0.004) | (0.005) | (0.006) | (0.027) | (0.022) | (0.010) |
| Inside treated CA | -0.043*** | -0.038*** | -0.048*** | -0.037*** | -0.062 | 0.029 | -0.024 |
| | (0.009) | (0.009) | (0.010) | (0.012) | (0.057) | (0.108) | (0.021) |
| Within 500m buffer of treated CA | -0.010** | -0.004 | -0.011** | 0.005 | 0.003 | 0.006 | -0.002 |
| | (0.004) | (0.004) | (0.005) | (0.005) | (0.030) | (0.023) | (0.013) |
| Number of bathrooms | 0.007*** | 0.007*** | 0.006*** | 0.013*** | 0.057*** | 0.059*** | 0.014*** |
| | (0.000) | (0.001) | (0.001) | (0.002) | (0.008) | (0.006) | (0.002) |
| Number of bedrooms | 0.166*** | 0.172*** | 0.169*** | 0.165*** | 0.170*** | 0.179*** | 0.158*** |
| | (0.002) | (0.004) | (0.005) | (0.005) | (0.014) | (0.011) | (0.006) |
| Number of bedrooms squared | -0.019*** | -0.020*** | -0.020*** | -0.019*** | -0.019*** | -0.019*** | -0.018*** |
| | (0.000) | (0.001) | (0.001) | (0.001) | (0.002) | (0.002) | (0.001) |
| Detached house | 0.254*** | 0.222*** | 0.211*** | 0.194*** | 0.235*** | 0.216*** | 0.193*** |
| | (0.003) | (0.005) | (0.008) | (0.007) | (0.015) | (0.014) | (0.007) |
| Semi-detached house | 0.119*** | 0.097*** | 0.088*** | 0.070*** | 0.082*** | 0.066*** | 0.073*** |
| | (0.003) | (0.004) | (0.007) | (0.006) | (0.014) | (0.012) | (0.006) |
| Terraced house/country cottage | 0.040*** | 0.026*** | 0.015** | 0.001 | 0.002 | -0.013 | -0.000 |
| | (0.003) | (0.004) | (0.006) | (0.006) | (0.013) | (0.012) | (0.006) |
| Bungalow | 0.311*** | 0.285*** | 0.281*** | 0.257*** | 0.292*** | 0.269*** | 0.257*** |
| | (0.003) | (0.006) | (0.008) | (0.009) | (0.019) | (0.016) | (0.009) |
| Floorsize ($m^2$) | 0.006*** | 0.006*** | 0.007*** | 0.007*** | 0.008*** | 0.007*** | 0.007*** |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| Floorsize squared | -0.000*** | -0.000*** | -0.000*** | -0.000*** | -0.000*** | -0.000*** | -0.000*** |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) |
| New property | 0.084*** | 0.087*** | 0.088*** | 0.088*** | 0.047** | 0.076*** | 0.077*** |
| | (0.002) | (0.004) | (0.005) | (0.006) | (0.024) | (0.017) | (0.006) |
| Leasehold | -0.054*** | -0.067*** | -0.065*** | -0.073*** | -0.100*** | -0.104*** | -0.070*** |
| | (0.003) | (0.004) | (0.006) | (0.006) | (0.014) | (0.012) | (0.006) |
| Single garage | 0.112*** | 0.097*** | 0.100*** | 0.097*** | 0.096*** | 0.097*** | 0.098*** |
| | (0.001) | (0.002) | (0.003) | (0.003) | (0.007) | (0.005) | (0.003) |
| Double garage | 0.190*** | 0.162*** | 0.161*** | 0.159*** | 0.160*** | 0.156*** | 0.158*** |
| | (0.002) | (0.003) | (0.005) | (0.005) | (0.015) | (0.010) | (0.005) |
| Parking space | 0.076*** | 0.063*** | 0.065*** | 0.061*** | 0.052*** | 0.049*** | 0.063*** |
| | (0.001) | (0.002) | (0.003) | (0.003) | (0.007) | (0.005) | (0.003) |
| Central heating | 0.089*** | 0.094*** | 0.098*** | 0.100*** | 0.085*** | 0.094*** | 0.095*** |
| | (0.001) | (0.002) | (0.003) | (0.003) | (0.007) | (0.007) | (0.003) |
| Building age: 10–19 years | -0.047*** | -0.063*** | -0.062*** | -0.075*** | -0.071*** | -0.068*** | -0.069*** |
| | (0.002) | (0.003) | (0.004) | (0.005) | (0.016) | (0.015) | (0.005) |
| Building age: 20–29 years | -0.079*** | -0.106*** | -0.104*** | -0.125*** | -0.133*** | -0.126*** | -0.113*** |
| | (0.002) | (0.005) | (0.007) | (0.008) | (0.026) | (0.021) | (0.007) |
| Building age: 30–39 years | -0.092*** | -0.127*** | -0.123*** | -0.150*** | -0.169*** | -0.141*** | -0.133*** |
| | (0.003) | (0.006) | (0.010) | (0.011) | (0.032) | (0.027) | (0.009) |
| Building age: 40–49 years | -0.104*** | -0.148*** | -0.142*** | -0.180*** | -0.199*** | -0.165*** | -0.158*** |
| | (0.004) | (0.008) | (0.012) | (0.013) | (0.036) | (0.031) | (0.011) |
| Building age: 50–59 years | -0.121*** | -0.171*** | -0.167*** | -0.207*** | -0.232*** | -0.204*** | -0.175*** |
| | (0.004) | (0.009) | (0.015) | (0.016) | (0.044) | (0.038) | (0.014) |
| Building age: 60–69 years | -0.135*** | -0.198*** | -0.194*** | -0.238*** | -0.320*** | -0.265*** | -0.215*** |
| | (0.005) | (0.011) | (0.019) | (0.020) | (0.051) | (0.042) | (0.018) |
| Building age: 70–79 years | -0.136*** | -0.213*** | -0.207*** | -0.263*** | -0.326*** | -0.273*** | -0.234*** |
| | (0.006) | (0.013) | (0.021) | (0.022) | (0.053) | (0.046) | (0.019) |
| Building age: 80–89 years | -0.132*** | -0.218*** | -0.213*** | -0.277*** | -0.339*** | -0.313*** | -0.243*** |
| | (0.007) | (0.014) | (0.023) | (0.024) | (0.062) | (0.054) | (0.021) |
| Building age: 90–99 years | -0.111*** | -0.208*** | -0.204*** | -0.280*** | -0.360*** | -0.304*** | -0.248*** |
| | (0.008) | (0.016) | (0.025) | (0.027) | (0.068) | (0.063) | (0.023) |

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) |
|---|---|---|---|---|---|---|---|
| | Log property transaction price | | | | | | |
| Building age: Over 100 years | -0.083*** | -0.176*** | -0.176*** | -0.261*** | -0.348*** | -0.284*** | -0.227*** |
| | (0.009) | (0.017) | (0.027) | (0.030) | (0.074) | (0.065) | (0.025) |
| Build date: 1900–1909 | 0.040*** | 0.121*** | 0.128*** | 0.208*** | 0.256*** | 0.222*** | 0.173*** |
| | (0.009) | (0.018) | (0.028) | (0.031) | (0.077) | (0.067) | (0.025) |
| Build date: 1910–1919 | 0.074*** | 0.153*** | 0.158*** | 0.226*** | 0.262*** | 0.256*** | 0.196*** |
| | (0.008) | (0.016) | (0.027) | (0.028) | (0.071) | (0.059) | (0.024) |
| Build date: 1920–1929 | 0.093*** | 0.157*** | 0.162*** | 0.215*** | 0.225*** | 0.189*** | 0.190*** |
| | (0.007) | (0.014) | (0.024) | (0.025) | (0.062) | (0.050) | (0.021) |
| Build date: 1930–1939 | 0.082*** | 0.128*** | 0.130*** | 0.168*** | 0.187*** | 0.163*** | 0.151*** |
| | (0.006) | (0.013) | (0.021) | (0.023) | (0.058) | (0.045) | (0.020) |
| Build date: 1940–1949 | 0.040*** | 0.078*** | 0.078*** | 0.111*** | 0.063 | 0.053 | 0.096*** |
| | (0.005) | (0.012) | (0.018) | (0.021) | (0.058) | (0.048) | (0.018) |
| Build date: 1950–1959 | 0.017*** | 0.033*** | 0.041*** | 0.057*** | 0.017 | -0.004 | 0.046*** |
| | (0.004) | (0.010) | (0.016) | (0.018) | (0.047) | (0.039) | (0.015) |
| Build date: 1960–1969 | 0.001 | 0.007 | 0.018 | 0.023 | -0.017 | -0.012 | 0.011 |
| | (0.004) | (0.009) | (0.013) | (0.015) | (0.044) | (0.037) | (0.013) |
| Build date: 1970–1979 | -0.015*** | -0.016** | -0.008 | -0.004 | -0.059 | -0.046 | -0.011 |
| | (0.003) | (0.007) | (0.011) | (0.012) | (0.042) | (0.033) | (0.011) |
| Build date: 1980–1989 | 0.013*** | 0.017*** | 0.025*** | 0.029*** | -0.023 | -0.010 | 0.024*** |
| | (0.003) | (0.006) | (0.008) | (0.010) | (0.038) | (0.029) | (0.008) |
| Build date: 1990–1999 | 0.022*** | 0.020*** | 0.022*** | 0.029*** | -0.020 | -0.008 | 0.017** |
| | (0.002) | (0.005) | (0.006) | (0.008) | (0.034) | (0.025) | (0.008) |
| Build date: pre 1900 | 0.098*** | 0.149*** | 0.162*** | 0.244*** | 0.312*** | 0.259*** | 0.216*** |
| | (0.009) | (0.018) | (0.029) | (0.031) | (0.081) | (0.070) | (0.026) |
| Location cont. | YES | YES | YES | YES | YES | YES | YES |
| Neighbourhood cont. | YES | YES | YES | YES | YES | YES | YES |
| Year effects | YES | YES | YES | YES | YES | YES | YES |
| Ward effects | YES | YES | - | - | - | - | - |
| Nearest treated CA effects | - | - | YES | YES | YES | YES | - |
| Matched CA effects | - | - | - | - | - | - | YES |
| Treatment group: CAs designated | 1996-2010 | 1996-2010 | 1996-2010 | 1996-2010 | 1996-2002 | 1996-2002 | 1996-2010 |
| Control group | Full England sample | Within 2km of treated CA | Within 2km of treated CA | Within 500m of pre-1996 CA & within 2km of treated CA | Within 500m of CA designated 1987-1995 & within 4km of treated CA | Within 500m of CA designated 2003-2010 & within 4km of treated CA | Within 500m of pre-1996 CA matched on propensity score |
| $R^2$ | 0.921 | 0.922 | 0.915 | 0.915 | 0.861 | 0.864 | 0.909 |
| AIC | -587375 | -156426 | -130469 | -67044 | -5410 | -8475 | -41206 |
| Observation | 1088k | 302k | 302k | 178k | 214k | 323k | 133k |

Notes: Extended presentation of Table 2 in the main paper. Standard errors in parentheses are clustered on location fixed effects. Conservation area control groups in columns (4)–(7) have separate fixed effects for the areas inside and outside a CA. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

*Time trend DD and distance gradient DD: Parametric estimates*

Tables A11 and A12 present the parametric estimates of specifications (14) and (15), which are discussed but not reported in section 5.2 of the main paper to save space.

**Tab. A11. Time trend DD effects**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | log property transaction price | | | | | | | | | |
| Inside treated CA × Post designation | 0.015 | 0.022 | 0.024 | 0.027 | -0.006 | 0.023 | 0.033 | 0.038* | 0.036 | 0.020 |
| | (0.015) | (0.015) | (0.015) | (0.017) | (0.018) | (0.023) | (0.021) | (0.023) | (0.024) | (0.024) |
| Within 500m buffer of treated CA | 0.006 | 0.013* | 0.015** | 0.020** | -0.007 | 0.013 | 0.017** | 0.022** | 0.017 | 0.009 |
| × Post designation | (0.007) | (0.007) | (0.007) | (0.008) | (0.012) | (0.008) | (0.008) | (0.009) | (0.010) | (0.014) |
| Inside treated CA × Years designated | 0.000 | -0.004 | -0.004 | -0.007** | -0.002 | -0.010 | -0.016* | -0.019* | -0.019* | -0.020* |
| | (0.003) | (0.003) | (0.003) | (0.003) | (0.003) | (0.010) | (0.009) | (0.010) | (0.010) | (0.011) |
| Inside treated CA × Years designated$^2$ | | | | | | -0.001 | -0.001 | -0.001 | -0.001 | -0.002* |
| | | | | | | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) |
| Inside treated CA × Post designation | 0.003 | 0.007** | 0.008** | 0.009** | 0.008* | 0.020 | 0.026** | 0.032** | 0.031** | 0.031** |
| × Years designated | (0.003) | (0.003) | (0.003) | (0.004) | (0.004) | (0.014) | (0.012) | (0.013) | (0.013) | (0.014) |
| Inside treated CA × Post Designation | | | | | | 0.000 | 0.000 | 0.000 | 0.000 | 0.001 |
| × Years designated$^2$ | | | | | | (0.001) | (0.001) | (0.001) | (0.001) | (0.001) |
| Within 500m of treated CA | 0.002 | -0.002* | -0.002* | -0.004*** | -0.001 | -0.001 | -0.004 | -0.007* | -0.004 | -0.009 |
| × Years designated | (0.001) | (0.001) | (0.001) | (0.001) | (0.002) | (0.004) | (0.004) | (0.004) | (0.005) | (0.007) |
| Within 500m of treated CA | | | | | | -0.000 | -0.000 | -0.000 | 0.000 | -0.001 |
| × Years designated$^2$ | | | | | | (0.000) | (0.000) | (0.000) | (0.000) | (0.001) |
| Within 500m of treated CA | 0.001 | 0.004*** | 0.004*** | 0.005*** | 0.003 | 0.003 | 0.007 | 0.011** | 0.008 | 0.009 |
| × Post designation ×Years des. | (0.002) | (0.001) | (0.001) | (0.002) | (0.003) | (0.005) | (0.005) | (0.005) | (0.006) | (0.010) |
| Within 500m of treated CA | | | | | | 0.000 | 0.000 | 0.000 | -0.000 | 0.001 |
| × Post designation × Years des.$^2$ | | | | | | (0.000) | (0.000) | (0.000) | (0.000) | (0.001) |
| Hedonic controls | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Location controls | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Neighbourhood controls | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Year effects | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Ward effects | YES | YES | - | - | - | YES | YES | - | - | - |
| Nearest treated CA effects | - | - | YES | YES | - | - | - | YES | YES | - |
| Matched CA effects | - | - | - | - | YES | - | - | - | - | YES |
| Control group as in Table 2, column | (1) | (2) | (3) | (4) | (7) | (1) | (2) | (3) | (4) | (7) |
| $R^2$ | 0.920 | 0.921 | 0.912 | 0.914 | 0.907 | 0.920 | 0.921 | 0.912 | 0.914 | 0.907 |
| AIC | -547,688 | -147,818 | -120,160 | -64,425 | -39,321 | -548,078 | -147,839 | -120,191 | -64,467 | -39,329 |
| Observations | 995k | 277k | 277k | 164k | 123k | 995k | 277k | 277k | 164k | 123k |

Notes: Standard errors in parentheses are clustered on the location fixed effects. Conservation area control groups in columns (4)–(7) have separate fixed effects for the areas inside and outside a CA. Observations dropped if years designated falls outside of range -10 years:+10 years. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

**Tab. A12. Distance gradient DD effects**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | log property transaction price | | | | | | | | | |
| Within 1400m of treated CA × | 0.027*** | 0.014 | 0.012 | 0.008 | -0.003 | 0.026** | 0.014 | 0.012 | 0.008 | -0.005 |
| Post designation | (0.010) | (0.010) | (0.011) | (0.011) | (0.014) | (0.011) | (0.012) | (0.012) | (0.012) | (0.015) |
| Within 1400m of treated CA × | -0.057 | -0.032 | -0.030 | -0.029 | -0.070 | -0.096 | -0.046 | -0.040 | -0.040 | -0.118 |
| Distance to boundary x Post des. | (0.081) | (0.075) | (0.080) | (0.077) | (0.068) | (0.156) | (0.154) | (0.162) | (0.157) | (0.143) |
| Within 1400m of treated CA × | | | | | | -0.059 | -0.017 | -0.018 | -0.017 | -0.099 |
| Distance to boundary$^2$ × Post des. | | | | | | (0.132) | (0.131) | (0.140) | (0.136) | (0.130) |
| Outside treated CA × Post desig- | 0.004 | 0.005 | 0.005 | 0.004 | 0.010 | 0.009 | 0.009 | 0.008 | 0.009 | 0.016 |
| nation | (0.010) | (0.010) | (0.010) | (0.009) | (0.011) | (0.012) | (0.012) | (0.011) | (0.011) | (0.012) |
| Outside treated CA × Distance to | 0.039 | 0.016 | 0.013 | 0.011 | 0.046 | 0.064 | 0.014 | 0.013 | 0.004 | 0.080 |
| boundary × Post des. | (0.081) | (0.075) | (0.080) | (0.078) | (0.069) | (0.157) | (0.155) | (0.163) | (0.159) | (0.145) |
| Outside treated CA × Distance to | | | | | | 0.070 | 0.028 | 0.025 | 0.029 | 0.109 |
| boundary$^2$ × Post des. | | | | | | (0.133) | (0.132) | (0.140) | (0.136) | (0.130) |
| Hedonic controls | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Location controls | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Neighbourhood controls | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Year effects | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Ward effects | YES | YES | - | - | - | YES | YES | - | - | - |
| Nearest treated CA effects | - | - | YES | YES | - | - | - | YES | YES | - |
| Matched CA effects | - | - | - | - | YES | - | - | - | - | YES |
| Control group | Full England sample | Within 2km of treated CA | Within 2km of treated CA | Within 1.4km of CA designated before 1996 & within 2km of treated CA | Within 1.4km of pre-1996 CA matched on propensity score | Full England sample | Within 2km of treated CA | Within 2km of treated CA | Within 1.4km of CA designated before 1996 & within 2km of treated CA | Within 1.4km of pre-1996 CA matched on propensity score |
| $R^2$ | 0.921 | 0.922 | 0.915 | 0.914 | 0.905 | 0.921 | 0.922 | 0.915 | 0.914 | 0.921 |
| AIC | -587,538 | -156,448 | -130,478 | -118,076 | -101,076 | -587,533 | -156,444 | -130,478 | -118,074 | -587,538 |
| Observation | 1088k | 302k | 302k | 281k | 327k | 1088k | 302k | 302k | 281k | 327k |

Notes: Standard errors in parentheses are clustered on the location fixed effects. * $p < 0.10$, ** $p < 0.05$, *** $p < 0$

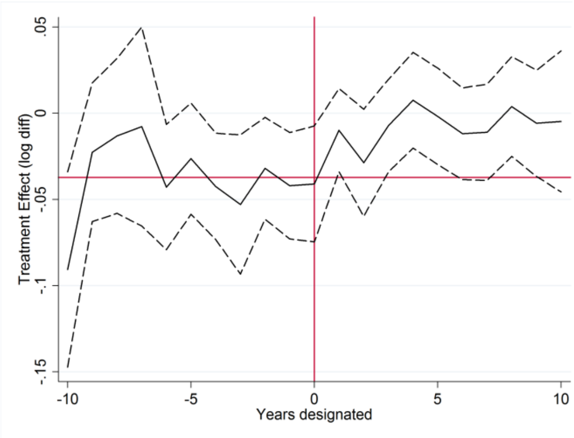
*Semi-parametric temporal and spatial treatment effects*

Figure A3 reports the results for the semi-parametric estimation of the temporal effects of designation using equation (A19). We compare the bin estimates for the naïve DD in the left panels to the matched CA control group in the right panels. The left charts show that the post-period internal and external estimates deviate significantly from the pre-period mean (hence the significant DD estimates) but that this is driven by a general upward trend. This corroborates the results in Table A11, column (1), where no significant discontinuity nor shift in trend for the naïve control group exists and hence the advantages of the time trend DD over the standard DD method is highlighted. The charts in the right panels also corroborate the evidence presented using the parametric trends equations in the main paper (Figure 1). Specifically, they show that for the internal effects the post-treatment estimates tend not to deviate significantly from the pre-treatment effects but that there are upward shifts in the trend when compared to the pre-treatment trend. For the external effects there is a general upward trend in the less carefully matched control groups and a downward trend in the stronger control groups but no shift in the trend at the designation date.

**Fig A3. Semi-parametric temporal bins estimates**

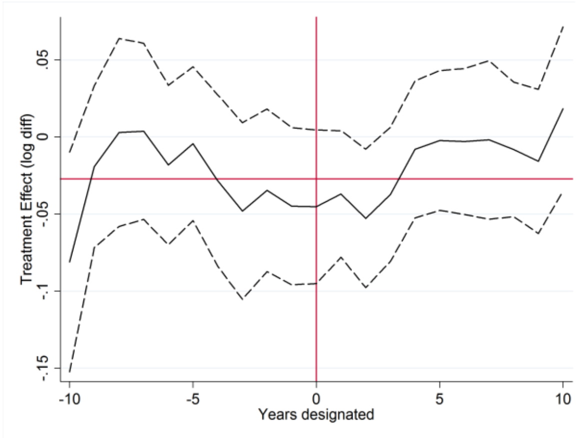
Internal effects: Full dataset

Control group as in Table A10 column (1)



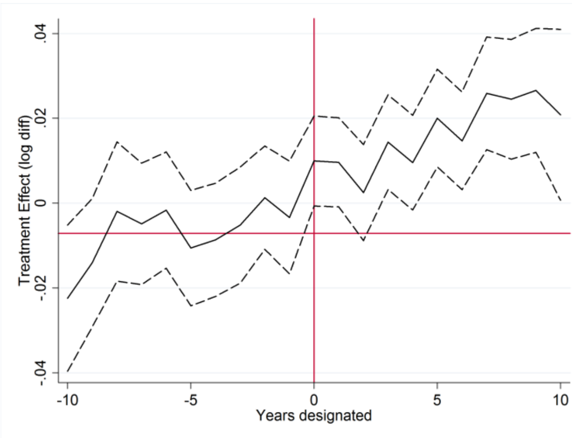

Internal effects: Matched CA

Control group as in Table A10 column (7)



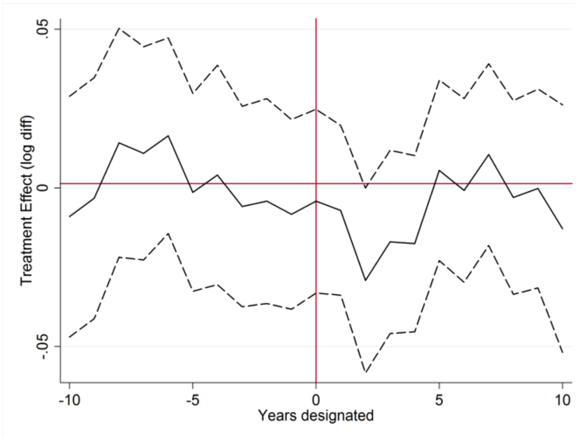

External effects: Full dataset

Control group as in Table A10 column (1)




External effects: Matched CA

Control group as in Table A10 column (7)

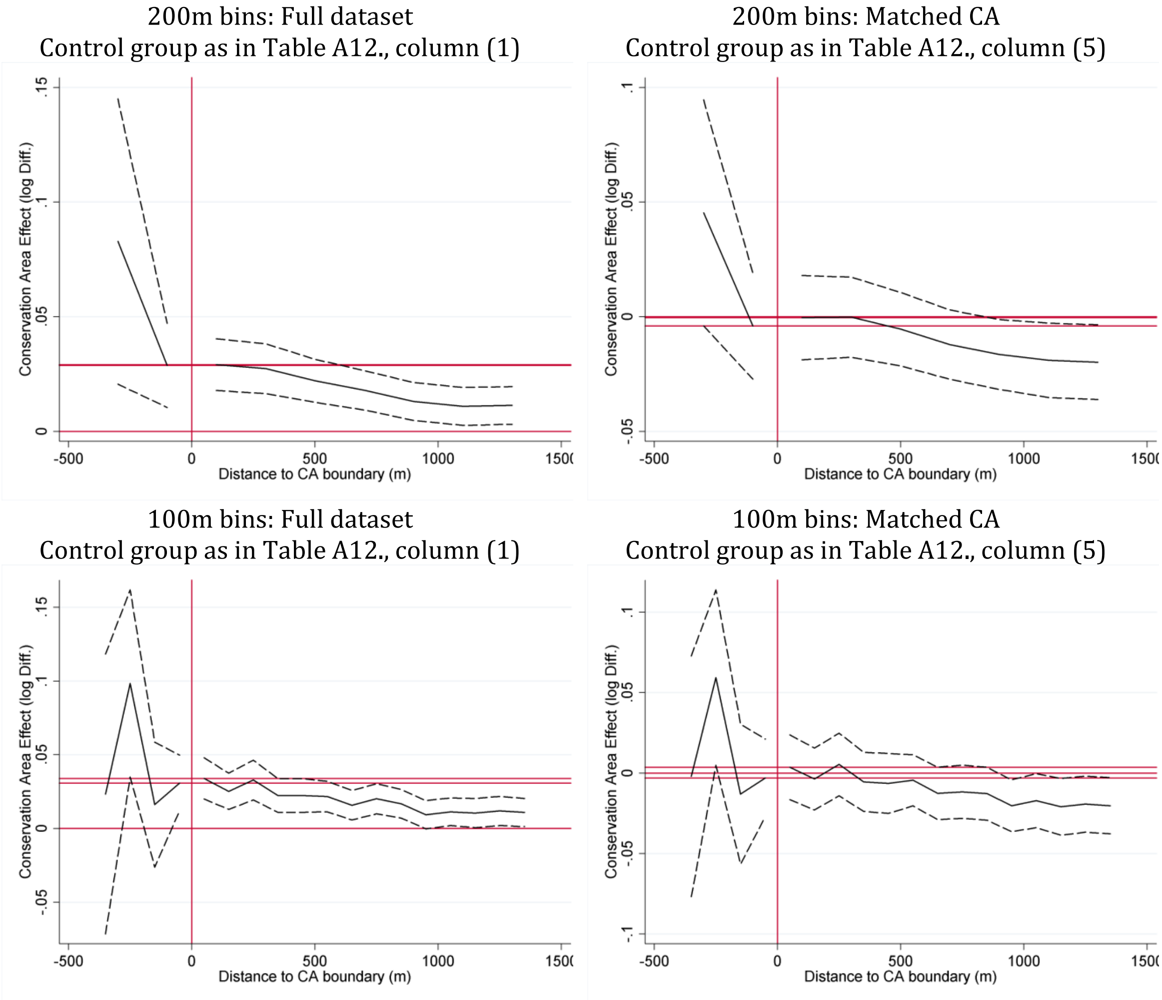



Notes: The solid black line plots the estimated differences between treatment group and control group against year since designation date using equation (A19). The dashed lines indicate the 5% confidence intervals. The left charts show results for the control group used in column (1) of appendix Table A10. The right charts show results for the control group used in column (7) of appendix Table A10. The horizontal red line illustrates the mean of the pre-treatment estimates.

Figure A4 demonstrates the semi-parametric distance gradient DD effects using different bin sizes of 100m and 200m using appendix equation (A20). These semi-parametric charts closely resemble their parametric counterparts. Notably, there is no significant and positive effect in the first bin outside the CA when using the preferred specification of column (7) from Table A10. This is consistent with the parametric findings and baseline DD findings that there is no significant external policy effect and that our second hypothesis cannot be accepted. There is, however, one significant bin inside the CA at 200–300m. This provides some support for the idea that heritage externalities are stronger deeper within the CAs such that there may be a positive policy effect. This effect then declines to zero for the deepest bin of greater than 300m.

In general, the large functional flexibility of the bin models comes at the cost of large standard errors since they are demanding in terms of degrees of freedom. While these estimates are informative with respect to the spatial pattern of heritage externalities, it is difficult to statistically affirm the existence of policy effects.

**Fig A4. Semi-parametric spatial bins estimates**




Notes: The solid black line plots the difference-in-differences treatment effect at different distances from the CA boundary using appendix equation (A20). The dashed lines indicate the 5% confidence intervals. The left charts show results for the control group used Table A12., column (1). The right charts show results for the control group used in Table A12., column (5). The horizontal red lines illustrate the mean of the pre-treatment estimates, the final pre-period bin and the first post-period bin.

# 6 Literature

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