Making – or Picking – Winners: Evidence of Internal and External Price Effects in Historic Preservation Policies

Douglas S. Noonan
Associate Professor
School of Public Policy
Georgia Institute of Technology
Atlanta, GA 30332-0345
Doug.Noonan@pubpolicy.gatech.edu

Douglas J. Krupka
Senior Research Associate
Institute for the Study of Labor (IZA)
Schaumberg-Lippe Straße
D-53113, Bonn, Germany
dkrupka@iza.org

Abstract:
Much has been written identifying property price effects of historic preservation policies. Little attention has been paid to the possible policy endogeneity in hedonic price models. This paper outlines a general case of land use regulation in the presence of externalities and then demonstrates the usefulness of the model in an instrumental-variables estimation of a hedonic price analysis – with an application to historic preservation in Chicago. The theoretical model casts doubt on previous results concerning price effects of preservation policies. The comparative statics identify some determinants of regulation that seem, on the face of it, most unlikely to also belong in a hedonic price equation. The analysis employs these determinants as instruments for endogenous regulatory treatment in a hedonic price analysis. OLS estimation of the hedonic offers results consistent with much of previous literature, namely that property values are higher for historic landmarks. In the 2SLS hedonic, robust estimates of the "own" price effect of historic designation are shown to be large and negative (approx. -27%) for homes in landmark districts. Further, significant and substantively important (positive) external price effects of landmark designations are found. The paper concludes with a discussion of the policy implications of these findings for historic preservation.

This document contains demographic information from Geolytics ™ East Brunswick New Jersey.

I. Introduction

The preservation of historic properties and districts is an increasingly popular form of land-market intervention, especially for local governments. Federally, the National Register of Historic Places (NRHP) comprises over 1.4 million buildings and objects on over 80,000 listed properties.[1] More locally, over 2,000 local historic district commissions exist in the U.S., many arising after the National Historic Preservation Act of 1966. Such policies are typically justified on the grounds of preserving beneficial external effect of historical properties. Other benefits often cited include stabilizing local land markets and conferring subsidies to historic property owners.

Attempts to quantify the costs of such policies have centered on the price effects of historical protection on the "protected" properties. A handful of studies have attempted to estimate spatial spillovers. These studies, however, have all been plagued by the endogeneity of the protection, where higher quality or exemplar historic properties (and districts) tend to receive special regulatory protection. The problem of endogenous policy treatment poses a challenge to those using hedonic price methods in other contexts as well. The sizable literature on the economics of historic preservation has generally found a positive price effect for officially designated historic buildings, often quite large, while lacking a substantive theory to explain why that might be case. Historic preservation advocates often cite these studies despite the absence of a formal model of the regulators in this context and the inability of these price effects to be interpreted causally. The present analysis addresses both of these shortcomings.

We first outline a general case of land use regulation in the presence of externalities Then, we demonstrate the usefulness of the model in an instrumental-variables estimation of a hedonic price analysis – with an application to historic preservation in Chicago. The model of

[1] From the National Register of Historic Places Official Website (http://www.nps.gov/nr/about.htm). For additional discussion, see Schuster (2002) and Swaim (2003).

restrictions on development, owing to concerns about externalities caused by existing developments, identifies theoretical conditions leading to the regulatory takings. The analysis employs these determinants as instruments for endogenous regulatory "treatment" in a hedonic price analysis. This approach demonstrates the importance of addressing endogenous policy variables in standard hedonic approaches. It also controls directly for historical quality by using the novel Chicago Historical Resources Survey, another innovation in the literature.

II. Background

Governments at the local, state, and federal level implement a variety of policy tools in the name of historic preservation. These policies typically target historic objects, buildings, or districts. These policies typically reflect multiple objectives like actual preservation of cultural resources, local economic development, and compensation of private preservation efforts. Officially designated landmarks and heritage resources can enjoy a form of "certification of quality" by government authorities, various forms of marketing by public agencies, subsidized capital and operating expenses, special public services such as public transit tours, and other advantages designed to promote tourism. Some preservation policies also involve government caretaking or owning of resources, in the form of public museums, monuments, or parks.

The imposition of binding restrictions on private property owners is largely a local affair. Federal regulations (e.g., National Historic Preservation Act of 1966, Antiquities Act of 1906) are either voluntary, subsidy-based, or primarily deployed in the context of federally owned property. Many state and local preservation policies give teeth where the federal rules have few. This paper considers the city of Chicago where, since 1968, the Commission on Chicago Landmarks has recommended landmark designations to the City Council. The stated purposes of the ordinance cover a wide range of topics like preservation, urban renewal, tourism, welfare, and character. By April 2008, 259 individual landmarks and 50 historic districts (comprising over 9,000 properties) had been designated. (42 individual landmarks, 7 districts, and roughly 4,500 included properties were added between 2005 and 2008). It is important to note that designation is not voluntary and property owners can appeal designation on hardship grounds. Once designated, alterations or construction affecting the landmark must be reviewed and approved by the Commission (part of the city's Department of Planning and Development). The City also offers several financial incentives for landmark property owners, depending on the type of landmark property. For instance, owner-occupied residences can receive waivers of building permit fees and a 12-year freeze on property taxes.

With such numerous objectives, evaluating the effectiveness of the policy poses a serious challenge. Establishing the counterfactual (i.e., what would happen in the absence of the policy) is complicated by the heterogeneous implementation and the possibility that policy treatments depend on outcome variables. For example, evaluating the effect of designation on a property's sale price becomes complicated when designations "follow the market" (Schaeffer and Millerick 1991) and merely reflect the additional value already recognized in the market, when designations promote urban renewal by targeting areas already in decline, or perhaps both at different times. Making causal inferences about the effects of preservation policies by merely observing how outcomes correlate with designation is thus quite problematic.

Insofar as preservation policies restrict property use, property values can be expected to decline and sales prices would reflect this. A negative marginal price in a hedonic analysis would reflect this kind of takings. On the other hand, access to financial assistance, such as subsidized loans or fee exemptions, should increase property values and hence prices. The effect of merely honorific designation on prices is less straightforward, possibly having no effect or perhaps conferring cachet or generating higher visibility to designated properties.

Other, more external impacts may be expected as well. If preservation policies or landmark designation bring more stability to a neighborhood by restricting changes in buildings or neighborhood character, property prices may rise because such stability mitigates investment risk for property owners. Landmark designation and historic preservation may also yield rather intangible benefits to the community – which might be reflected in property prices. Ahlfeldt and Maennig (forthcoming) argue that preservation policies conserve the "fabric of built heritage" and add value by producing variety. Asabere and Huffman (1994) suggest that historic preservation policies may solve a market failure in "providing a sense of unity with the past," a public good. Landmark designation may strengthen the "social fabric" of a community, enhancing local property values and prices (New York Landmarks Conservancy 1977). Schaeffer and Millerick (1991) noted how historic designation can indicate "public commitment" to the neighborhood. Numerous other observers cite landmark designation policies as "catalyzing" rehabilitation of neighboring communities (Listokin et al. 1998, Coulson and Leichenko 2001, Rypkema 1994).

Mason (2005) reviews a sizable literature estimating impacts of historic preservation policies – usually focusing on price impacts.[2] Many studies find evidence to support strong and positive impacts of these policies and fuel local efforts for still more. Recent studies have measured price premiums using a repeat-sales hedonics methodology (Coulson and Lahr 2005, Noonan 2007). The repeat-sales method suffers from possible sample selection bias and the possibility that appreciation rates affect the likelihood of designation. Unfortunately, designations are hardly exogenous to property values, especially when improving or protecting land values is ostensibly one of the goals of the policy intervention. Ideally, the empirical approach would correct the impact estimates for the bias resulting when the expected impacts

[2] Interestingly, the prominent impact studies by economists and preservationists alike focus on price impacts and economic development impacts – not on impacts on historic resources actually preserved.

affect the likelihood and type of policy intervention. This analysis uses a simple model to explain which areas receive the intervention and integrates several data sources to instrument for the endogenous policy variable in the hedonic price equation. This formal economic model of the determinants of historic preservation policymaking will be the first of its kind as far as we know. The first-stage estimates of determinants of designations offer valuable insights into the economics of local real estate regulation. Accordingly, the 2SLS estimator offers more robust evidence of causal (price) impacts historic preservation policies than most previous studies' methods and data permit.

III. Theoretical Model

In order to get consistent estimates of the policy effect of preservation policy, we attempt an instrumental variables estimation. In order to motivate this approach, we develop a simple model of the designation process. We imagine a historic preservation regulator maximizing his administrative utility function with respect to the restrictiveness of his preservation interventions, r, which we treat as continuous.[3] The regulator has direct preferences over r and the utilities of the property owner (P) and of other affected stakeholders or neighbors (N). Thus, the regulator will choose the restrictiveness of his intervention into the property market (his preservation policy) in order to maximize:

1) $$u(r;z) = U\left(P(r;z), N(r;z), r; z\right),$$

where U_N, $U_P > 0$ and $U_r < 0$. The regulator's utility is rising in the welfare of the (regulated) property owner and in the welfare of the neighbors. This is the case because he cares about high property values or because residents apply political and administrative pressure on the regulator

[3] Obviously, preservation status is discrete. Noonan and Krupka (2008) deal with this aspect of policy explicitly. Generally, the "restrictiveness" can be thought of as a continuous latent variable that includes the degree of curtailed transformation rights and stringency of enforcement that makes r appear continuous in practice even if designation status itself is dichotomous.

to increase their own utility. The regulator balances the property owner's and neighbors' opposing interests in restrictiveness. Property owners prefer to have fewer restrictions ($P_r < 0$) in order to preserve the option of redevelopment.[4] Neighbors favor more restrictions ($N_r > 0$), assuming those restrictions do not restrict *their own* options. Neighbors are expected to value restrictions on nearby properties because this reduces the risk of attractive properties being redeveloped in undesirable ways. The regulator's direct preferences over the restrictiveness arise from administrative costs of the program such as the costs of monitoring and enforcing compliance, as well as the administrator's internalization of the subsidies or tax breaks that local regulation stipulates be provided as compensation to owners of restricted properties. The direct preferences over the exogenous factors z will not be important to the model because they are not choice variables for the administrator.[5] The role of exogenous z will become important below.

The regulator optimizes by setting the marginal utility of restrictiveness to zero:

2) $$F(r,z) = U_P P_r + U_N N_r + U_r = 0 .$$

For interior solutions, equation (2) holds at the optimum and thus implicitly defines r^* as a function of z. We assume that the second-order condition (that $F_r < 0$) is satisfied so that equation (2) implicitly defines $r^*(z)$ as the utility-maximizing level of restrictiveness. The partial effects of any of these exogenous characteristics on the optimal level of restriction, r^*, are thus:

3) $$\frac{\partial r^*}{\partial z} = -\left(U_P P_{rz} + U_N N_{rz} + U_{rz}\right) / F_r$$

[4] The regulation might compensate restricted owners with subsidies or signal their property's quality to potential buyers. However, owners would prefer to have these advantages without the restrictions.

[5] These exogenous factors will include characteristics of individual properties, owners, or neighborhoods and will be discussed below. The preferences over z will vary depending on which exogenous factor is considered.

where $F_r = U_P P_{rr} + U_N N_{rr} + U_{rr} < 0$ is implied by the assumption that the second-order condition is satisfied.[6]

Equation (3) helps us understand how independent variation in a host of exogenous neighborhood or property characteristics will affect the optimal level of restrictiveness. In general, such a factor will increase the level of restrictiveness whenever $U_P P_{rz} + U_N N_{rz} + U_{rz} > 0$, which means that it increases the marginal utility of restrictiveness. The term is basically the sum of the effect of *z* on the marginal utilities of each stakeholder, weighted by the regulator's weight of each stakeholder's utility. For a variable *z* to be a valid instrument in a hedonic price model, it must have a non-zero effect in equation (3), but not affect property values.[7] This implies that $P_z = P_{rz} = 0$, while N_{rz} and U_{rz} do not both equal zero. That is, neighbors and regulators might "care" about the factor, but owners and potential buyers will not care.

The historic nature of the property will obviously not serve as such an instrument. While neighbors might like the external effects of the historic quality ($N_z > 0$), owners might either find the outdated structure onerous ($P_z < 0$) or quaint ($P_z > 0$). More restrictions on historic properties will increase neighbors' utility ($N_{rz} \geq 0$, because reduced uncertainty about the persistence of the positive externalities will benefit risk-averse neighbors) but could make living in a historic property worse ($P_{rz} < 0$, because modifications will be more difficult). From the above, it is apparent that historic quality serves as a poor instrument for preservation: it has an indeterminate effect on preservation, and likely has a direct effect on housing prices. Similarly, neighborhood demographic characteristics will likely affect both preservation decisions *and* housing prices in the neighborhood.

---

[6] A sufficient condition for this condition to hold are that administrative costs increase more-than-linearly, that the added benefit to neighbors is decreasing in restrictiveness, and that the costs imposed on property owners by restrictions increases about linearly.

[7] This condition holds even under other regulator objective functions, such as on that sets U_P very close or equal to zero as Glaeser (2006) suspects describes the Landmarks Preservation Commission in New York.

While neither historic quality nor neighborhood demographics will likely serve as valid instruments for historic designation, the interaction of them likely will. Neighborhood demographics that increase the benefits from designation *when historic externalities are present* will increase the likelihood that a historic property is formally designated. For instance, more owner-occupied units in the neighborhood (recipients of district tax breaks) will increase the likelihood that historic neighborhoods will be preserved, while having little effect on the probability that newer neighborhoods are so preserved. Conversely, having more owner-occupants in the area in 1970 will not affect the utility or disutility for owners *in the 1990s* of living near outdated properties. The identification assumption is thus that z is some vector of interactions between the historic quality of the property being sold or its environs and key neighborhood demographics circa 1970. Such interactions are assumed to have no effect on owner utility ($P_z = P_{rz} = 0$), but important effects on neighbor utility ($N_{rz} > 0$). The reasonableness of assuming these interactions (i.e., supply and demand "shifters" for designation) do not belong in a hedonic price equation is further supported by including the uninteracted historical quality, historic density, and various current neighborhood demographics in the hedonic and in using lagged neighborhood demographics measured decades before the property sale. Nonetheless, relevant diagnostic tests for instrument validity are still warranted.

A feature of historic preservation policy in Chicago and elsewhere is that protected properties can be bundled together into landmark districts. This is a detail away from which we have abstracted considerably. Landmark districts offer regulators another policy instrument for making preservation decisions. The decision to include a property in a landmark district, however, still follows a similar structure as that outlined above. Administrative costs associated with a building in a district, along with the attendant internal and external effects, are likely to differ from individual designation decisions. Despite these differences, the same factors, z, discussed above can serve as instruments for either type of designation. Due to data limitations, this analysis emphasizes the district designations.

IV. Data and Method

The following analysis combines data from many sources. First, property data come from actual sales data recorded in the Multiple Listing Service (MLS) for all single-family attached residential property sales in the city of Chicago during the 1990s. This type of property composed the bulk of property sales (and roughly 75% of the housing units) in the city at the end of our data period, and the MLS serves as the information clearinghouse for most arms-length housing sales in the city. The data include 71,275 attached home sales in Chicago from 1990 – 1999. MLS tracks many property attributes such as the address, numbers and types of rooms, and parking availability. For the square footage and year built variables, which are missing in many observations, several fixes are considered. Ultimately, for simplicity and because Noonan (2007) shows negligible gains for more sophisticated approaches, simple imputation based on the exogenous variables listed in Table 1 is employed.[8] List-wise deletion takes care of missing values for other variables in the regressions that follow. The final sample used is slightly less than 60,000.

Second, the City of Chicago's Landmarks Division in the Department of Planning and Development provide information on the landmarks (City of Chicago 2004). Information such as the addresses, dates of construction and designation, architect and architectural style, and historic themes are available for the individual landmarks and historic districts in the city. Added to the official landmarks data is the Chicago Historical Resources Survey (CHRS). Starting in 1983, historians from the Landmarks Commission inventoried the half million

[8] Missing values for "square footage" and "year built" are imputed using several auxiliary regressions following the same approach used in Noonan (2007).

properties in Chicago's city limits. The survey report describes the methodology in greater detail (Commission on Chicago Landmarks 1996). Ultimately, field work obtained detailed information from a final sample of 17,366 historically significant properties. The database of CHRS properties contains information on addresses, architects, significance and maintenance, and construction dates (http://www.cityofchicago.org/Landmarks/CHRS.html). The CHRS surveyors also assessed each CHRS property for its historic value or integrity. Very significant and well preserved properties were given codes of either red or orange. Sold properties thus fall into one of three categories depending on whether the sold property is a CHRS red or orange property (*RO*), is given lesser historical quality codes (*OTH*) but is still significant enough to be in the CHRS, or it was overlooked by the CHRS. Table 1 shows the overlap between the MLS attached-home sales in the CHRS database and the properties designated as landmarks. Obviously, Chicago's landmark designating process involves more than just historical quality, especially for district designation.

Third, the analysis also uses a variety of other geographic data for the city including Chicago's community areas and Census TIGER files. To link properties to their block-group level Census variables, the Geolytics™ dataset is employed to produce boundary-constant neighborhood demographics for 1990-2000. The Geolytics™ Neighborhood Change Database is used to get tract-level measures of demographics and housing in 1970 for use as instruments.

Table 1: Overlap between CHRS and Landmark Designations in Sales Data

		Is a landmark: landmark building	Is a landmark: landmark district	Not a landmark	Total
In CHRS	"red" or "orange"	176	548	1,685	2,409
In CHRS	not "red" or "orange"	1	358	903	1,261
	Not in CHRS	147	1,090	66,381	67,618
	Total	323	1,996	68,969	71,288

Table 2: Definitions and descriptive statistics

Variable	Definition	Mean	Std. Dev.
lnP	ln (real sales price, adjusted to 1 January 2000 $ using Chicago's housing CPI deflator)	11.898	0.659
DISTRICT	in a landmark district at time of sale?	0.036	0.186
LANDMARK	in a designated landmark building (not a district) at time of sale?	0.004	0.060
CountLmk	number of landmark buildings in block-group, exclusive of property's own status, at time of sale	0.328	0.748
DistShare	share of block-group's land area inside a landmark district, exclusive of property's own district, at time of sale	0.033	0.096
named	sale property is in building in CHRS that has a building name (e.g., "The Overton Building")	0.009	0.094
OTH	sale property is in CHRS not as "red" or "orange"	0.009	0.095
RO	sale property is coded "red" or "orange" in CHRS	0.044	0.205
CHRS100	count of CHRS properties within 100 meters	4.174	8.158
CHRS250	count of CHRS properties within 250 meters	27.079	38.886
CHRS500	count of CHRS properties within 500 meters	96.138	115.992
lnarea	ln (area of unit in $feet^2$) (see footnote 9)	7.103	0.417
yearbuilt	year built (see footnote 9)	1956.792	26.924
unitbldg	number of units in the building	150.405	231.928
rooms	number of rooms	4.743	1.690
bedrms	number of bedrooms	1.866	0.801
baths	number of baths	1.553	0.659
mbbth	master bathroom dummy	0.489	0.500
fireplace	number of fireplaces	0.313	0.515
garage	garage dummy	0.368	0.482
parking	garage or parking dummy	0.884	0.320
parkspot	parking spot dummy	0.172	0.377
waterfront	on the waterfront	0.072	0.258
distCBD	distance to CBD in km	6.953	5.023
distLake	distance to Lake Michigan in km	1.989	3.125
distwater	distance to closest water (river, lake) feature, in km	0.876	0.804
distCTA	distance to closest CTA rail line in km	0.737	0.695
distpark	distance to closest park in km	0.416	0.332
northside	on the north side of the city	0.916	0.277
latitude	decimal degrees north	41.929	0.050
income	median household income (in $1000s), block-group, 1990	33.083	21.964
college	percent with a college degree, block-group, 1990	0.476	0.224
medValue	median house value (in $1000s), block-group, 1990	219.804	168.199
popdens	population density (1000s/km2), block-group, 1990	33.923	23.893
white	percent white, block-group, 1990	0.721	0.234
medyrblt	median year built for residences, block-group, 1990	1953.018	13.458
new const.	percent of housing units built in last 10 years, block-group, 1990	0.094	0.150
saleyear	year of sale	1995.375	2.812
OwnOcc70	percent of housing units occupied by owner, tract, 1970	0.281	0.322
VacRat70	percent of housing units vacant, tract, 1970	0.087	0.059
LongRes70	percent of housing units with occupants moving in before 1960, tract, 1970	0.197	0.110
NewOne70	percent of housing units built in last 1 year, tract, 1970	0.043	0.066

The empirical model common to hedonic property analyses (Rosen 1974) takes the semi-log form:

4) $\ln P = \alpha + \beta X + \delta DESIGNATION + \varepsilon_{it}$

where X is a vector of property attributes and neighborhood quality measures, and *DESIGNATION* represents designation status, the discrete analogue to the index r from the theoretical model above. The problem of endogeneity arises if, for example, designation status tends to be conferred on properties with unexpectedly high (or low) sale prices. Given the explicit concern about impacts on designees' sale prices apparent in Chicago's landmark ordinance, the possibility of endogenous designation is worth exploring. As shown above, valid instruments will be those variables in z that predict r but do not belong in X. The first-stage of the two-stage least squares regression essentially estimates r^* using X and instruments z that are excluded from X in estimating the second-stage, equation (4). This provides consistent estimates of δ and permits a better understanding of why landmark designations occur where they do. Table 1 lists the variables used in estimating equation (4) along with their definitions and descriptive statistics.

Identifying price effects of landmark preservation policies in this setting poses noteworthy challenges and opportunities. First, due to the rarity of individual building designations (even in a landmarks program as vibrant as Chicago's), the data only permit robust estimations of district designations' price effects. The final sample has 284 sales in individually designated landmark buildings, 272 of which are in five different buildings. This makes it difficult to distinguish between the effect of building-specific unobservables and the individual designation effect. Fortunately, enough variation exists for district designations. Second, although previous research tends to focus on the effects of designation on a property's own price (Coulson and Leichenko 2001 and Ahlfeldt and Maennig 2009 are notable exceptions), the external effects of historic preservation – effects of designation on nearby property prices – remain arguably the most important effect of the policy. Even if the own-price effect of designation is negative, preservation policies can still be justified on positive externality grounds. Including measures of proximity to landmarks in *X* in equation (4) allows the recovery of amenity (or disamenity) effects of designation. This avoids the limited overlap between the home sales data and the landmarks inventory because the nearby landmarks include *all* landmarks (e.g., Douglas Tomb, St. Ignatius High School, Robie House), regardless of their housing a home sale during the 1990s. The external price effects of landmark building designation are thus recoverable, even if its own-price effects are not.

V. Results

A. OLS Results

First, the results of simple OLS regressions are presented. Table 3 reports the main results from these regressions across a number of specifications of the hedonic price equation. Column 1 of Table 3 presents the results from a fairly naïve regression of landmark district status and property characteristics on log real sales price. These results suggest large price premiums (approximately 25%) for homes in preserved districts.[9] (The premium rises to 60% in a totally unconditional model where only *DISTRICT* is included and all other hedonic attributes are omitted.) Although homes in districts sell for a premium over comparable homes, and homes in districts tend to have nicer attributes than non-district homes, districts are inherently spatial and thus geographic controls become vital. Column 2 adds geographic and local neighborhood demographic controls. As a result, the price premium for properties in landmark districts falls to roughly 2% (with a p-value of 0.06). The coefficients on the property characteristics are

[9] Following Halvorsen and Palmquist (1980), percent effects of dummy variables like designation status are derived from the expression: $\exp(\beta) - 1$.

unsurprising. Controls for geographic position of the structure relative to major features of the landscape and census block-group demographics also exhibit generally expected signs. Nicer homes are in districts, and districts are in nicer neighborhoods. In this sense, these historic preservation districts appear to be "picking winners." Failing to control for both housing attributes and geography can severely distort the effects of landmark districts. Controlling for them, the premium falls much closer to zero.

While the effect of designation on the price of the designated home is certainly important, landmark designation usually has other aims. In some senses, the primary rationale for preservation is that the threatened historic properties embody positive externalities overlooked by private interests when considering redevelopment. These historical spillovers should be captured at least partly by neighboring property values. Column 3 includes counts of (individually designated) landmarks in the home's block group and a measure of the share of the home's block group's land area that is occupied by another landmark district. Column 3 shows sizable effects of proximity to landmark districts and rapidly diminishing returns to nearby landmark buildings. A ten percent increase in the nearby land area within a district is associated with a 2.1% rise in property values. Having one landmark nearby is associated positively with price, but additional individual landmarks beyond that actually diminish property values. Properties sold inside of districts receive a roughly 2% premium after controlling for proximity to other nearby designations.

Even though the results in column 3 of Table 3 show a smaller price effect of districts than the prevailing literature on price effects of historic preservation efforts[10], one important

[10] Notably, Noonan's (2007) Table 3 shows positive price effects for Chicago districts, and Noonan (2007). Coulson and Lahr (2005) find 14 – 23% effects for Memphis districts. Coulson and Leichenko (2001) and Leichenko et al. (2001) find 5 – 20% price effects for individual designations in Texas. Coulson and Leichenko (2001), and Ahlfeldt and Maennig (2009) estimate (universally) positive price effects from proximity to other designated properties. The OLS model in Noonan (2007) for Chicago data finds a 9% price effect for *landmark* and a 5% effect for *district*.

source of bias in the estimator is the possibility that the landmark variables are correlated with unobserved historic quality. The landmarks commission may be selecting properties to designate based on characteristics not controlled for in column 3. As designation itself need not actually affect the historic quality – it seeks to prevent the decline of existing historic quality, not create new quality – the risk of "picking winners" rather than "making winners" is particularly high here. If historic quality has a positive or negative market valuation, and designated properties are more "historic" than non-designated structures, then part of the designation effect reflects a pre-existing historic value that was not created by designation, only correlated with it. The model in column 4 leverages the measures of historic quality available for the roughly 17,000 properties deemed sufficiently historically significant to include in the CHRS. Measures of the properties own historical quality (its color code ranking and whether it has a known name like "The Smith Building") and measures of the nearby density of historical properties (counts of CHRS properties within several buffers) are included in the model for column 4. As a result, the own-price effect of *DISTRICT* gets much stronger, the price effect of nearby districts actually grows more positive, and the external price effect of nearby landmark buildings remains virtually unchanged. The bias from omitting historic significance suggests that neighborhoods with high density of historic significance are less valued by the market and thus downwardly bias the district effect. Controlling for nearby historical properties does little to affect the price effects of proximity to individual landmarks but, like the bias for *DISTRICT*, removes a downward bias on the effect of nearby landmark districts. Interestingly, CHRS properties generally sell for a premium and that premium is slightly less for red- or orange-coded buildings. This likely arises from the CHRS surveyors downgrading color rankings for properties that have undergone renovations, which are presumably valuable to owners. Named buildings appear to sell for substantial premium. Being nearby CHRS properties has mixed effects on property values.

Homes sold with more CHRS properties very nearby are generally unaffected, those in the middle buffer range suffer a price penalty, and homes embedded in a larger (500m) community with more historic properties receive a small premium.

[TABLE 3 HERE]

The sensitivity of the implicit prices of landmark variables to controls for historical quality marks an important contribution to the previous work, which has generally failed to include objective measures of historical quality distinct from official designation (Noonan 2007, Coulson and Lahr 2005). In column 4, the inclusion of controls for a building's historic quality has a significant impact on the *DISTRICT* coefficient.[11] For landmark districts in Chicago, the omitted-variable bias appears most strong when geographic and neighborhood quality variables are absent. Neighborhood historic quality may predict which properties receive landmark status. There is evidence of external effects of a historic preservation policy, a major justification for the policies, although the nonlinearity complicates matters.

B. 2SLS Results

While the results in Table 3 are informative, the discussion in the theoretical section above suggests that taking the OLS coefficients as policy effects – even in the face of so many control variables – is not necessarily justified. Many unobserved factors that affect the sales price might also affect the probability that a property is designated. Chief among these would be a poorly observed historic quality or maintenance level in older properties. While Noonan (2007) and Coulson and Lahr (2005) attempt to deal with this problem via a repeat-sales or first-differenced model, the approach here uses instrumental variables. Table 4 presents estimates of

[11] If column 4 in Table 3 also included a *LANDMARK* variable, the price premium for properties in individually designated buildings is not significantly different from zero and other effects are essentially the same. The *LANDMARK* coefficient is also insignificant in a column 3 model.

a hedonic price equation similar to the model presented in column 4 of Table 3 estimated using a two-step efficient GMM IV estimator (with standard errors robust to arbitrary heteroskedasticity). As described above, the instruments are selected from the interaction of a neighborhood demand (for historic preservation) shifter and a measure of historic quality around the property being sold. The neighborhood demand variable is percent of housing that is owner-occupied measured at the tract-level in the 1970 census, which immediately followed the passage of the Landmarks Ordinance but predates sales by at least two decades. The historic quality variables are (a) *named*, an indicator for the property's historic significance, and (b) *CHRS100*, indicating the historic density around the home,, both measured by the CHRS that largely predates property sales. Conditional on the many controls for neighborhood quality and historic density, the 2SLS hedonic assumes that the interactions are valid exclusions and that they predict the likelihood of being designated in a landmark district. The (difference-in-Sargan) C statistic confirms that *DISTRICT* is endogenous in the OLS price equation at the 0.01% level.

[TABLE 4 HERE]

The first column of Table 4 presents results for the price equation for the two-stage GMM IV estimator of the preferred model. The other columns provide sensitivity tests and are discussed in Section V.C. In column 1, the instruments predict *DISTRICT* status well, with named homes or homes nearby more historic resources being more likely to have been designated in landmark districts if the share of owner-occupied housing in the tract in 1970 was higher. This is as expected, where higher owner-occupancy rates predict greater demand for district designations. Diagnostic tests for weak instruments, discussed in more detail below, all point to strong instruments. Most of the coefficients on the exogenous variables do not change very much from column 4 of Table 3 to column 1 of Table 4. Controlling for the endogeneity in *DISTRICT* does lead to larger premiums for CHRS homes (of any color code) and to smaller price effects for nearby landmark districts. Having more CHRS properties within 100 meters is now also an amenity. One interpretation is that unobservables correlated with designation in the OLS model, and purged in the 2SLS model, are also correlated with designees' historical quality and proximity to other landmarks.

The effects of district designation in this model differ markedly from the OLS estimates. Designations affect price in at least two ways: (internally) affecting designees' prices and (externally) affecting neighbors. First, there is no evidence of positive own-price effects from district designation. These estimates in column 1 of Table 4 suggest that the effect of being included in a historic district on sales price is negative 19% (with a 95% confidence interval ranging from -29% to -7%). The large district effect likely reflects several forces. On the one hand, inclusion in a historic district restricts redevelopment options of owners (and buyers), which should lower the value of the property. On the other hand, district designation may offer many benefits, like tax benefits and possibly a kind of certification of (or signal for) the property's cachet. For attached housing in Chicago, at least, the tax benefits are outweighed by the restrictions on renovation. The cachet effects also appear minimal given the model with excellent controls for historical quality. Buildings with names in the CHRS sell for a 6% premium on top of a 5-8% premium for being in the CHRS. Furthermore, the stability that district designation brings to the neighborhood's overall character (in terms of the types of land uses and buildings' external appearance) may be seen as disamenities by buyers in districts. Homes in districts may relatively lack access to modern urban conveniences like shopping, parking, and other mixed uses.

Second, external price effects of landmarks might affect those near to landmarks. Conditional on the historical nature of the neighborhood, important price effects from proximity to officially designated landmarks are shown in Table 4. There are clearly external price effects of more landmark buildings and districts in a home's neighborhood above and beyond the neighborhood's collection of historic buildings. Having a district (that does not contain the home) in the block group at the time of sale is associated with higher property values proportional to the size of that district. Every 10% of the block group occupied by a district predicts a 2% increase in sale price *for all homes in that block group*. This significant positive spillover likely arises for many of the same reasons outlined earlier. Being near a stable historic district with its intangible benefits but none of the restrictions is an important local amenity.[12] Moreover, insofar as district designation imposes constraints on the supply of new housing within the district, demand may be diverted to areas just outside of the district boundaries. Both improving nearby neighborhood quality and shifting demand to these areas can account for the positive *DistShare* effect in the 2SLS model.[13]

Having an individually designated landmark building (that does not contain the sold home) in the home's block group confers a modest (3%) price premium for the first landmark. That premium reflects a rapidly diminishing return to landmark buildings nearby such that homes with two or more landmarks in their block group actually sell at a sizeable discount from

---

[12] Families search for a certain kind of neighborhood might be attracted by landmark districts for their old-time neighborhood "feel" and stronger "social fabric" (New York Landmarks Conservancy 1977). Time-constrained families may restrict their housing search to nearby such neighborhoods, increasing the demand for homes just outside of districts. Other intangible benefits for the area have been remarked on by Schaeffer and Millerick (1991) and Asabere and Huffman (1994a).

[13] If being near districted properties increases prices for non-districted properties, then presumably it would also increase prices for districted ones as well. The coefficient on *DISTRICT*, consistently negative across models, suggests that any such positive external effect is outweighed by negative effects associated with being subject to district regulations, on average. This interpretation of the results in Table 4, critical for appreciating the relative magnitude of the negative internal and positive external price effects, suggests that district designations have weaker "amenity" affects than their redevelopment constraint effects.

otherwise comparable properties. The story about the external effects of historical quality and landmark preservations is thus somewhat mixed. Being in historically significant communities and near landmark districts suggests price premiums, while the price advantages of close proximity to individual landmark buildings become disadvantages when landmark density is high. The external benefits of individual landmarks likely depend on neighborhood context. Regardless, there is no clear evidence that designating individual landmarks carries universally positive spillovers, while being in or near landmark districts is associated with higher prices.

The external effects of historical quality and preservation policies are also evident in column 1 of Table 4. Having more historical properties in the larger community (between 250 and 500 meters) increases sale price, more historical properties in in the middle buffer distance (between 100 and 250 meters) decreases price, and more CHRS buildings in the closest buffer is not significantly related to price . This is all on top of the strong positive effect of median housing age of the observation's block group as part of the control variables.

C. Robustness

To assess the robustness of the 2SLS estimator in column 1 of Table 4, the validity of the exclusion restrictions and the strength of those instruments must be examined. Appropriate instruments must not belong in the price equation and also be closely correlated with the endogenous (*DISTRICT*) variable. The theory outlined in section III suggests that the interaction between historic quality measures and key 1970s demographics should predict designation yet not also belong in the price equation. Conditional on the many controls already in the hedonic specification, there is no reason to suspect these interaction terms are relevant hedonic attributes. The model in column 1 of Table 4 is overidentified (two instruments, one endogenous regressor), so that tests of the overidentifying restrictions – or of instrument exogeneity – are possible. Table 4 presents Hansen's *J* statistic. Although robust to arbitrary heteroskedasticity, Hansen's *J* may lack power against the null that the instruments are valid (i.e., uncorrelated with the error term in equation (4)).[14] The *J* statistic in all the models in Table 4 is not significant at conventional levels, consistent with exogeneous instruments.

The strength of the instruments in the first stage is supported by the diagnostic statistics shown in Table 4. Table 4 reports the first-stage coefficients for the IVs. Column 1's model exhibits very strong IVs as discussed above. Table 4 also reports the heteroskedasticity-robust Kleibergen-Paap statistic. This statistic addresses concerns that weak IVs may bias the second-stage estimates (Stock and Yogo 2005). The large values, 53.0 in column 1, exceed the double-digit "benchmark" often mentioned, thus indicating very strong instruments and should allay concerns about irrelevant IVs.

Another check on the robustness of the IV approach employs an alternative set of IVs. The second-stage results should remain mostly unchanged when a different valid set of IVs is used. Column 2 of Table 4 shows the results of such an estimation. There, the two instruments are (a) the interaction between the property's historic quality (the *named* variable, taken from the CHRS and hopefully minimally sensitive to any post-designation and pre-survey maintenance effects) and residential turnover (share of the households in the same home five years earlier) in the tract circa 1970, and (b) the interaction between historic density (measured by *CHRS100*) and new construction (share of the housing stock built in the last year) in the tract measured in 1970. The idea is that historically notable neighborhoods or buildings in areas with less turnover will be less likely to become designated in a district since the relatively stable neighborhood requires less protection. The first-stage results in column 2 bear out this expectation. The *J* statistic fails to reject the exogeneity assumption and the Kleibergen-Paap statistic again points to very strong instruments. The 23% price discount observed in Column 2 is quite close to the column 1 results

[14] The concern about over-rejecting the null in the presence of clustered errors (Hoxby and Paserman 1998) with a standard overidentification statistic lends further confidence to the *J* statistics reported in Table 4.

using different instruments, lending still more confidence to the results. Estimates for the other regressors are substantively similar as well. Armed with four potential instruments, the model in column 3 reports the GMM estimator with all four instruments. Again, the instruments collectively appear strong and exogenous, and the coefficients of interest change minimally.

Using the larger instrument set allows testing of another interesting hypothesis. A second endogenous regressor, the years elapsed since the property was initially designated in a district, is added to the model in column 3. Here, the IV diagnostics are a bit more suspect: the *J* statistic of 6.83 has a p-value of 0.03 casting doubt on the validity of the instruments, although the Kleibergen-Paap statistic of 49.45 suggests strong instruments. The price effect for *DISTRICT* is -0.11 (z = -1.89) and the time-elapsed interaction coefficient is -0.023 and is statistically significant (z = -4.01). These results point to the possibility that the district discount or "penalty" grows larger over time, perhaps as properties become further outdated. The implicit prices for the exogenous variables are qualitatively similar to those in column 3, except for the *DistShare* effect, which become essentially zero.

Finally, the possibility of building-specific unobservables is explored. One interpretation of the endogeneity evident in the OLS model (column 4, Table 3) is that unobservables remain correlated with district designation. While a repeat-sales estimator can address this, this comes at a high cost in terms of drastically reduced sample size and the maintained assumption that time-varying unobservables play no role. Noonan (2007) nonetheless estimates a repeat-sales model and accordingly loses the ability to identify the effects of designation on the price of the sold property itself. The 2SLS approach here avoids the reduced sample but lacks controls for time-invariant unobservables. Building-specific unobservables can be factored out via a fixed-effects model (at the cost of only 3,249 observations that are singletons for their building or street address). The addition of building fixed effects eliminates the building-level unobservables correlated with *DISTRICT*, but policy endogeneity may yet remain (e.g., if rapidly appreciating buildings tended to become designated).

The fixed-effect model identifies district effects by using effectively only observations from buildings that had sales before and after a district designation. Thus, different results in the fixed-effect model and those reported in Table 4 may derive either from controlling for important building-specific unobservables, identification using far fewer districts, or both.[15] Where district designations vary rarely within building, this approach suffers from many of the same limitations as Noonan (2007). Nonetheless, it presents an interesting robustness check. In the fixed-effect context, the instruments need to predict the timing of a sale (before or after designation) rather than whether a building will become designated. The interaction between *RO*, the change in block-group income from 1980 to 1990, and *saleyear* captures this risk of designation (*RO* × Δ*income*) and timing of sale. A second instrument uses a different demand-shifter, percent college educated, but is otherwise the same.

The results of the fixed-effect estimator appear in column 4 of Table 4 and provide further support for the model in column 1. The results for the property attributes (e.g., *rooms*, *baths*) are quite similar. The geographic, neighborhood, and CHRS variables are all dropped from this model because they are building-specific and time-invariant. Interestingly, the instrumented price effect of *DISTRICT* in this model is -37% (z = -2.78). While this reinforces the earlier results, the effects of proximity to other landmarks reverse. The price effect of nearby landmarks becomes much weaker and convex rather than concave. In a similar reversal, *DistShare* has a large negative effect (-0.64) that is statistically significant. Hansen's *J* of 0.28 offers little reason to doubt the exogeneity of these instruments. The instruments both positively

[15] The subsample receiving a *DISTRICT* treatment shrinks from 2,143 units in 18 different districts in Table 3 and the rest of Table 4 to 217 units in just 3 different districts.

and significantly predict district status in the first stage as expected, although their Kleibergen-Paap statistic of 22.32 raises some concern about bias from the relatively weak instruments.

Despite the consistency of results and the strength of the diagnostic statistics reported in Table 4, there is still some concern for the sensitivity of these results. Alternative instruments with weaker theoretical justification and typically more troublesome diagnostic statistics are available and yield substantively different results. The low power of Hanson's *J* test also advises caution. Moreover, clustering errors at the building level greatly inflates the standard errors, especially for the *DISTRICT* variable and the variables from the CHRS, leaving many of the effects insignificant. (The external effects of landmark designations remain significant.) A more cautious interpretation of these findings is just that no robust evidence of positive own-price effects from landmarks is found.

VI. Conclusion

After developing a simple theoretical model that casts doubt on the exogeneity of policy variables and helps to identify some plausible instruments, this paper demonstrates the value of correcting for policy endogeneity in hedonic price analyses by using a more robust estimator with richer data than previous studies of the implicit price of historic landmark designation. OLS estimation of the hedonic price model offers results consistent with much of the previous literature, namely that property values are higher in historic districts. The rich data allow unprecedented controls of historic quality, neighborhood quality, and property attributes. Surprisingly, controlling for historical observables has an upward influence on the estimated price effect of *DISTRICT* in OLS. The external effects of district designations are positively affected by better controls for neighborhood historic quality (and individual designations' spillover is virtually unchanged) in OLS. The better test for this kind of omitted variable bias makes use of the 2SLS model, however. Unfortunately, omitting the historic attributes prevents estimation of the 2SLS model without replacement instruments. The vital role of historical quality in the OLS and 2SLS models demonstrates how careful analysis of the impact of historic preservation policies depends on controls for existing historic quality.[16]

Most critically, the endogeneity of district designation status is explored using instruments derived from a simple model of regulator behavior. The empirical results provide some robust evidence that previous methods may be particularly vulnerable to endogeneity and omitted variable biases in this context. Although Noonan (2007) attempts to overcome the omitted-variable problem with a repeat-sales model, akin to mitigating endogeneity with a difference-in-difference estimator, our approach here explicitly accounts for the endogeneity in a IV framework. The IV approach has an advantage in being able to identify implicit prices for designation when designations rarely occur between observed sales. The variable of interest, historic district designation status, is found to be endogenous. The IV approach also allows for the explicit estimation of the determinants of designation in the sample. For a variety of instrument sets that leverage the historic quality measures from the CHRS, the first-stage results point to strong instruments whose predicted impacts are consistent with theory. In the 2SLS hedonic, robust estimates of the "own" price effect of historic designation are shown to be large and negative for landmark districts.

The effects of designation on nearby properties are shown to be statistically significant and substantively important. Unlike previous literature, the number of historic landmark buildings designated in a neighborhood appears nonlinearly associated with prices even in the 2SLS framework. The first nearby landmark has a positive spillover, but subsequent neighborhood landmarks appear to be disamenities. Proximity to districts also affects price. In

[16] Identifying a policy effect without controls for historical quality poses a serious challenge for researchers. In some contexts, like the NRHP but not Chicago's landmarks, eligibility is tied to a threshold structure age like "50 years". Such arbitrary cutoffs suggest that a regression discontinuity design might identify policy effects even with very limited information about historic quality. We leave that to future research.

the aggregate, these external effects are quite substantial. For Chicago, roughly two-thirds of sales are in block groups with no other landmark districts or buildings – leaving a third of the properties in the neighborhood of landmarks. 19% of the units in the sample have a district in their neighborhood, conferring sizable price premiums. (With *DistShare*'s mean of 0.20 among those units neighboring districts, they enjoy an average +4% spillover.). Of all sales, 13% have one landmark building in their block group, also conferring sizeable premiums. Importantly, however, over 8% of sold units have two or more landmark buildings and thus many properties "near" landmark buildings actually experience negative spillovers from these designations. These negative spillovers can be economically important; condo and townhome prices positively correlate with landmark density.[17]

While the OLS price effects presented here are noticeably smaller than have been found elsewhere, none of the previous work has directly addressed policy endogeneity. Correcting for the endogeneity in the hedonic price model yields large, negative price effects. Controlling for historic quality can affect the substantial external price effects as well. Caution is warranted in interpreting these price premiums. First, there is still some possibility that these effects are artifacts of unobserved quality. Second, the hedonic price describes the marginal price effect of designation, not a welfare effect. In first-stage hedonics, changing prices may result as much from altering demand (by displacing it elsewhere) or shifts in supply. Given the nature of landmark district designation (i.e., preserving neighborhoods and hence constraining supply), the

[17] As a very crude estimate, using coefficients in column 1 of Table 4 and assuming that the sample of sales is representative of Chicago's housing, removing all landmark buildings would lower home values by $263 on average (winners gain $14,320 on average while losers lose $4,899 on average and greatly outnumber the winners). Similarly, removing all districts would lower home values by $45 on average (winners gain $59,203 average while losers lose $10,804 on average and again greatly outnumber the winners). Seen in this redistributive light, the historic preservation policy appears to concentrate the harms of designation while spreading the gains widely. The combined effect of landmark buildings and districts shown here has a small net positive effect ($309 is roughly 0.2% of mean home value). This net effect masks how roughly 7% of the units would gain by the absence of landmarks by 13% of their property value, but they are outnumbered by 4:1 by those whose home values would suffer without the landmarks by 4%. Large values are at stake for the 35% of the units that are either winners or losers in Chicago.

lower prices for homes in districts suggests that the restrictions on property use are indeed lowering demand for the housing assets. These lower prices may be indicative of a takings common to preservation laws where a few individuals suffer costly encumbrances on their property for the sake of positive external benefits. An interesting extension for future research would be to explore the progressivity of such a policy given the positive correlation between property value and designations.

Evaluating a preservation policy involves several aspects. First, there is the effect of the policy on the preserved buildings – including effects on historical quality preserved and property values. Next, there is the effect of the preservation policy on the surrounding properties and neighborhood – including on neighborhood dynamics and property values. If preservation policy has increased the supply of the historic resource in a city, then some of the credit for the external effects of historic structures belongs to the policy. Preservation policy, however, might not increase the total amount of the historic resource especially if it is perceived as a taking. Preserving heritage that already exists is not the same as increasing supply. As Turnbull (2002) has shown, the threat of "preservation" can speed owners towards redevelopment, which could lead to an overall *decrease* in historic resources (even if the policy itself effectively preserves the properties it does designate). These more dynamic effects of preservation policy on the stock of historic properties in a city have received no attention to our knowledge, mostly because good data on the stock of historic resources in an area do not exist. These dynamic effects could play critical roles in the general efficiency of preservation programs and on the optimal administration of these programs when they are enacted. Even a more robust approach to estimating "own" and external price effects of landmark preservation, such as that presented here, does little to answer the larger question of "how do preservation policies affect the stock of historic buildings?"

References:

Ahlfeldt, G. M. and W. Maennig (2009). "Substitutability and complementarity of urban amenities: External effects of built heritage in Berlin." Forthcoming in *Real Estate Economics*.

Asabere, P. K., & Huffman, F. E. (1994a). The value discounts associated with historic façade easements. *The Appraisal Journal*, *62*(2), 270-277.

Asabere, P. K. & Huffman, F. E. (1994b). Historic designation and residential market values. *The Appraisal Journal*, *62*(3), 396-401.

Benson, V. O., & Klein, R. (1988). The impact of historic districting on property values. *The Appraisal Journal*, *56*(2), 223-232.

City of Chicago. (2004). Chicago landmarks home page. Retrieved October 6, 2004 from http://www.cityofchicago.org/Landmarks/

Clapp, J. M., & Giaccomotto, C. (1998). Price indices based on the hedonic repeat-sales method: Application to the housing market. *Journal of Real Estate Finance and Economics*, 16(1), 5-26.

Clark, D. E., & Herrin, W. E. (1997). Historical preservation and home sale prices: Evidence from the Sacramento housing market. *Review of Regional Studies*, *27*, 29-48.

Commission on Chicago Landmarks. (1996). *Chicago Historic Resources Survey: An Inventory of Architecturally and Historically Significant Structures*. Chicago: Department of Planning and Development.

Commission on Chicago Landmarks. (2006). Landmarks ordinance with rules and regulations of the Commission on Chicago Landmarks [Electronic version]. Chicago: City of Chicago.

Coulson, N. E., & Lahr, M. L. (2005). Gracing the land of Elvis and Beale Street: Historic designation and property values in Memphis. *Real Estate Economics*, *33*(3), 487-507.

Coulson, N. E., & Leichenko, R. M. (2001). The internal and external impact of historical designation on property values. *Journal of Real Estate Finance and Economics*, *23*(1), 113-124.

Glaeser, Ed. (2006). Bonfire of the landmarks. *New York Sun* opinion 29 November 2006. Downloaded http://www.nysun.com/article/44261 on 30 November 2006.

Halvorsen, Robert and Raymond Palmquist. (1980). The interpretation of dummy variables in semilogarithmic equations. *American Economic Review*, 70(3): 474-475.

Leichenko, R. M., Coulson, N. E., & Listokin, D. (2001). Historic preservation and residential property values: An analysis of Texas cities. *Urban Studies*, *38*(11), 1973-1987.

Listokin, D., Listokin, B., & Lahr, M. (1998). The contributions of historic preservation to housing and economic development. *Housing Policy Debate*, 9(3), 431-478.

Little, R. J. A. (1992). Regression with missing X's: A review. *Journal of the American Statistical Associ*ation, 87(420), 1227-1237.

Mason, Randall. 2005. "Economics and Historic Preservation: A Guide and Review of the Literature." Discussion Paper. Brookings Institution Metropolitan Policy Program. Available at: http://www.brookings.edu/metro/pubs/20050926_preservation.pdf . Last accessed 23 February 2008.

New York Landmarks Conservancy. 1977. *The Impacts of Historic Designation—Summary.* Study conducted by Raymond, Parish, Pine and Weiner, Inc. New York Landmarks Conservancy, New York.

Noonan, D. S. (2007). Finding an Impact of Preservation Policies: Price Effects of Historic Landmarks on Attached Homes in Chicago 1990-1999. *Economic Development Quarterly*. 21(1).

Noonan and Krupka (2008). "Determinants of historic and cultural landmark designation: why we preserve what we preserve." IZA discussion paper no. 3777. available at: http://ssrn.com/abstract=1293549 .

Rosen, S. (1974). Hedonic prices and implicit markets: Product differentiation in perfect competition. *Journal of Political Economy*, 82(1), 34-55.

Royston, P. (2005). Multiple imputation of missing values: Update of ice. *Stata Journal*, 5(4), 527-536.

Rypkema, D. D. (1994). The economics of historic preservation: A community leader's guide. Washington, DC: National Trust for Historic Preservation.

Schaeffer, P. V., & Millerick, C. A. (1991). The impact of historic district designation on property values: An empirical study. *Economic Development Quarterly*, *5*(4), 301-312.

Schuster, J. Mark. 2002. "Making a List and Checking it Twice: The List as a Tool of Historic Preservation." Working Paper 03-03. Harris School of Public Policy Studies, University of Chicago. 25 November 2005.

Swaim, Richard. 2003. "Politics and Policymaking: Tax Credits and Historic Preservation." *Journal of Arts, Management, Law and Society*. 33 (1): 32 – 39.

Turnbull, Geoffrey K. (2002). Land Development under the Threat of Taking. *Southern Economic Journal* 69(2), 290-308.

Table 3: OLS Regression Results

	1		2		3		4	
	Coef.	t	Coef.	t	Coef.	t	Coef.	t
DISTRICT	0.2209***	21.64	0.0156*	1.92	0.0217***	2.65	0.1139***	10.22
CountLmk					0.0708***	12.45	0.0729***	12.73
$CountLmk^2$					-0.0318***	-16.12	-0.0314***	-15.88
DistShare					0.2131***	13.54	0.3935***	20.55
named							0.0418***	3.64
OTH							0.0330***	3.55
RO							0.0219**	2.38
CHRS100							-0.0002	-0.90
CHRS250							-0.0014***	-15.48
CHRS500							0.0001***	4.69
lnarea	0.6835***	17.69	0.5497***	16.12	0.5471***	16.07	0.5434***	16.03
yearbuilt	0.0009***	12.64	0.0002***	2.89	0.0002***	3.07	0.0003***	4.35
unitbldg	0.0002***	13.16	-0.0001***	-5.24	-0.0001***	-5.87	-0.0001***	-7.10
$unitbldg^2$	-1.4E-08***	-6.13	1.3E-08***	4.58	1.4E-08***	4.86	1.6E-08***	5.53
rooms	-0.0087***	-3.45	0.0167***	4.31	0.0170***	4.33	0.0173***	4.39
bedrooms	-0.0122*	-1.74	0.0799***	5.73	0.0802***	5.74	0.0802***	5.73
baths	0.2545***	24.32	0.1752***	17.68	0.1733***	17.53	0.1708***	17.38
mbbth	0.1168***	22.70	0.0610***	16.48	0.0600***	16.20	0.0599***	16.29
fireplace	0.1390***	22.51	0.0508***	11.25	0.0521***	11.61	0.0538***	12.09
garage	-0.0003	-0.05	-0.0250***	-6.29	-0.0226***	-5.69	-0.0236***	-5.99
parking	0.1552***	20.62	0.1035***	18.58	0.1050***	18.85	0.1052***	18.92
parkspot	-0.1198***	-20.63	-0.0707***	-16.75	-0.0708***	-16.77	-0.0730***	-17.36
saleyear	0.0300***	33.83	0.0390***	58.48	0.0398***	59.84	0.0393***	59.30
waterfront			0.0093*	1.72	0.0068	1.27	0.0042	0.78
distCBD			-0.0709***	-22.61	-0.0745***	-23.45	-0.0761***	-23.62
$distCBD^2$			0.0014***	20.51	0.0014***	18.55	0.0014***	18.63
distLake			0.0223***	10.03	0.0242***	10.92	0.0266***	11.84
$distLake^2$			0.0000	0.42	0.0002*	1.75	0.0001	0.78
distwater			0.0268***	5.20	0.0237***	4.56	0.0335***	6.22
$distwater^2$			-0.0083***	-7.44	-0.0085***	-7.57	-0.0100***	-8.72
distCTA			0.0408***	8.40	0.0472***	9.38	0.0378***	7.55
$distCTA^2$			-0.0051***	-4.96	-0.0070***	-6.16	-0.0062***	-5.58
distpark			-0.1697***	-11.77	-0.1725***	-12.01	-0.1462***	-10.21
$distpark^2$			0.1841***	15.12	0.1917***	15.84	0.1628***	13.49
northside			122.7497***	5.23	70.2282***	3.02	59.8161**	2.56
latitude			2.0346***	7.08	1.3450***	4.72	1.0167***	3.52
north.×latitude			-2.9267***	-5.22	-1.6721***	-3.01	-1.4229**	-2.55
income			0.0012***	8.88	0.0007***	5.50	0.0004***	3.28
medyrblt			-0.0063***	-37.52	-0.0058***	-33.67	-0.0057***	-32.58
medValue			0.0000***	4.27	0.0000***	3.40	0.0001***	7.95
white			0.2590***	24.12	0.2534***	23.58	0.2454***	22.90
popdens			-0.0001*	-1.92	-0.0004***	-5.52	-0.0004***	-5.07
college			0.0476***	3.91	0.0543***	4.45	0.0906***	7.20
new const.			0.4103***	33.51	0.3968***	32.40	0.3689***	30.03
constant	-55.2298***	-29.78	-143.6637***	-12.02	-117.4465***	-9.93	-103.1579***	-8.61
R^2	0.6108		0.7938		0.7955		0.7975	

*, **, *** indicates significance at the 0.1, 0.05, 0.01 levels, respectively

TABLE 4: 2SLS Results

	1			2			3			4		
	coef.		z	coef.		z	coef.		z	coef.		z
DISTRICT	-0.2092	***	-3.04	-0.2607	***	-4.26	-0.2862	***	-5.13	-0.4698	***	-2.78
CountLmk	0.0653	***	10.73	0.0636	***	10.67	0.0641	***	10.75	-0.0236	*	-1.71
$CountLmk^2$	-0.0310	***	-15.51	-0.0307	***	-15.39	-0.0311	***	-15.54	0.0084	**	2.51
DistShare	0.2092	***	4.77	0.1790	***	4.47	0.1659	***	4.44	-0.6380	***	-3.55
named	0.0545	***	4.46	0.0557	***	4.56	0.0570	***	4.62			
OTH	0.0795	***	5.68	0.0865	***	6.56	0.0918	***	7.15			
RO	0.0431	***	4.20	0.0465	***	4.59	0.0486	***	4.85			
CHRS100	0.0013	***	3.19	0.0016	***	4.07	0.0017	***	4.68			
CHRS250	-0.0008	***	-5.13	-0.0007	***	-4.85	-0.0007	***	-4.74			
CHRS500	0.0001	***	4.97	0.0001	***	5.11	0.0001	***	4.91			
lnarea	0.5492	***	16.04	0.5500	***	16.09	0.5508	***	16.08	0.2588	***	12.63
yearbuilt	0.0001	**	1.96	0.0001	*	1.65	0.0001	*	1.67	0.0000		-0.15
unitbldg	-0.0001	***	-6.67	-0.0001	***	-6.63	-0.0001	***	-6.39	0.0000		0.95
$unitbldg^2$	1.6E-08	***	4.97	1.6E-08	***	4.88	1.6E-08	***	4.76	-9.7E-10		-0.28
rooms	0.0177	***	4.43	0.0177	***	4.45	0.0177	***	4.44	0.0166	***	4.03
bedrooms	0.0795	***	5.68	0.0794	***	5.69	0.0793	***	5.68	0.1409	***	9.05
baths	0.1742	***	17.51	0.1748	***	17.61	0.1750	***	17.61	0.1680	***	15.33
mbbth	0.0588	***	15.63	0.0585	***	15.52	0.0588	***	15.57	0.0413	***	13.47
fireplace	0.0544	***	12.03	0.0545	***	11.98	0.0548	***	12.03	0.0029		0.73
garage	-0.0257	***	-6.42	-0.0255	***	-6.36	-0.0260	***	-6.47	-0.0578	***	-17.53
parking	0.1055	***	18.76	0.1052	***	18.64	0.1048	***	18.54	0.0904	***	19.38
parkspot	-0.0737	***	-17.43	-0.0736	***	-17.34	-0.0735	***	-17.31	-0.0807	***	-25.74
saleyear	0.0399	***	58.87	0.0399	***	59.08	0.0399	***	59.01	0.0396	***	66.76
waterfront	0.0133	**	2.34	0.0144	**	2.53	0.0163	***	2.89			
distCBD	-0.0808	***	-25.01	-0.0793	***	-24.54	-0.0872	***	-27.27			
$distCBD^2$	0.0014	***	18.82	0.0014	***	18.73	0.0014	***	19.17			
distLake	0.0291	***	12.84	0.0283	***	12.53	0.0325	***	14.41			
$distLake^2$	0.0001	***	1.27	0.0001		1.13	0.0002	***	1.81			
distwater	0.0297	***	5.47	0.0285	***	5.20	0.0292	***	5.35			
$distwater^2$	-0.0100	***	-8.72	-0.0097	***	-8.42	-0.0105	***	-9.11			
distCTA	0.0502	***	9.20	0.0510	***	9.52	0.0564	***	10.66			
$distCTA^2$	-0.0081	***	-7.07	-0.0080	***	-6.97	-0.0092	***	-8.12			
distpark	-0.1536	***	-10.56	-0.1546	***	-10.62	-0.1556	***	-10.68			
$distpark^2$	0.1700	***	13.85	0.1710	***	13.92	0.1727	***	14.05			
northside	12.4060		0.53	26.4439		1.12	-42.5372	*	-1.82			
latitude	0.5760	**	1.99	0.7829	***	2.70	-0.0057		-0.02			
north.×latitude	-0.2914		-0.52	-0.6267		-1.11	1.0204	*	1.83			
income	0.0012	***	5.46	0.0000	***	6.44	0.0000	***	6.98			
medyrblt	-0.0058	***	-32.67	-0.0058	***	-32.63	-0.0058	***	-32.65			
medValue	0.0001	***	7.70	0.0000	***	7.57	0.0000	***	7.78			
white	0.2358	***	21.60	0.2361	***	21.69	0.2271	***	20.87			
popdens	-0.0005	***	-6.03	0.0000	***	-6.00	0.0000	***	-6.59			
college	0.0494	***	3.18	0.0427	***	2.87	0.0407	***	2.79			
new const.	0.3970	***	29.63	0.4009	***	29.67	0.4032	***	30.49			
constant	-85.4505	***	-7.11	-93.9996	***	-7.79	-61.0558	***	-5.10			
N	59642			59642			59642			56623		
centered R^2	0.7932			0.7917			0.7909			0.6043		

Kleibergen-Paap rk Wald F statistic	53.044			259.811			156.710			22.316		
Hansen J	0.829		p=0.36	1.675		p=0.20	5.067		p=0.17	0.285		p=0.59
First-stage results			t						t			t
CHRS100× OwnOcc70	0.0095	***	7.19				0.0072	***	5.89			
named× VacRat70	0.0534	**	2.41				0.0380	*	1.82			
CHRS100× LngRes70				0.0270	***	15.94	0.0241	***	12.9			
named× NewOne70				-1.0588	***	-14.99	-0.8108	***	-16.9			
RO×ΔCol70× saleyear										0.0207	***	6.40
RO×ΔInc70× saleyear										0.0099	***	6.66