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NEIGHBORHOOD DYNAMICS AND PRICE EFFECTS OF SUPERFUND SITE CLEAN-UP

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Table of Contents

Executive Summary

The existence of hazardous waste sites is an important problem in many older, built-out urban areas. Several authors have shown that property values around these sites are depressed, and are usually surrounded by undesirable neighborhoods. The reasons for this collocation are relatively straightforward: hazardous waste sites are unpleasant neighbors, so properties in the adjoining areas must sell for less. This combination of low prices and dirty environment attracts residents who value clean environments less, usually because they are poor. The problems with large concentrations of poor residents are well documented, and these concentrations further lower property values.

Local governments have looked to hazardous waste site clean-up as a way to improve property values. Clean-ups of Superfund sites are performed under the guidance of the Environmental Protection Agency, through its authority under the Comprehensive Environmental Response, Compensation, and Liability Act (CERCLA). The costs of such remediation activity have been widely publicized and easily accounted, but the benefits have been harder to quantify. Of considerable interest to local governments is the effect of clean-up activity on local property values, since property taxes are these governments' main revenue source. Increases in property values are also of academic interest, since they are the primary measure that urban and environmental economists use in quantifying the benefits of such remediation.

Most estimates of the Superfund site price effect use data on house sales collected before the clean-up activity has finished and estimate a price effect based on these data. From this estimate, a benefit to clean-up cost can be computed. This estimate may be biased, however, if unobserved characteristics of the area (which would not change following clean-up activity) also affect prices. Some studies correct for this by examining how property values *change* when clean-ups occur. These studies are preferable, since they offer both a better estimate of the price effect of an existing site, and a direct measure of the benefit of site remediation.

We contend that these standard measures of this price effect are not appropriate because the clean-up of a site will also induce neighborhood change. Because different families may be more willing to live in an area after remediation, they will out-bid the area's original residents for the homes near the site, and thus the composition of the neighborhood will change. Since neighborhood composition has been shown to have strong effects on real estate prices, such neighborhood transition will create indirect effects of the site remediation. Similar indirect effects will exist because of reinvestment in the area surrounding the clean-up site.

Census data is used to investigate the effect of site remediation on median housing values, housing stock characteristics and neighborhood composition. We compute the direct, or "pure," price effect of the clean-up, and find that cleaning up a Superfund site directly increases home values by 2 to 5 percent. This is consistent with the rest of the literature. However, we are able to go further and compute the indirect effects, which we find to be quite substantial. As much as 50 percent of the total effect of an EPA clean-up comes through the indirect channels: induced neighborhood transition and housing reinvestment or construction.

These results have several important implications. First, they inform our interpretation of the environmental justice of the process by which poor residents are exposed to hazardous wastes. In our most flexible models, we show that after clean-ups, richer families tend to move into the remediated areas, pushing the poorer original residents out, possibly to other dirty areas. This shows that targeting environmental remediation towards favored groups will be at least partially offset by these groups sorting out of the area that has been cleaned-up. If the original poor residents are mostly renters, the clean-up will have benefited them very little or even hurt them by forcing them to undertake costly moves.

Second, the results offer a better understanding of the likely results of environmental remediation on the surrounding areas. Remediation not only makes the area more desirable (thus raising home values), but also induces further investment and immigration of more "desirable" populations. Both of these induced effects will further increase home values. These indirect effects are substantively important. At least some portion of this reinvestment and relocation will probably come at the expense of other areas, so the indirect effects should be used only cautiously in cost-benefit analysis. However, local governments interested in the likely effect on property value (and property tax receipts) will care less about these offsetting effect in other areas.

Our approach is to observe census block groups in 1990 and 2000, noting which block groups were in the vicinity of a Superfund site clean-up. We are able to estimate a system of equations (as opposed to the simpler one-equation models used in much of the literature) taking into account the causal feedback between housing values, housing stock investment, neighborhood composition and EPA clean-ups. From these estimates, it is possible to compute both consistent estimates of the "pure" price effect of the clean-up, and the indirect effects. Most of the literature focuses on the pure effect.

With our system of equations approach, we are able to go further and compute the indirect effects. As noted above these effects are found to be quite substantial. The indirect effects are also quite stable across model specifications.

I. Introduction

Hedonic price analysis is often used to estimate the implicit price of structural or neighborhood characteristics. There is a temptation in hedonic analysis to interpret the coefficients identified in the first-stage price regression from cross-sectional variation as predictors of within-unit variation over time. This temptation is especially strong in the case of environmental variables, such as hazardous waste sites. These coefficients may be biased, however, by unobservable factors correlated with the presence of hazardous waste sites, which affect price. Some researchers have used repeat-sales panel data to control for time-invariant omitted factors, observing the change in price as hazardous waste sites are cleaned up.

What both the cross-sectional and panel approaches are unable to address is the fact that urban populations are mobile, and that the housing stock is malleable. Neighborhood sorting occurs and reinvestment takes place as hazardous waste sites are cleaned up. If neighborhood characteristics affect price, the effect of clean-ups on price *through* neighborhood transition and reinvestment will be important parts of the total effect of hazardous waste clean-ups. Partial equilibrium cross-sectional studies or panel studies that neglect the effect of environmental quality on neighborhood composition and investment decisions may miss these effects.

This paper uses panel data to estimate a system of equations that allows for endogeneity among prices, neighborhood characteristics and housing stock variables. Although the data are not ideal, our estimates of the direct effect of hazardous waste clean-ups closely resemble those estimated in other studies. However, our system of equations estimates allow us to compute the indirect effects as well, which improves our understanding of expected price changes. We find that these indirect effects are substantively significant.

The implications of these results are twofold. First, that sorting does appear to occur because of clean-up activity has implications for the environmental justice literature. Cross-sectional models of minority group exposure may not be adequate given these dynamic responses of housing markets to environmental improvements. Second, the relative magnitude of the indirect effects suggests that these effects should be considered carefully in cost-benefit analyses of hazardous waste clean-ups, or any other policy intervention that might cause neighborhood transition.

The rest of the paper runs as follows: Section II briefly reviews the relevant literatures. Section III lays out the empirical model, derives the total effect of a clean-up and describes the data. Section IV presents the results and Section V concludes.

II. Analyzing the Effects of Superfund Sites

The location of NPL sites is a major policy issue, especially in terms of the equity of exposure. "Toxic Wastes and Race in the United States" (United Church of Christ, 1987) provided the first systematic analysis of the location of hazardous waste NPL sites. This evidence helped launch the environmental justice movement by providing a static look at zip codes nationwide and found that minorities were disproportionately exposed to hazardous waste. Several similar studies subsequently produced varying results (Hird 1993, Zimmerman 1993, Yandle and Burton 1996, Baden and Coursey 2002). Much of this research had methodological issues concerning the use and interpretation of geographic data (Baden et al. 2005).

These static analyses are unable to address the causal mechanism through which possible environmental injustices arise. Been (1997), Anderton *et al*. (1994) and Baden and Coursey (2002) all cast doubt on hypothesis that siting follows race. Hamilton (1995) found that facility expansion decisions were more likely to occur in neighborhoods where collective political action was weakest, but found no independent effect of race. While these studies examine sorting around existing sites, to date no studies have rigorously investigated neighborhood transition in the wake of Superfund site remediation. If there are price effects associated with environmental remediation, then neighborhood sorting should follow. While this dynamic has yet to be observed following siting (Ringquist 2006), we offer new evidence on sorting following remediation.

The hedonic literature concerning price effects of Superfund NPL sites is sizable. While some studies, such as Greenberg and Hughes (1992) use simple means comparisons to draw inferences about the effects of environmental hazards on property values, the majority of economic studies estimate price effects from first-stage hedonic regressions. Our review of 11 studies[1] that have produced either price estimates for immediate vicinity effects or price gradient effects that are comparable across studies gives us a sense of what the direct effects of site proximity are. Six of

[1] Clark and Nieves (1994), Dale et al. (1999), Gayer et al. (2000), Ketkar (1992), Kiel (1995), Kiel and Williams (2005), Kiel and Zabel (2001), Kohlhase (1991), McCluskey and Rausser (2003), Mendelsohn et al. (1992) and Michaels and Smith (1990).

these studies report adjacency effects, which range from 1 percent to –12 percent of the total property value. The rest of the papers estimate price gradients around these sites. These gradients range from insignificant to about six percent per mile away from the hazardous waste site. In these studies, the price gradient becomes insignificant between one and six miles from the site. In light of the trend of increased data and computing availability, only Ketkar (1992) uses aggregated census data to specify the dependent variable. In that New Jersey sample, hazardous waste sites account for 2 percent lower median housing values per municipality.

III. Theory and Empirical Methodology

Theoretical Underpinnings

In general, hedonic studies use cross-sectional data to estimate a first stage equation of the form:

$$P = \beta_0 + \beta_E E + \beta_S S + \beta_N N + \beta_G G + \varepsilon_1, \qquad (1)$$

where E measures environmental goods, S is a vector of structural characteristics of a property, N is a set of neighborhood demographic characteristics, and G is a set of geographical amenities (such as distance to CBD).

Because we are interested in the within-unit effects of changes in E, a first-differences approach with panel data is useful (Mendelsohn *et al.* 1992). This approach also has the advantage of differencing out any time-invariant omitted variables that might bias the coefficients in equation one. Another important advantage of using a first-difference approach (especially considering our sample subsumes hundreds of local housing markets) is that all time-invariant local market idiosyncrasies, such as the local rent gradient, also cancel out, assuming constant prices on these characteristics.[2] By looking at changes over time, we observe not only cross-sectional variation in S and N, but also observe how these variables respond to changes in the environmental good, something that is impossible with cross-sectional data. First differencing, we arrive at:

$$\dot{P} = \beta_E \dot{E} + \beta_S \dot{S} + \beta_N \dot{N} + \varepsilon_2, \qquad (2)$$

where $\dot{X} = X_t - X_{t-1}$. In this paper, we acknowledge that in this equation neighborhood demographics ($\dot{N}$) and housing stock characteristics ($\dot{S}$) are likely not only to affect prices ($\dot{P}$), but also respond to prices, as well as to each other. If demographic groups differ in their demand for environmental quality, or people build different housing in cleaner areas, then $\dot{N}$ and $\dot{S}$ will also respond to changes in environmental quality. Thus, a properly specified estimation of equation (2) would also include corrections for this endogeneity:

[2] For local rent gradients and climatic amenities, we relax this constant price assumption in some specifications.

$$\dot{S} = \gamma_E \dot{E} + \gamma_N \dot{N} + \gamma_P \dot{P} + \varepsilon_3 \qquad (3)$$

$$\dot{N} = \delta_E \dot{E} + \delta_S \dot{S} + \delta_P \dot{P} + \varepsilon_4 \,. \qquad (4)$$

In equation (3), the substitution away from land and towards capital when land prices are high suggests that changes in price of land will have a causal effect on $\dot{S}$. Neighborhood demographics, such as family size and income, will also affect the equilibrium quantity and quality of the housing stock if the demand for housing is related to these demographics. $\dot{E}$ may also have its own direct effect, but the partial correlation would depend on whether housing and the environmental good are complements or substitutes.

In equation (4), demographic groups' differing demands for E may cause them to sort into neighborhoods according to their willingness to pay for these attributes (Diamond and Tolley, 1982). Similar arguments hold for both the inclusion of $\dot{S}$ and $\dot{P}$ in equation (4): richer residents' higher willingness to pay for housing may lead them to sort into neighborhoods with nicer houses and higher prices, at least when the capital stock is somewhat inelastic.

We are specifically interested in the effect of $\dot{E}$, especially when E changes due to policy intervention, as in the case of hazardous waste site clean-ups. Once the system of equations (2)-(4) is accepted, the total effect of a clean-up ($\dot{E}$) can be seen to depend not solely on its direct effect β_E in equation (2), but also on its indirect effects. Converting the system to matrix notation, assuming G does not change, totally differentiating and dividing through by $d\dot{E}$, we get:

$$\begin{bmatrix} 1 & -\beta_S & -\beta_N \\ -\gamma_P & 1 & -\gamma_N \\ -\delta_P & -\delta_S & 1 \end{bmatrix} \begin{bmatrix} d\dot{P}/d\dot{E} \\ d\dot{S}/d\dot{E} \\ d\dot{N}/d\dot{E} \end{bmatrix} = \begin{bmatrix} \beta_E \\ \gamma_E \\ \delta_E \end{bmatrix}. \qquad (5)$$

W can then use Cramer's Rule to obtain the total effect of a change in E:

$$\frac{d\dot{P}}{d\dot{E}} = \frac{\beta_E + \beta_S\gamma_E + \beta_N\delta_E + \beta_S\gamma_N\delta_E + \beta_N\delta_S\gamma_E - \beta_E\gamma_N\delta_S}{1 - \beta_S(\gamma_N\delta_P + \gamma_P) - \beta_N(\delta_S\gamma_P + \delta_P) - \gamma_N\delta_S} \,. \qquad (6)$$

The first term in the numerator is the direct effect, while the rest of the terms are indirect effects. Under the assumption of no endogeneity in equation (2), this total derivative reduces to the first three terms in the numerator. From equation (6), it is clear that the total effect of a hazardous waste clean-up consists of more than simply its direct effect.

To identify the parameters in equation (5), we estimate the system of equations (2)-(4). As mentioned above, using panel data and a first-difference approach eliminates any time-invariant factors that may be correlated with our independent variables. This makes the search for instruments in our estimation of equations (2)-(4) much less arduous. As described more fully below, in general, we use lagged differences (1980-1990) and twice-lagged levels (1980) as instruments for first-differences in S or N.

In this framework, the preferred data set would include a national sample of properties and a rich set of housing and resident characteristics over time. The two most obvious candidates (the American Housing Survey and the Public Use Micro Sample) only provide geographic information at the county-level. Since the effects of hazardous waste sites have been found to be highly localized (Hite *et al.* 2001, Mendelsohn *et al.* 1992), such large geographic scales are inadequate for our purposes.

In the absence of national microdata, we use aggregate measures of housing and population characteristics at the neighborhood (block group) level. Using block-group averages and medians, we wish to see how neighborhood transitions induced by site clean-ups affect total changes in prices. There are some advantages to this level of aggregation (Goodman 1977). Coulton *et al.* (2004) show that the block group matches survey respondents perceptions of "neighborhood" better than other available level of aggregation. We use U.S. census data from 1980, 1990 and 2000, processed by Geolytics, Inc. so that block-group boundaries do not change from decade to decade. This geographic consistency across years enables panel data analysis. We treat block groups, the smallest level of aggregation for which our data are available, as the unit of analysis in the first-difference approach.

The use of aggregated data, even at the neighborhood level, limits our ability to infer price effects at the individual level. Nonetheless, some hedonic research has shown that estimates using aggregate data produce reasonably accurate results

(Freeman 1979, Nelson 1979, O'Byrne *et al.* 1985).[3] Moreover, the median housing value in a neighborhood is of considerable policy import. Learning more about the effects of clean-ups on this neighborhood measure is informative, even if it does not recover the true underlying hedonic price. The results based on such aggregate measures can be viewed in an epidemiological light: the effects of average exposure on average outcomes, while not the ideal, are nonetheless interesting.

Variables and Descriptive Statistics

Data from several sources are combined to estimate the model. The results are presented in section IV, emphasizing the estimation of equation (2). $\dot{P}$ is the change in the block group's log of median house value from 1990 to 2000. Our instrument for $\dot{P}$ is the 1980 log of median home value.

Our variable of interested is $\dot{E}$, which represents EPA clean-up activity over the 1990's. Derived from public EPA data (EPA 2003), this variable equals one if a block group contains a site that was deleted or partially deleted from the NPL during the 1990s. is the most complete and final designation of a hazardous waste site, indicating that the EPA is satisfied that the site has been cleaned enough to pose no further health risk. We do not instrument for $\dot{E}$. While Viscusi and Hamilton (1999) find that neighborhood characteristics affect clean-ups, all of these characteristics are level effects, which will be differenced out of our system. Likewise, Gayer (2000) found no conclusive evidence that clean-ups depend on $\dot{S}$, $\dot{N}$ or $\dot{P}$.

$\dot{S}$ is a vector of housing characteristics expected to affect prices at an individual as well as an aggregate level. We estimate two different models. In the more parsimonious estimations, we use a restricted vector, $\dot{S}_1$. $\dot{S}_1$ includes changes in five variables: median year built of housing units, average number of rooms per unit, percent of housing units with gas or electric heating, housing density (housing units per square mile) and the percent of units in small buildings (containing four or fewer housing units). More extensive estimations use an additional vector of

[3] See Shultz and King (2001) for additional review of the use of aggregated Census data in hedonics.

endogenous variables, $\dot{S}_2$, which includes changes in three more variables: percent of housing units with complete plumbing, the average number of bedrooms and the percent of housing units that are stand-alone. The instruments for these variables are both the 1980 levels of each of these variables and the change from 1980 to 1990 of these variables.[4]

Neighborhood demographic characteristics ($\dot{N}$) are also split into a parsimonious set ($\dot{N}_1$) and a more extended set ($\dot{N}_2$) of variables. $\dot{N}_1$ includes the changes in the following variables: the log of the neighborhood median household income, the percent population that is white but not Hispanic, the percent population aged 25 or older who have completed at least a bachelor's degree, the percent population below 1.5 times the poverty level and the percent of the population employed in manufacturing, warehousing, transportation or utilities industries. $\dot{N}_2$ includes the changes of six more variables: the percent renter-occupied housing, percent population aged under 18 years, the average commute time for people working outside the home, the percent of households who do not have a vehicle available, population density and the average people per housing unit. As with the $\dot{S}$ vector, these variables are instrumented for with twice lagged levels and once lagged differences.[5]

Time-invariant components of G will cancel out in the first difference estimation. In alternative estimations of the price equation, the assumption of constant hedonic prices for these characteristics is relaxed. These include a natural amenity index computed at the county level by the USDA ERS (USDA 1999), a set of metropolitan fixed effects, and a set of interactions between the MSA dummies and distance to CBD, which was derived from various Census TIGER files and the

[4] One exception to this is the percent of units in structures containing four or less units. Because of coding changes between census years, the twice lagged value and once lagged difference are not available. Instead we instrument with the twice lagged level and once lagged difference of the percent of housing units in structures with *nine* or less housing units.
[5] Two exceptions remain. Because of the changes in definition of race and ethnicity between the 1990 and 2000 censuses, the percent white variable might not be precisely comparable across these decades. Also, while in 1990 and 2000 the data on education are reported for person's aged 25 or older, in 1980 they are reported for those 18 and over, so again, the instruments are not exactly identical in definition to the endogenous variables.

National Atlas of the United States (2004). By including these time-invariant factors in our price model, it allows housing price trends to vary according to climate and topography, across MSAs and within MSAs according to location within the urban geography. Table 1 presents the variable names, descriptions and descriptive statistics of all the variables described above.

TABLE 1. VARIABLE NAMES, DESCRIPTIONS AND DESCRIPTIVE STATISTICS

Vector	Name	Description[a]	Mean	Standard Deviation
$\dot{P}$	Price	log of median value, owner-occupied housing	0.3589	0.322
$\dot{E}$	Clean-up, own/adjacent	own or adjacent block group has an NPL site deleted from list	0.0053	0.073
$\dot{S}_1$	Year built	median year structure was built	3.3446	20.190
	Rooms	average number of rooms in housing units	0.1058	0.457
	Utility heat	percent housing units with gas or electric heat	0.0492	0.102
	Housing density	housing units per mile^2	52.8743	702.253
	Small structures	percent housing units sharing structure with 4 or less housing units	0.0004	0.080
$\dot{S}_2$	Plumbing	percent housing units with complete plumbing	-0.0007	0.028
	Bedrooms	average number of bedrooms in housing units	0.0071	0.243
	Solo unit	percent housing units not sharing structure with any other housing units	0.0031	0.091
$\dot{N}_1$	Income	log of median household income	0.3343	0.231
	White	percent non-Hispanic white population	-0.0678	0.11
	College	percent population age 25+ with at least college degree	0.0486	0.079
	Poor	percent population with income under 1.5 poverty line	-0.0031	0.099
	Blue collar	percent workers employed in “industrial” sectors	-0.0622	0.087
$\dot{N}_2$	Renter	percent occupied housing units that are renter-occupied	-0.0053	0.096
	Children	percent population aged 18 or younger	-0.0023	0.062
	Commute	average travel time for those working outside of home	2.092	4.943
	Walker	percent of households with no vehicle available	-0.0056	0.067
	Population density	people per mile^2	199.2871	2117.182
	Household size	people per housing unit	-0.0099	4.966
G	Natural amenities scale	county-level amenity index (composed of topography, temperatures, humidity and sunlight)	1.0601	3.209
	MSA dummies	Fixed effect for each MSA		
	MSA × distance	MSA-specific log of distance to historic city center		

[a] All variables are measured as changes from 1990 to 2000, except for level variables in *G*.

IV. Results

Table 2 summarizes the results of estimating several alternative specifications of the neighborhood transition and NPL clean-up system. Model 0 is just a first-differenced price equation (equation (2)), where $\dot{S}$ and $\dot{N}$ are treated as exogenous. Model 1 refers to the basic system in equations (2)-(4), including the $\dot{P}$ equation and 10 equations for vectors $\dot{S}_1$ and $\dot{N}_1$. Model 2 extends this system to also include endogenous $\dot{S}_2$ and $\dot{N}_2$ vectors. Models 3 and 4 add the levels of *G* to Models 1 and 2, respectively.

Estimates of Model 0 show median housing value rising with changes in structural and demographic characteristics of the neighborhood. The initial first-difference price regression offers results generally consistent with expectations and previous literature. The effects of these controls are relatively stable across models. NPL site clean-up is associated with a 6 percent rise in prices in Model 0. This direct effect declines by less than a percentage point as we move from a simple first-difference estimation in Model 0 to the system of equations in Model 1 and 2. The direct effect is reduced considerably when the effects of *G* variables are allowed to vary over time, as in Models 3 and 4. The direct effect drops to about 2 percent. Apparently some of the effect of clean-ups is correlated with shifts in housing markets' price gradients or fixed effects. If clean-ups tend to occur in high growth cities, then Superfund site remediation could be capturing this effect when gradients are presumed static. Overall, these results are near the midpoint of the findings referenced in Section II.

The effects of clean-ups on neighborhood composition and housing stock are presented in the bottom panel of Table 2. In many cases, NPL site clean-ups are associated with significant changes in structural and demographic changes in the neighborhood. Allowing prices of geographic attributes to vary over time alters the estimated effects of clean-ups on the other endogenous variables. Some consistent effects are evident, however. Clean-ups appear associated with larger shares of housing with gas or electric heat, of nonwhite population, of children and with the average commute time.

TABLE 2. RESULTS FOR $\dot{P}$ EQUATION AND SELECTED RESULTS FOR OTHER EQUATIONS, INCLUDING ADJACENTS

Model	0a	1a	2a	3a	4a
Vectors included:	$\dot{S}_1,\ \dot{N}_1$	$\dot{S}_1,\ \dot{N}_1$	$\dot{S}_1,\ \dot{S}_2,\ \dot{N}_1,\ \dot{N}_2$	$\dot{S}_1,\ \dot{N}_1,\ G$	$\dot{S}_1,\ \dot{S}_2,\ \dot{N}_1,\ \dot{N}_2, G$
N:	198,640	195,692	195,942	195,313	195,293
First- differenced variables:	β	β	β	β	β
Year built	0.001***	0.001***	0.0005***	0.001***	0.001***
Rooms	0.097***	0.300***	0.428***	0.244***	0.342***
Utility heat (%)	0.256***	-0.803***	-0.750***	-0.047**	-0.225***
Housing density[a]	<0.0001	0.029***	0.014***	0.026***	-0.001***
Small structures (%)	-0.369***	0.155***	0.024	-0.013	0.037
Plumbing (%)			-0.183***		-0.265***
Bedrooms			-0.275***		-0.258***
Solo unit (%)			-0.007		-0.235
Income	0.237***	0.642***	0.627***	0.484***	0.480***
White (%)	0.302***	0.583***	0.384***	-0.180***	0.169***
College (%)	0.211***	0.992***	1.042***	1.220***	1.152***
Poor (%)	-0.159***	-0.153***	0.040**	-0.119***	-0.040***
Blue collar (%)	0.103***	-0.451***	-0.536***	-0.319***	-0.329***
Renter (%)			-0.445***		-0.495***
Children (%)			-0.260***		-0.069***
Commute			0.004***		0.002***
No vehicle (%)			0.159***		-0.026
Population density[a]			-0.024***		0.017***
Household size			-0.006***		-0.003*
Clean-up, own/adjacent (β_E)	0.059***	0.054***	0.050***	0.021**	0.018**
Natural amenity scale				-0.008***	-0.007***
MSA & distance interactions				Yes	Yes
constant	0.269***	0.609***	0.287***	0.730***	0.563***
Dependent variable:	Partial effect of "Clean-up in or adjacent" by equation, i.e., γ_E or δ_E				
Year built		-0.098	0.511	0.617**	1.117***
Rooms		-0.029**	-0.048***	0.040***	0.040***
Utility heat (%)		0.008***	0.010***	0.014***	0.013***
Housing density		23.533	-8.151	-27.734	-23.762
Small structures (%)		-0.003	-0.003	-0.001	-0.002
Plumbing (%)			-0.001		-0.001
Bedrooms			-0.016*		0.008
Solo unit (%)			-0.003		-0.0001
Income		-0.006	-0.002	0.018*	0.023***
White (%)		-0.024***	-0.019***	-0.002	-0.004
College (%)		0.008***	-0.0005	0.002	0.002
Poor (%)		0.014***	0.003	0.004	0.003
Blue collar (%)		-0.012***	-0.004*	0.002	0.002
Renter (%)			0.021***		0.001
Children (%)			0.005**		0.005**
Commute			-0.476***		-0.388***
No vehicle (%)			0.008***		0.002
Population density			66.152		41.498
Household size			0.362**		0.226

[a] measured as 1000s/mi^2

***, **, * for p<0.01, <0.05, <0.10, respectively

The full price effect of an NPL clean-up can be calculated by solving for $d\dot{P}/d\dot{E}$ in equation (5). These effects appear in Table 3 along with direct price effects reprinted from Table 2. In each of the models estimated, the full effect exceeds the direct effect, typically by two percentage points. Relative to a direct price effect on the order of 2 – 5 percent, the indirect price effects of $\dot{E}$ are substantial. Housing markets and residential sorting mechanisms appear responsive to changes in environmental quality, in ways important to prices. Regardless, direct price estimates, from neighborhood-level hedonic analysis (as in Model 0) or from systems models (as in Model 4), capture only part of the effect of clean-ups on prices.

TABLE 3. SUMMARY OF DIRECT AND FULL PRICE EFFECTS OF CLEAN-UP ACTIONS

Model	Vectors Included	$\dot{E}$	Direct Effect ∂P/∂E	Full Effect dP/dE
1a	$\dot{S}_1$, $\dot{N}_1$	clean-up in own or adjacent block group	0.0544	0.0751
2a	$\dot{S}_1$, $\dot{S}_2$, $\dot{N}_1$, $\dot{N}_2$	clean-up in own or adjacent block group	0.0503	0.0805
3a	$\dot{S}_1$, $\dot{N}_1$, G	clean-up in own or adjacent block group	0.0211	0.0432
4a	$\dot{S}_1$, $\dot{S}_2$, $\dot{N}_1$, $\dot{N}_2$, G	clean-up in own or adjacent block group	0.0179	0.0461

V. Discussion

In this paper, we consider the price effects of changes in environmental quality in two important dimensions often overlooked in the literature. First, we explicitly model neighborhoods (block groups) as panel data in a first-difference model. This allows for better controls of omitted variables and allows explicit estimation of within-observation covariation in prices and environmental change. Second, we treat important attributes of the neighborhood (*P*, *S* and *N*) as simultaneously determined. This allows us to estimate the direct and indirect pathways through which changes in environmental quality can affect prices. The evidence suggests that there are indeed substantial indirect effects on prices through induced changes in *N* and *S*.

While hedonic prices may be relatively easy to compute, using these estimates as predictions of policy effects requires great care. Hedonic prices derived from variation in environmental quality (*E*) across units are often interpreted as marginal willingness to pay to improve *E*. This marginal price, β, clears the market when households choose among properties with varying environmental quality. Yet, many unobserved attributes of housing likely correlate with *E*. Repeat-sales using panel data can help researchers avoid attributing price effects of these unobservables to policy interventions. More importantly, as the results here suggest, even unbiased estimates of β may be inappropriate for predicting the price effects of a change in *E*. An estimated β that explains between-observation variation in price may be a poor predictor of within-observation price changes in response to changes in *E*. Shocks to *E* are likely to induce shifts in housing and other markets, and the joint determination of several important variables like price and neighborhood composition. An estimator that reflects the partial price effect, holding key neighborhood composition variables fixed, may overlook significant changes in those variables induced by the policy intervention.

Estimating richer models of the joint determination of prices, neighborhood composition, and environmental quality offers important insight into these indirect effects. How these indirect effects should be used in, say, a cost-benefit analysis depends on the context. If a clean-up attracts housing investment or high-income families, some of that investment and in-migration is coming at the expense of other areas. Thus, these indirect effects should be used judiciously by policy-makers interested in efficiency. More local interests may care less about effects in other areas or markets.

These findings have important implications for the environmental justice debate. EPA clean-ups induce neighborhood change, or sorting, as local housing markets adjust to changes in urban environments. In our most flexible model (4), remediations tended to occur in neighborhoods with rising income and more children. As wealthy families settle in, no longer deterred from living near a disamenity they were particularly sensitive to, housing prices rise still further as the neighborhood improves other dimensions. This adds to the literature on siting, and our empirical results contribute to the growing literature on neighborhood transition and waste sites (Ringquist 2006). It also suggests that attempts to target clean-ups at favored demographics might not lead to more just outcomes. As is common, individual and market behavior can undermine the best of policy intentions.

The present research invites further inquiry into simultaneous neighborhood and environmental change. A more robust system would allow for endogeneity in listing and remediation of NPL sites. A general equilibrium approach might also model other important markets, such as the labor market, to fully assess the expected price changes associated with remediation. Recent applications to air quality (e.g., Bayer et al. 2003, Smith et al. 2004) demonstrate the utility of general equilibrium models in examining joint environmental and neighborhood change. Certainly micro-level data would allow for more useful estimates and validation of our findings in local markets. Whether price effects of NPL sites vary across sites or metropolitan areas, perhaps using a random coefficients framework, warrants additional attention. Given the national scope of the policy and, at times, the date, a meta-analytic also approach my yield fruitful summaries of general trends in NPL impacts across various markets and regions. More generally, the approach taken here in response to challenges in using hedonic estimates for public policy can be applied to a variety of policy arenas.

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