# Willingness to pay for clean air:

# Evidence from quasi-experiment in Japan


**Abstract**

Over the past decades ambient concentrations of suspended particulate matter (SPM) have dramatically reduced in metropolitan areas in Japan. This paper analyzes a property-level panel dataset for 1992-2015 to measure the extent to which the improved air quality was capitalized into residential land prices. Single-equation estimates show no evidence of the capitalization. To address potential endogeneity biases, I instrument SPM concentrations with municipalities' designation status under the Automobile NOx/PM Law (ANPL), and find that the elasticity of residential land prices with respect to SPM concentration is -0.57. Using IV approach, I also find that the improvements in SPM concentrations increased inflows of migrations. These results are confirmed with a battery of robustness checks.



Keywords: Instrumental variable; hedonic pricing; urban air pollution; suspended particulate matter

JEL classification: Q51, Q53, R48

## 1. Introduction

Motor vehicles are a major source of urban air pollution across the world. Exposure to high concentrations of vehicular air pollution is associated with a wide range of health concerns including decreased lung function, increased risk of respiratory symptoms, and even premature death. Poor air quality also has adverse effects on long-term growth for local economies by keeping high-skilled human capital and firms away (Glaeser et al., 2001).

Old-vehicle scrappage programs have been introduced in many countries as a means to improve urban air quality. The reason for targeting old vehicles is that they account for a disproportionate share of vehicular emissions. Vehicles from older model years have less sophisticated (and also worn-out) pollution-control equipment. Removing the most polluting vehicles from the roads is thought to have a material positive effect on air quality.

Yet, little has been known about the environmental benefits of vehicle scrappage programs. Sandler (2012) analyzed a Vehicle Buyback Program in California for 1996–2010, finding that the program was effective in reducing vehicular emissions only in the short run. Li et al. (2013) examined the extent to which the 2009 US Cash-for-Clunker scheme reduced total gasoline consumption and carbon dioxide emissions, concluding that the scheme was not cost-effective. However, no prior research has explored some fundamental questions: do vehicle scrappage programs improve air quality and welfare, and by how much? This paper fills this gap by analyzing a diesel replacement program in Japan.

In 1992, Japan introduced an Automobile NOx Control Law (ANCL) to improve urban air quality.[1] The ultimate goal of the ANCL was for all monitors in designated areas to meet the national standard on nitrogen dioxide ($NO_2$). A key intervention to achieve this goal was to ban old diesel fleets that do not meet the emission standards under the ANCL from registering in designated areas after certain grace periods. The ban has applied to trucks, buses, special-use vehicles, and diesel passenger cars, particularly targeting old diesel trucks. The ban was introduced in a staggered way during 2004–2015, with the exact year of compliance being determined by the first year of registration of the vehicle. Owners of vehicles that are not in compliance could be sentenced to up to six months in jail, or be required to pay a fine of up to US$2,700.

Under the diesel replacement program, approximately 2.6 million non-compliant vehicles were replaced with newer vehicles (Iwata and Arimura, 2009). Trucks (56%) have been dominant, followed by diesel passenger cars (30%), special use vehicles (11%), and buses (2%). The Ministry of Economy, Trade and Industry (2008) suggested that 70% of replaced vehicles under the program were scrapped, and 28% were exported overseas, indicating that diesel clunkers were eliminated from the designated areas without substantially impairing air quality in other regions of Japan.

[1] NOx refers to nitrogen oxides, mainly consisting of nitric oxide (NO) and nitrogen dioxide ($NO_2$). We focus on $NO_2$, the most prevalent form of NOx in the atmosphere. Japan's national standard has been set in terms of $NO_2$.

The objective of the current paper is twofold. First, we estimate the causal effect of the diesel replacement program under the ANCL on ambient concentrations of $NO_2$. Second, we estimate the willingness to pay (WTP) for clean air under the framework of the hedonic pricing model. For the first task, we utilize a panel dataset at the monitor level for 1981–2015. Our identification strategy involves a difference-in-differences (DD) procedure: comparing temporal changes in air quality before and after the intervention between designated ("treatment") and non-designated ("control") areas. We also examine the dynamics of the program's effects. For the second task, we use a property-level panel dataset for 1992–2015 to estimate the residential land price-pollution gradient. Our instrumental variable approach exploits switchover of designation status under the ANCL to induce an exogenous change in $NO_2$ concentrations.

A key empirical issue is to establish a credible counterfactual air quality in the absence of the program. One challenge is that the intervention could influence air quality in the control group, as new compliant vehicles could be driven outside designated areas. Carefully examining the extent and scope of the spatial leakages of the intervention, we find that neighboring municipalities benefitted from the program *vis-à-vis* more distant municipalities.[2] Therefore, we exclude neighboring municipalities from the control group throughout the analysis, and assume that spatial leakages are negligibly small in other municipalities. This

[2] Details will be discussed in Section 3.

assumption is consistent with the freight survey by Tokyo Metropolitan Area Transport Planning Council (2005), which revealed that inter-regional transports by diesel fleets are limited in Japan.

Another challenge is to disentangle the effects of the diesel replacement program from the effects of other treatments and happenings affecting air quality in polluted areas. We focus on a Diesel Vehicle Driving Regulation enacted by the prefectures in the designated areas and other local measures such as clean vehicle promotions, traffic flow controls and public awareness campaigns that, to the best of our knowledge, pose the most significant threat. Through careful assessments, we find no evidence that these interventions had noticeable pollution-reducing effects.[3] Despite this, our estimates could be upwards biased given that it is impossible to fully control for other factors influencing air quality in the designated areas. Our estimates should thus be seen as an upper bound on the environmental benefits of the program.

We find that the diesel replacement program under the ANCL yielded considerable environmental benefits. The program effects kicked in three years after the ANCL was enacted in 1992 and accelerated over the next 10 years, peaking around 2004 when the first compliance requirements started. The program on average led to a 4.2% reduction in the annual mean ambient concentration of $NO_2$ for the years 1992–2015, equal to 11% of the total

[3] Details will be discussed in Section 4.

decrease. The program was also responsible for 24% of the total reduction in $NO_2$ concentrations at previously non-attained monitors in the designated areas. Our IV estimates suggest that the elasticity of residential land prices with respect to $NO_2$ concentrations is –0.55. This implies that the WTP for the improvements in air quality induced by the program was approximately US$11 billion. Utilizing Arimura and Iwata (2008)'s cost estimates, we conclude that the net benefit of the intervention under the ANCL was around US$3 billion.

Since "dieselgate" in September 2015, bans on diesel vehicles have very much been on the agenda, especially in European cities. In February 2018, a German court ruled that Stuttgart can ban diesel cars from driving in downtown areas to improve air quality (Bennhold, 2018). This ruling could ultimately roll out across Europe. The mayors of Paris, Madrid, and Athens have agreed to completely outlaw diesel vehicles from the centers of their cities by 2025. The United Kingdom has considered the introduction of a diesel scrappage scheme (Butcher, 2018). Japan's geographically-focused diesel replacement program offers one potential model.

## 2. Background

### *a. $NO_2$*

$NO_2$ is soluble in water, reddish-brown in colour, and a strong oxidant. On a global scale, emissions of $NO_2$ from natural sources (e.g., lightning, forest fires, and bacterial and volcanic activity) far outweigh man-made emissions. Because natural emissions are distributed over

the entire surface of the earth, however, background atmospheric concentrations are low. The major source of anthropogenic emissions of $NO_2$ into the atmosphere is the combustion of fossil fuels. In urban areas, automobile exhaust gas emissions are typically the major source (OECD, 2014).

Exposure to high concentrations of $NO_2$ can endanger human health and living environments. Recent studies of asthmatic children and adults have demonstrated associations between $NO_2$ and symptoms of asthma, reduced response to bronchodilators, and reductions in lung function (Guarnieri and Balmes, 2014). Epidemiological studies have revealed an association between exposure to high concentrations of $NO_2$ and mortality. Faustini et al. (2014) undertook a meta-analysis of 23 papers published during 2004–2011 and concluded that the long-term effects of $NO_2$ on mortality are as great as those of $PM_{2.5}$. $NO_2$ also plays an important role in the formation of acid rain and photochemical smog.

### *b. The intervention*

As a result of discontent about deteriorating urban air quality and the deregulation of the national $NO_2$ standard in 1978, in the 1980s and 1990s residents in some polluted areas sued the national government, expressway companies, and automakers for health damages induced from vehicular air pollution. The lawsuits led to official recognition that vehicular air pollution, including $NO_2$, was responsible for health damages. In some cases, the government

and expressway companies were ordered to implement measures to reduce emissions from road transport. In response, Japan's government enacted the ANCL in June 1992.

The ANCL stipulated that $NO_2$ must be below 0.06 ppm in terms of the $98^{th}$-percentile of the daily mean throughout the year for all monitors in the "designated" areas shown in Figure 1. As of 2015, 197 municipalities in Tokyo, Saitama, Chiba, Kanagawa, Osaka, Hyogo, Aichi, and Mie prefectures had been designated. This is not all of the municipalities in these prefectures. The designation criteria are (i) vehicle travel density, (ii) vehicle ownership rate, and (iii) ambient concentration of NOx. If all of these criteria were more than three times the national average, a municipality would be designated after hearings involving the prefectural governor and officials. Designations were enacted in 1992 and 2001 (only). Once imposed, no designation has been lifted.

<Figure 1>

The diesel replacement program was a key component of the ANCL. Diesel engines operate at a higher temperature and pressure than gasoline engines, generating more NOx. Despite their unwanted fumes, diesel is quite popular for trucks and buses due to superior performance in terms of fuel efficiency, durability, horsepower, and fuel expenses.[4] In

[4] In 2000, diesel accounted for 91% and 99% of total fuel consumption by trucks and buses in Japan, respectively (Ministry of Land, Infrastructure, Transport and Tourism, 2000).

addition, low turnover of trucks and buses mean that the fleets can consist of quite old vehicles. As of 2000, diesel vehicles accounted for 81% of NOx emissions from motor vehicles and were responsible for 52% of total man-made NOx emissions in the designated areas (Ministry of Environment, 2002).

The diesel replacement program was designed to ensure that the most polluted vehicles would be expelled from the designated areas. The ANCL has stipulated the NOx emission performance standard to be met for targeted vehicles. The emission standards for trucks, buses, and special-use vehicles are 0.48 g/km for gross vehicle weights of less than 1.7 tons, 0.63 g/km for 1.7–2.5tons, and 5.9 g/kWh for more than 2.5 tons, respectively. The emission standard for diesel passenger vehicles is 0.48 g/km regardless of weight. These emission standards have remained constant over time. Under the program, registration of vehicles that do not meet the ANCL emission standard would be banned after certain grace periods.

The registration restriction was introduced in a staggered way during 2004–2015, determined by the first year of registration of the vehicle. No compliance was required before the year 2004. The compliance years were arbitrarily set such that older vehicles have shorter grace periods. For example, non-compliant standard trucks first registered in 1998 or before were not allowed to be registered in designated areas as of 2004, whereas the compliance for those registered in 2002 was the year 2012.

### *c. Ambient concentration of $NO_2$*

Table I shows the average ambient concentration of $NO_2$ during 1981–2015 in the ANCL-designated areas. Column 1 shows that as of 2015, there was 623 air quality monitoring stations in these areas. Column 2–4 show that the designated areas are highly populated. As of 2015, 197 municipalities were designated. These municipalities accounted for 4% of Japan's land area and were home to 53 million people, 41% of the national total.

<Table I>

Columns 5–7 of Table I show the ambient concentration of $NO_2$ in terms of three different measurements. As of 2015, the annual mean, $98^{th}$-percentile daily average, and average hourly maximum were 0.017 ppm, 0.034 ppm, and 0.072 ppm, respectively. The WHO air quality guideline for $NO_2$ is 0.075 ppm in terms of the annual mean, indicating that the air quality in Japan's urban areas is now high. Column 8 shows that only one monitor remained non-attained as of 2015. The bottom part of the table demonstrates that air quality in the designated areas has improved over the post-intervention period.

## 3. Empirical strategy

### *a. Effects of the intervention on $NO_2$ concentration*

Our identification strategy involves a DD procedure, building off prior research (e.g., Wolff, 2014). We estimate the average treatment effect ($\alpha_1$) using the following specification:

$$\ln N_{m,t} = \alpha_0 + \alpha_1 Treatment_{c,t} + \delta \boldsymbol{X}_{m,t} + \rho \boldsymbol{P}_{c,t} + \theta_m + \gamma_t + \varepsilon_{m,t} \quad (1)$$

*m* stands for monitor, *c* stands for municipality, and *t* is year. *N* is $NO_2$ concentration. *Treatment* represents designation status under the ANCL: 1 if a municipality is designated, and 0 otherwise. $\boldsymbol{X}$ is a vector of monitor-level determinants such as weather conditions and monitor-specific time trends. $\boldsymbol{P}$ is a vector of municipality-level variables including socioeconomic and spatial factors. $\theta$ is monitor fixed effects, and $\gamma$ time fixed effects. $\varepsilon$ is an error term. We present heteroskedasticity-robust standard errors that are clustered at the municipality level to account for potential serial correlation (Bertrand et al., 2004).

The DD approach relies on two identification assumptions. One is the stable unit treatment value assumption. A key challenge is that new compliant vehicles are probably driven outside designated areas. Without considering this, $\alpha_1$ could be biased upward. In order to gauge the extent and scope of potential spatial leakages, we estimate equation (1) using control groups differentiated by their distance from designated areas. Table II reports the results. When all non-designated areas are included, the result suggests that the point estimate of Treatment is –0.024 at the 5% significance level. Next, we estimate a specification where the control group is restricted to municipalities contiguous to the designated areas only ("Neighbors 1"). The average treatment effect is zero, implying that $NO_2$ concentrations in Neighbors 1 declined at the same rate as the treatment groups as a result of the program. Thus, it appears that Neighbors 1 benefitted from spatial leakages. We also find a similar result in a specification

in which the control group is restricted to municipalities next to Neighbors 1 (called "Neighbors 2").

<Table II>

On the other hand, we find a noticeable average treatment effect (–0.031) for a specification in which the control group is restricted to municipalities that are next to Neighbors 2 (called "Neighbors 3"). This finding remains when more distant municipalities are utilized as the control group. Therefore, we exclude neighboring municipalities (Neighbors 1 and Neighbors 2) from the control groups throughout the analysis, and assume that spatial leakages are negligibly small in other municipalities.

This assumption is consistent with a freight survey by the Tokyo Metropolitan Area Transport Planning Council (2005) which provided large-scale data for freight travel between the origin and destination at the facility level. The survey covered Tokyo, Kanagawa, Saitama, and Chiba prefectures, comparable to the largest designated areas under the ANCL (Figure 1). It revealed that as of 2003, (i) intra-regional freight accounted for 79% of total freights originated from Tokyo metropolitan areas,[5] (ii) 94% of intra-regional freights were

[5] This is mainly because economic activities in those areas are highly concentrated. For example, as of 2001, one third of gross domestic product and household consumption was concentrated in the Tokyo metropolitan area. This area accounted for only 4% of the total land areas of Japan.

distributed by trucks, and (iii) marine and rail transports accounted for 30% of inter-regional freights. These indicate that inter-regional truck transport is limited, making spatial leakages less relevant.

Another identification assumption under the DD approach is the parallel trend assumption (PTA). Figure 2 shows the trends of annual mean ambient concentration of $NO_2$ in the treatment and control groups. The red dotted line is a counterfactual drawn by shifting the line for the control group for 1992–2015 upward in parallel. That $NO_2$ trends for the pre-intervention period between treatment and control groups seem to be common provides some support in favor of the PTA. However, there are other things that could affect pollution levels, besides the diesel replacement program. We are also concerned about the possibility that a convergence process may be at work (Stern et al., 2017). Given that the treatment group had higher initial $NO_2$ levels, their counterfactual $NO_2$ concentrations may be expected to decline at a faster rate than for the control group. We address these concerns in Section 4.

<Figure 2>

The discrete specification of equation (1) does not allow for estimating how quickly $NO_2$ concentrations declined after the law was enacted, and whether this impact increased, stabilized, or shrunk over time. It is also interesting to see how owners of non-compliant vehicles responded before the compliance started in 2004. Following Gehrsitz (2017), we estimate the following specification:

$$\ln N_{m,t} = \omega_t(Treatment_c \times \gamma_t) + \delta \boldsymbol{X}_{m,t} + \rho \boldsymbol{P}_{c,t} + \theta_m + \gamma_t + \varepsilon_{m,t} \quad (2)$$

where $Treatment_c$ is here a time-invariant dummy. The other elements are identical to equation (1). $\omega_t$ indicates the extent to which the $NO_2$ concentrations in designated municipalities relative to 1992 differed from that for the control group (i.e., no dummy interaction is included for 1992). A larger negative coefficient during the post-treatment period implies greater pollution-reducing effects of the intervention.

### *b. Effects of clean air on residential land prices*

Drawing on the hedonic pricing model proposed by Rosen (1974), we proceed to estimate the implicit price of air quality embedded in residential land values. The land price can be expressed as $P(\boldsymbol{z}) = P(z_1, z_2, \ldots\ldots, z_n)$, where $\boldsymbol{z}$ is a vector of attributes such as land characteristics (e.g., area), neighborhood factors (e.g., crime rate, school quality, access to train stations), and environmental amenities (e.g., air quality). The first-order condition in the hedonic set-up is that the marginal willingness to pay (MWTP) for the attribute will equal the marginal rate of substitution, holding all other attributes fixed.

We estimate the following hedonic pricing function:

$$\ln L_{h,t} = \beta_0 + \beta_1 \ln N_{h,t} + \gamma \boldsymbol{W}_{h,t} + \theta \boldsymbol{D}_{c,t} + \omega_h + \rho_t + u_{h,t} \quad (3)$$

where *h* is property, *c* is municipality, and *t* stands for year. *L* is residential land price. *N* is ambient $NO_2$ concentration. $\boldsymbol{W}$ is a vector of property-level dummies reflecting distance to the

nearest train station, area, use, building structure, utilities (e.g., water, gas, sewage), regulations, building-to-property site ratio, floor-to-area ratio, and property-specific time trends. $\boldsymbol{D}$ includes log per capita income and log population density at the municipality level. $\omega$ and $\rho$ are property- and year-fixed effects. $u$ is an error term.

Our focus is identifying $\beta_1$. A key identification issue is that omitted variables might influence $N$ and $L$ simultaneously. An important factor is the development of transport networks. Newly-built roads, for example, improve accessibility, likely resulting in an increase in property values and emissions from road transport. However, we are not able to control for the impacts of some key variables – such as vehicle flows – owing to data limitations.

To overcome potential sources of endogeneity, we instrument within-variation in log $N$ with designation status under the ANCL (*Treatment*) to induce exogenous changes in $NO_2$ concentration. Our IV strategy is similar to that of Chay and Greenstone (2005) and Bento et al. (2015), who employed nonattainment status under the Clean Air Act Amendments as an instrument for the temporal change in PM concentration. Other studies use alternative IV strategies. Examples include employment in agriculture and mining industries (Hanna, 2007), distant sources of pollution (Bayer et al., 2009), and flue gas desulphurization at power plants (Luechinger, 2009).

A potential concern is that the standard errors might be serially correlated. The simplest approach is a clustering adjustment. However, clustering in IV estimation makes asymptotic theory a poor guide to the finite-sample distribution (Cameron and Miller, 2015). Young (2017) studied a sample of 1400 IV regressions in 32 papers published in the journals of the American Economic Association and found that when standard errors are adjusted for clustering, the rejection rate of the null hypothesis at the 1% level for the IV estimator (30%) is much lower than that for OLS (51%).

Approaches to clustering standard errors vary among hedonics studies. Chay and Greenstone (2005) focused on non-clustered heteroskedastic-robust standard errors, whereas Bento et al. (2015) adjusted standard errors for clustering at the county level. Hanna (2007) and Bayer et al. (2009) reported non-clustered standard errors. We report both heteroskedastic-robust standard errors and robust standard errors clustered at the property level. Property-level clustering is motivated by the fact that the property fixed effects ($\omega_h$) in equation (3) control for correlation of errors between any properties, but not serial correlation for a property.

### *c. Variable construction and data*

Definitions and data sources are listed in an Appendix. Ambient concentration of $NO_2$ is recorded every hour throughout each year by either absorption photometry or chemiluminescence methods. The hourly data are converted to daily, monthly, and annual averages. The sample is limited to monitors that operated for at least 70% of the year.

Weather conditions include temperature, precipitation, and winds. We use the weather information from the nearest meteorological station to each pollution monitor. As of 2015, there were 127 meteorological stations across the country.

Socioeconomic controls include log household size, the log size of the manufacturing industry, the operation of new thermal power plants, log vehicle registrations, and log average incomes. As of 2013, 390 thermal power plants utilizing fossil fuel generation were operating in Japan. We use information on the date of the opening of each thermal power plant, assuming that they have continued to operate since that date. For a small number of controls, we interpolate (and if necessary, extrapolate) values for missing years.

Consideration of spatial factors is crucial, as improvements or degradations in air quality in one municipality may be partially attributed to spatial proximity to nearby municipalities, due to wind and other factors (Auffhammer et al., 2009). To deal with spatial issues, we include the natural log of neighbors' pollution ($NP$), defined as follows:

$$\ln NP_{c,t} = \ln \sum_{i=1}^{n} N_{i,t}^{c} \frac{B_i^c}{TB_c}$$

where $c$ stands for municipality, $i$ stands for the municipality contiguous to municipality $c$, and $t$ stands for year. $N$ is the annual mean of ambient $NO_2$ concentration for municipality $i$. $TB$ is the length of the total border of municipality $c$, and $B$ the length of the border between

municipalities $i$ and $c$. We use shapefile spatial data (as of October 2016) provided by the Environmental Systems Research Institute to calculate $B$ and $TB$.

For the analysis of the land price-pollution gradient, the main data source is the Published Land Prices (or *Koji-Chika*), compiled annually by the Ministry of Land, Infrastructure, Transport and Tourism (MLITT). Residential land prices per square meter are based on assessments by qualified real estate appraisers. A sample is chosen by the MLITT so that the land price represents the average price in the surrounding areas. We proxy the pollution level at the property by the reading at the nearest monitor for air pollution.

We use assessed land prices, rather than transaction prices, for several reasons. The MLITT record assessed land prices for the same properties repeatedly, enabling us to construct a longitudinal panel dataset. Second, assessed land prices seek to largely reflect transaction prices. Third, assessed land prices are widely analyzed in the literature (Tabuchi and Yoshida, 2000). Finally, transaction data are available only from 2005.

Our sample includes the period of Japan's municipal mergers, initiated in 1995. Throughout, we use data based on the municipal codes as of October 2016. To avoid the influence of overconfidence and speculation during Japan's real estate bubble period, we restrict the sample period to 1992–2015 when analyzing the land price-pollution gradient. We use balanced panels.

## 4. Results

### *a. Effects of the intervention on $NO_2$ concentration*

Table III reports the results for equation (1), for monitor-level panel datasets for 1981–2015. The overall goodness-of-fit of the regressions is 0.60–0.80, sufficient to conduct an econometric analysis. Column 1 shows that the point estimate for the treatment variable is –0.042. The 95% confidence interval ranges from –0.065 to –0.019. The results of the point estimate suggest that designation under the ANCL led to a reduction in annual mean ambient concentration of $NO_2$ of around 4.2% on average for the years 1992–2015. This equates to around 11% of the total decrease over the period.

<Table III>

The base of Table III reports the average treatment effects on the $98^{th}$-percentile daily average and hourly maximum of $NO_2$ concentration. The point estimate for the $98^{th}$-percentile daily average (–0.050) is of particular interest, because it allows us to assess the extent to which the intervention contributed to the ultimate target of the ANCL: all monitors in the designated areas meeting the national standard on $NO_2$. The estimate suggests that the intervention led to around 27 monitors that had previously violated the national standard becoming compliant over 1992–2015. This amounts to 24% of the total reduction in non-attained monitors in the designated areas.

The timings of the designation under the ANCL were 1992 and 2001. Column 2 of Table III reports separate effects for the two separate designation commencements. We find similar effects for each. Column 3 focuses on non-attained monitors. This is motivated by the fact that a majority of monitors in designated areas already complied with the national standard (see Table I). Without considering this, the true program effects could be potentially underestimated (Auffhammer et al., 2009). To examine heterogeneous treatment effects among monitors, we introduce an interaction term of Treatment with Nonattainment, a dummy variable taking the value one if a monitor violated the national standard as of 1991. The marginal effect for non-attained monitors (–0.076) is more than double that for attained monitors (–0.031).

Column 4 examines if the location of monitors is relevant for the size of average treatment effects. We interact Treatment with Roadside, a dummy taking the value one if a monitor is located roadside, and zero otherwise. It is expected that a larger effect would emerge for roadside monitors, as found by Wolff (2014) and Gehrsitz (2017) for Germany. In accordance with expectation, the point estimate suggests that the pollution-reducing effect for roadside monitors is 25% larger than that for monitors in the remaining areas. However, the difference is statistically insignificant.

Figure 3 reports estimates of equation (2). We see that air pollution was initially higher in designated areas *vis-à-vis* non-designated areas (relative to what was to prevail in 1992). The average coefficients of treatment leads between 1981 and 1988 is 0.042. Such differences disappeared a few years before the intervention.

<Figure 3>

Post-intervention, the estimates in Figure 3 suggest that the effects kicked in after three years and accelerated over the next 10 years, peaking around 2004 when the first compliance requirements started. This makes sense, as the owners of non-compliant vehicles did not have much incentive to replace their polluted cars during the early periods. It appears that they become more responsive to the program as the compliance requirements approached. Interestingly, the effects appear to have gradually diminished and stabilized after the compliance period began.

***b. Robustness***

Our baseline estimate suggests that the diesel replacement program on average improved air quality by 4.2% over the years 1992–2015. However, this estimate might be biased upward owing to the failure to disentangle the program effects from the effects of other treatments and happenings affecting air quality only for the designated areas. First, we directly estimate the impacts of a Diesel Vehicle Driving Regulation (DVDR) enacted by the prefectures in the

designated areas, which is probably the most significant threat. The DVDR prohibited diesel trucks and buses registered in non-designated areas from entering designated areas unless they met their new emission standards. The DVDR was enforced in 2003 for Tokyo, Saitama, Chiba, and Kanagawa, in 2004 for Hyogo, and in 2009 for Osaka.

For the examination of the influence of the DVDR, we take advantage of the fact that it was only introduced in some ANCL-designated areas, and adopt the following specification:

$$\ln N_{m,t} = \alpha_0^* + \alpha_1^* DVDR_{c,t} + \delta \boldsymbol{X}_{m,t} + \rho \boldsymbol{P}_{c,t} + \theta_m + \gamma_t + \varepsilon_{m,t} \quad (4)$$

where *m* stands for monitor, *c* stands for municipality, and *t* is year. *DVDR* is a dummy variable taking the value of one if a municipality is subject to the *DVDR*, and 0 otherwise. The other elements are identical to equation (1). The sample is restricted to the designated areas for 1992–2015. Bear in mind that in this analysis the treatment group is the designated areas where the DVDR was introduced, whereas the control group is the designated areas where the DVDR was not introduced. $\alpha_1^* < 0$ implies that the DVDR had pollution-reducing effects. Table IV shows the results. Overall, we find no evidence that the DVDR improved air quality. Pollution-reducing effects of the DVDR are not detectable, even if the equation is estimated by prefecture.

<Table IV>

Other local measures are also relevant to consider, including clean vehicle promotions, traffic flow controls, and public awareness campaigns. To address this concern, we utilize the fact that some urban cities were not designated under the ANCL, although they were as polluted as the designated areas. These polluted cities may well have undertaken similar other measures.[6] Thus, estimating the specification where the control group is restricted to these polluted cities allows for disentangling the program effects form the effects of local measures. We also undertake a placebo test where we pretend that non-designated urban cities were treated to see the extent to which the local measures improved air quality.

Columns 5 and 6 of Table III report the results. Column 5 restricts the control group to urban cities with more than 500,000 inhabitants as of 2015. We find that the point estimate hardly changes from the baseline estimate (Column 1). This result also suggests that our findings on the program effects on air quality are not likely to be biased by an ongoing convergence process, given that a similar result is obtained when focusing on municipalities with similar initial conditions. Column 6 restricts the sample to non-designated areas and uses a time-varying dummy (*Placebo*) that takes the value one from 1992 for municipalities that had more than 500,000 inhabitants as of 2015 (equivalent to the control group in column 5). We find no

[6] For example, Fukuoka, one of the non-designated urban cities, $NO_2$ level was 0.027 ppm as of 1992, equivalent to that in the designated areas, and two monitors in that city violated the national standard. To meet the national standard for all monitors, Fukuoka government had introduced clean vehicle promotions, traffic flow controls, and public awareness campaigns since 1993 (Fukuoka city government, 2014).

evidence of a negative effect for this placebo variable. The same conclusions are reached for the 98$^{th}$-percentile daily average and the hourly maximum measures.

Our estimates might be picking up a reduction in $NO_2$ emissions from industry, induced by government interventions in the designated areas. As far as we are aware, the most important intervention was NOx emission controls introduced in 1981 for designated plants in some industrial areas in Tokyo, Osaka, and Kanagawa. These are still in effect. However, we are not concerned about the influence of this intervention, as it was in place throughout our whole sample period. We are not aware of any other major interventions that aimed for a reduction in $NO_2$ emissions from stationary sources in the designated areas.

### *c. Effects on residential land prices*

Table V presents the results for single-equation estimation of equation (3). All specifications analyze property-level panel data covering 5,501 properties for 1992–2015, amounting to 105,548 observations. The overall goodness-of-fit of the regression is 0.70. Column 1 reports a specification in which year- and property-fixed effects are controlled for, column 2 accounts for property-specific time trends, and column 3 adds municipality covariates and time-variant property-specific factors. We report heteroscedasticity-robust standard errors for columns 1–3, and standard errors clustered at the property-level in column 4. These four specifications are used for IV and reduced-form estimations as well (Tables VI and VII).

<Table V>

The single-equation estimates in Table V suggest that capitalization of pollution in lower house prices is not detectable, similar to the findings of Chay and Greenstone (2005) and Hanna (2007). We find a *positive* coefficient for all specifications, although with small magnitudes (between 0.018 and 0.053). The estimates are consistent with arguments that ignoring the potential for endogenous pollution causes the estimate to be biased upward.

Table VI reports the results for IV estimation of equation (3). Panel A shows that the first stages of the IV regressions are satisfactory. As expected, the coefficients for *Treatment* are negative and statistically significant. The *F* statistic on the instrument exceeds 30 in all specifications, allowing us to reject the null hypothesis of the Stock-Yogo weak instrument test of 10% maximal IV size. Panel B shows that, contrary to single-equation estimates (Table V), the IV approach allows us to detect clear evidence of the land price-pollution gradient. Column 3 suggests that, holding other factors constant, a 1% decrease in $NO_2$ concentration is capitalized into a 0.55% increase in residential land prices. However, Column 4 shows that clustering substantially increases the standard error, leading to a failure to reject the hypothesis of no land price-pollution gradient at the conventional significance levels.

<Table VI>

Careful attention should be paid to the interpretation of our IV estimates. In this hedonic analysis, the sample periods have been restricted to 1992–2015 to circumvent the influence of Japan's earlier bubble economy. Our IV estimates thus identify the land price-pollution gradient only for the second treatment group designated in 2001 (see Figure 1). The IV estimates may well also be valid for the first treatment group, owing to the similar characteristics between the two groups. As explained in Section 2, the designation areas were chosen based on common criteria. In addition, as shown in Table III, we find evidence that both designations yield similar treatment effects.

Table VII reports the reduced-form estimates of the program effects on residential land prices. Overall, the results are consistent with the IV estimates. Designation status under the ANCL is associated with a 1–2.3% hike in residential land prices. These estimates are all statistically distinguishable from zero, except for Column 4. The regression fits (0.69–0.71) are better than those for IV estimates (0.50–0.65). The reduced-form estimates imply that the intervention improved air quality, resulting in the capitalization of air pollution into residential land prices.

<Table VII>

### *d. Benefits of the intervention*

Our estimates suggest that the diesel replacement program under the ANCL caused residential land prices in designated areas to increase by 0.24% on average. Assuming the rate is constant, we calculate the benefit of the intervention by taking the difference between mean land prices (in 2015 yen) in all designated areas and its counterfactual (the mean land prices multiplied by (1/1.0024)). Converting into US$ using the 2015 current exchange rate and discounting to the year 2004 using a discount rate of 3%, we obtain an estimate of the benefit of the intervention of US$4.2 per square meter.[7] The annual survey on property tax indicates that the total taxable land area for residences in the designated areas was 3.5 billion square meters. This implies that the WTP for the reduction in $NO_2$ concentration caused by the intervention under the ANCL was approximately US$11 billion.

The benefits stand against the cost of the intervention. The largest costs likely emanate from the early retirement of non-compliant vehicles. Arimura and Iwata (2008) analyzed administrative vehicle registration data to calculate the compliance costs incurred as a result of the intervention. They measured compliance costs by taking the difference between actual replacement costs and a counterfactual. They used disaggregated data to estimate the accelerated replacement period of the vehicles involved and monetize the counterfactual costs. Their estimates suggest that the aggregate costs of early replacement were around US$8

[7] For the purpose of comparison with Arimura and Iwata's cost estimates, we follow their discounting methods.

billion. When combined with our findings, we can conclude that the intervention under the ANCL produced net social benefits of around US$3 billion.

Our estimates have several limitations. Our hedonic analysis assumed that individuals are homogenous in tastes and income. In a heterogeneous population, the hedonic pricing function is not equivalent to the bid price function, potentially resulting in an overestimation of the MWTP (Kanemoto, 1988). Our analysis also assumed that individuals can move without cost. Improvements in air quality in designated areas might induce individuals residing in non-designated areas to move in. However, if movement involves costs, the MWTP could be underestimated (Bayer et al., 2009). Finally, our calculations do not consider benefits emanating from other factors, such as improvements in vehicle safety and comfortability.

## 5. Conclusion

Over the past decades, vehicle scrappage programs have been introduced in many countries. Recently, a nationwide diesel scrappage scheme has been seriously considered in the United Kingdom to mitigate air pollution. In this paper, we have examined whether vehicle scrappage programs can improve air quality and welfare by analyzing a diesel replacement program implemented in Japan under the 1992 Automobile NOx Control Law (ANCL). The results suggest that the intervention led to a 4.2% reduction in annual mean ambient concentration of $NO_2$, yielding net social benefits of around US$3 billion.

The unique feature of the diesel replacement program under the ANCL is that it takes a regulation-based approach: owners of polluted vehicles were required to replace earlier than they were intending through registration restrictions. This contrasts to incentive-based programs such as the 2009 US Cash-for-Clunkers scheme. The sizable environmental benefits of the ANCL found in this study imply that a regulation-based scrappage program can be worthwhile to consider, particularly when the abatement targets are geographically focused.

The ANCL has not required non-compliant vehicles to be scrapped, leading to a considerable amounts of clunkers being exported overseas. This may damage air quality in destination countries. Empirical evidence on clunker (and pollution) leakage effects to overseas jurisdictions via used-vehicle exports, has been quite limited (Davis and Kahn, 2010). Examination of pollution leakage across borders due to the ANCL would be an avenue for future research.

**Table I**
**Ambient concentration of $NO_2$ in designated areas**

| Year | Monitor | Municipality | Population (millions) | Area ($km^2$) | Ambient concentration (ppm) | | | Non-attained monitors |
|---|---|---|---|---|---|---|---|---|
| | | | | | Annual mean | $98^{th}$-percentile | Hourly maximum | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| 1981 | 439 | 163 | 40.9 | 12,359 | 0.026 | 0.051 | 0.129 | 106 |
| 1982 | 456 | 164 | 41.4 | 12,433 | 0.026 | 0.049 | 0.121 | 85 |
| 1983 | 473 | 165 | 41.8 | 12,463 | 0.025 | 0.047 | 0.114 | 69 |
| 1984 | 480 | 167 | 42.3 | 12,526 | 0.025 | 0.050 | 0.121 | 97 |
| 1985 | 491 | 171 | 42.9 | 12,641 | 0.025 | 0.047 | 0.115 | 73 |
| 1986 | 506 | 173 | 43.6 | 12,806 | 0.025 | 0.049 | 0.120 | 91 |
| 1987 | 520 | 176 | 44.2 | 13,031 | 0.027 | 0.055 | 0.131 | 168 |
| 1988 | 534 | 177 | 44.5 | 13,066 | 0.027 | 0.052 | 0.120 | 133 |
| 1989 | 544 | 178 | 44.9 | 13,155 | 0.028 | 0.053 | 0.119 | 150 |
| 1990 | 554 | 179 | 45.3 | 13,183 | 0.027 | 0.054 | 0.124 | 173 |
| 1991 | 564 | 181 | 45.7 | 13,381 | 0.029 | 0.055 | 0.124 | 174 |
| 1992 | 579 | 181 | 45.9 | 13,384 | 0.028 | 0.051 | 0.115 | 114 |
| 1993 | 590 | 181 | 46.1 | 13,390 | 0.028 | 0.052 | 0.117 | 155 |
| 1994 | 602 | 182 | 46.2 | 13,396 | 0.028 | 0.052 | 0.120 | 158 |
| 1995 | 612 | 184 | 46.6 | 13,487 | 0.029 | 0.052 | 0.124 | 128 |
| 1996 | 621 | 189 | 47.5 | 13,580 | 0.029 | 0.053 | 0.117 | 166 |
| 1997 | 629 | 192 | 48.2 | 13,636 | 0.029 | 0.054 | 0.117 | 182 |
| 1998 | 641 | 193 | 48.5 | 13,661 | 0.028 | 0.055 | 0.115 | 196 |
| 1999 | 650 | 194 | 49.0 | 13,676 | 0.027 | 0.049 | 0.102 | 91 |
| 2000 | 653 | 194 | 49.2 | 13,676 | 0.027 | 0.049 | 0.105 | 83 |
| 2001 | 655 | 193 | 49.3 | 13,665 | 0.027 | 0.049 | 0.105 | 87 |
| 2002 | 663 | 194 | 49.8 | 13,680 | 0.026 | 0.048 | 0.104 | 78 |
| 2003 | 666 | 193 | 49.8 | 13,671 | 0.026 | 0.046 | 0.098 | 53 |
| 2004 | 666 | 193 | 50.1 | 13,672 | 0.025 | 0.046 | 0.096 | 43 |
| 2005 | 671 | 193 | 50.3 | 13,675 | 0.024 | 0.044 | 0.099 | 35 |
| 2006 | 668 | 194 | 50.6 | 13,689 | 0.024 | 0.044 | 0.093 | 38 |
| 2007 | 660 | 194 | 50.9 | 13,690 | 0.022 | 0.042 | 0.087 | 22 |
| 2008 | 662 | 194 | 51.2 | 13,691 | 0.021 | 0.039 | 0.089 | 19 |
| 2009 | 665 | 194 | 51.7 | 13,687 | 0.020 | 0.039 | 0.083 | 17 |
| 2010 | 648 | 195 | 52.0 | 13,703 | 0.019 | 0.038 | 0.079 | 10 |
| 2011 | 643 | 195 | 52.1 | 13,704 | 0.019 | 0.037 | 0.076 | 3 |
| 2012 | 633 | 195 | 52.2 | 13,706 | 0.018 | 0.037 | 0.074 | 4 |
| 2013 | 628 | 196 | 52.4 | 13,748 | 0.018 | 0.037 | 0.077 | 4 |
| 2014 | 628 | 197 | 52.5 | 13,777 | 0.017 | 0.034 | 0.072 | 2 |
| 2015 | 623 | 197 | 52.7 | 13,777 | 0.017 | 0.034 | 0.072 | 1 |
| *Temporal changes during*: | | | | | | | | |
| Pre-intervention periods (1981–1991) | | | | | 9% | 7% | –3% | 64% |
| Post-intervention periods (1992–2015) | | | | | –39% | –34% | –37% | –99% |

*Notes*: The sample is limited to monitors that operated for at least 6,000 hours in a year. A monitor would be regarded as non-attained if ambient concentration of $NO_2$ exceeded 0.06 ppm in terms of daily average for $98^{th}$-percentile of days. See Figure 1 for the designated areas. ppm = parts per million.
*Sources*: National Institute for Environmental Studies, Ministry of Internal Affairs and Communications.

**Table II**

**Examination of spillover effects to neighboring municipalities**

| Dependent variable: Ln annual mean ambient $NO_2$ concentration $(ppm)_{m,t}$ | | | | |
|---|---|---|---|---|
| Control group: | Coefficient on $Treatment_{c,t}$ | Standard error | $R^2$ | Observations |
| All | –0.024** | (0.011) | 0.69 | 26,182 |
| *Only municipalities contiguous to:* | | | | |
| Designated areas (Neighbors 1) | 0.000 | (0.012) | 0.79 | 13,810 |
| Neighbors 1 (Neighbors 2) | –0.001 | (0.011) | 0.79 | 13,009 |
| Neighbors 2 (Neighbors 3) | –0.031** | (0.014) | 0.79 | 13,087 |
| Neighbors 3 (Neighbors 4) | –0.032*** | (0.012) | 0.80 | 12,614 |
| Remaining municipalities | –0.031*** | (0.011) | 0.70 | 21,102 |

*Notes*: *m*, *c*, and *t* stand for monitor, municipality, and year, respectively. Neighboring municipalities are mutually exclusive. Log wind, temperature, log precipitation, log passenger cars, log trucks, log household, log manufacturing employment, thermal power plants, log income per capita, log neighbors' pollution, year fixed effects, monitor fixed effects, and monitor-specific time trends are controlled for. All specifications are based on a balanced panel for 1981–2015. Standard errors are heteroscedasticity-robust and clustered at the municipality level.

***, **, and * indicate statistical significance at 1%, 5%, and 10%, respectively.

**Table III**
**Effects of the ANCL on $NO_2$ concentration**

| Dependent variable: Ln annual mean $NO_2$ concentration (ppm)$_{m,t}$ | | | | | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Treatment$_{c,t}$ | –0.042*** | | –0.031** | –0.040*** | –0.041*** | |
| | (0.011) | | (0.012) | (0.012) | (0.015) | |
| Treatment 1992$_{c,t}$ | | –0.045*** | | | | |
| | | (0.015) | | | | |
| Treatment 2001$_{c,t}$ | | –0.049*** | | | | |
| | | (0.014) | | | | |
| Treatment$_{c,t}$ × Nonattainment$_m$ | | | –0.045** | | | |
| | | | (0.018) | | | |
| Treatment$_{c,t}$ × Roadside$_m$ | | | | –0.010 | | |
| | | | | (0.016) | | |
| Placebo$_{c,t}$ | | | | | | 0.020 |
| | | | | | | (0.031) |
| $R^2$ | 0.68 | 0.68 | 0.68 | 0.68 | 0.79 | 0.58 |
| Year fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Monitor fixed effects | Yes | Yes | Yes | Yes | Yes | Yes |
| Monitor-specific time trends | Yes | Yes | Yes | Yes | Yes | Yes |
| Control variables | Yes | Yes | Yes | Yes | Yes | Yes |
| Monitors | 727 | 727 | 727 | 727 | 424 | 479 |
| Observations | 22,897 | 22,897 | 22,897 | 22,897 | 14,152 | 14,407 |
| Samples | Full | Full | Full | Full | Control group: Urban cities only | Non-designated areas only |
| *Average treatment effects when $NO_2$ concentration is measured by:* | | | | | | |
| 98$^{th}$-percentile daily average | –0.050*** | - | - | - | –0.053*** | 0.038 |
| Hourly maximum | –0.063*** | - | - | - | –0.056*** | 0.024 |

*Notes*: Subscripts *m*, *c*, and *t* stand for monitor, municipality, and year, respectively. All specifications utilize a balanced panel for 1981–2015. "Treatment" is a time-varying dummy that takes the value one if a municipality is designated under the ANCL. "Placebo" is a time-varying dummy variable that takes the value one from 1992 if a municipality in non-designated areas has more than 500,000 inhabitants as of 2015. "Nonattainment" is a time-invariant dummy that takes the value one if a monitor violated the national standard on $NO_2$ as of 1991. "Roadside" is a time-invariant dummy that takes the value one if a monitor is located roadside. Control variables include log wind, temperature, log precipitation, log passenger cars, log trucks, log household, log manufacturing employment, thermal power plants, log income per capita, and log neighbors' pollution. Standard errors are heteroscedasticity-robust and clustered at the municipality level.
***, **, and * indicate statistical significance at 1%, 5%, and 10%, respectively.

**Table IV**
**Impacts of the Diesel Vehicle Control Law**

| Dependent variable: Ln annual mean ambient $NO_2$ concentration $(ppm)_{m,t}$ | | | | |
|---|---|---|---|---|
| Treatment group: | Coefficient on $DVDR_{c,t}$ | Standard error | $R^2$ | Observations |
| All | 0.014 | 0.013 | 0.86 | 15,036 |
| Tokyo | 0.027 | 0.026 | 0.85 | 5,679 |
| Kanagawa | –0.010 | 0.028 | 0.86 | 5,806 |
| Saitama | 0.024 | 0.025 | 0.84 | 5,516 |
| Chiba | 0.036 | 0.022 | 0.81 | 5,842 |
| Osaka | 0.033 | 0.022 | 0.85 | 6,213 |
| Hyogo | 0.053** | 0.022 | 0.82 | 4,925 |

*Notes*: Subscripts *m*, *c* and *t* stand for monitor, municipality, and year, respectively. All specifications are based on a balanced panel for 1992–2015. Log wind, temperature, log precipitation, log passenger cars, log trucks, log household, log manufacturing employment, thermal power plants, log income per capita, log neighbors' pollution, year fixed effects, monitor fixed effects, and monitor-specific time trends are controlled for. Standard errors are heteroscedasticity-robust and clustered at the municipality level.
***, **, and * indicate statistical significance at 1%, 5%, and 10%, respectively.

**Table V**
**Results for single-equation estimation**

| Dependent variable: Ln residential average land price $(yen/m^2)_{h,t}$ | | | | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| Ln annual mean $NO_2$ concentration$(ppm)_{h,t}$ | 0.053*** | 0.053*** | 0.018*** | 0.018* |
| | (0.004) | (0.004) | (0.004) | (0.009) |
| $R^2$ | 0.69 | 0.69 | 0.71 | 0.71 |
| Year fixed effects | Yes | Yes | Yes | Yes |
| Property fixed effects | Yes | Yes | Yes | Yes |
| Property-specific time trends | No | Yes | Yes | Yes |
| Municipality covariates | No | No | Yes | Yes |
| Temporal change in property-specific factors | No | No | Yes | Yes |
| Properties | 5,501 | 5,501 | 5,501 | 5,501 |
| Observations | 105,548 | 105,548 | 105,548 | 105,548 |

*Notes*: Subscripts *h* and *t* stand for property and year. All specifications use property-level panel data for 1992–2015. Municipality covariates include log per capita income and log population density. Temporal change in property-specific factors relate to areas, use, building structure, utilities, distance to the nearest train station, regulation imposed on the property, building-to-property site ratio, and floor-to-area ratio. Columns 1–3 report heteroscedasticity-robust standard errors. Column 4 reports standard errors that are robust to heteroscedasticity and clustered at the property level.

***, **, and * indicate statistical significance at 1%, 5%, and 10%, respectively.

**Table VI**
**Instrumental variables estimates of effects of clean air on land prices**

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| *Panel A: First stage* | | | | |
| Dependent variable: Ln annual mean $NO_2$ concentration (ppm)$_{h,t}$ | | | | |
| Treatment$_{c,t}$ | –0.020*** | –0.020*** | –0.015*** | –0.015*** |
| | (0.004) | (0.004) | (0.004) | (0.004) |
| Cragg-Donald Wald *F* statistic | 58.61 | 55.87 | 34.96 | 34.96 |
| $R^2$ | 0.76 | 0.76 | 0.76 | 0.76 |
| *Panel B: Second stage* | | | | |
| Dependent variable: Ln residential land price (yen/$m^2$)$_{h,t}$ | | | | |
| Ln annual mean $NO_2$ concentration$_{h,t}$ | –0.931*** | –0.829*** | –0.549** | –0.549 |
| | (0.183) | (0.183) | (0.215) | (0.675) |
| $R^2$ | 0.50 | 0.53 | 0.65 | 0.65 |
| Year fixed effects | Yes | Yes | Yes | Yes |
| Property fixed effects | Yes | Yes | Yes | Yes |
| Property-specific time trends | No | Yes | Yes | Yes |
| Municipality covariates | No | No | Yes | Yes |
| Temporal change in property-specific factors | No | No | Yes | Yes |
| Properties | 5,501 | 5,501 | 5,501 | 5,501 |
| Observations | 105,548 | 105,548 | 105,548 | 105,548 |

*Notes*: Subscripts *h*, *c*, and *t* stand for property, municipality, and year, respectively. All specifications use property-level panel dataset for 1992–2015. The null of a weak instrument is rejected if the *F* statistic on the instrument exceeds the Stock–Yogo critical value (16.38). Municipality covariates include log per capita income and log population density. Temporal change in property-specific factors relate to land area, land use, building structure, utilities, distance to the nearest train station, regulation imposed on the land, building-to-land ratio, and floor-to-area ratio. Columns 1–3 report heteroscedasticity-robust standard errors. Column 4 reports standard errors that are robust to heteroscedasticity and clustered at the property level.
***, **, and * indicate statistical significance at 1%, 5%, and 10%, respectively.

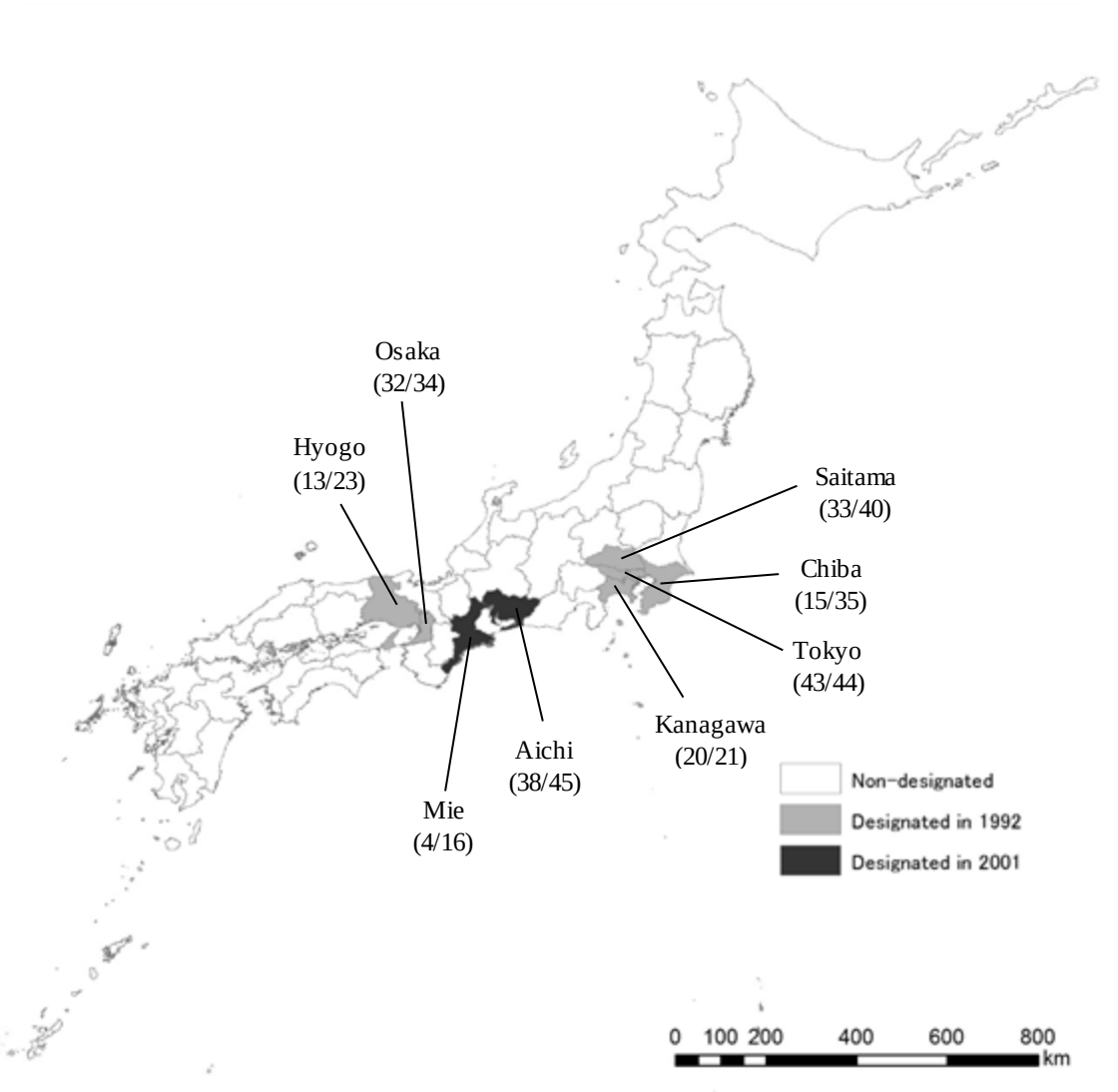
**Table VII**
**Reduced-form estimates of effects of the intervention on land prices**

| Dependent variable: Ln residential average land price (yen/m$^2$)$_{h,t}$ | | | | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| Treatment$_{c,t}$ | 0.023*** | 0.020*** | 0.010*** | 0.010 |
| | (0.004) | (0.004) | (0.004) | (0.012) |
| $R^2$ | 0.69 | 0.69 | 0.71 | 0.71 |
| Year fixed effects | Yes | Yes | Yes | Yes |
| Property fixed effects | Yes | Yes | Yes | Yes |
| Property-specific time trends | No | Yes | Yes | Yes |
| Municipality covariates | No | No | Yes | Yes |
| Temporal change in property-specific factors | No | No | Yes | Yes |
| Properties | 5,501 | 5,501 | 5,501 | 5,501 |
| Observations | 105,548 | 105,548 | 105,548 | 105,548 |

*Notes*: Subscripts *h* and *t* stand for property and year. All specifications use property-level panel data for 1992–2015. Municipality covariates include log per capita income and log population density. Temporal change in property-specific factors relate to areas, use, building structure, utilities, distance to the nearest train station, regulation imposed on the property, building-to-property site ratio, and floor-to-area ratio. Columns 1–3 report heteroscedasticity-robust standard errors. Column 4 reports standard errors that are robust to heteroscedasticity and clustered at the property level.

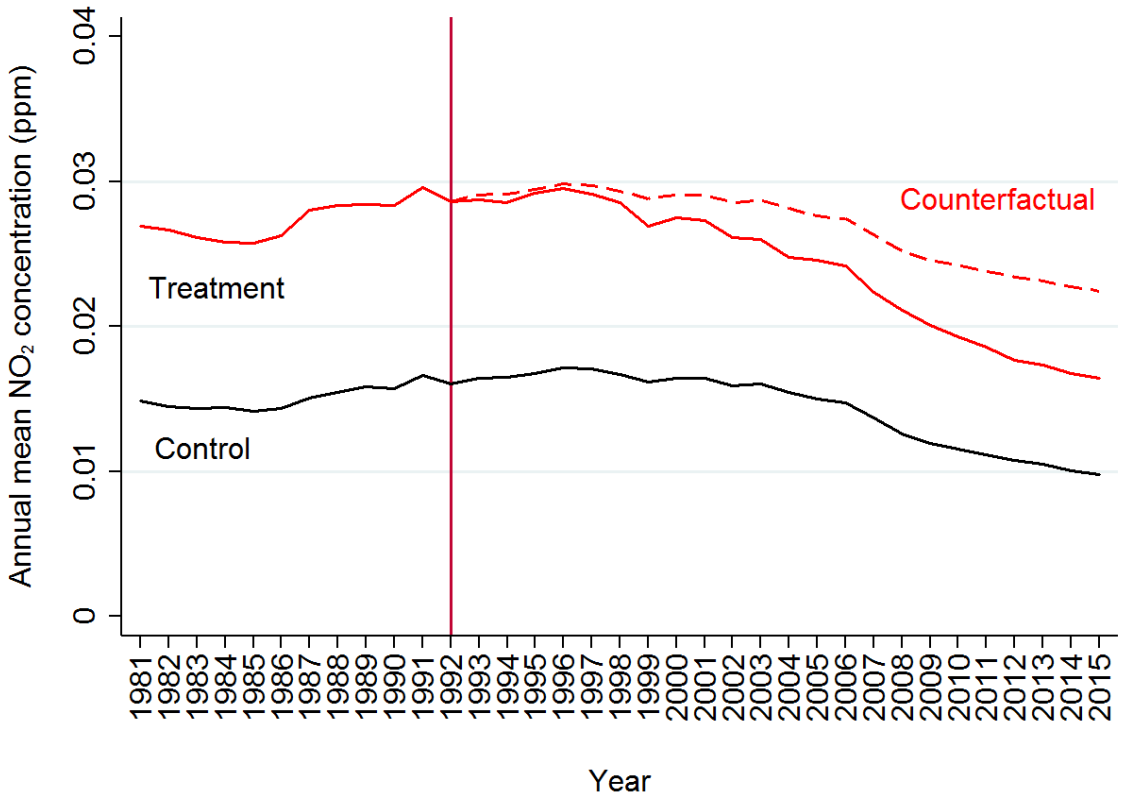
***, **, and * indicate statistical significance at 1%, 5%, and 10%, respectively.

**Figure 1**
**Designated areas under the Automobile NOx Control Law**



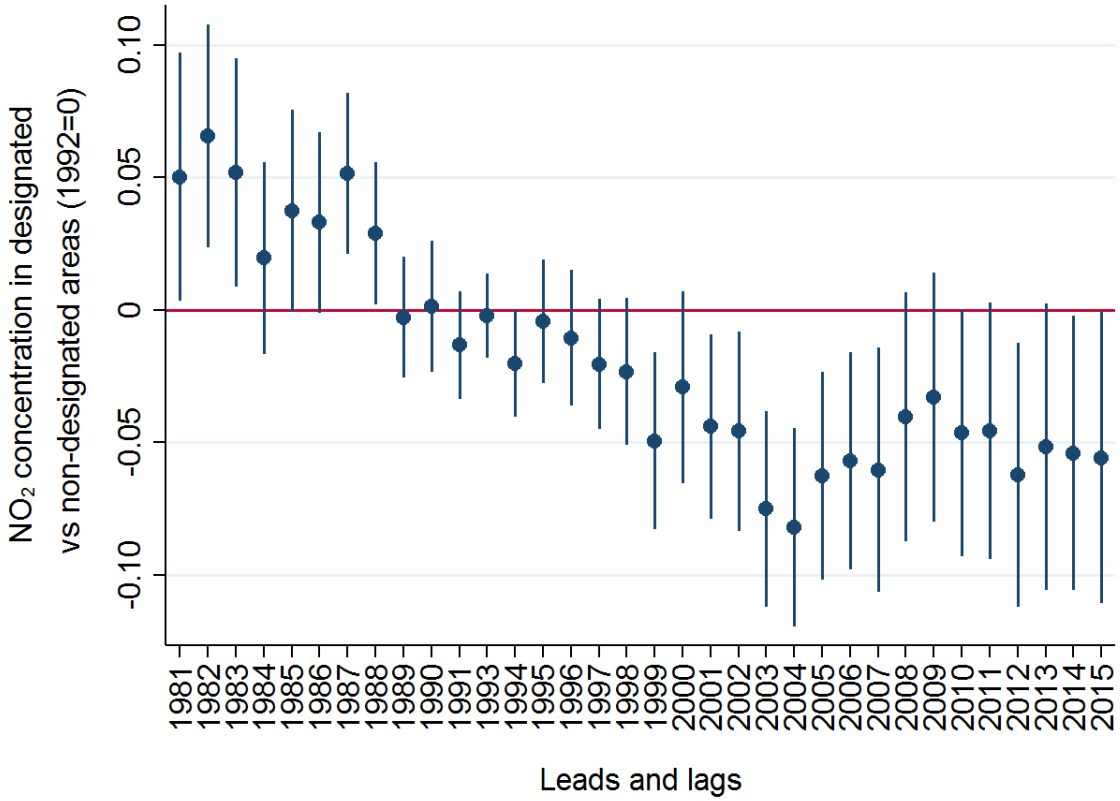

*Notes*: The first number in the brackets stands for the number of designated municipalities. The second refers to the number of monitored municipalities.

**Figure 2**
**Annual mean ambient concentration of $NO_2$, 1981–2015**




*Notes*: Sample is restricted to monitors in place and reporting for the full period 1981–2015. Red line shows the annual mean ambient concentration of monitors located in the designated areas. Black line shows the counterpart of monitors in the non-designated areas. Red dotted line represents a counterfactual drawn by shifting upward the line of control groups for 1992–2015 in parallel. The year 1992 is when the Automobile NOx Control Law (ANCL) was enacted.

**Figure 3**
**Estimates of lead and lag effects**




*Notes*: Estimation results for equation (2). Points show the point estimate of $\omega_t$ and vertical bands represent 95% confidential intervals. The dataset is a balanced panel covering 326 monitors for 1981–2015. Observations = 23,261. Other variables (log wind, temperature, log precipitation, log passenger cars, log trucks, log household, log manufacturing employment, thermal power plants, log income per capita, log neighbors' pollution, year fixed effects, monitor fixed effects, and monitor-specific time trends) are controlled for. Standard errors are heteroscedasticity-robust and clustered at the municipality level.

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## Appendix
## Variable descriptions

*Ambient concentration of $NO_2$*: Monitor-level $NO_2$ concentration measured by annual mean, $98^{th}$-percentile of daily average, and hourly maximum. Unit: parts per million (ppm). National Institute for Environmental Studies (NIES).

*Residential land price*: Property-level assessment value of land used for residence, yen/$m^2$. Ministry of Land, Infrastructure, Transport and Tourism (MLITT).

*Treatment*: Time-varying dummy variable, equal to 1 if municipality was designated under the ANCL. 0 otherwise. Ministry of the Environment.

*Temperature*: Average temperature measured every 10 minutes at meteorological station nearest to pollution monitor, degree Celsius. Japan Meteorological Agency (JMA).

*Wind*: Average wind velocity measured every 10 minutes at the nearest meteorological station, meters per second. JMA.

*Precipitation*: Total precipitation at the nearest meteorological station, millimeters. JMA.

*Income per capita*: Total taxable income divided by population in municipality, thousand yen. Ministry of Internal Affairs and Communications (MIAC).

*Household*: Number of households in municipality. MIAC.

*Manufacturing employment*: Number of people employed in secondary industry in a municipality. MIAC.

*Passenger cars*: Number of passenger vehicles registered as of March in a municipality. MLITT.

*Trucks*: Number of trucks registered as of March in a municipality. MLITT.

*Thermal power plants*: Time-varying dummy variable, equal to 1 if a municipality has one or more thermal power plant (fossil-fuel based) under operation. 0 otherwise. MLITT.

*Neighbors' pollution*: Average of annual mean $NO_2$ concentration for municipalities contiguous to municipality, ppm. Weighted by length of border in meters. NIES, Environmental Systems Research Institute.

*Population density*: Population per square meter of municipality. MIAC.

*Temporal change in property-specific factors*: Time-varying dummy variables, equal to 1 if property experiences a change from previous year in terms of distance to the nearest train station, areas, use, building structure, utilities, regulations, building-to-property site ratio, and floor-to-area ratio, respectively. 0 otherwise. Note that these are separate variables. MLITT.