LEED Adopters: Public Procurement and the Private Supply of "Green" Buildings*

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Governments often use their considerable purchasing power to promote policy objectives. However, it remains unclear whether or when public procurement preferences can stimulate private markets for innovative products and services. We examine municipal government procurement policies that mandate the construction of "green" buildings. We ask whether these policies aimed at government buildings actually influence private-sector building practices; specifically, the adoption of the U.S. Green Building Council's LEED standard. Using a combination of matching methods, panel-data and instrumental variables we find that municipal green building policies not only stimulate the supply of professionals investing in the required expertise, but also have spillover effects that spur private sector demand for green buildings. These findings suggest a link between product compatibility and environmental sustainability; specifically, new environmental standards may suffer from "excess inertia" if adoption requires coordinated investments, for example between builders and real estate professionals. Government procurement policies are one way to break these deadlocks.

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Governments increasingly use their purchasing power to promote environmental policy objectives. For example, the US Environmental Protection Agency (EPA) has developed environmentally preferable purchasing guidelines for goods ranging from paint, paper and cleaning supplies to lumber and electricity, and many state and local governments have taken similar steps.[1] The European Union's green public procurement policy is predicated on the rationale of stimulating market supply, noting that "a significant demand from public authorities for 'greener' goods will create or enlarge markets for environmentally friendly products and services…[and] stimulate the use of green standards in private procurement."[2]

Since government purchases account for 10-15 percent of GDP in developed countries, green procurement policies could have a substantial impact on the environment. However, many of these policies have a broader goal of encouraging private adoption of similar environmental standards and policies. In principle, governments' green procurement policies could lead to private adoption of similar standards, stimulate supplier markets for more environmentally benign products and generally encourage firms and individuals to prioritize environmental concerns in their procurement decisions. But do public procurement policies have these intended impacts, and if so why? To our knowledge, our study is the first to examine this question.

We examine the diffusion of the US Green Building Council's LEED standard for sustainable building practices. Specifically, we ask whether private-sector developers and real-estate professionals are more likely to seek LEED certification in cities with a municipal green building policy that applies *only* to new public buildings (i.e., a green procurement policy). We find that LEED adoption by real-estate developers is 80 percent

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[1] See, for example, the model Green Purchasing Guidelines circulated by the National Association of State Procurement Officials: http://www.naspo.org/content.cfm/id/green_guide. EPA Guidelines from Executive Order 13101.

[2] Commission of the European Communities. 2008. Public procurement for a better environment. Brussels, July 16. Communication from the commission to the European Parliament, the Council, the European Economic and Social Committee and the Committee of the Regions, page. 2. http://eur-lex.europa.eu/LexUriServ/LexUriServ.do?uri=COM:2008:0400:FIN:EN:PDF accessed January 2011

greater in municipalities with a public green-building policy than in a matched control sample of cities of similar size and demographic characteristics (including measures of "green preference" such as voting on environmental ballot initiatives and Toyota Prius ownership rates). Further analysis reveals that the impact of these municipal procurement policies on private sector procurement does not stop at the city line. Specifically, we find more LEED adoption among "neighbor cities" that border a city that adopts a green building policy, compared to these neighboring cities' own set of matched controls. The large magnitude of these neighbor-city effects suggest that private LEED adoption is not purely an effort to pre-empt local regulations, a motive described in Maxwell, Lyon and Hackett (2000), or seek favors in the in zoning and permitting process.

To explain the link between public green procurement policies and the diffusion of the LEED standard among private developers, we consider two possible mechanisms. First, government procurement may provide a boost in demand for the services of LEED accredited real-estate professionals (APs) – architects, contractors and other professionals who have passed an exam certifying their knowledge of LEED building principles – causing more professionals to seek that credential and reducing the marginal cost of LEED adoption for private developers. Alternatively, the market for green buildings may exhibit excess inertia (Farrell and Saloner 1986), where developers' willingness to adopt LEED depends on the availability of local real-estate professionals who are familiar with the standard, and local professionals' willingness to make LEED-specific investments is contingent on demand from developers. For cities stuck (perhaps temporarily) in a low-LEED equilibrium, government procurement policies may jump-start the development of specialized input markets by providing a guaranteed demand for LEED accredited real-professionals, thereby reducing the risk of investing in standard-specific human capital.

We find two pieces of evidence consistent with the coordination failure or excess inertia theory of spillovers from public to private LEED procurement. First, while real estate professionals could presumably cover the fixed costs of accreditation purely from private demand in larger cities, we find that the estimated impact of a public green procurement policy is increasing in city-size (measured as population or total non-residential construction activity). And second, we use green policy adoption in distant cities as an instrument for the number of LEED APs in nearby cities to estimate the causal impact of LEED APs on private developers' LEED adoption rates. The results suggest that the supply of LEED APs is a salient factor in private decisions to pursue LEED certification.

Our study is among the first to examine how the private sector responds to public green procurement policies. Corts (2010) shows that government procurement of "flex fuel" vehicles increases the supply of ethanol at local filling stations. We extend his results by asking whether the increased supply of complementary goods reflects coordination or capacity building, and by measuring the "same side" spillovers in private adoption of the government procured good (i.e. an increase in private LEED building that is analogous to increased private purchasing of "flex fuel" vehicles).

Overall, our findings suggest that government purchasing policies can stimulate private adoption of green building practices. While this could occur through a wide variety of mechanisms – such as increasing local awareness of the benefits of green building (i.e., moral suasion) or encouraging the adoption of a particular measurement system – we emphasize the idea that governments may break deadlocks that emerge when coordinated investments are required to adopt a common standard. That is, governments may take the role of "lead adopter."

The balance of the paper is organized as follows: Section I outlines a simple framework for analyzing the impact of green building procurement policies, and describes the USGBC Leadership in Energy and Environmental Design (LEED) standard. Section II describes our data, measures, and empirical methods. Section III describes the empirical results, and Section IV offers concluding remarks.

I. Background

A. Government Green Procurement Policies

Government purchasing guidelines often use price preferences or quantity targets (typically called set-asides) to reward products that meet environmental criteria such as incorporating recycled content, exhibiting pollution levels well below regulatory limits, or exceeding voluntary energy efficiency standards. When the government is a major customer, these policies can have a significant direct environmental impact due to the government's own procurement decisions. Governments can also use green procurement policies to signal concern for the environment when regulatory intervention is costly or infeasible.

When the government is not a major customer, the impact of government green-procurement policies will depend on how government purchasing interacts with private sector procurement decisions.[3] In practice, governments recognize this, and design policies that they hope will "influence the behavior of other socio-economic actors by setting the example, and by sending clear signals to the market-place" (OECD, 2000, p. 20). For example, when the Governor of Massachusetts established an environmental purchasing policy for state agencies in 2009, one of his stated objectives was to "encourage manufacturers and service providers to incorporate environmental and sustainability considerations into their products and operations locally, nationally, and even globally."[4] Similarly, one of the priorities of the United Kingdom government's "sustainable procurement" initiative is "stimulating the market to develop more sustainable solutions."[5]

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[3] Marron (2003) estimates that government purchases account for less than 20 percent total expenditures in *all* non-defense product categories.

[4] Deval L. Patrick. 2009. Executive Order No. 515 Establishing an Environmental Purchasing Policy, October 27. http://www.mass.gov/Agov3/docs/Executive%20Orders/executive_order_515.pdf

[5] United Kingdom Office of Government Commerce (OGC). 2010. Sustainable Procurement and Operations on the Government Estate: Government Delivery Plan Update – December 2009. http://www.ogc.gov.uk/documents/Delivery_plan_Dec09.pdf

In principle, the choice of government procurement rules can influence private purchasing behavior through either supply or demand channels. Moreover, the private response to government procurement rules might either reinforce or counteract the direct impacts of a government green purchasing policy. Figure 1 provides a simple framework, based on Marron (2003), for categorizing these potential impacts.

On the supply side, procurement policies can lead to greener private purchasing when there are significant scale economies in key input markets, so an initial government purchase reduces the average cost of serving the marginal private customer. Government purchasing preferences may also help suppliers overcome "chicken and egg" coordination problems by pushing the market towards a particular standard. For example, we show how municipal green-building policies led architects and builders to pursue LEED credentials, which presumably stimulates demand in other complementary markets (e.g., for energy- and water efficient products). This "coordination failure" story is based on the central two-sided markets assumption that private customers and input suppliers cannot somehow internalize the benefits of making a coordinated investment in the same platform, perhaps because of the risk that standard-specific investment will be stranded or under-utilized. Finally, government procurement rules may lead to increased competition and innovation on favored product attributes. For example, Siemens (2003) suggests that a preference for the Energy Star label in government computer purchasing led to increased innovation in energy efficient electronics.

There is no guarantee, however, green procurement policies will have a positive impact on private adoption. For instance, government procurement programs favoring environmental products may "crowd out" private purchases of these same products if supply is inelastic and consumers are willing to substitute to "browner" alternatives due to the price pressure caused by the government procurement program. Thus, the potential benefits associated with green procurement programs can be eroded or even eliminated in some settings.[6]

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[6] While we could find no procurement examples, there is some evidence that the supply of green power is inelastic, so government subsidies for green electricity are primarily spent on marketing and advertising these higher-priced services to end consumers, as opposed to investing in new generation facilities (Rader 2008).

On the demand side, private purchasing can reinforce green procurement rules when the government policy increases the visibility or credibility of a green product (or label), or if the policy sets a moral example that private purchasers choose to follow. We expect these "informational" demand-side effects to be most salient when the green product or label has minimal market share (so demonstration effects are particularly important) and when private customers already have other incentives to adopt greener products (e.g., because of energy-cost savings).

On the demand side, public procurement may crowd-out private demand if consumers come to perceive that the public sector is already "doing enough" to support the underlying policy goals (e.g. through minority set-asides or green procurement). Finally, when procurement rules define a sharp cutoff between "green" and "brown" products, and the cost of green provision is higher, private-sector supply of environmental goods may become concentrated just above that threshold. This seems especially likely when procurement policies are based on voluntary standards developed by firms with strong incentives to pre-empt more stringent regulation (Lyon and Maxwell 1999; King and Lenox 2000; Reid and Toffel 2009). Interestingly, this suggests that government purchasing rules should sometimes avoid specifying particular private standards, particularly in cases where there are questions about the motives of the standards developers or the stringency of the private certification.[7]

In practice, the importance of each of supply and demand-side channels described above will depend on specific features of the relevant product market. Our empirical analysis will focus on real estate development from 2001 to 2008. There are several reasons to expect that private purchasing will respond positively to a public green procurement policy in this setting.

First, government is an especially large customer in the market for real estate. Using BEA

[7] Cabral and Kretschmer (2007) develop a formal model of a similar story in the context of compatibility standards. Government agencies may even have strong incentives to adopt these voluntary consensus standards (see OMB Circular A-199 and NIST database).

data from 2002, Marron (2003) shows that 26.3 percent of all "Maintenance and Repair Construction" spending comes from federal, state, and local government. This product category receives the largest share of government procurement, except for munitions. Second, builders can realize direct benefits from green investments that produce energy savings or increases tenants' willingness-to-pay (Eichholtz, Kok and Quigley 2009). And third, our sample covers a period when LEED was just emerging as the dominant green building standard (see Figure 2). When there is no established standard, government procurement policies may help to solve coordination problems. Moreover, the literature on technology diffusion suggests that at the beginning of any S-curve the costs of adoption may be low (or even negative) for the marginal technology adopter.

While each of these factors suggests that we should observe a positive correlation between public procurement and private adoption of green building standards, they also suggest that we should be cautious about extrapolating our findings to settings with mature standards and technologies, few direct benefits or a small share of government purchases. Nevertheless, given the institutional characteristics of the market for green buildings between 2001- and 2008, we propose to empirically test the following set of predictions:

> DEVELOPER CERTIFICATION: Government green procurement policies will stimulate the adoption of green building practices by the private developers.

Note that this prediction could be true for a variety reasons, including demonstration effects, moral suasion, declining marginal costs of adoption (due to scale effects) or in response to anticipated regulatory changes. However, if government green building procurement policies exert no influence on the private costs and benefits of green building aside from favorable regulatory treatment (e.g., preferential treatment in zoning or municipal inspections), there should be no spillover to neighboring cities.

> PROFESSIONAL ACCREDITATION: Government LEED procurement policies will stimulate LEED accreditation by nearby real estate professionals.

If real-estate professionals are confident that their green-building human capital will be observed and rewarded in the market, they might invest in this know-how without any government encouragement or formal certification program. However, uncertainty about how and whether the market will observe and reward “green building” creates a possibility of stranded investment, and a possibility of government procurement spillovers.

> TWO-SIDED COORDINATION: If green building procurement policies solve a “chicken and egg” coordination problem between developers and real-estate professionals, the supply of LEED APs will have a positive causal impact on the rate of LEED registrations for new buildings.

Below, we develop a strategy for measuring the causal effect of accredited professionals on LEED registrations that uses policy adoption in distant cities as an instrument for the number of LEED APs in nearby cities.

B. LEED Certification & Accreditation

Leadership in Energy and Environmental Design (LEED) is a green building certification program developed and administered by the non-profit US Green Building Council (USGBC). The program was started in 1998, and initially focused on rating the environmental attributes of new construction. It has since added rating schemes for commercial and retail interior design, new homes, neighborhoods, and the renovation of existing buildings.

The LEED rating system for new buildings awards points for incorporating specific design elements or achieving environmental performance thresholds in eight project categories.[8] All certified projects must achieve a minimum number of points in each category and

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[8] The LEED point categories are: Location and Planning, Sustainable Sites, Water Efficiency, Energy and Atmosphere, Materials and Resources, Indoor Environmental Quality, Innovation and Design and Regional Priority.

increasing total point thresholds qualify projects for increasingly prestigious certification levels: Certified, Silver, Gold, and Platinum.

The costs of LEED certification will vary by type of project and certification level, and are primarily related to coordinating the required design elements and using more expensive materials and technologies. We could find no systematic data on these construction-related costs, which vary between buildings. Activities required to achieve some LEED points seem relatively cheap (e.g., installing bike racks), while others are quite expensive (e.g., remediating a brown-field site). The administrative costs of LEED certification are small by comparison, amounting to roughly $450-600 to register a project with USGBC and an additional $2,000 certification fee. Some developers also choose to hire a consultant to provide guidance on the LEED-eligibility of particular design choices and procurement decisions, and to prepare the LEED application.

For a commercial building, the benefits of LEED can accrue from reduced operating costs and/or increased rents and occupancy rates. Engineering estimates from a study of 121 LEED certified projects that volunteered data on energy use suggest these buildings consume 25-30 percent less energy than the national average for comparable projects (Turner and Frankel 2008). However, LEED certification emphasizes design elements rather than energy consumption, and several observers have suggested that more work is needed to understand whether LEED certified buildings actually deliver long-term environmental benefits. As for revenues, Eichholtz, Kok and Quigley (2009) find that LEED certified buildings charge 3% higher rents (with an additional 2.5% for Silver) and have higher sale prices and occupancy rates.

The LEED certification process begins with the developer registering a project with USGBC, which "serves as a declaration of intent to certify" the building, provides access to LEED information and tools, and lists the projects in the publicly-available online

LEED project database.[9] Once the construction or renovations have been completed, the certification application is submitted, reviewed, and approved, the applicant is sent a plaque (often displayed in the lobby in commercial buildings) and the project becomes eligible for inclusion in the online LEED database of certified projects.

While the LEED system debuted in 1998, it did not achieve significant scale until the second half of the 2000's. Figure 2 illustrates the number of new LEED registrations per year from 2000 to 2007. This figure only surpassed 1,000 annual registrations in 2005, and jumped to 4,000 in 2007 (the peak of the real estate cycle). The growth in LEED registrations reflects several factors, including the addition of new certification programs for additional building categories (e.g., homes and renovations), increased awareness of the program, and a growing installed base of LEED accredited professionals. Figure 2 also shows that federal, state, and local governments have been significant LEED adopters since the program began.

II. Data

To assess the impact of municipal green-building procurement policies on private LEED adoption, we collected data on 735 California cities from 2001 to 2008. We selected California because it has the largest economy of any US state, and the greatest number of cities that have adopted a green-building policy. Our dataset combines information from a variety of sources. We obtained measures of LEED diffusion from the USGBC, data on non-residential construction starts from McGraw Hill, city-level demographic data from the US Census, and we hand-collected data on the municipal adoption of green-building policies. Summary statistics are presented in Table 1.

Our main explanatory variable indicates whether a focal city (or a neighboring city that shares a common border) had adopted a municipal green building policy in the current

[9] Green Building Certification Institute, "LEED for New Construction: Registering a Project," http://www.gbci.org/main-nav/building-certification/certification-guide/LEED-for-New-Construction/Project-Registration/registration.aspx, accessed January 2011.

calendar year. We gathered information on the adoption of municipal green-building policies by hand, starting from lists compiled by the USGBC and the U.S. Department of Energy-funded Database of State Incentives for Renewables and Efficiency (DSIRE).[10] Our broad search identified 155 US cities that had adopted some type of green building ordinance by 2008. Forty of these municipalities were located in California, though we exclude from our analysis six cities whose regulations impose green building mandates on private-sector development.[11]

Municipal green-building policies vary in a number of dimensions, including the types of structures affected (by size, owner, and use); whether they cover new buildings or also renovations; and how they measure environmental performance. We gathered as much detail on individual policies as we could through city web sites and the online library of municipal codes.[12] Our research suggests that 87 percent of all green-building polices contained a purchasing rule: a requirement that new public projects adhere to some type of environmental standard, and ninety percent of these specified the LEED standard.

We code an indicator variable *Green Policy* to equal "1" if a city adopted a green procurement policy by 2008, and "0" otherwise. We code *Green Neighbor* to equal "1" for cities that do not adopt a green procurement policy but have an adjacent city (i.e., a city with a common border) that does adopt a green procurement policy, and "0" otherwise. Table 1 shows that four percent of the cities in our estimation sample adopted a municipal green building policy by 2008, and 15 percent of the cities in our data had a green neighbor.

A. Measurement

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[10] We acknowledge the excellent research assistance provided by Mark Stout in completing this task. The DSIRE list of state and local incentives is available at http://www.dsireusa.org/ and the USGBC list can be found at http://www.usgbc.org/PublicPolicy/SearchPublicPolicies.aspx?PageID=1776 .
[11] Table A1 provides a complete list of the adopter cities.
[12] Available at www.municode.com

We use two main outcome variables to measure the diffusion of LEED within the private sector, and one variable to measure government LEED adoption. All of our outcomes are based on data obtained from the US Green Building Council. Our unit of analysis is the city (or city-year), where we define cities in terms of a Census Place, which was chosen as the geographical unit that most closely resembles the political unit of a municipality.

LEED Registrations is a count of new privately-owned non-residential or mulit-unit residential buildings that registered for LEED certification between 2001 and 2008. This variable captures LEED adoption by local real-estate developers. Table 1 shows that re were between 0 and 99 *LEED Registrations* across all cities in our estimation sample (which excludes Los Angeles, San Francisco, San Diego and San Jose).[13] The average city in our sample saw two new LEED registered buildings over this time period. While *LEED Registrations* should reflect private developers' intention to use green-building practices, it is only the first step towards certification. The USGBC encourages projects to register early, since many decisions that will influence certification levels must be taken at early stages of the overall development process. Because the lag from registration to certification may be several years, and the LEED standard was diffusing rapidly toward the end of our sample period, a count of certified buildings would exclude a large number of projects in our data set.[14,15]

As a second outcome variable, we create a count of *Government Registrations*, in order to verify that municipal government green procurement policies actually lead to an increase in government LEED procurement. This variable is a count of new non-residential structures owned by a local government that registered for LEED certification between 2001 and 2008. The cities in our sample registered between zero and twelve new

[13] We exclude the four largest cities in California when calculating these summary statistics, since they could not be matched (and are therefore excluded from the analysis below) and tend to distort the sample averages due to their extreme size.
[14] For the buildings where we have certification data, the average lag between registration and certification is between 2 and 3 years. Anecdotal evidence suggest that few registered buildings fail to certify at some level.
[15] A second drawback of relying on *LEED Registrations* or *LEED certifications* is that they do not contain any information on the environmental impact of certification, a topic we leave to future research.

buildings, with an average of 0.3 LEED registered buildings per city between 2001 and 2008.

Our final outcome measure captures LEED-specific human capital investments by local real-estate professionals. *LEED Accredited Professionals* is a cumulative count of building industry professionals (e.g., architects and general contractors) who pass the USGBC's LEED accreditation exam between 2001 and 2008. This exam certifies that a real estate professional has knowledge of green building practices in general, and the LEED standard in particular. In 2004, it cost roughly $350 to take this test. We link new LEED APs to cities according to their business address maintained in the USGBC directory of LEED Accredited Professionals. By 2008, between zero and 416 LEED APs were located in the cities in our estimation sample, with an average of 7.5 accredited professionals per city.

Construction activity: To control for variation in the underlying rate of new building activity, we purchased data on new building starts from McGraw Hill's Dodge Construction Reports. The control variable *New Buildings* is a cumulative count of non-residential construction starts between 2003 and 2007 (the years we could afford to purchase). The mean number of new non-residential construction starts for a city in our estimation sample from 2003 to 2007 was 26.21. Since this variable is highly skewed and strongly correlated with city population ($\sigma = 0.88$), we also report the number of new buildings per capita in Table 1.

Demographics: For each city in the analysis, we collected from the 2000 U.S. Census[16] *Population* (measured in units of 10,000), *Income* (median household income in $10,000's), and *College* (the share of adults with some college education).

Environmental Preferences: In addition to these standard demographic variables, we collected several novel measures of a city's preference for environmental sustainability. First, we calculated *Green Ballot Share* as the share of citizens' own votes in favor of

[16] Matching political jurisdictions to census data was done at the level of the Census Place.

statewide ballot initiatives addressing environmental quality (Kahn 2002).[17] As indicated in Table 1, these ballot questions received support from an average of 61 percent of each city's citizenry. Second, we calculated the Toyota Prius market share in 2008 based on ZIP code level vehicle registration data from RL Polk (Kahn and Vaughn 2009). We aggregate these registration data to the city-level to create the variable *Prius2008,* which has a mean of 0.54 percent.[18] Finally, we use scores created by the League of Conservations voters for each member of the California State Senate and House of Representatives to proxy for the environmental preference of cities in their districts. These scores range from zero (poor performance on LCV issues) to one-hundred (strong alignment with LCV), with an average near 50 for both House and Senate across all cities in our estimation sample.

B. Matching and Covariate Balance

In the first part of our empirical analysis, we use the Coarsened Exact Matching (CEM) approach described by Iacus, King and Porro (2009) to examine the reduced-form impact of green procurement policies on *LEED Registrations* and *Accreditations*. This approach assumes that after stratifying and re-weighting the data to account for the distribution of observed exogenous variables, the endogenous treatment variables (i.e., *Green Policy* or *Green Neighbor*) are as good as randomly assigned. Intuitively, CEM is just a method of pre-processing a dataset before running a weighted least-squares regression. One begins by "coarsening" (discretizing) the variables in order to construct a multi-dimensional histogram. The next step is to discard observations from any cell that does not contain *both* treated and control observations. Finally, the units are weighted such that a weight of "1" is assigned to each treated unit, and T_i/C_i to each control observation in cell i (where T_i and C_i are the number of treatment and control observations in the i^{th} stratum of the multi-dimensional histogram respectively).

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[17] Using data from University of California's Statewide Database (http://swdb.berkeley.edu/), we calculated the proportion of votes in favor of various environmental ballot initiatives during 1996-2000 within the Census Place that best corresponded to each city.

[18] The highest Prius registration rate is 3.74 percent in Portola Valley (just west of Palo Alto).

Iacus, King and Porro (2009) describe several advantages of CEM. First, it is transparent and easier to implement than propensity score balancing. Second, CEM ensures that the re-weighted control sample matches *all* of the sample moments of the treated sample (not just the means). Third, applying CEM to a subset of observables will not lead to greater imbalance in other variables. Fourth, unlike conventional regression control methods, CEM does not rely on modeling assumptions to extrapolate counterfactual outcomes to regions of the parameter space where there are no data on controls. Finally, Monte Carlo tests and comparisons to experimental data suggest that CEM outperforms alternative matching estimators that rely on the same assumption of exogenous treatment conditional on observables.

We use CEM to construct two matched samples: one consisting of *Green Policy* adopters and their quasi-control group, and another consisting of *Green Neighbors* and their quasi-control group. For the adopters, we match on *Population* and *Prius2008,* which yields a matched group consisting of 25 adopters and 180 controls. While we would prefer to add more covariates to the matching procedure, this leads to a treatment sample with fewer than 20 cities and does not measurably improve the quality of the match. For the *Green Neighbors*, we match on *Population*, *Prius2008*, *New Buildings* and *Income*. Because the neighbor cities are smaller and more numerous, this more stringent CEM procedure still leaves an estimation sample of 81 *Green Neighbors* and 377 matched control cities.

Table 2 illustrates how CEM dramatically improves the balance in the means of the treatment and control samples. Each row in the table reports means for the treatment and control cities in a particular sample, and a T-statistic from regressing each covariate on the treatment dummy (i.e. *Green Policy* or *Green Neighbor*). The leftmost panel in Table 2 compares all cities that adopt a green-building policy (excluding the four largest) to the full set of potential controls (i.e., all other cities in California) using unweighted OLS regressions.[19] Not surprisingly, we find that cities adopting a green-building policy are

[19] All four of the largest cities in California (Los Angeles, San Diego, San Jose and San Francisco) adopted a green building procurement policy. Including these cities in the analysis leads to a dramatic increase in imbalance, and a similarly large increase in the results presented below.

larger, greener, wealthier, and better educated than the potential controls. There is a statistically significant difference in the means of each variable except for a per-capita measure of new construction activity.

The middle panel in Table 2 compares CEM weighted means for the matched sample of *Green Policy* adopters and their controls. Note that matching on *Population* and *Prius2008* excludes four cities from the treatment group, dropping its size to just 25 municipalities. Since we already dropped the four largest cities from our treatment group, these newly excluded cities were primarily municipalities with very high levels of Prius ownership (e.g., Berkeley and Santa Monica), as can be seen by the 0.15 percentage point drop in *Prius2008* among the adopters. Since we used the distribution of *Population* and *Prius2008* to construct the match, we should observe no difference in the means of these variables across treatment and control cities. However, Table 2 shows that matching on these two dimensions removes differences in the means of all observables across the two sub-samples.

The rightmost panel in Table 2 compares means for neighboring cities and their matched controls. The treated cities in this comparison are smaller and slightly less green than their neighbors who adopt a green building policy. Once again, the matching and reweighting removes observable differences in the means of most covariates. We do observe a statistically significant difference in the means of LCV Senate. However, it is not surprising that we should reject the null hypothesis of no difference in one of our 18 tests, and we find no difference in the means of *LCV House* or *Green Ballot Share*, our other political proxies for environmental preference.

C. LEED Diffusion

While the majority of our control variables are cross-sectional, it is possible to create panel data using the dates for policy adoption and the LEED outcome variables. Figure 3 illustrates how the CEM-weighted means of our main outcome variables evolved between 2001-2008. The figure consists of four bar graphs, with the top panel (Part A) showing trends for matched policy-adopters and their controls, and the bottom panel (Part B) illustrating the same trends for the matched neighbors and their controls.

All four of the graphs in Figure 3 illustrate the same rapid acceleration in LEED diffusion that we observed in Figure 2. And in all four cases, the effect is more pronounced for *Green Policy* adopters (or *Green Neighbors*) than for the matched control sample. (These patterns are even more striking if we do not use the CEM weights, since there are relatively more small cities in the matched control samples, and he weighting procedure makes these small markets less important). We also observe a small "bump" in accreditation for both treated and control cities in 2004, which was likely driven by anticipated changes in the USGBC exam that increased the costs of becoming a LEED AP. The next section of the paper shows that the different patterns observed for treatment and control cities in Figure 3 are statistically significant, using both cross-sectional and panel regressions, before turning to an instrumental variables framework to sort out the causal linkages between *LEED Registration* and *LEED Accreditation*.

III. Results

A. Cross-sectional Models

We begin our empirical analysis with a cross-sectional comparison of cumulative LEED adoption in the treatment and control cities. Our matching approach creates a matched group with treatment and control cities that are balanced with respect to all of the observable covariates we associate with policy adoption. Under the assumption that assignment to the treatment group is independent of potential outcomes conditional on observables, a simple t-test is sufficient to estimate the causal impact of the green building procurement policy. Since adding controls is more familiar, and may lead to increased precision, we use OLS regression instead of a t-test.[20] Specifically, we estimate the following linear regression:

$$\textbf{(1)} \qquad Y_i = \alpha_i + \beta \cdot GreenPolicy_i + \boldsymbol{\gamma} \cdot X_i + \varepsilon_i$$

where Y_i is the number of new *LEED Registrations* for non-municipal commercial buildings in city i in 2008, and X_i represents a set of controls (*Prius 2008, Green Ballot Share, Population, New Buildings, College,* and *Income*). As described above, the city-size and demographic variables were obtained from the 2000 Census. The Prius registration data are from 2008, while the green ballot measure is a cumulative score based on voting patterns across several prior years. We are interested in the coefficient β, which estimates the difference in the number of *LEED Registrations, Government Registrations* or *LEED Accreditations* in policy-adopting cities and in their control cities (or alternatively, between those in adjacent neighboring cities and their control cities).

We estimate this model using weighted OLS regression. As stressed in Angrist and Pishke (2008), OLS provides the best linear approximation to the conditional expectation function, regardless of the fact that Y_i is a count variable.[21] The results are presented in Table 3. The first three columns report estimates from weighted OLS regressions that compare CEM-matched *Green Policy* adopter cities to their controls. The outcome in the first column is *LEED Registrations*. We find a statistically significant increase of 7.8 registrations in cities with a green-building policy. Since the mean count of LEED Registrations is 7.9, this estimate represents a 98 percent increase in LEED adoption. One way to test the extent to which the matching procedure is working well is to assess whether the estimates are sensitive to the choice of control variables (as noted above, the controls are included mainly to increase precision). We find that the estimated treatment effect changes very little as we omit various groups of controls. Not surprisingly, we also

[20] It should be emphasized, however, that we do not use the control variables to extrapolate potential outcomes to regions of the parameter space where there are very few treated or untreated units.

[21] Estimating a model with an exponential conditional expectation function (i.e. Poisson with a robust co-variance matrix) produces similar results.

find that the effect grows significantly larger if we do not perform the CEM matching and weighting.

The second column of Table 3 shows that there are an average of 1.6 more Government LEED Registrations in cities adopting a green-building procurement policy. This is not surprising, since "green building" is the stated policy goal, and 90 percent of these policies use LEED as the relevant yardstick. Nevertheless, it is reassuring to see a large and statistically significant impact.

The third column in Table 3 shows that there is an increase of 13.09 *LEED Accreditations* in *Green Policy* adopting cities, relative to the matched controls. This corresponds to an increase of roughly 33 percent beyond the mean of 39.74, but is not statistically significant. Once again, the result is highly robust to specification, and grows large (and statistically significant) if we include the largest policy-adopting cities that do not have a CEM match. One reason why the *LEED Accreditation* result is weaker than the *LEED Registration* result may be that real estate professionals work out of surrounding communities (as we discuss in detail below).

The next three columns in Table 3 focus on cities that share a border with a *Green Policy* adopter, and report results of the matched *Green Neighbors* and their controls. We examine the policy impact on neighboring cities for three reasons. First, the neighboring city sample may address lingering concerns about omitted variables (e.g., tastes for greenness) that could influence both policy adoption and LEED diffusion. Second, the neighbors provide a larger and more representative sample of "treated" cities. Finally, the presence or absence of neighboring city effects is informative about the underlying mechanisms that link public green procurement policies to private adoption of LEED. In particular, if the effect of *Green Policy* adoption in adopting cities in mainly driven by unobserved (to the analyst) regulatory or zoning preferences for LEED projects, we would expect much smaller effects in adjacent neighbor cities that do not adopt a policy.

The third column in Table 3 shows a statistically significant increase of 0.8 *LEED Registrations* among neighbors relative to the matched controls. This translates into a marginal effect of 72 percent when normalized by the baseline registration rate of 1.11 buildings per year. This is quite close to the 84 percent marginal effect for *Green Policy* adopters based on our estimates in the first column.[22] Again, these results are robust to specification, and grow larger as we relax the matching criteria. The very similar marginal effects (of policy-creating cities and the neighboring cities) indicates that effects of the procurement policies do not stop at the city line. From these findings, we conclude that link between green building procurement policies and the private sector adoption of green buildings are not solely due to preferential treatment of green buildings by city-level zoning or permitting officials. Instead, our results imply that these procurement policies are more likely to be reducing marginal costs of green building, and the green building infrastructure that emerges (e.g., architects' and builders' increasing expertise in green building practices) benefits not only the policy-adopting city but also spills over to neighboring cities. This interpretation of the neighbor-city effects is also consistent with our finding (in column five) that there are positive *Government Registration* effects in neighboring cities that do not themselves adopt a green-building procurement policy, but who might respond to the emergence of a green-building infrastructure based on the LEED measurement system.

Finally, the rightmost column in Table 3 presents weighted OLS estimates of the impact of a *Green Neighbor* on *LEED Accreditations*. We find a statistically significant increase of 3.9 LEED APs, or roughly 74 percent. As noted above, if the market for architects and contractors is regional, these results may explain the weaker impact of policy adoption on *LEED Accreditations* in the policy-adopting cities.

B. Policy Adoption Hazards

[22] We also get similar estimates from unreported Poisson regressions, where the hazard ratio for *Green Policy* (column one) is 1.80, or an 80 percent marginal effect, and the hazard ratio for *Green Neighbor* (column four) is 1.67, or a 67 percent marginal effect.

One concern with a causal interpretation of the results in Table 3 is the potential for reverse causality: an active community of LEED APs, or a growing stock of LEED registered buildings, may promote the adoption of green-building procurement policies. We address this concern by running logit models of the hazard of policy adoption. If policy adoption is driving the differences reported in Table 3, we should find no correlation between the installed base of LEED APs or registered buildings and the adoption of a *Green Policy*. Table 4 reports estimates from the following specification:

(2) $\text{logit}(\Pr(\textit{Green Policy}))= \gamma + \lambda_t + \beta\, Y_{it} + X_i$

where Y_{it} is either *LEED Registrations* or *LEED Accreditations*, λ_t is a full set of calendar-year dummies and X_i is a vector of city-level controls. We estimate this model on the sample of adopters and matched controls, keeping city-year observations for *Green Policy* adopters during or prior to the year of the policy change, and defining a new dependent variable that equals one in the year the city adopts the green procurement policy.

The first two columns in Table 4 show results from a CEM-weighted logit model comparing policy adopters to their matched control cities. We find no evidence that of a large surge in LEED APs or registrations prior to Green Policy adoption. In fact, the only measure that is positively associated with policy adoption is an annual measure of construction activity. These results lend support to a causal interpretation of the estimates in Table 3, as opposed to stories of reverse causation (private LEED adoption sways public-policy makers) or regulatory capture (LEED APs lobby for a self-serving building code).

The third and fourth columns in Table 4 focus on cities that adopt a green-building procurement policy. For these cities, we find evidence that larger cities with a stronger taste for the environment (as measured by *Prius 2008*) adopt a green-building policy sooner. However, we find no correlations between the installed base of LEED buildings or LEED APs and the timing of policy adoption. The effects of *Population* and *Prius2008* in these regressions are interesting in its own right. In particular, given the evidence of spillover effects in Table 3, the hazard results suggest that promoters of green certification outside California could usefully focus on municipal governments in large "green" cities as a key constituency.

C. Panel Models

While our CEM matching strategy is fundamentally cross-sectional, we can nevertheless exploit the panel nature of the policy-adoption and outcome measures to estimate models that compare LEED diffusion in treatment and control cities, before versus after the adoption of a green-procurement policy. The results of this approach are presented in Table 5 and Figure 4.

The first panel in Table 5 (Part A) presents results from a serried of pooled cross sectional OLS regressions, using the following specification

$$\textbf{(3)}\ Y_{it} = \alpha + \beta_1\ GreenPolicyEver_i + \beta_2\ GreenPolicyNow_{it} + \boldsymbol{\gamma} \cdot X_i + \lambda_t + \varepsilon_{it}$$

We interpret β_1 as a selection effect that captures the average pre-treatment difference in outcomes between treatment and control cities, while β_2 is a marginal effect that reflects average difference after the green-building policy is adopted. In Table 5, we see that for all three outcome variables and both matched samples our estimates of β_1 are small and statistically indistinguishable from zero, while estimates of β_2 are large, positive and statistically significant. Overall, these results are consistent with patterns observed in Figure 3, where there is little difference between treatment and control cities in the early years of our sample because adoption rates were quite low, but LEED adoption accelerated more quickly for treated cities in the second half of our sample. The bottom panel in Table 5 (Part B) introduces a city fixed-effect, which absorbs both $GreenPolicyEver_i$ and all time-invariant city-level covariates X_i. However, this produces little or no change in our estimates of the marginal effect β_2 based on the pooled cross-sectional specification of equation (3).

Our last set of panel models relax the functional-form assumption that treatment leads to a one-time boost in the rate of LEED registration or accreditation. Specifically, we estimate the following flexibly parameterized model:

$$\textbf{(4)} \qquad Y_{it} = \alpha_i + \beta_y \cdot GreenPolicy_{iy} + \boldsymbol{\gamma} \cdot X_{it} + \delta_t + \varepsilon_{it}$$

where Y_{it} is the number of new *LEED Registrations* for non-municipal commercial buildings in city i in year t; α_i is a complete set of city-specific intercepts; δ_t is a complete set of year dummies to capture secular trends in LEED diffusion; and X_{it} are a set of exogenous time-varying controls (in our case, *New Buildings*, extrapolated to fill missing years in 2001, 2002 and 2008). We are interested in the β_y coefficients, which measure the change in *LEED Registrations* for policy-adopters (or adjacent neighbors) at y years from the adoption date. Since a complete set of β_y would be co-linear with city-specific intercepts for treated municipalities, we normalize the treatment and control cities to have identical fitted values (up to the fixed-effects) one-year before the policy is adopted by omitting β_{-1}. The CEM weights are retained in estimation.

We present the results of estimating (4) by graphing the β_y coefficients and their 95-percent confidence intervals in Figure 4. The top panel shows estimates for LEED Registrations in policy-adopting cities, and the bottom panel shows estimates for LEED Accreditations in neighboring cities. The flat line connecting years -5 to -1 in both figures once again illustrates that there is no evidence of a diverging trend in the treatment and control cities during the period before the policy change. In particular, we cannot reject the joint null hypothesis that all of the β_y coefficients for $y<0$ are jointly zero in either Part A ($p=0.69$) or Part B ($p=0.53$). This is a standard test of the diff-in-diffs maintained assumption that changes in the control sample means provide a valid estimate of counterfactual changes for the treated sample means. A second pattern revealed in Figure 4 is that treated cities gradually diverge from controls, rather than experiencing a sudden jump in LEED adoption following the policy change.

D. Instrumental Variables

Thus far, we have provided evidence that green-policy adoption is correlated with an increase both LEED registered buildings and LEED APs. This is consistent with the explanation that green-building procurement policies can break a deadlock among building professionals, who are reluctant to become APs without evidence of demand, and real-estate developers, who are reluctant to register for LEED unless there are professionals available. However, simply examining the reduced-form correlation between policy adoption, Registrations and APs obscures the underlying two-way causal links between APs and Registrations that are the basis of this explanation. This section proposes a set of instrumental variables that can isolate these relationships.

To estimate causal impact of LEED APs on *LEED Registrations* we need a variable that is correlated with the supply of APs and uncorrelated with unobserved factors linked to registrations. We propose to use green-policy adoption in "distant cities" as our instrument. Specifically, we instrument for the number of APs in all cities within 25 miles of a focal city using the number of Green Policies adopted by cities located between 25 and 50 miles away from the center of the focal city.

This instrument is motivated by the idea that markets for the services of real-estate professionals may be more diffuse than green tastes or policy impacts. Figure 5 provides some evidence on the spatial distribution of this market in the form of a histogram of the distance between architects and general contractors and the projects they work on, based on our McGraw Hill project level construction starts data. While both contractors and architects tend to work on projects close to their office address, the median project distance is 28 miles, and the 75^{th} percentile of the project-distance distribution is roughly 75 miles. Thus, we might expect local professionals to respond to distant green building policies that have no direct impact on the decisions of real-estate developers, other than through the supply of LEED APs.

To isolate the impact of LEED Registrations on APs, we require a variable that is correlated with the number of LEED projects, but uncorrelated with unobserved drivers of local real-estate professionals' decisions to seek accreditation. Building on the ideas in

Corts (2010), we propose using *New Buildings* as an instrument for *LEED Registrations*. The underlying idea is that as the number of buildings increases, so does the probability of having one or more LEED projects that could stimulate the local LEED AP accreditation rate. Since the number of new buildings is clearly exogenous to the decision to seek LEED accreditation, our main concern with this approach is the level of real estate activity may be correlated with other factors (e.g. city size or tastes for the environment) that enter the *Accreditations* decision. However, since our estimates are conditional on *Population* and *Prius 2008*, the key assumption is that variation in the *intensity* of development between 2003 and 2007 increases the rate of *LEED Registration* (e.g. because of competition among developers) without otherwise altering the incentive to seek LEED accreditation.

Table 6 presents our instrumental variables results. All of the IV models contain unreported controls for *Population*, *Income*, *College*, *Prius 2008* and *Green Ballot Share*.The leftmost panel uses our first IV strategy (i.e. using policy adoption as an instrument for *LEED Accreditations*) on the sample of neighbors and matched controls. In the first column, we report OLS estimates of the correlation between neighbor city LEED APs in adjacent neighbor cities *LEED Registrations* in a focal city. The second column uses *Green Neighbor* as an instrument of the supply of APs. We find a very strong first-stage relationship between policy adoption and the supply of APs, and no change relative to the OLS estimate for the relationship between APs and Registrations.

Columns 3 and 4 in Table 6 use the estimation strategy described above, where we instrument for neighboring city with distant green policy adoption. In column 3, we see that the correlation between LEED APs within 25 miles of a focal city and *LEED Registrations*, is somewhat larger than the neighbor city correlation reported in column 1. This is consistent with our idea of a diffuse market for real-estate services. Column 4 presents our IV estimates, which show a strong first-stage correlation between distant green policies and nearby *LEED Accreditations*, and a strong positive impact of nearby APs on *LEED Registrations*.

The last two columns in Table 6 examine the causal impact of *LEED Registrations* on *LEED Accreditations* using *New Buildings* as an instrument for Registrations. Once again, we find a strong first-stage relationship, and a positive impact of LEED building rates on the supply of APs. As with the earlier sets of IV results, we cannot reject the hypothesis that a simple OLS regression would provide unbiased estimates of the underlying structural parameters.

The goal of our instrumental variables analyses was to provide further support for the hypothesis that green-building procurement policies can stimulate private demand for green building by helping local markets over come "excess inertia" that occurs when neither developers nor local real-estate professionals wish to make the first investment in a new standard. We found evidence of a positive causal relationship between the supply of LEED APs and the rate of LEED Registration, and similarly in the other direction. This is a necessary condition for the existence of a "chicken and egg" dilemma in new standards adoption. Moreover, the results showing how distant green procurement policies can influence local LEED Registration rates through the supply of nearby APs point to the importance of supply-side spillovers in the diffusion of LEED.

E. Coordination vs. capacity-Building

Though we have focused on the idea that government procurement policies that incorpoate LEED can foster coordination between private developers and LEED APs, an alternative story that involves no underlying coordination problem is that government adoption simply provides sufficient demand that real-estate professionals can cover the fixed costs of accreditation. This explanation suggests that the marginal effect of a municipal procurement policy should be declining in city-size, since private demand for LEED will be more likely to cover these fixed costs in larger markets. Thus, the first four columns in Table 7 interact two measure of market size (*Population* and *New Buildings*) with the *Green Policy* and *Green Neighbor* treatment dummies in an OLS regression with the number of LEED APs as the outcome. All four models suggest that the impact of policy adoption on professional accreditation is actually increasing with city-size. This is consistent with our coordination story, or demand-side explanations that emphasize awareness or learning, but not the fixed-cost capacity-building explanation.

In the last two columns of Table 7, we use our distant policy adoption instruments to examine how the impact of APs on *LEED Registrations* varies with city-size. Once again, our results suggest that the marginal effect of an increase in LEED APs is increasing in market size, which is inconsistent with a story where the main effect of government green procurement is to help early APs cover their fixed costs. Overall, we take the results in Table 7 as evidence in support of the idea that government green building procurement policies help jump-start both the demand for and supply-of specialized inputs (i.e. LEED APs) to the green-building industry.

F. Limitations

In considering the generalizability of our results, it is important to consider two factors. First, we find that government procurement rules stimulated a market, construction services, where governments are an especially large purchaser. Further research is required to examine the extent to which such rules exhibit similar effects in other markets where governments represent a smaller share of demand. Second, since LEED was just emerging as the *de facto* standard for green-building certification during our sample period, the marginal private adopter is likely to have high marginal benefits (or low costs) of going green once verification becomes possible. Thus, we would expect to find similar market-stimulating effects of government procurement of other goods and services in contexts where common standards have yet to emerge.

6. Conclusions

This paper provides evidence that public procurement policies can influence private sector purchasing decisions in a way that reinforces underlying policy goals. Given the relative scale of public and private purchasing, this may be a necessary condition for public procurement guidelines to have substantive impacts (e.g., on the scale of regulatory policy).

While there is a substantial economic literature asking whether public investments "crowd out" private spending (e.g. Goolsbee (2000) on government R&D, or Hoxby (1996) on public and private education), we find few studies of government spending "crowding in" private investment by, for example, acting as a focal adopter that "tips" the market towards a particular standard or certification scheme. Yet that is often a goal of socially motivated procurement policies, such as "buy green" initiatives.

This paper looks for evidence of "crowding in" using data on private-sector diffusion of the LEED green-building certification program following the adoption of municipal bylaws that require public construction to follow green building practices (i.e. green-building procurement policies). This is admittedly a case where one might expect such reinforcing spillover effects, since LEED was rapidly emerging as the *de facto* standard for green building certification, and many private developers could reasonably expect that "green building" would yield direct economic benefits in the form of energy savings and increased demand. Nevertheless, we find that cities with a municipal green-building policy have roughly 80 percent more LEED registrations by 2008, compared to a matched control sample that has similar size, demographics and tastes for environmentalism (as proxied by voting behavior and Toyota Prius ownership).

Our analysis is subject to several caveats. First, Despite our efforts to construct a well-matched control sample using the new methods developed by Iacus et al (2009), there is clearly room a concern that our estimates are biased upwards because of an omitted taste for greenness that is correlated with both municipal procurement policies and private LEED adoption. However, we are somewhat comforted by finding similar "crowding in" effects in a sample of neighboring cities that do not themselves adopt a green-building policy. We also check and find no evidence of a divergence in LEED adoption between ""treated" cities (either policy adopters or adjacent neighbors) and their matched controls prior to the change in procurement policy. These findings provide evidence against stories of reverse causation or policy adoption by municipalities that are "captured" by greener elements of the real estate profession. Our preferred explanation for the papers main results is that green procurement policies produced a combination of moral suasion, increased awareness and fostered the development of complementary markets for specialized inputs (i.e. LEED APs).

As second caveat for this study is that we do not measure the environmental impacts of increased LEED adoption (or even the final certification of all registered buildings). Engineering studies suggest that LEED certification is correlated with increased energy efficiency. However, those estimates are based on data from a self-selected sample of LEED certified buildings. We hope to extend this research by examining the impact of public green-building policies on certification levels and perhaps environmental impacts.

Finally, since our findings suggest that government procurement policies can catalyze the adoption of a privately developed certification scheme, one might ask whether governments typically choose "the right" standard? In the case of LEED, it is not clear whether municipal green-building policies promoted lock-in to a particular standard (the leading alternative was the EPA's Energy Star label), or increasing returns simply led private and public actors to coalesce around the most popular measurement system at the time. Nevertheless, our *LEED Accreditation* results show that government purchasing policies can promote standard-specific investments by various third parties (e.g. architects, contractors and suppliers of green building materials). This both points to procurement policies as an effective policy tool, and highlights the potential dangers of lock-in to a government-selected standard (particularly if it was developed by firms hoping to pre-empt more stringent regulation). The question of how government should be involved in the *ex ante* development of voluntary standards that might later provide the basis for procurement policies is an intriguing topic for future research.

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Table 1. Summary Statistics

Variable Name	Definition	Mean	SD	Min	Max
LEED Registrations	Private LEED Registered Buildings (2001-08 cumulative)	1.93	6.25	0.00	99
LEED Accreditations	LEED Accredited Professionals (2001-08 cumulative)	7.51	27.38	0.00	416
Gov't Registrations	Gov't LEED Registered Buildings (2001-08 cumulative)	0.29	0.94	0.00	12
Green Policy	Focal City Adopted Green Building Policy (dummy)	0.04	0.19	0.00	1.00
Green Neighbor	Adjacent City Adopted Green Building Policy (dummy)	0.15	0.36	0.00	1.00
Prius 2008	Toyota Prius as percent of all car registrations	0.54	0.59	0.00	3.74
Green Ballot Share	Mean voter share supporting green ballot measures	0.61	0.15	0.20	1.00
LCV Senate	State Senator's LCV[1] score	46.54	40.88	0.00	100
LCV House	State Representative's LCV[1] score	49.56	39.97	0.00	100
Population	City Population (10,000's)	2.99	5.10	0.00	46.15
New Buildings	Non-residential construction starts (2003-07 cumulative)	26.21	54.71	0.00	869
Buildings per Capita	New Buildings / Population	12.06	18.42	0.00	204
College	Percent college educated	0.23	0.17	0.01	0.89
Income	Median household income	4.80	2.17	0.00	20.00

Notes: Summary statistics for cross-section of 735 California cities (excluding Los Angeles, San Diego, San Jose and San Francisco). [1]LCV = League of Conservation Voters.

Table 2. Covariate Balance in Full and Matched Samples

	Full Sample No weights			Matched Sample Weighted			Matched Sample Weighted		
	Adopters	*Others*	*T-stat*	*Adopters*	*Controls*	*T-stat*	*Neighbors*	*Controls*	*T-stat*
Prius 2008	0.93	0.53	3.62	0.78	0.76	0.17	0.66	0.66	0.04
Green Ballot Share	0.72	0.60	4.35	0.71	0.68	1.02	0.69	0.66	1.71
LCV Senate	68.69	45.58	3.00	67.72	59.43	0.81	67.38	55.10	2.11
LCV House	69.00	48.53	2.72	64.24	60.05	0.34	64.75	59.52	0.81
Population	14.36	2.53	13.68	13.91	13.76	0.04	4.18	4.10	0.13
New Buildings	140.79	21.59	12.64	142.4	108.92	0.83	23.09	23.2	0.03
Buildings per Capita	10.62	12.20	0.45	10.91	9.67	0.71	7.93	7.28	0.42
College	0.35	0.22	4.09	0.34	0.33	0.13	0.3	0.27	1.13
Income	5.58	4.77	1.97	5.72	5.77	0.12	5.57	5.51	0.17
Cities	29	697		25	177		81	372	

Notes: Left panel reports means of each covariate and T-statistic from unweighted OLS regression of X on *Green Policy* dummy. Middle and right panels columns CEM-weighted means of each covariate and T-statistic from CEM-weighted OLS regression of X on *Green Policy* dummy (middle panel) or *Green Neighbor* dummy (right panel). CEM weights are described in Iacus, King and Porro (2009) and discussed in text.

Table 3. Effects of Green Building Procurement Policies on LEED Registrations

	Adopters and Matched Controls			Neighbors and Matched Controls		
Outcome	*Registrations*	*Gov't*	*Accreditations*	*Registrations*	*Gov't*	*Accreditations*
Green Policy	7.80	1.59	13.09			
	(3.49)**	(0.48)***	(15.68)			
Green Neighbor				0.80	0.26	3.90
				(0.36)**	(0.10)**	(1.26)***
Prius 2008	-3.53	0.10	-2.53	0.38	0.10	1.43
	(4.60)	(0.50)	(20.46)	(0.43)	(0.17)	(1.48)
New Buildings	0.06	0.01	0.24	0.02	0.01	0.13
	(0.02)**	(0.00)*	(0.12)**	(0.01)***	(0.00)**	(0.03)***
College	0.45	-0.00	0.93	0.04	-0.00	0.24
	(0.24)*	(0.02)	(0.97)	(0.02)*	(0.01)	(0.08)***
Income	-1.22	0.09	3.10	-0.26	-0.01	-0.95
	(1.17)	(0.12)	(4.65)	(0.12)**	(0.03)	(0.47)**
Green Ballot Share	0.12	-0.00	0.26	0.02	0.00	-0.02
	(0.12)	(0.02)	(0.44)	(0.02)	(0.00)	(0.04)
Senate LCV	0.01	0.00	0.24	-0.00	0.00	0.02
	(0.05)	(0.01)	(0.21)	(0.00)	(0.00)	(0.02)
House LCV	0.01	0.00	-0.25	-0.01	-0.00	-0.03
	(0.04)	(0.01)	(0.18)	(0.01)	(0.00)	(0.01)**
CEM Weights	Y	Y	Y	Y	Y	Y
Observations	202	202	202	453	453	453
R-squared	0.49	0.28	0.35	0.17	0.1	0.35
Mean DV	7.92	0.86	39.74	1.11	0.2	5.24
Marginal Effect	98%	185%	33%	72%	130%	74%

Notes: OLS regressions with robust standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city. Number of treated and control units in the matched samples are displayed in Table 2.

Table 4. Logistic Hazard Models of Green Policy Adoption

	Adopters and Matched Controls		Green Policy Adopters Only	
Outcome	Green Policy		Green Policy	
Unit of Anlaysis	City-year		City-Year	
LEED Registrations	0.05		-0.07	
	(0.04)		(0.06)	
LEED Accreditations		-0.00		-0.00
		(0.01)		(0.01)
New Buildings (Annual)	0.02	0.03	0.00	0.00
	(0.01)*	(0.01)**	(0.01)	(0.01)
Prius2008	-0.08	-0.48	3.39	3.27
	(1.50)	(1.45)	(1.43)**	(1.39)**
Green Ballot Share	0.02	0.02	-0.08	-0.08
	(0.03)	(0.03)	(0.05)	(0.05)*
LCV Senate	0.01	0.02	-0.00	0.00
	(0.01)	(0.01)	(0.01)	(0.01)
LCV House	-0.01	-0.01	0.01	0.01
	(0.01)	(0.01)	(0.01)	(0.01)
Population	-0.04	-0.04	0.08	0.08
	(0.05)	(0.04)	(0.03)***	(0.03)**
College	-0.01	0.01	-0.12	-0.12
	(0.04)	(0.04)	(0.09)	(0.08)
Income	0.12	0.05	0.40	0.39
	(0.17)	(0.16)	(0.41)	(0.41)
CEM Weights	Y	Y	N	N
Observations	1164	1164	135	135

Notes: Robust standard errors (clustered on city) in parentheses; *** p<0.01, ** p<0.05, * p<0.10.

Table 5.A: Pooled Cross-Sectional OLS Regressions

	Adopters and Matched Controls			Neighbors and Matched Controls		
Outcome	*Registrations*	*Gov't*	*Accreditations*	*Registrations*	*Gov't*	*Accreditations*
Green Policy (Ever)	-0.14	0.06	-1.96			
	(0.35)	(0.05)	(1.23)			
Green Policy (Now)	2.13	0.29	10.82			
	(0.78)***	(0.11)***	(3.67)***			
Green Neighbor (Ever)				-0.02	0.00	-0.13
				(0.02)	(0.01)	(0.07)*
Green Neighbor (Now)				0.14	0.04	1.08
				(0.05)***	(0.02)**	(0.25)***
Additional Controls	Y	Y	Y	Y	Y	Y
City Fixed Effects	N	N	N	N	N	N
Year Fixed Effects	Y	Y	Y	Y	Y	Y
Observations	1616	1616	1616	3624	3624	3624
R-squared	0.28	0.18	0.40	0.13	0.05	0.22

Notes: Robust standard errors (clustered on city) in parentheses; *** p<0.01, ** p<0.05, * p<0.10. Additional unreported controls are Prius 2008, Green Ballot Shane, LCV Senate, LCV House, Annual New Buildings, Population, College and Income. These models do not include CEM weights.

Table 5.B: City-Fixed Effects OLS Regressions

	Adopters and Matched Controls			Neighbors and Matched Controls		
Outcome	*Registrations*	*Gov't*	*Accreditations*	*Registrations*	*Gov't*	*Accreditations*
Green Policy (Now)	2.34	0.27	11.11			
	(0.78)***	(0.11)**	(3.64)***			
Green Neighbor (Now)				0.13	0.04	1.34
				(0.05)***	(0.03)	(0.25)***
Additional Controls	N	N	N	N	N	N
City Fixed Effects	Y	Y	Y	Y	Y	Y
Year Fixed Effects	Y	Y	Y	Y	Y	Y
Observations	1640	1640	1640	3664	3664	3664
CA Cities	205	205	205	458	458	458
R-squared	0.16	0.10	0.25	0.09	0.03	0.16

Notes: Robust standard errors (clustered on city) in parentheses; *** p<0.01, ** p<0.05, * p<0.10. All regressions control for Annual New Buildings and do not include CEM weights.

Table 6: Instrumental Variable Models

	Neighbors and Matched Controls		All Cities w/o Green Policy		All Cities w/o Green Policy	
Outcome	*Registrations*		*Registrations*		*Accreditations*	
log(Adjacent APs+1)	0.21	0.21				
	(0.06)***	(0.08)***				
log(APs w/in 25 Miles)			0.36	0.59		
			(0.17)**	(0.25)**		
Registrations					2.07	2.61
					(0.28)***	(0.99)***
First Stage Coefficients and Statistics						
Green Neighbor		3.53				
		(0.21)***				
log(Policies 25 to 50 Miles)				1.21		
				(0.13)***		
New Buildings						0.04
						(0.01)***
F-Test of Excluded IVs		285.00		90.12		24.25
N	453	453	244	244	244	244
R-squared	0.28	0.28	0.45	0.45	0.51	0.50

Notes: IV regressions with robust standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.10. Unit of analysis is a city. The sample for columns 3-6 includes all cities with more than 20,000 inhabitants that do not adopt a green policy. All models include controls for Population, income, College, Green Ballot Share, and Prius 208.

Table 7: City Size Interaction Effects

	Neighbors and Matched Controls		Neighbors and Matched Controls		All Cities w/o Green Policy	
Oucome	*APs*	*APs*	*APs*	*APs*	*Buildings*	*Buildings*
Specification	*OLS*	*OLS*	*OLS*	*OLS*	*IV*	*IV*
Green Policy	-33.77	-39.00				
	(17.14)*	(10.97)***				
Green Policy * Population	3.14					
	(1.21)**					
Green Policy * New Buildings		0.39				
		(0.08)***				
Green Neighbor			3.05	0.16		
			(1.88)	(1.24)		
Green Neighbor * Population			0.20			
			(0.27)			
Green Neighbor * New Buildings				0.16		
				(0.06)***		
log(APs w/in 25 Miles)					0.56	0.40
					(0.24)**	(0.22)*
log(Aps 25) * Population					0.05	
					(0.02)***	
log(Aps 25) * New Buildings						0.01
						(0.00)***
First Stage F-Statistics						
log(Policies 25 to 50 Miles)					44.96***	55.16***
log(Policies 25) * Population					110.03***	
log(Policies 25) * Buildings						47.11***
Observations (Cities)	202	202	453	453	244	244
R-squared	0.6	0.58	0.18	0.22	0.47	0.49

Notes: Robust standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.10. Additional unreported controls are Prius 2008, Green Ballot Shane, LCV Senate, LCV House, Annual New Buildings, Population, College and Income. OLS models use CEM weights.

Figure 1: Possible Impacts of Green Procurement Policies

Type of Interaction	Private Purchasing Becomes: Greener	Private Purchasing Becomes: Less Green
Supply	Scale economies Coordination / Lead adopter Induced innovation	Crowding out (inputs)
Demand	Learning /Awareness Moral suasion	Threshold effects Crowding out (benefits)

Figure 2: New LEED Registrations by Year and Building Owner Type

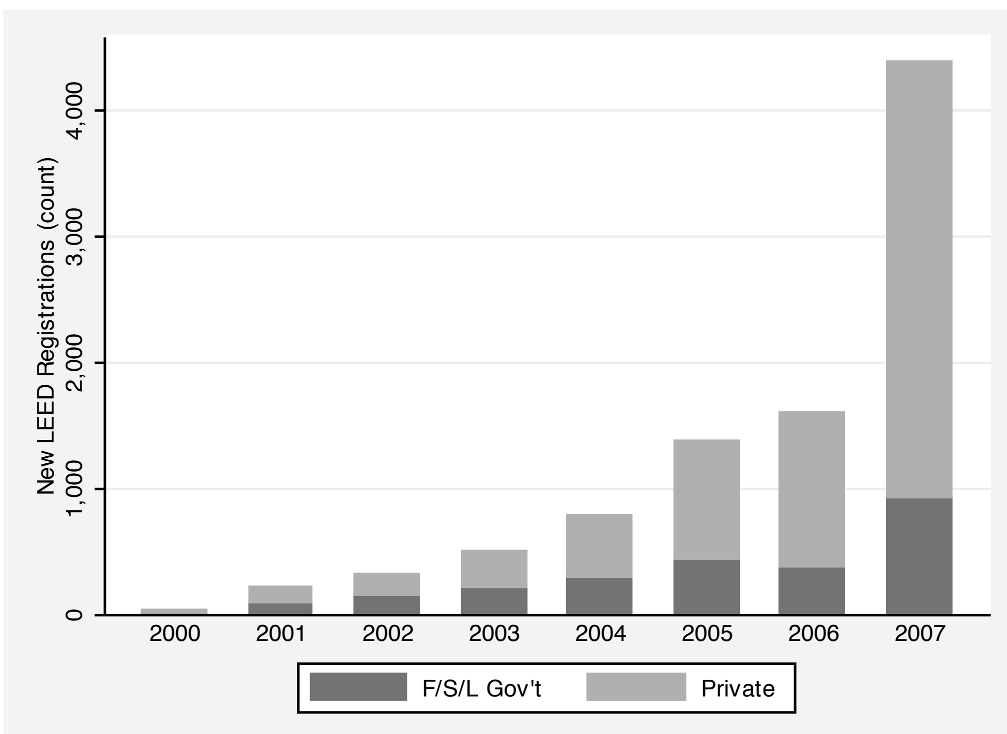
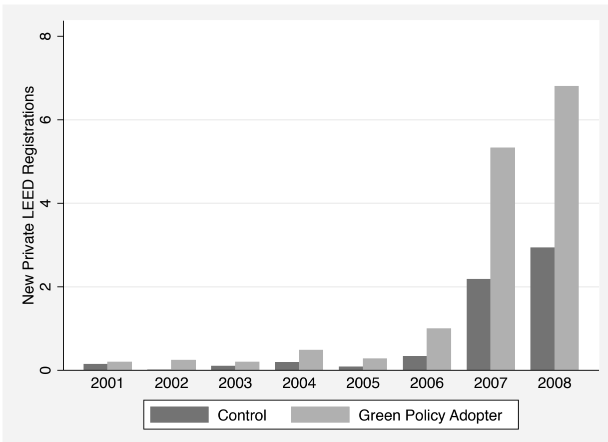
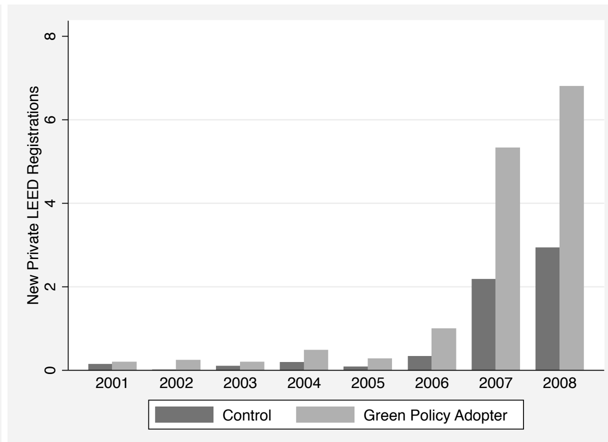


Figure 3: Registration and Accreditation Trends

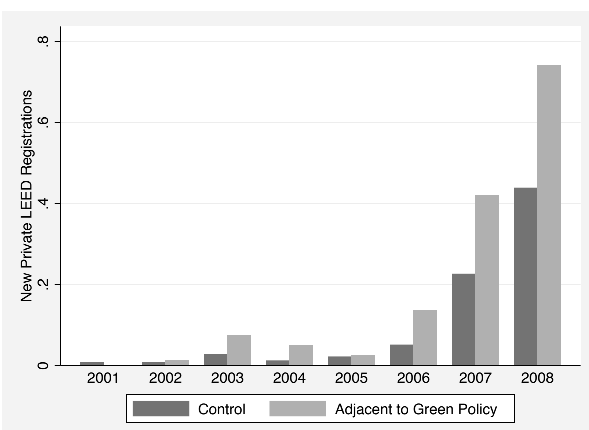
Part A: Green Policy Adopters and Matched Controls

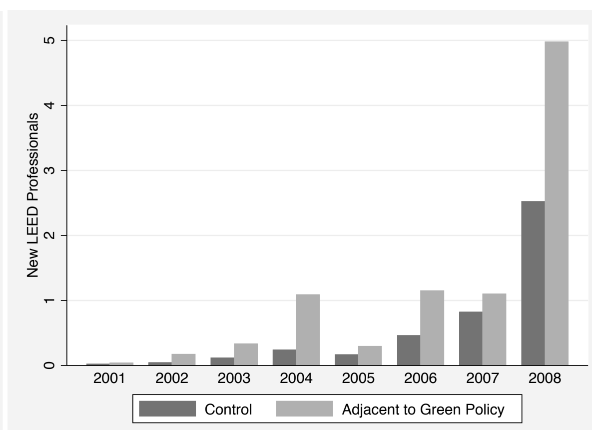




Registrations *Accreditations*

Part B: Adjacent Neighbors and Matched Controls



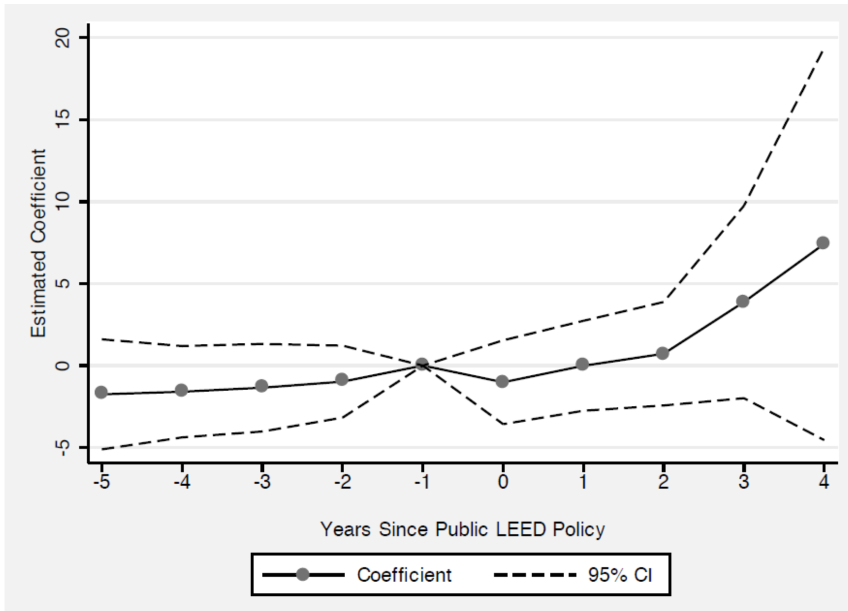


Registrations *Accreditations*

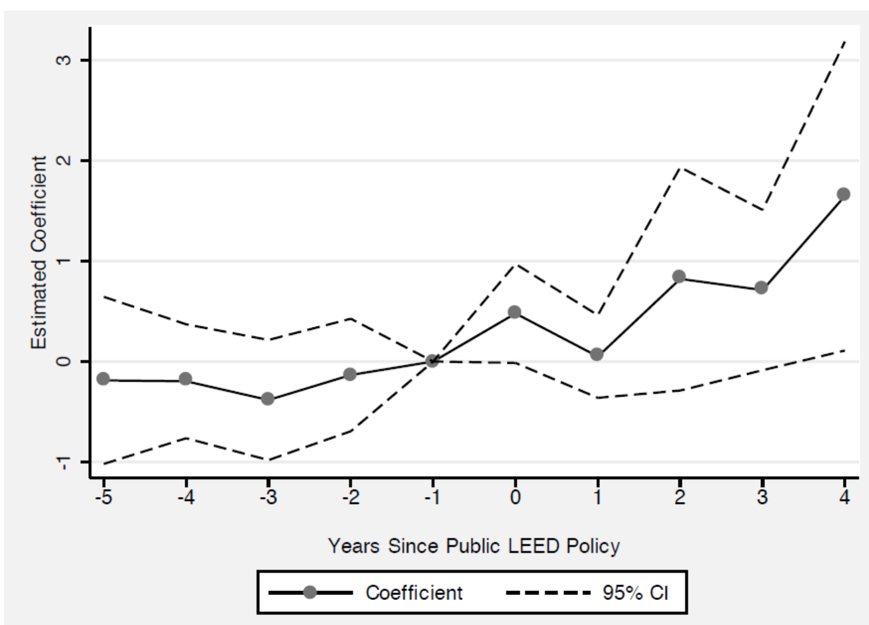
Notes: All figures are based on CEM-weighted annual means.

Figure 4: Annual Treatment Effects of Green Building Procurement Policies

Part A: Registrations in Policy Adopting Cities (Relative to Matched Controls)



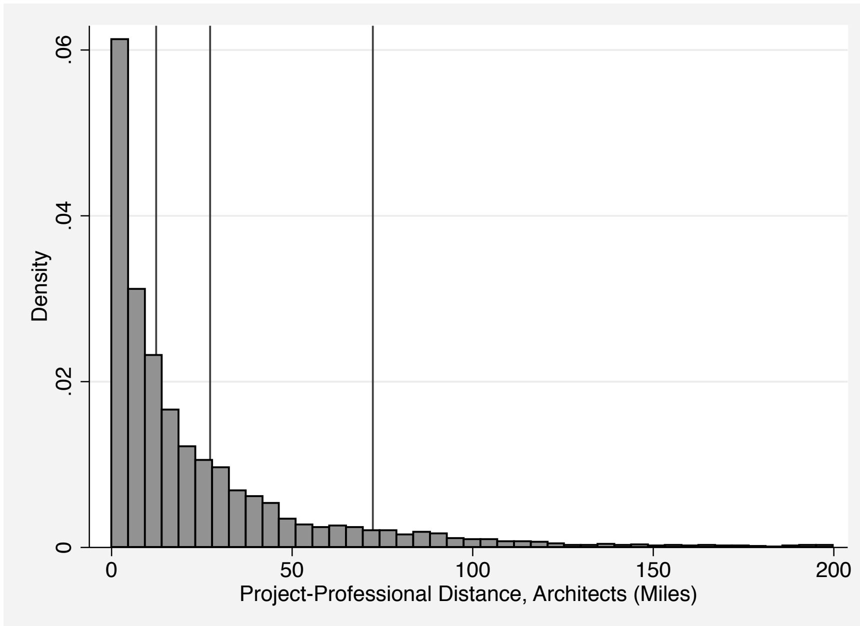
Part B: Accreditations in Adjacent Neighbor Cities (Relative to Matched Controls)



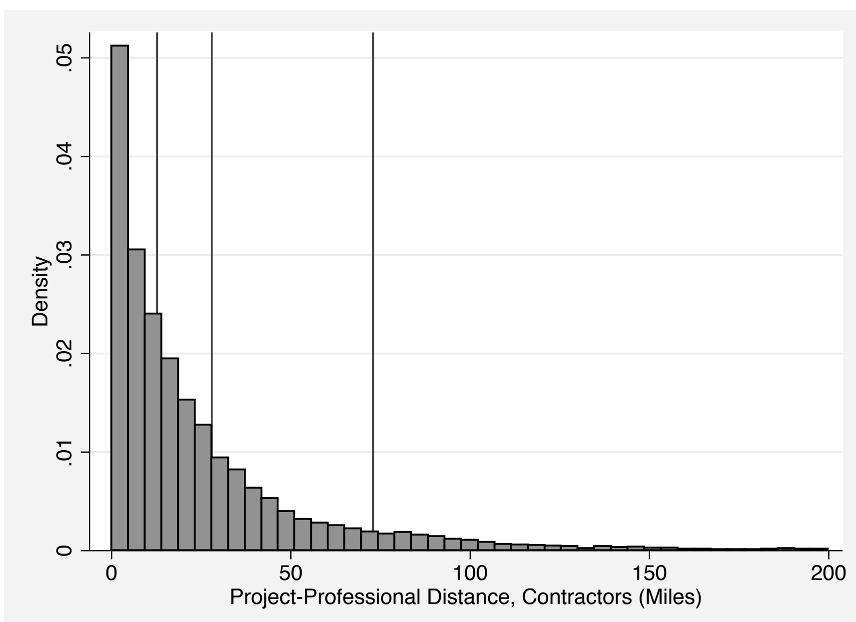
Notes: Coefficients and 95 percent confidence intervals obtained from a city-year fixed effect OLS regression. See text for details.

Figure 5: Size of Local Labor Markets for Real Estate Professionals

Part A: Distance from Architect Address to Building Address



Part B: Distance from General Contractor Address to Building Address



Notes: Vertical bars represent 25th, 50th and 75th percentile values in the empirical distance distribution.

Appendix

Table A1: List of California Cities with a Green Building Policy by 2008

	City Name	*Matched*	*Population*
1	Los Angeles	No	369.49
2	San Diego	No	122.34
3	San Jose	No	89.50
4	San Francisco	No	77.67
5	Long Beach	Yes	46.15
6	Sacramento	Yes	40.70
7	Oakland	No	39.95
8	Anaheim	Yes	32.80
9	Stockton	Yes	24.38
10	Fremont	Yes	20.34
11	Glendale	Yes	19.50
12	Santa Clarita	Yes	15.07
13	Santa Rosa	Yes	14.76
14	Irvine	Yes	14.31
15	Sunnyvale	Yes	13.18
16	Corona	Yes	12.50
17	Costa Mesa	Yes	10.87
18	Berkeley	No	10.27
19	Santa Clara	Yes	10.24
20	Ventura	No	10.09
21	Richmond	Yes	9.92
22	Santa Barbara	Yes	9.23
23	Santa Monica	No	8.41
24	San Leandro	Yes	7.95
25	Carlsbad	Yes	7.82
26	Livermore	Yes	7.33
27	Alameda	Yes	7.23
28	Temecula	Yes	5.77
29	La Mesa	Yes	5.47
30	Cupertino	Yes	5.05
31	West Hollywood	Yes	3.57
32	Dublin	Yes	3.00
33	Cotati	Yes	0.65