LEED Adopters:
Public Procurement and Private Certification

Timothy Simcoe

Boston University
School of Management
595 Commonwealth Ave.
Boston, MA 02215
(617) 358-5725
tsimcoe@bu.edu

Michael W. Toffel

Harvard Business School
Morgan Hall 497
Boston, MA 02163
(617) 384-8043
mtoffel@hbs.edu

1. Introduction

Governments increasingly use purchasing rules 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. Many state and local governments have taken similar steps.[1]

Since government purchases account for 10-15 percent of GDP in most developed countries, green procurement policies could have substantial direct effects on the environment. However, compared to a blanket regulatory intervention, green procurement rules will have little impact unless government purchasing can somehow influence the remaining 85 to 90 percent of private-sector trade.

In principle, green procurement policies might either encourage or crowd out environmentally responsible private sector purchasing. In some cases, government

[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.

guidelines may stimulate private demand through moral suasion, or by legitimating a particular eco-label or certification scheme. Given scale economies in green production, government procurement preferences might also reduce the marginal costs of private adoption. On the other hand, when the marginal costs of green production are increasing and private agents view "brown" products as a close substitute, green procurement policies would simply displace private investment in environmental goods.

This paper is an empirical case study of the interaction between local government procurement rules and private-sector adoption of green building practices. Specifically, we examine the diffusion of the US Green Building Council's LEED standard for sustainable building practices, and ask whether private real-estate developers 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 private sector LEED adoption is roughly 80 percent greater in municipalities with a green-building policy than in a matched control sample with similar size and demographic characteristics (including measures of "green preference" such as voting on environmental ballot initiatives and Toyota Prius ownership rates). We also find more LEED adoption among "neighbor cities" that share a border with a city that adopts a green building policy (relative to their own set of matched controls). While our primary measure of LEED adoption is registration of a new buildings with the US Green Building Council (USGBC), we also examine the number of LEED accredited professionals: architects, contractors and other professionals who have passed a exam certifying their knowledge of LEED building principles. We find a substantial increase in LEED APs at cities that adopt a municipal green building policy and their neighbors (though these results are only statistically significant in the larger sample of neighboring cities).

Our findings suggest that municipal government purchasing policies can stimulate private adoption of green building practices. However, they do not imply similar effects in all categories of green procurement. Governments are an especially large purchaser of construction services. And 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.

Nevertheless, we know of no other research that documents a private-sector response to public procurement policies, and our findings may also shed light on the underlying mechanisms. In particular, the large neighboring-city effects suggest that private LEED adoption is not purely an effort to pre-empt local regulators, as in Maxwell, Lyon and Hackett (2000). Rather, the neighboring-city effects and the increase in LEED APs suggest that local procurement policies sway private decisions by increasing local awareness of the benefits of green building (i.e. moral suasion) and jump-starting the development of specialized input markets.

The balance of the paper proceeds as follows: Section 2 describes a simple economic framework for analyzing the link between government procurement policy and voluntary private certification programs. Section 3 describes the USGBC Leadership in Energy and Environmental Design (LEED) standard for green building practices. Section 4 describes our data, measures, and empirical methods. Section 5 presents the results, and Section 6 offers some concluding remarks.

2. Some Economics of Green Procurement

Government purchasing guidelines often use price preferences or quantity targets (typically called set-asides) to reward products that demonstrate a level of "greenness" by incorporating recycled materials, exceeding official pollution standards or qualifying for environmental labels. If the government is a major customer, these policies may have a significant environmental impact. Green-procurement may also be used to signal concern for the environment when regulatory intervention is costly or infeasible.

When the government is not a major customer, the impact of green-procurement policies will depend on how government purchasing interacts with private sector procurement decisions.[2] 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).

In principle, the choice of government procurement rules can influence private purchasing behavior through either supply or demand channels, and might either reinforce or counteract the direct impact of a green purchasing policy. Figure 1 builds on a simple framework proposed by Marron (2003).

Figure 1: How Public Procurement Affects Private Purchasing

Private Purchasing Becomes

		Greener	**Less Green**
Type of Interaction	**Supply**	Scale Economies Standardization Induced innovation	Crowding Out
	Demand	Awareness Moral suasion	Threshold effects

On the supply side, procurement policies can lead to greener private purchasing if there are significant scale economies, so an initial government purchase reduces the average cost of serving additional private customers. 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-efficient materials). Finally, government procurement rules may lead to increased competition and innovation

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

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.

It is also possible for supply-side interactions to counteract the direct effects of a green-procurement program. In particular, if the supply of green product is inelastic and/or private actors see "brown" goods as a close substitute, green purchasing policies may simply "crowd out" private investment. While we could find no procurement examples, there is some evidence that government subsidies for green electricity are primarily spent on marketing and advertising these higher-priced services to end consumers (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 important) and private customers already have an incentive to adopt greener products (e.g. because of energy-cost savings).

Finally, when procurement rules define a sharp cutoff between "green" and "brown" products, private-sector demand 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 be biased away from private standards.[3]

In practice, the importance of each mechanism depicted in Figure 1 will depend on specific features of the relevant product market. Our empirical analysis will focus on real-

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[3] 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).

estate development from 2001 to 2008. There are several reasons to expect that private purchasing will reinforce a green procurement policy in this setting. First, government is an especially large customer. Using BEA data from 2002, Marron (2003) shows that 21.4 percent of all "Maintenance and Repair Construction" spending comes from state and local governments and 4.9 percent from the federal government. This is the largest share of any product category after munitions. Second, builders can realize direct benefits from green investments that produce energy savings or increases tenants' willingness-to-pay. Finally, we study a period when LEED was just emerging as the dominant green building standard. Thus, the costs of LEED adoption may be low (or even negative) for the marginal private builder; a situation that is unlikely to persist as the standard diffuses widely.

Given these institutional characteristics of the market for green buildings between 2001- and 2008, we will statistically test the following set of predictions:

> H1: Public-sector green-building procurement policies lead to increased private adoption of green-building certification (i.e. local LEED adoption).
>
> H2: Public-sector green-building procurement policies lead to increased entry in the market for complementary inputs (i.e. increased LEED accreditation by local real-estate professionals).
>
> H3a/b: If green procurement polices harness scale economies or moral suasion, private adoption will increase in neighboring cities. If public procurement crowds out private investment or the benefits of green certification are highly localized (e.g. due to preferential treatment in zoning or municipal inspection), there will be no spillover to neighboring cities.

3. LEED Certification

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 measuring the environmental impact 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.[4] All certified projects must achieve a certain number of points in each category. However, there are four certification levels – Certified, Silver, Gold and Platinum – and projects reach higher levels by accumulating more LEED points.

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 on these construction-related costs, which will vary from one building to the next. Some LEED points seem relatively cheap (e.g. for installing bike racks), while others are quite expensive (e.g. brown-field site remediation). The administrative costs of LEED certification are small relative to the costs of new construction. It costs roughly $450-600 to register a project, and $2,000 to complete a certification.

For a commercial building, the benefits of LEED may come 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 rather than energy consumption, and several observers suggest that more work is needed to

[4] 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.

understand whether these buildings actually deliver long-term environmental benefits. On the financial side, 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.

While the LEED system debuted in 1998, it did not achieve a 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 presumably reflects a variety of 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 1 also shows that Federal State and Local governments have been significant LEED adopters since the program began.

4. Data, Measures and Methods

To analyze he impact of municipal green-building 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, including measures of LEED diffusion from the USGBC; hand-collected data on the municipal adoption of green-building policies; data on non-residential construction starts from McGraw Hill; and city-level demographic data from the US census. Summary statistics are presented in Table 1.

4.1 Outcome Variables

We use two main outcome variables to measure the diffusion of LEED. Our first outcome is LEED *Registrations*, a count of new privately-owned buildings that registered for LEED certification between 2001 and 2008. These data come from the US Green Building

Council. The cities in our estimation sample registered between 0 and 99 new buildings for LEED between 2001 and 2008, with a mean of just less than 2 projects.[5]

While *LEED Registrations* should reflect private developers' intention to use green-building practices, the variable has two drawbacks.[6] First, registration does not imply that a building is ultimately certified. Rather, it is a low cost first-step towards LEED 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. And since LEED was diffusing rapidly towards the end of our sample period, focusing on certified buildings would exclude a large number of projects (though we hope to revisit this question with more recent data in the near future).[7] The second drawback of *LEED Registrations* is that it does not contain any information on the environmental impact of certification, a topic we leave to future research.

Our second outcome variable is *LEED Accreditations*, 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 through their business address, as maintained in the USGBC directory of LEED accredited professionals. The cities in our estimation sample were home to between zero and 416 LEED APs by 2008, with an average of 7.5 accredited professionals per city.

4.2 Explanatory Variables

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[5] We exclude the four largest cities in calculating these summary statistics, since they cold not be matched (and are therefore excluded from the main analysis below) and tend to distort the sample averages due to their extreme size. Those cities are Los Angeles, San Diego, San Jose and San Francisco.

[6] We exclude public-sector projects since they could be directly influenced by the municipal green-building policy (though including them does not change our results).

[7] 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.

Municipal policies: 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 calendar quarter. We gathered information on the adoption of municipal green-building policies by hand, starting from lists compiled by the USGBC and the DSIRE database of government incentives for energy efficiency.[8] 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 six cities that have regulate private green building from our analysis.[9]

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.[10] Our research suggests that 87 percent of all green-building polices contained what we would call a "purchasing rule" (i.e. a commitment that new public projects would adhere to some type of environmental standards). Ninety percent of these policies used LEED points and/or certification levels as their measure of environmental performance.

We define an indicator variable *Green Policy that* equals "1" if a city adopted a green procurement policy by 2008. Similarly, we set the *Green Neighbor* indicator variable to one if *Green Policy* equals zero, but an adjacent city (i.e. a city with a common border) does adopt a green procurement policy. 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.

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[8] 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 .

[9] Table A1 provides a complete list of the adopter cities.

[10] Available at www.municode.com

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.

Tastes and Demographics: For each city in the analysis, we collected a variety of demographic data from the 2000 U.S. Census.[11] The census variables used below are *Population* (measure in units of 10,000), *Income* (median household income in $10,000's), and *College* (the share of adults with some college education).

In addition to these standard demographic variables, we collected two novel measures of a city's preference for environmental sustainability. First, we observed citizens' own votes in favor of statewide ballot initiatives addressing environmental quality (Kahn 2002).[12] The variable *Green Ballot Share* shows an average support of 61 percent for these ballot questions. 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 *Prius 2008,* which has a mean of 0.54 percent.[13]

4.3 Matching and Balance Tests

Our main statistical method is the Coarsened Exact Matching (CEM) approach described by Iacus, King and Porro (2009). This approach assumes that after stratifying and re-

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[11] Matching political jurisdictions to census data was done at the level of the Census Place.
[12] 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 199x-200x within the Census Place that best corresponded to each city.
[13] The highest Prius registration rate is 3.74 percent in Portola Valley (just west of Palo Alto).

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" the observables 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, one assigns a weight of "1" 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).

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 control samples: one for *Green Policy* adopters and the other for *Green Neighbors*. For the adopters, we match on *Population* and *Prius2008*. 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 *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 (weighted) control cities.

Table 2 demonstrates how the CEM matching dramatically improves the balance in the means of the treatment and control samples. The leftmost two columns in Table 2 compare 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).[14] Not surprisingly, we find that cities adopting a green-building policy are larger, greener, richer and better educated than the potential controls. Beneath the individual sample means, Table 2 reports a likelihood ratio test for the joint null hypothesis that the means of all 7 variables are equal in the treatment and control samples. This test strongly rejects that hypothesis (p=0.02) for the raw data.

The middle two columns in Table 2 compare weighted means for *Green Policy* adopters and their matched 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 *Prius2008* (e.g. Berkeley and Santa Monica), as can be seen by the 0.15 percentage point drop in that variable. Note the substantially improved match in the means of *Population* and *New Buildings*. Remarkably, matching on *Population* and *Prius2008* substantially narrows the difference in means for education, income and green-voting measures. A likelihood ratio test does reject the null hypothesis that the means of all observables are equal across the *Green Policy* adopting and matched control samples (p=0.81).

The last two columns in Table 2 compare 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 CEM weighted means for neighbors and their matched controls are very similar, and a likelihood ratio test does not reject the null hypothesis that all of these means are equal (p=0.16).

[14] 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.

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. Table 3 illustrates trends for the key policy and outcome variables over our 2001-2008 sample period. The top panel (3.A) shows trends for policy-adopters and matched controls, while the bottom panel (3.B) reports the same statistics for neighbors and their matched controls.

The first column in Table 3.A shows that the earliest green-building procurement policies in our estimation sample were adopted in 2003, and that the number of cities with a *Green Policy* increased steadily thereafter. The first column in 3.B shows that there had already been "neighbor adoptions" in 2001 and 2002 (for cities adjacent to Los Angeles, San Francisco, San Diego or San Jose). Also, the proportion of *Green Neighbor* cities grows faster than the actual *Green Policy* adopters, since larger cities were among the first to adopt green procurement policies.

The middle two columns in Tables 3.A and 3.B examine *LEED Registrations*. These began quite low, but start to grow dramatically around 2006, and much more so for cities in the two treatment groups. The final two columns in Tables 3.A and 3.B show a similar pattern for *LEED Accreditations* (a change in the LEED AP exam in early 2005 may explain the apparent dip in accreditation at that time). Overall, Table 3 illustrates rapid growth in LEED adoption at the end of our sample period. Thus, while our main specification is cross-sectional, we obtain similar results from a difference-in-differences regression with city fixed-effects, as we illustrate below.

5. Results

Our main results are presented in Table 4. The first two columns report estimates from weighted OLS regressions that compare *Green Policy* adopter cities to their CEM-matched controls. The outcome in the first column is *LEED Registrations*. We find a statistically significant increase of 6.6 registrations in cities with a green-building policy. Since the mean count of LEED Registrations is 7.9, this is an increase of roughly 84 percent. This result is relatively stable as controls are added or removed, and grows significantly larger if we do not perform the CEM matching and weighting.

In the second column of Table 4, we find an increase of 11.55 *LEED Accreditations* in *Green Policy* adopting cities, relative to the matched controls. This corresponds to an increase of roughly 30 percent, 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. This is consistent with the results in the final column of Table 3, and something that we plan to look into using the McGraw Hill data, which provides contact information for the key professionals on each new building.

In the last two columns of Table 4, we focus on cities that share a border with a *Green Policy* adopter, and the CEM matched controls for these *Green Neighbors*. 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 green-ness) 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.

The third column in Table 3 shows a statistically significant increase of 0.77 *LEED Registrations* relative to the CEM matched controls. Since the *Green Neighbor* cities are typically smaller than policy adopters, this translates into a 69 percent increase, which is quite close to the estimated marginal effect for the *Green Policy* adopters. In fact, if we estimate the same regressions using a Poisson specification (so the coefficients may be interpreted as elasticities), the *Green Policy* effect is 0.59 and the *Green Neighbor* effect is 0.51. Again, these results are robust to specification, and grow larger as we relax the matching criteria.

Finally, the fourth column in Table 4 presents weighted OLS estimates of the impact of a *Green Neighbor* on *LEED Accreditations*. We find a statistically significant increase of 4.1 LEED APs, or roughly 77 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.

5.1 City Fixed Effects

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 difference-in-differences models, which compare LEED diffusion in treatment and control cities, before versus after the adoption of a green-procurement policy. Specifically, we estimate the following flexibly parameterized model:

$$\textbf{(1)} \qquad Y_{it} = \alpha_i + \beta_y \cdot GreenPolicy_{iy} + \boldsymbol{\gamma} \cdot X_{it} + \delta_t + \varepsilon_{it}$$

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 difference-in-difference results by graphing the β_y coefficients and their 95-percent confidence intervals for LEED Registrations in policy-adopting cities (Figure 3) and for LEED Accreditations in neighboring cities (Figure 4). These figures illustrate two very robust patterns in our panel data. First, the flat line connecting years -5 to -1 shows 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 Figure 3 ($p=0.69$) or Figure 4 ($p=0.53$). 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.

The second pattern revealed in Figures 3 and 4 is that "treated" cities gradually diverge from controls, rather than experiencing a sudden jump in LEED adoption following the policy change. Thus, if we replace the β_y coefficients by a single dummy that equals "1" for all years when the green-procurement policy is in place, the estimate on this parameter for *LEED Registrations* in policy-adopting cities is $\beta=1.45$ ($p=0.10$). However, if we redefine "treatment" as the period two or more years after policy-adoption, the effect increases to $\beta=3.9$ ($p=0.09$). Similarly, the coefficient on LEED Accreditation in neighboring cities increases from 0.74 ($p=0.01$) to 1.03 ($p=0.02$).

Finally, we note that the basic pattern of these difference-in-difference figures remains unchanged, but statistical significance increases dramatically, if we drop the CEM weights and/or relax the matching procedure (as is common in a great deal of applied empirical work).

5.2 The Hazard of Policy Adoption

Our final set of empirical results focus on the timing of green building policy adoption within the set of adopting cities. Specifically, we estimate logistic hazard models by keeping city-year observations for *Green Policy* adopters in or before 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. Using this new sample and dependent variable, we estimate logit models of green-building policy adoption (including a full set of calendar year dummies) and report the marginal effects.

Table 5 contains two main results. First, there is no sign of a surge in *LEED Registrations* or *LEED Accreditations* prior to the change in policy. While this is what we might expect based on the evidence in Figures 3 and 4, it provides additional evidence against stories of reverse causation (private LEED adoption sways public-policy makers) or regulatory capture (LEED APs lobby for a self-serving building code).

The second finding in Table 5 is that within the sub-sample of policy adopters, larger and greener cities adopt their municipal green-building policies sooner. This reinforces the importance of including *Population* and *Prius2008* in the CEM procedure. It is also interesting in its own right. Given the evidence of spillover effects in Table 3, the hazard model results suggest that promoters of green certification outside California could usefully focus on municipal governments in large cities as a key constituency.

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	New LEED Registered Buildings (01-08)	1.93	6.25	0.00	99
LEED Accreditations	New LEED Accredited Professionals (01-08)	7.51	27.38	0.00	416
Green Policy	Focal City Adopted Green Building Policy	0.04	0.19	0.00	1.00
Green Neighbor	Adjacent City Adopted Green Building Policy	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 share supporting green ballot measures	0.61	0.15	0.20	1.00
Population	Population (10,000's)	2.99	5.10	0.00	46.15
New Buildings	Non-residential construction starts (03-07)	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).

Table 2. Covariate Balancing Tests

	Full Sample No weights		Matched Sample Weighted		Matched Sample Weighted	
	Adopters	*Controls*	*Adopters*	*Controls*	*Neighbors*	*Controls*
Prius 2008	0.93	0.53	0.78	0.76	0.66	0.66
Green Ballot Share	0.72	0.60	0.71	0.68	0.69	0.66
Population	14.36	2.52	13.91	13.72	4.18	4.09
New Buildings	140.79	21.50	142.40	108.59	23.09	23.15
Buildings per Capita	10.62	12.11	10.91	9.65	7.93	7.28
College	0.35	0.23	0.34	0.33	0.30	0.27
Income	5.58	4.77	5.72	5.78	5.57	5.51
Chi-square(7)	17.21		3.76		10.45	
p-val	0.02		0.81		0.16	
Cities	29	706	25	180	81	377

Notes: First two columns report unweighted means for cross-section of 735 California cities (excluding Los Angeles, San Diego, San Jose and San Francisco). Middle columns report weighted means for matched sample (based on Population and Prius 2008). The CEM weights are described in Iacus , King and Porro (2009). Final two columns report weighted means for neighbors and matched controls.

Table 3.A: Outcome Trends for Adopters and Matched Controls

	Green Policy	LEED Registrations		LEED Accreditations	
Year	*Adopters*	*Adopters*	*Controls*	*Adopters*	*Controls*
2001	0.00	0.20	0.03	0.16	0.03
2002	0.00	0.24	0.04	1.24	0.17
2003	0.04	0.20	0.07	1.92	0.34
2004	0.28	0.48	0.08	6.80	1.36
2005	0.36	0.28	0.02	2.24	0.32
2006	0.52	1.00	0.17	7.32	1.31
2007	0.72	5.32	0.80	10.68	1.55
2008	1.00	6.80	1.17	29.60	5.77

Table 3.B: Outcome Trends for Neighbors and Matched Controls

	Green Policy	LEED Registrations		LEED Accreditations	
	Neighbors	*Neighbors*	*Controls*	*Neighbors*	*Controls*
2001	0.07	0.00	0.01	0.04	0.01
2002	0.38	0.01	0.00	0.17	0.02
2003	0.52	0.07	0.01	0.33	0.03
2004	0.60	0.05	0.01	1.09	0.10
2005	0.73	0.02	0.01	0.30	0.07
2006	0.83	0.14	0.03	1.15	0.17
2007	0.91	0.42	0.12	1.10	0.23
2008	1.00	0.74	0.18	4.98	0.81

Notes: Each column reports un-weighted annual means for the matched samples. Sample sizes are as reported in Table 2: Adopters=25; Neighbors=81; Adopter Controls=180, and Neighbor Controls=377.

Table 4. Effects of Green Building Procurement Policies on LEED Registrations

	Adopters and Matched Controls		Neighbors and Matched Controls	
Outcome	*Registrations*	*Accreditations*	*Registrations*	*Accreditations*
Green Policy	6.63	11.55		
	(3.20)**	(13.79)		
Green Neighbor			0.77	4.06
			(0.38)**	(1.32)***
Prius 2008	-6.39	-17.80	0.20	1.31
	(4.23)	(21.26)	(0.42)	(1.55)
Green Ballot Share	20.78	23.00	0.47	-6.62
	(12.09)*	(48.43)	(1.06)	(3.91)*
Population	-0.49	-1.22	-0.00	0.25
	(0.24)**	(1.20)	(0.06)	(0.19)
New Buildings	0.09	0.33	0.02	0.09
	(0.02)***	(0.13)**	(0.01)***	(0. 30)***
College	44.01	117.83	4.45	25.34
	(21.96)**	(103.68)	(2.38)*	(7.35)***
Income	-1.24	2.79	-0.27	-0.89
	(1.16)	(4.79)	(0.14)*	(0.46)*
Observations	202	202	453	453
R-squared	0.55	0.36	0.15	0.34
Mean DV	7.92	39.74	1.11	5.24

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 5. Hazard Models of Green Procurement Policy Adoption

	Adoption Hazards	
Outcome	Green Policy	
Specification	Logit (Marginal FX)	
LEED Registrations	-0.01	
	(0.01)	
LEED Accreditations		-0.00
		(0.00)
Prius 2008	0.35	0.35
	(0.12)***	(0.12)***
Green Ballot Share	-0.62	-0.63
	(0.43)	(0.43)
Population	0.01	0.01
	(0.00)***	(0.00)***
New Buildings	0.00	0.00
	(0.01)	(0.01)
College	-1.09	-1.16
	(0.70)	(0.68)*
Income	0.04	0.04
	(0.03)	(0.03)
Year Effects	Yes	Yes
Observations	135	135

Notes: Robust standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.10. Table reports estimates from logit models (or equivalently, logsitic hazard models) with city-year as unit of analysis and only green-policy adopters in the estimation sample.

Figure 2: LEED Registrations by Year and Building Owner Type

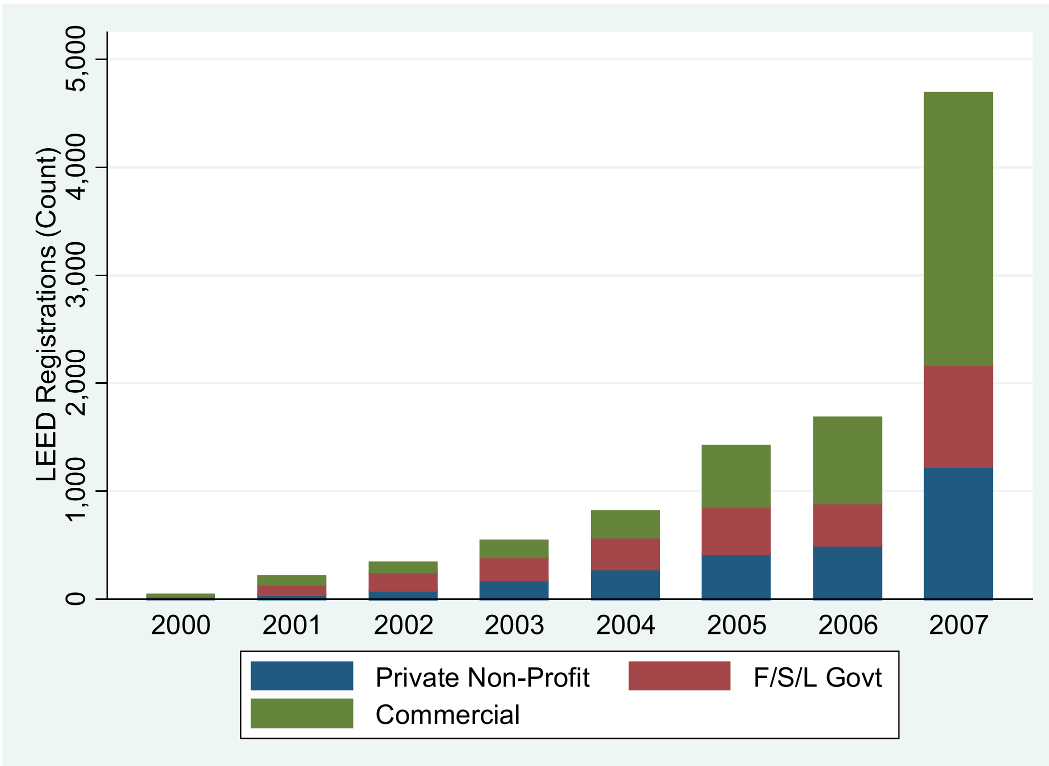
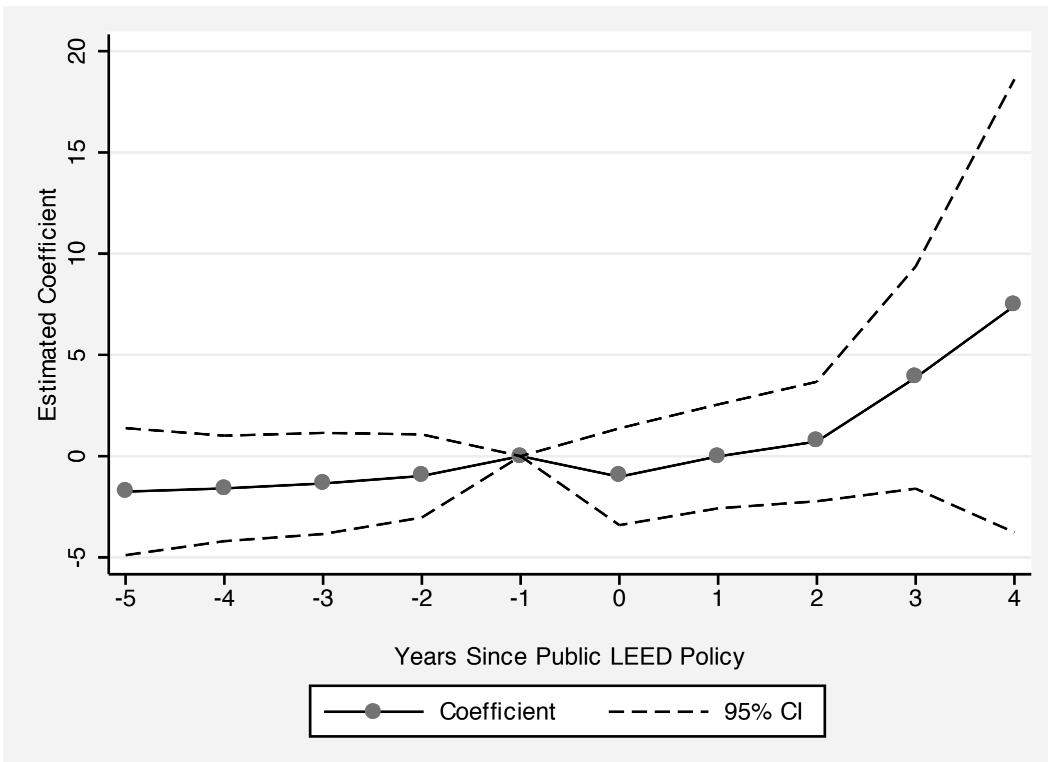


Figure 3: Annual Treatment Effects on LEED Registrations for Cities that Adopt a Green Building Procurement Policy



Notes: Robust standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.10. First two columns are OLS with city as unit of analysis. Last two columns are logit models (or equivalently, logsitic hazard models) with city-year as unit of analysis and only green-policy adopters in the estimation sample.

Figure 4: Annual Treatment Effects on LEED Accreditations for Neighboring Cities

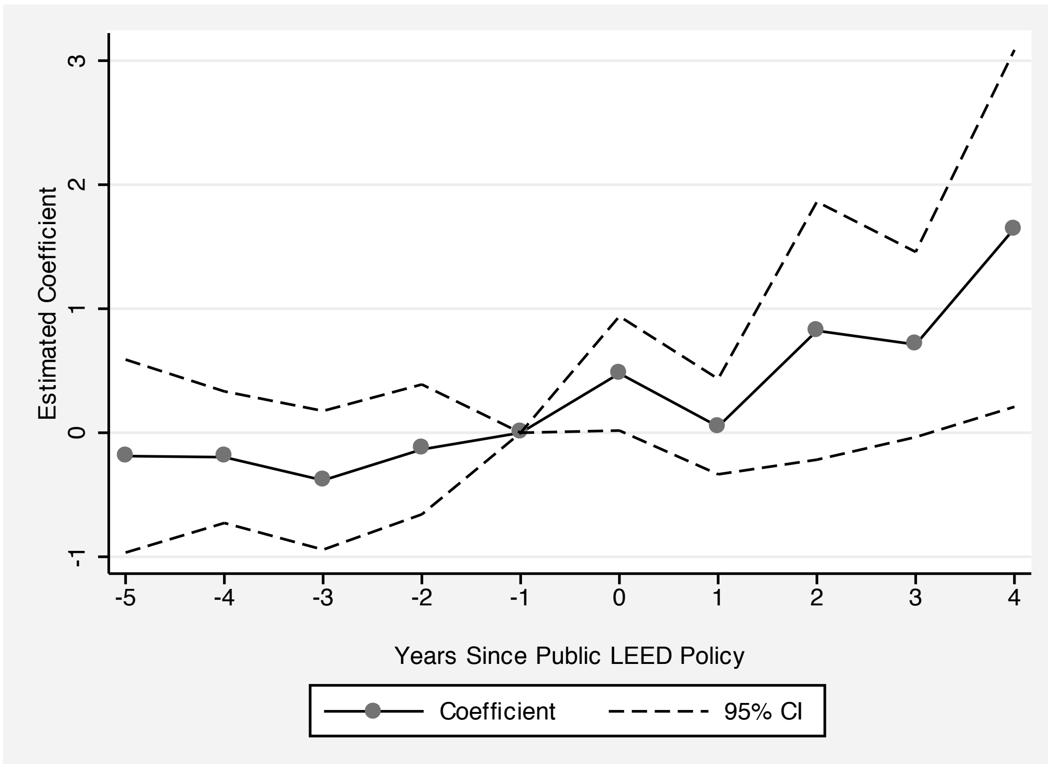


Table A1: List of California Cities with a Green Building Policy by 2008

	City Name	*Matched*	*Population*
1	Los Angeles	0	369.49
2	San Diego	0	122.34
3	San Jose	0	89.50
4	San Francisco	0	77.67
5	Long Beach	1	46.15
6	Sacramento	1	40.70
7	Oakland	0	39.95
8	Anaheim	1	32.80
9	Stockton	1	24.38
10	Fremont	1	20.34
11	Glendale	1	19.50
12	Santa Clarita	1	15.07
13	Santa Rosa	1	14.76
14	Irvine	1	14.31
15	Sunnyvale	1	13.18
16	Corona	1	12.50
17	Costa Mesa	1	10.87
18	Berkeley	0	10.27
19	Santa Clara	1	10.24
20	Ventura	0	10.09
21	Richmond	1	9.92
22	Santa Barbara	1	9.23
23	Santa Monica	0	8.41
24	San Leandro	1	7.95
25	Carlsbad	1	7.82
26	Livermore	1	7.33
27	Alameda	1	7.23
28	Temecula	1	5.77
29	La Mesa	1	5.47
30	Cupertino	1	5.05
31	West Hollywood	1	3.57
32	Dublin	1	3.00
33	Cotati	1	0.65