

Self- and Social Signaling: Evidence from Solar Adoption in California*

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Abstract

Prosocial behavior, engaging in costly actions that confer benefits on others, plays a role in many economic contexts and has been explained by altruism, social pressure, expectations of fairness and reciprocity, and signaling. We examine prosocial behavior signaling in the residential solar context, which allows us to distinguish the likely role of self-signaling, proving to yourself that you are the person you aspire to be, and social signaling from alternative explanations, including warm glow. Self-signaling is separately identified from social signaling by the exogenous visibility of potential solar arrays. We show that Democrats are more likely to adopt non-visible solar arrays when the composition of their neighborhood is more Republican, our evidence of self-signaling. The effect is economically meaningful: a one-standard-deviation increase in the share of Republican neighbors increases Democratic households' willingness to pay for solar by \$305. However, this effect is offset if the solar installation would be visible: both Republican and Democratic households are less likely to adopt visible solar, a net-negative social signal. Both the self- and social signaling effects are crowded out by the private benefits of installing solar.

Keywords: Prosocial behavior, self-signaling, social signaling, solar, public goods, energy policy, externalities, conformity, voter affiliation

JEL Codes: D64, H41, Q40, D83.

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1 Introduction

Public policy often relies on private prosocial behavior, such as adopting clean energy technologies, to supplement public goods. Designing optimal subsidies and interventions requires identifying the underlying behavioral mechanisms: whether individuals act from a desire to affirm who they are to themselves, to signal who they are to others, or both. Disentangling these incentives in field settings remains difficult, leaving an empirical gap in how policy interacts with these motivators. We overcome this hurdle by studying residential rooftop solar adoption, a high-stakes, real-world setting that provides private financial benefits, environmental benefits, and a potentially visible signal of environmental or pro-social identity.

To separate these motives empirically, we develop a conceptual model building on [Bénabou and Tirole \(2011\)](#) in which households may adopt solar not only because of private financial benefits and environmental preferences, but also because adoption helps them affirm a prosocial identity to themselves or communicate that identity to others. Households who value being a prosocial type but who live in neighborhoods dominated by those who are less likely to be of a high prosocial type, proxied for by Republican voter registration ([VanLange et al., 2012](#)), receive a self-signal identity restoration effect, increasing adoption. Our model also allows for a non-conformity penalty in settings of value misalignment ([Shayo, 2020](#)), consistent with polarization, and privately captured benefits of solar may “crowd-out” both forms of signaling.

Our empirical analysis combines household-level characteristics, voter registration, solar adoption data, and Google Project Sunroof measures of expected solar value and roof orientation, which we use to measure potential installation visibility from the street. Our main empirical strategy leverages within-blockgroup plausibly exogenous variation in the political composition of very local neighbors, the expected private value of solar, and the visibility of potential solar installations to separately identify effects consistent with self-signaling and social signaling (conformity) motives. Low-visibility solar adoption helps to

isolate the behavior consistent with self-signaling from social signaling. We find three key causal relationships.

First, consistent with self-signaling of one’s environmentally-oriented prosocial bona fides, we find that non-Republicans are more likely to adopt solar that is not visible when they live in a more Republican neighborhood.¹ A one-standard-deviation increase in the Republican share of a Democrat’s neighborhood generates an equivalent willingness-to-pay to adopt solar on average of \$305, or 0.9% of the average cost of an installation. The public good provided in our setting by this effect—cleaner air and lower grid emissions—is increasing in the self-signal, because solar displaces fossil electricity generation. We find that this effect increases overall installed capacity by 11.2 megawatts (2.8%) in our sample, decreasing environmental externalities by \$31.8 million in net present value (NPV).²

Second, we show that visible installations carry an adoption penalty that is stronger as a neighborhood becomes more Republican, likely due to costs of non-conformity. Democratic households have an average willingness to pay of \$124 to avoid this effect, separate from aesthetic values. The associated change in negative environmental externalities exceeds \$6.7 million in our sample, even when not accounting for the fact that lower adoption rates of visible installations will also slow diffusion (Bollinger et al., 2022).

Third, we provide evidence that both effects can be “crowded-out” by the private benefits of adopting solar, akin to the experimental findings of Dubé et al. (2017). Removing the crowding-out effect on what we interpret as the self-signal would increase total capacity by 0.67 megawatts (0.17%), decreasing the NPV of externalities by \$1.9 million. Eliminating the crowding-out effect on what we interpret as the social signal would result in 89 fewer installations and decrease total capacity by 0.198 megawatts (0.05%) (NPV of \$566,348).

The differential patterns of adoption by visibility and peer political affiliation indicates the importance of non-pecuniary product attributes. Our model best explains the observed

¹We find a positive effect, although insignificant, for Republicans as well, who may also have a motivation to self-signal their prosocial values.

²Our sample comprises 13% of California single-family detached households and spans a variety of counties including predominantly rural, suburban, and suburban-urban.

behavior, and alternative mechanisms have difficulty producing all of our findings: a *positive* correlation between adoption of solar and share Republican neighbors for non-visible solar, negative correlations between adoption and share Republican neighbors for visible solar, and an attenuating effect on both when private benefits are higher. We contribute to the growing literature on self- and social signaling, providing real-world, supportive evidence for their role in leading to these adoption patterns.

2 Related Literature

Theoretical notions of self-signaling and altruism are well-established in the literature (Bem, 1972; Thaler and Shefrin, 1981; Andreoni, 1989; Bodner and Prelec, 2003; Bénabou and Tirole, 2011), as are broader notions examining consumer norms and identity (Akerlof and Kranton, 2000; Peattie, 2010; Perez-Truglia, 2018; Pryor et al., 2019; Yan et al., 2021; Karing, 2024). Empirical work has shown that individuals gain utility from self-signals, largely by modeling self-signaling as the component of impure altruism that can be crowded out through private incentives (Gneezy et al., 2012; Grossman, 2015; Dubé et al., 2017), the costs of engaging in the prosocial activity (Garcia-Rada et al., 2022), or satiation (Tonin and Vlassopoulos, 2013).³ Our evidence for self-signaling is the combination of the local context affecting adoption rates for non-visible installations with the private benefit mitigating the effect.

Previous empirical evidence relies on experimental variation generated in a controlled setting (Dixon and Mikolon, 2021; Gallus and Heikensten, 2020; Mechtenberg et al., 2024). Self- and social signals are also difficult to disentangle, since social recognition implies visibility and the effects of visibility need not be uniform in sign, as in social comparisons (Butera et al., 2022; Allcott and Kessler, 2019) or when facing a polarized “dual audience” (Perez-Truglia and Cruces, 2017). In a field experiment related to this work, Bursztyn et al. (2020) show

³Specifically, this literature models impure altruism as an additive, independent private utility component. Changes in the pecuniary payoff of the action do not affect the impure altruism term by construction. Bénabou and Tirole (2006) note “the need to go beyond the standard dichotomy between “pure” and “impure” altruism and distinguish... between fixed preferences... and motives that relate to what a person’s behavior *says* about him or her, which will depend on the informational and economic context...”.

that individuals are willing to forego significant benefits to maintain a particular self-image, but that visibility of the decision to other participants reduced their willingness to do so. Nearly all previous work measuring the role of self- and social signaling focuses on low-stakes experiments with small monetary payoffs, limited audience size, low-cost effort, and short duration, all of which may lead to a lack of generalizability to real-world contexts (Levitt and List, 2007).

3 Conceptual Framework

Our model of solar adoption includes four additively separable sources of utility: outcome utility (pecuniary benefits net of cost, plus impure altruism (Andreoni, 1989; Ottoni-Wilhelm et al., 2017)), self-signal utility (Bodner and Prelec, 2003; Bénabou and Tirole, 2006, 2011), social signal utility, and a non-conformity penalty (Akerlof and Kranton, 2000; Carpenter and Myers, 2010; Perez-Truglia and Troiano, 2018). This section presents the basic framework of the model and some of the implications. See Appendix A for details.

Key variables are household political affiliation A , neighborhood Republican affiliation $\bar{A}$, expected bill savings less installation cost EV , installation visibility Vis , and unobserved household type $\theta \in \{\theta^h, \theta^l\}$. A household may know its type “in the gut” but not “in the mind,” (Bodner and Prelec, 2003) and so infers it from its own actions via self-diagnosis, valued differently by political affiliation: $V(\theta|A)$, with $V(\theta^h|D) - V(\theta^l|D) \geq V(\theta^h|R) - V(\theta^l|R)$. The difference in total utility from adopting versus not adopting is ΔU .

A household self-diagnoses by updating a prior ρ_0 over its own type to a posterior $\hat{\rho}$ upon adopting (or not), via Bayes’ rule conditional on Pr_H and Pr_L — each type’s adoption probability in equilibrium. The resulting self-signal utility of adopting relative to not adopting is therefore a function of $V(\theta)$, Pr_H , Pr_L , and ρ_0 : it is increasing in $V(\theta^h|A) - V(\theta^l|A)$ and in $Pr_H - Pr_L$, since a larger gap between types’ adoption rates makes the action more diagnostic of type, and decreasing in ρ_0 , since a household already convinced it is θ^h learns little new

from its own decision. Despite some model simplifications, our equilibrium results mirror those in [Bénabou and Tirole \(2011\)](#).

Table 1: Key Model Comparative Statics and Testable Implications. See Appendix A for proofs and derivations.

Mechanism	Comparative Static	Sign	Intuitive Driver
1. Self-Signal Identity Restoration	$\frac{d\Delta U}{dA} \Big _{Vis=0}$	> 0	Republican neighbors decrease priors: Living in a more Republican area lowers baseline self-priors (ρ_0). In infrequent-adoption settings, this creates an “identity deficit” that increases the value of adopting prosocial technologies to restore self-signal utility. (<i>Stronger for Democrats</i>)
2. Self-Signal Crowding-Out	$\frac{d^2\Delta U}{dA dEV} \Big _{Vis=0}$	< 0	Private payoffs diminish signal: Private payoffs ($EV \uparrow$) push θ^l types off the flattest part of the logit S-curve, decreasing the separation between types and diluting self-signal strength. (<i>Stronger for Democrats</i>)
3. Social Non-Conformity Penalty	$\frac{d\Delta U}{dA} \Big _{Vis=1} - \frac{d\Delta U}{dA} \Big _{Vis=0}$	< 0	Visible non-conformity: Except in extremely high $\bar{A}$ neighborhoods, visible adoption in a more Republican neighborhood signals explicit divergence from local norms, creating social non-conformity that depresses adoption when $ V^{nc} $ is large.
4. Social Non-Conformity Crowding Out	$\frac{d^2\Delta U}{dA dEV} \Big _{Vis=1} - \frac{d^2\Delta U}{dA dEV} \Big _{Vis=0}$	> 0	Private payoffs attenuate non-conformity penalty: Except in extremely high $\bar{A}$ neighborhoods, higher private payoffs allow observers to attribute visible adoption to financial gain rather than type or political affiliation, providing “cover” that reduces the non-conformity penalty

We assume ρ_0 is decreasing in $\bar{A}$: living among neighbors whose affiliation suggests they are θ^l lowers a household’s own prior belief in being θ^h ([Allport, 1954](#); [Turner and Onorato, 1999](#); [Turner et al., 2008](#)). This yields a *self-signal identity restoration* effect where greater $\bar{A}$ (lower ρ_0) raises adoption’s expected utility, crowded out by higher private payoffs EV ,

which compress $Pr_H - Pr_L$ and erode adoption’s signaling value (testable implications 1 and 2, under conditions detailed in Appendix A).

When an installation is visible, adoption also carries social signal utility, governed by an analogous prior $s_0 = 1 - \bar{A}$ over others’ perception of the household’s type, updated the same way. This is paired with a non-conformity penalty of opposite valence, $Vis \cdot V^{nc} \cdot |\hat{s} - s_0|$, $V^{nc} < 0$, increasing in $\bar{A}$ for visible installations only. Because both depend on visibility, we estimate only their net effect, which is negative; it is also crowded out by EV (implications 3-4, under conditions detailed in Appendix A). Table 1 summarizes all four testable implications.

4 Empirical Approach

4.1 Data

To test our predictions, we assemble a dataset of home locations and attributes, household characteristics, voter affiliation, and solar potential. From CoreLogic, we extract 994,454 owner-occupied, single-family detached homes built prior to 2014 across seven California counties.⁴ For each home, we observe the year built, lot size, home size, assessor value, bedrooms, and stories. We winsorize lot size and assessor value at 0.1% and merge household income and number of children from InfoUSA by address and with voter registration data from the California Secretary of State. Identifying the two most active voters between 2010–2014, we classify households as unregistered, Republican (both voters registered Republican), or Democratic/mixed (all others). For each household, we calculate the Republican share (excluding the index household) of the local neighborhood defined alternatively as census block-by-street, blockgroup-by-street, and block (our preferred measure).⁵

⁴Alameda, Contra Costa, Orange, San Diego, San Luis Obispo, Solano, and Yolo.

⁵These definitions of neighborhood are finer than the colloquial usage of neighborhood, with average sizes of 7.5, 14.1, and 18.9. They intend to capture immediate neighbors with whom one is most likely to interact. We provide a robustness check in which we measure neighborhood as the four spatially nearest neighbors as well.

Table 2: **Summary Statistics:** Summary statistics for variables used in analysis.

	All			Non-adopter			Adopter		
	Median	Mean	SD	Median	Mean	SD	Median	Mean	SD
Voter Registration									
Registered Democrat/Mixed hh	1.00	0.57	0.50	1.00	0.56	0.50	1.00	0.57	0.50
Registered Republican hh	0.00	0.21	0.40	0.00	0.20	0.40	0.00	0.25	0.43
Unregistered Voter hh	0.00	0.23	0.42	0.00	0.23	0.42	0.00	0.18	0.38
Neighborhood									
Republican block-by-street	0.17	0.20	0.19	0.17	0.20	0.19	0.21	0.24	0.21
Republican blockgroup-by-street	0.18	0.20	0.16	0.17	0.20	0.16	0.22	0.24	0.17
Republican block	0.18	0.20	0.14	0.18	0.20	0.14	0.23	0.24	0.15
Neighborhood by Affiliation									
Republican block-by-street: Democrat/mixed hh	0.14	0.18	0.19	0.14	0.18	0.18	0.20	0.22	0.20
Republican block-by-street: Republican hh	0.25	0.27	0.20	0.25	0.26	0.20	0.29	0.30	0.21
Republican block-by-street: Unregistered hh	0.15	0.19	0.19	0.14	0.18	0.19	0.20	0.22	0.20
Republican blockgroup-by-street: Democrat/mixed hh	0.15	0.18	0.15	0.15	0.18	0.15	0.20	0.22	0.17
Republican blockgroup-by-street: Republican hh	0.25	0.27	0.17	0.25	0.26	0.17	0.29	0.30	0.17
Republican blockgroup-by-street: Unregistered hh	0.16	0.19	0.16	0.16	0.18	0.15	0.20	0.22	0.17
Republican block: Democrat/mixed hh	0.16	0.18	0.14	0.15	0.18	0.14	0.21	0.22	0.15
Republican block: Republican hh	0.26	0.27	0.14	0.25	0.26	0.14	0.29	0.30	0.14
Republican block: Unregistered hh	0.17	0.19	0.14	0.17	0.19	0.14	0.21	0.22	0.15
Household and House Characteristics									
Children	0.00	0.31	0.60	0.00	0.30	0.59	0.00	0.40	0.69
Income (\$1,000)	89.33	108.59	75.86	87.00	106.04	74.50	112.75	134.12	84.09
Expected Value (\$1,000)	5.61	5.53	0.89	5.59	5.51	0.90	5.77	5.73	0.79
Home Size (1,000 sq. ft)	1.61	1.81	10.31	1.58	1.77	10.81	2.03	2.22	0.99
Bedrooms	3.00	3.30	1.00	3.00	3.27	1.00	4.00	3.62	0.97
Lot size (1,000 sq. ft) [†]	6.51	12.37	55.45	6.48	11.69	53.15	7.42	19.19	74.24
1(Year built > 1990)	0.00	0.10	0.30	0.00	0.10	0.29	0.00	0.15	0.36
1(Stories > 1)	0.00	0.23	0.42	0.00	0.22	0.42	0.00	0.27	0.44
Assessor value (\$100k) [†]	3.19	4.16	4.27	3.10	4.05	4.18	4.15	5.33	4.88
Visibility									
Visible (degrees from closest street)	0.39	0.41	0.29	0.39	0.41	0.29	0.38	0.40	0.29
Visible (side of street)	0.00	0.24	0.43	0.00	0.25	0.43	0.00	0.23	0.42
Visibility = 0	0.00	0.12	0.33	0.00	0.12	0.33	0.00	0.14	0.35

[†] Winsorized at .1%

We measure household-level variation in expected solar returns using Google Project Sunroof, which estimates solar photovoltaic performance via satellite imagery.⁶ The platform assumes a \$150 monthly bill and reports the 2016 NPV of a 15-year investment after installation costs. The variation in solar payoff that we utilize stems from differences across homes in roof profile and irradiance. Due to policy changes after 2016, we restrict the study period to 1999–2016.⁷ We use Sunroof coordinates to match Sunroof data to the nearest home, excluding the top 5% of matches by distance. Data on solar adoption come from the LBNL Tracking the Sun database, which provides near-universal coverage through 2016.

Visibility is a continuous $[0, 1]$ measure of the extent to which solar panels, if installed, would be visible from the street. For each house, we draw 360 radii of 65 meters from the roof center; a radius counts as visible if it intersects the nearest street line and the panel’s pitch and azimuth (from Google Sunroof) render it visible. See Appendix B for details and visual examples. Our final sample consists of 905,028 households, including 74,663 (8.2%) adopters (see Table 2 for summary statistics).

⁶<http://sunroof.withgoogle.com>.

⁷Our study period begins prior to the period for which we have voter activity data which is used to capture affiliation. We provide robustness checks that (1) exclude pre-2012 adopters from the sample, and (2) use only households that were last sold prior to 2005, both ensuring that voter data is accurate to the relevant households during the bulk of the study period. Expected solar returns differ over years of the study period due to changing (but uniformly applied) incentives and electricity rates. In this respect, our measure of expected value is a suitability index with a dollar value set at 2016 levels.

4.2 Estimation

We estimate a parameterized version of the difference in adoption utility and non-adoption utility, ΔU , adding a logistic error term $\Delta \varepsilon_i$:

$$\begin{aligned} \Delta U_i &= \Delta u_i^o + \Delta u_i^p + \Delta u_i^s + \Delta u_i^{nc} + X_i \cdot \omega_1 \\ &+ \sum_{a \in \{R, U\}} X_i \cdot 1\{A_i = a\} \cdot \omega_{2a} + \psi_{b(i)g(i)} + \Delta \varepsilon_i, \end{aligned} \quad (1)$$

in which we define the outcome utility, self-signaling utility, and the sum of the social signaling and non-conformity utility, respectively, as:

$$\begin{aligned} \Delta u_i^o &\equiv \alpha_0 + \sum_{a \in \{R, U\}} \alpha_{1t} 1\{A_i = a\} + \alpha_2 EV_i + \sum_{a \in \{R, U\}} \alpha_{3a} EV_i \cdot 1\{A_i = a\} \\ &+ \alpha_4 Vis_i + \sum_{a \in \{R, U\}} \alpha_{5a} Vis_i \cdot 1\{A_i = a\} + \alpha_6 Vis_i \cdot EV_i, \\ \Delta u_i^p &\equiv \beta_1 \bar{A}_i + \sum_{a \in \{R, U\}} \beta_{2a} \bar{A}_i \cdot 1\{A_i = a\} + \beta_3 \bar{A}_i \cdot EV_i + \sum_{a \in \{R, U\}} \beta_{4a} \bar{A}_i \cdot EV_i \cdot 1\{A_i = a\}, \\ \Delta u_i^s + \Delta u_i^{nc} &\equiv \gamma_1 Vis_i \cdot \bar{A}_i + \sum_{a \in \{R, U\}} \gamma_{2a} Vis_i \cdot \bar{A}_i \cdot 1\{A_i = a\} + \gamma_3 Vis_i \cdot \bar{A}_i \cdot EV_i. \end{aligned} \quad (2)$$

We re-center $\bar{A}_i$ (the fraction of all other households in the neighborhood in our sample that are registered Republican) and EV_i (economic value, defined as the Google Sunroof-derived financial payoff from adopting solar) on the sample means for ease of interpretation. We combine Democratic and mixed-affiliation households such that affiliation a can take values of D/m , R , or U ; indicating the household head(s) are Democratic/mixed, Republican, or unregistered, respectively. The omitted category is D/m and interaction terms are relative to a Democratic/mixed-affiliation household.

In our main specification, Vis_i is the $[0, 1]$ visibility measure. We include it in the outcome utility since there can be a dis-amenity for visible solar due to aesthetic concerns. The interactions allow it to differ by voter registration and the economic value of installing. The

X_i consists of controls for the number of children, household income, house size, lot size, lot size squared, an indicator for post-1990 construction, assessor value, number of bedrooms, and stories; we allow these controls to have different effects for households of different political affiliation. The $\psi_{b(i)g(i)}$ are census-blockgroup fixed effects that differ by whether a household is registered to vote. This controls for potentially heterogeneous correlated unobservables for unregistered status (e.g., non-citizens). Let $\Delta\tilde{U}_i = \Delta U_i - \Delta\varepsilon_i$. Adoption probabilities take the familiar logit closed form $\Lambda(\Delta\tilde{U}_i)$ and we estimate the model using maximum likelihood.

We connect our testable implications in Table 1 to our empirical specification by noting that for Democratic/mixed-affiliation households:

- If $\beta_1 > 0$, so that adoption is increasing in the Republican affiliation of neighbors, then we have evidence consistent with a positive self-signal identity restoration effect
- If $\beta_3 < 0$, so that the self-signal identity restoration effect is attenuated by increased *EV*, then we have evidence consistent with self-signal crowding-out.
- If $\gamma_1 < 0$, so that visible adoption is decreasing in the Republican affiliation of neighbors, then we have evidence consistent with a negative social non-conformity penalty.
- If $\gamma_3 > 0$, so that the negative social non-conformity penalty is attenuated in *EV*, then we have evidence consistent with social non-conformity crowding-out.

The β_{2a} and γ_{2a} terms for $a \in \{R, U\}$ allow us to test whether the Republican and unregistered households differ in the size of these effects, relative to Democratic/mixed households. We allow the crowding-out effect of self-signaling to differ by political affiliation as well.

These testable hypotheses have the potential to show that the data are consistent with the theoretical predictions from the model. We will discuss alternative mechanisms in Section 5.2.

4.3 Identification

To separate the causal effects consistent with self- and social signaling, we need exogenous variation in three parameters: economic value, neighborhood political affiliation, and visibility. Variation in economic value results from variation in effective solar irradiance through roof orientation, variation in peer political affiliation variation results from the Republican share of the local neighborhood, and variation in visibility results from home orientation. The biggest threats to identification are unobservable factors that affect both adoption and these variables. One mechanism that could lead to such factors would be sorting.

Economic value and political affiliation vary geographically and may therefore be correlated with unobservable determinants of solar adoption. We include blockgroup-level fixed effects interacted with presence of voter registration throughout, so that all parameters are identified by comparing households to others in the same local area, as in the “nearest-neighbor research design” common in the neighborhood effects literature (Bayer et al., 2008; Linden and Rockoff, 2008; Blume et al., 2011; Bayer et al., 2025). Our identification assumption is that unobserved determinants of adoption are uncorrelated with economic value, the political affiliation of the very nearest neighbors, and potential visibility, conditional on extensive observed controls (all interacted with voter affiliation) and blockgroup-by-registered-voter fixed effects. Sorting is thus assumed to occur at the blockgroup level or higher.⁸

While we cannot test for the presence of correlated unobservables directly, we follow Bayer et al. (2008) and estimate correlations of observables between households within a neighborhood, both unconditional and conditional on blockgroup-by-registration fixed effects

⁸See McCartney and Shah (2022), who show in real estate agent interviews that buyers in Los Angeles target larger blockgroup-sized areas rather than specific blocks or streets, making the extremely local endogenous sorting of Manski (1993) unlikely. HOAs are unlikely to exist at the sub-blockgroup scale, and the California Solar Rights Act precludes HOA regulation of solar.

(see Appendix C).⁹ Most household characteristics, including voter affiliation, are uncorrelated once we condition on our fixed effects (Table C.1). Income and house characteristics remain correlated even with our fixed effects (Table C.2), as expected given the “tract housing” nature of homes in our sample; we include these observables as controls, each interacted with voter affiliation.

There remains the concern that sorting on unobservables might happen within the neighborhood level. Sorting may be direct, with households that have unobserved tastes for solar seeking homes with greater irradiance (not captured by our extensive controls) or greater (or lesser) potential visibility. Sorting may also be indirect, where tastes for solar coincide with unobserved tastes for home characteristics correlated with economic value, Republican neighbors, or potentially visible installations—a selection on unobservables problem.

It could be the case that Democrats in more-Republican neighborhoods have stronger unobserved tastes for solar if their homes are less energy-efficient, raising solar’s payoff. Beyond letting all controls (including income and year built) differ by affiliation, we show that Democrats in a blockgroup’s *most* Republican neighborhood do not differ significantly in year built (a proxy for energy efficiency) from those in its *least* Republican neighborhood (Table C.3). To test whether sorting by side-of-street is a concern (which is highly correlated with visibility), we show in a hedonic regression that side-of-street is uncorrelated with both home value and political affiliation (Table C.4, C.5).

5 Results

In Table 3, we present the results for our main coefficients of interest.¹⁰ Throughout, our preferred estimates define neighborhood at the census block-level shown in column (3). Some blockgroups have zero observed installations, which results in dropping 86,942 observations.

⁹We sample a single pair of households randomly selected from each neighborhood, then estimate the correlation between them conditional on the blockgroup fixed effects.

¹⁰Table D.1 in the Appendix shows the full results including covariates.

We perform a range of other robustness checks (Appendix D) including estimating a linear probability model (Table D.2) which preserves the complete sample.

Outcome utility: We find that Republican-registered households have a lower baseline utility for solar, consistent with having a smaller net valuation of being type θ^h . The main effect of visibility is negative for registered-voting households and is not significantly stronger among Republicans. Unregistered voters appear to be an exception, with no negative utility from adopting visible solar. While our results do not allow us to disentangle mechanisms, aesthetic concerns (separate from social signals) are one likely explanation, even among individuals who wish to signal environmental *bona fides*. We calculate the equivalent willingness-to-pay (WTP) using the visibility result and the coefficient on EV (in \$1,000's).¹¹ The equivalent WTP to avoid fully visible panels for a Democratic/mixed household in an average neighborhood is $\frac{-0.091}{-0.192} \times \$1,000 = \$472$, approximately 1.5% of the average system cost.

Self-signal: Adoption is increasing in Republican affiliation of neighbors if $\beta_1 > 0$. For Democratic/mixed households with average EV and zero visibility, the estimated effect in column (3) is positive and statistically-significant at 0.405. We interpret this as causal evidence that Democratic/mixed adoption is affected by the voter affiliation of neighbors, consistent with self-signaling.¹²

Republican households ($\beta_1 + \beta_{2R}$) receive less utility (-0.230, significant at the 10% level) from self-signaling with a point estimate of $0.405 - 0.230 = 0.175$; and we fail to reject the null of no effect for Republican households with a p-value of 0.114. The same result holds in columns (1) and (2). This is consistent with our model, where the utility from self-diagnosing as a “high” type is only weakly stronger for Democrat/mixed than for

¹¹Our EV measure is a snapshot dollar value of solar taken in 2016 where adoptions are counted from 1999-2016. We consider our WTP measures to be a noisy, but informative, estimate.

¹²We explore heterogeneity over income in Table D.3, finding a stronger effect among lower-income households.

Republican. We calculate an equivalent WTP of \$305 for Democratic/mixed households per one-standard-deviation increase (0.15) from the mean (0.20) in the Republican composition of the neighborhood, representing 0.94% of the average system cost and 5.5% of the average *EV*. This WTP is 2.4 times that of the Republican WTP for a similar increase.

Table 3: Main Results

Dependent Variable: Model: Neighborhood definition:	(1) Block x Street	Adopt (2) Blockgroup x Street	(3) Block
<i>Variables</i>			
Republican hh	-0.105* (0.053)	-0.105* (0.053)	-0.100+ (0.053)
Expected value	0.188*** (0.017)	0.191*** (0.017)	0.192*** (0.018)
Expected value $\times$ Republican hh	0.0006 (0.016)	0.002 (0.016)	0.006 (0.017)
Expected value $\times$ Unregistered voter hh	-0.040 (0.026)	-0.041 (0.026)	-0.043 (0.026)
Visible (frac of degrees from closest street) (α_4)	-0.093*** (0.019)	-0.092*** (0.019)	-0.091*** (0.019)
Visible (frac of degrees from closest street) $\times$ Republican hh	-0.006 (0.036)	-0.006 (0.038)	-0.020 (0.040)
Visible (frac of degrees from closest street) $\times$ Unregistered voter hh	0.081* (0.039)	0.079* (0.039)	0.084* (0.039)
Visible (frac of degrees from closest street) $\times$ Expected value	-0.003 (0.018)	-0.006 (0.018)	-0.004 (0.018)
Self-signal utility β			
Republican nbhd (β_1)	0.188*** (0.050)	0.270*** (0.068)	0.405*** (0.088)
Republican nbhd $\times$ Republican hh (β_{2R})	-0.136+ (0.080)	-0.178+ (0.101)	-0.230+ (0.127)
Republican nbhd $\times$ Unregistered voter hh	-0.025 (0.108)	-0.093 (0.131)	0.019 (0.174)
Republican nbhd $\times$ Expected value (β_3)	-0.170** (0.052)	-0.245*** (0.070)	-0.294*** (0.083)
Republican nbhd $\times$ Expected value $\times$ Republican hh	0.032 (0.064)	0.039 (0.081)	0.019 (0.097)
Republican nbhd $\times$ Expected value $\times$ Unregistered voter hh	0.067 (0.078)	0.055 (0.099)	0.134 (0.120)
Social signal utility γ			
Visible (frac of degrees from closest street) $\times$ Republican nbhd (γ_1)	-0.199* (0.093)	-0.261* (0.113)	-0.328* (0.132)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Republican hh (γ_{2R})	0.165 (0.161)	0.208 (0.209)	0.416 (0.254)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Unregistered voter hh	0.168 (0.203)	0.266 (0.235)	0.009 (0.278)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Expected value (γ_3)	0.182* (0.079)	0.259* (0.101)	0.205+ (0.107)
<i>Fixed-effects</i>			
Blockgroup-Registered voter	Yes	Yes	Yes
<i>Fit statistics</i>			
Observations	818,086	818,086	818,086
Pseudo R ²	0.10481	0.10482	0.10486
Log-Likelihood	-223,398.9	-223,396.0	-223,385.9

Conley (2km) standard-errors in parentheses

Signif. Codes: ***, 0.001, **, 0.01, *, 0.05, +: 0.1

Abbreviations used: neighborhood (nbhd), household (hh)

To provide context, we set the neighborhood effect on the self-signal to zero for all voter affiliations and predict the change in the number of installations. Overall adoption declines by 3,855 installations representing 11.2 megawatts of capacity (2.8%). To calculate impacts on emissions, we merge the change in solar adoptions with marginal emissions factors from [Sexton et al. \(2021\)](#), which are expressed in dollars of reduced externalities per unit of solar capacity in a given zip code. Eliminating the self-signal results in a decrease in emissions valued at \$2.0 million per year, equivalent to \$31.8 million over the panel lifespan (discounted at 4%).¹³

The effect of Republican neighbors on adoption is diminishing in EV if $\beta_3 < 0$, indicating “crowding-out”. Results in Table 3 show that the effect is attenuated by EV with a point estimate of $\beta_3 = -0.294$. This is separate from any crowding-out in the “warm glow” captured by the coefficient on EV . In a somewhat surprising finding, we find no significant difference in the crowding-out effect for Republican (β_{4R}) and unregistered households (β_{4U}) relative to Democratic/mixed households. Under a scenario where the crowding-out effect is set to the mean EV for all voter affiliations, installations would increase by 305 and capacity by 0.67 megawatts (0.17%), reducing negative externalities by \$117,646 per year, with a NPV of \$1.91 million. Our results help explain previous research that shows Republican households respond more to the economic value of adopting solar relative to Democrats ([Dokshin and Gherghina, 2024](#)).

Social signal: Visible solar adoption is decreasing in Republican affiliation of neighbors if $\gamma_1 < 0$, controlling for the effect of visibility on the outcome utility (α_4). Our estimate for Democratic/mixed households, γ_1 , is -0.328 and is statistically significant. The effect on the social signal for Republicans is $\gamma_1 + \gamma_{2R} = 0.088$ and is not statistically significant from zero (p-value 0.67). A similar result applies for unregistered households. We interpret this as

¹³Nearly all is due to reduced CO2 emissions, with 10,745 tons/year valued at \$174/ton in 2017 dollars ([Rennert et al., 2022](#)).

causal evidence for a visibility penalty increasing in the Republican affiliation of neighbors, consistent with a social non-conformity penalty for Democratic/mixed households.

The penalty is substantial: the equivalent WTP to avoid the net social signal from a one-standard-deviation increase in Republican neighborhood affiliation for a Democratic/mixed household with an average-visibility and average- EV is $-0.145 \times \frac{-0.328}{(0.192-0.004)} = \$252$. For each Democratic/mixed household we calculate the equivalent value of the change in utility for each observation that would result from zeroing out the neighborhood effect on the net social signal. The average dollar equivalent for Democratic/mixed households is \$120. Eliminating the net social signal would result in 820 (+1.1%) additional adoptions, reducing “dollarized” emissions by \$411,459 per year with a NPV of \$6.7 million. This represents foregone benefits from solar adoption resulting from the neighborhood non-conformity penalty on the social signal.

We also find that as EV increases, the net social signal becomes less negative, since $\gamma_3 > 0$, demonstrating the crowding-out effect. Eliminating the crowding-out effect on the social signal would result in 89 fewer installations and a decrease in total capacity by 0.198 megawatts (0.05%). The reduction in emissions from the change in capacity is valued at \$34,859 with a NPV of \$566,348.

5.1 Robustness Checks

The causal interpretation of our results depends on us ruling out sorting. To further address sorting based on solar preferences, we first run our main analysis restricting the sample to homes that were last sold in 2005 or before, when residential solar was in its infancy (Table D.4). We find results that are very similar to the main results for all three neighborhood definitions.

Following the intuition of Altonji et al. (2005)—that the magnitude of selection on observables informs the potential for selection on unobservables—we also perform a stability test (Table D.5) showing that our estimated coefficients do not change significantly when we

add observable controls, even as R^2 increases substantially. We then provide a breakdown-point analysis as in Altonji et al. (2025) and Oster (2019), using the method of Diegert et al. (2025) (Table D.6). We find that unobservable drivers of adoption would need to explain 72% as much selection into a Republican neighborhood as all other observables combined (including income and home size, conditional on blockgroup) in order to overturn the direct effect of neighborhood affiliation on adoption of non-visible solar.¹⁴

If sorting on unobservables is still a remaining concern, we exploit even more local variation by estimating a robustness check that measures the neighborhood as the four nearest neighbors (4NN). The results shown in Table D.8 columns (1)-(2) are very similar to our main results that define neighborhood as “block-by-street” in Table 3. This also helps rule out confounding effects of solar marketing and sales targeting.¹⁵ While marketers could feasibly target on visibility and neighborhood (e.g., block), targeting at this hyper-local level (e.g., Democratic households with low potential visibility whose next-door neighbors are Republican) is unlikely.

We also include two robustness checks that address potential concerns with our measure of visibility. The visibility measure we use may not capture what neighbors come to know about each other’s installations. Since our visibility measure is based on visibility from the nearest street, even non-visible adoptions are not entirely *invisible* – they might be seen by backyard neighbors, or known through contractor referrals. Low levels of street visibility are a reasonable approximation for backyard neighbor or other similar visibility. To address this (and the possibility of non-linear effects of visibility), we discretize the visibility measure in quartiles and estimate a separate effect for each bin relative to a zero visibility bin. Table

¹⁴We calibrate the breakdown threshold by calculating the breakdown point for all observables *as if* they were unobserved. Only lot size explains more than 72% as much selection into Republican areas. Any unobserved selection would have to be greater than that of income and home size, which do not exceed the threshold (Table D.7).

¹⁵Marketers may target based on observables (income, political affiliation), on a neighborhood (higher-wealth or more Republican areas), or on the potential for a sale to generate further sales, such as visibility itself (Bollinger et al., 2022). In that case the main effect of visibility would be biased upwards. If marketers strategically targeted Democratic households with low potential visibility, we would expect a positive correlation between Republican neighborhoods and Democratic adoption at low visibility but a negative one at high visibility, confounding both our self- and social-signaling interpretations.

D.10 results show that in the lowest quartile of visibility (for visible installations), there is little difference in what we interpret as the self-signaling effect as for non-visible installations. Across all neighborhood definitions, the point estimates for the visibility bin interactions are larger for quartiles 2-4, and highest for quartile 4. The visibility effect is significant only in the most visible bins, with the least visible bins having no significant difference from the zero-visibility bin.

Since low levels of street visibility do not have a significant effect on adoption, we maintain our assumption that backyard neighbors and incidental visibility, both of which may be captured in our self-signal estimate, are not likely an important source of bias.¹⁶

Given the importance of the potential adopter’s *beliefs* about whether their adoption will be visible, we believe our measure of visibility is the most relevant. However, to check robustness to our definition of “visible,” we substitute our measure of visibility with a binary indicator for the north side of the street (Table D.9) and find our results largely unchanged.

Finally, we address concerns that households may chose to strategically site their solar, choosing non-optimal roof segments based on visibility. We check a sample of 100 adopters and compare the observed installation locations to Google Sunroof locations. Of the 100 adopters, 90 installed on at least one of the recommended roof segments. Of the 10 that installed off-recommended segments, only three opted to install on segments whose visibility from the nearest street is substantially different from the recommended roof segments.

These robustness checks help support the causal nature of our estimated effects, including the effect of Republican neighborhood affiliation, its interaction with visibility, and the interaction of these two variables with economic value. These first two effects support our belief that self- and social signaling are important factors in solar adoption, and their interaction with economic value is consistent with expected crowding out effects. We next address potential alternative mechanisms that might explain the effects.

¹⁶If significant, this would be a downward bias of the self-signal effect.

5.2 Alternative Mechanisms

We have provided empirical evidence consistent with positive self-signaling, a net negative social signal, and crowding out of both. Although we cannot entirely rule out all alternative mechanisms behind these results, none straightforwardly predict the estimated causal relationships between $\bar{A}$, Vis , EV , and adoption. We address plausible alternatives here.

Social norms and group identity (Akerlof and Kranton, 2000) are unlikely explanations for our results since they would instead predict lower adoption rates for non-visible installations in more Republican areas. Furthermore, while adopters with more Republican neighbors might self-diagnose as high types against a lower local threshold, this mechanism implies lower adoption utility in Republican areas as a lower threshold implies that less costly actions (e.g., Recycling) permit high-type self-diagnoses, lowering adoption utility—the opposite of our findings.

Altruism is also an unlikely explanation for our estimated effects. We use variation in neighborhood affiliation to identify the effects consistent with signaling, while separately allowing warm glow (impure altruism) and crowding out of the warm glow by EV . Thus, we need only assume that warm glow does not vary with the Republican affiliation of the neighborhood. It is hard to argue why it would. It could be that pure altruists might increase adoption to offset lower perceived aggregate public-good provision in Republican areas, but perceptions of aggregate solar adoption likely form over broader geographies than immediate neighborhoods (much broader than just one’s four nearest-neighbors). As an alternative explanation for the moderating effect of EV , it could be that the EV of adopting affects perceptions about others’ EV , which might reduce the perceived need to provide for the public good. However, this cannot explain why economic value would have a moderating effect on neighborhood affiliation that is of opposite direction depending on whether the installation is visible or not. Therefore, altruism cannot explain our findings.

It could also be the case that households choose to adopt solar to influence diffusion itself. This might be particularly relevant for visible installations. However, given that we

find a negative effect of visibility that does not differ for Republicans and non-Republican registered voters, this cannot explain the positive effect of a more Republican neighborhood on non-visible adoption for non-Republicans. It could partly explain the negative interaction effect with EV for the visible installations, if the perceived need to affect diffusion is smaller when one’s own EV is higher. That said, the relevant economic value for the need to affect diffusion is not one’s own, but rather of the peers one would wish to influence.

There might also be an effect of peer adoption on the household’s decision. Given the documented role of peer influence in solar diffusion (Bollinger and Gillingham, 2012; Bollinger et al., 2022), if there is a correlation between peer installations and any of our key regressors, we might be prone to omitted variable bias.

To address both households aiming to influence diffusion and peer adoption affecting households, we test a specification in which we exclude early adopters, who may be more motivated by the desire to influence the diffusion process. We then create a variable for all other households that measures the number of these early adopters in the peer group, to control for their impact via peer influence. Results are in Table D.11. We find a positive effect of the early peer adoptions, but the main results all hold.

Our preferred interpretation of the negative coefficient on the interaction between Republican share and EV (β_3) is that crowding out is due to monetary motivations dominating signaling by diluting the self-signal strength. An alternative mechanism could be that consumers who purchase higher value systems rely less on information and that neighborhoods with a higher Republican share have more information. However, it seems unlikely that neighborhoods with a higher Republican share have more information than neighborhoods with a lower Republican share.

To further support the self- and social signaling interpretations, we estimate the model on our sample after splitting at the median blockgroup density (Table D.12). The internalization of neighborhood affiliation as a prior belief for own-type should be stronger when houses are closer together and interactions are more frequent and salient. As expected, higher-density

blockgroups have larger estimated effects of Republican neighborhood (self-signal) and *Vis* interacted with Republican neighborhood (net social signal).

While we cannot entirely rule out every alternative, our reduced-form evidence demonstrates that public-good decisions respond to identity-relevant signals shaped by neighboring environments—consistent with both self- and social signaling.

6 Conclusion

We find empirical evidence demonstrating the causal role of the local social context in solar adoption, which provides evidence consistent with both self- and social signaling. We also show evidence of “crowding out” of both signals by the private payoff of adopting solar, also consistent with our model predictions.

The economic magnitudes of these effects are notable. The self-signal provides positive utility when the local context is dominated by a type less likely to adopt prosocial behavior. Most significant are the countervailing effects of visibility: while Republican neighborhoods provide greater self-signaling utility for Democratic/mixed households, visible installations incur an attenuating non-conformity penalty. Because self-signal crowding out is smaller in magnitude than the social signal effect, emphasizing privately-captured benefits could feasibly yield positive effects on adoption.

Our findings provide new insights into the motivations and conditions for prosocial behavior in general, especially in contexts where self-signals and social signals may be in opposition or politically polarized. Ignoring these motivations may confound estimates of willingness-to-pay for solar and may misattribute unresponsiveness to the private payoff stemming from the self-signal to irrationality or partisan bias. Our results point towards the value of developing attributes of solar that make it less conspicuous. This study informs where, and under what conditions, prosocial behavior is most likely, and helps clarify the

effects of increased polarization. Better understanding of the channels by which adopters update their priors may help point to effective policies.

Mandatory Generative AI Disclosure

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A Mathematical Appendix: A Model of Self- and Social Signaling

This appendix provides the formal structure, equilibrium definitions, underlying assumptions, and complete analytical derivations for the model of self- and social signaling through prosocial technology adoption.

A.1 Setup and Utility Framework

An agent considers adopting a prosocial technology (e.g. Solar). The adoption decision is given by $x \in \{0, 1\}$, in which $x = 1$ indicates adoption. Key factors that affect the utility gained from adopting are the individual's ideological identity, given by their voter affiliation, $A \in \{D, R\}$ indicating Democratic or Republican, the individual's prosocial type, $\theta \in \{\theta^h, \theta^l\}$ indicating high or low, the privately-captured lifetime expected pecuniary value of the installation net of installation costs, $EV \in \mathbb{R}$, the visibility of a (potential) solar installation to nearest public street, $Vis \in \{0, 1\}$, and the voter affiliation of their neighbors, $\bar{A} \in [0, 1]$, defined as the fraction of neighbors that are registered Republicans.

The agent has additively separable utility U over x comprising *outcome utility*, *self-signal utility*, *social signal utility* and a *non-conformity penalty*. We write U , exclusive of a private type 1 extreme value shock, as:

$$\begin{aligned} U(x, A, \bar{A}, EV, Vis, \theta) &= u^o(x, A, EV, Vis, \theta) + u^p(x, A, \bar{A}, EV) \\ &+ u^s(x, A, \bar{A}, EV, Vis) + u^{nc}(x, \bar{A}, EV, Vis) \end{aligned} \quad (3)$$

where each component is:

Outcome utility u^o Outcome utility includes the economic value of adopting (EV), impure altruism utility (Andreoni, 1989, 1990; Ottoni-Wilhelm et al., 2017) received by choosing to adopt solar based on the individual's type and voter affiliation, and the visibility of the potential solar installation (Vis), due to the household's aesthetic tastes.¹⁷

¹⁷We do not take a stand on whether EV crowds out impure altruism. Our results are robust to either assumption.

Self-signal utility u^ρ : Type “exerts influence at the moment of choice, but cannot be deduced merely by imagining what one might do in a given situation” (Bodner and Prelec, 2003) and as in Bénabou and Tirole (2011).¹⁸ The agent relies on previous, concrete actions taken to allow self-diagnosis of type. The expected utility derived from future self-diagnosis is the self-signal utility, which we write as:

$$u^\rho(x) = \rho(x) \cdot v^\rho(\theta^h|A) + (1 - \rho(x)) \cdot v^\rho(\theta^l|A), \quad (4)$$

where $v^\rho(\theta|A)$ is the psychological payoff of diagnosing oneself as type θ and ρ is the probability the agent is the high type. Action x leads the agent to update ρ in a way that may be influenced by internal factors, including own political affiliation A , and external factors, including $\bar{A}$. $v(\cdot)$ is scaled to units of outcome utility.

Social signal utility u^s : The utility derived from others’ perceptions of the agent is the social signal utility. We write u^s , expected social signal utility, analogous to u^ρ :

$$u^s(x, A, \bar{A}, EV, Vis) = Vis \cdot [s(x) \cdot v^s(\theta^h|A) + (1 - s(x))v(\theta^l|A)] \quad (5)$$

where s is the probability that others believe the agent to be the high type, and $v^s(\theta|A)$ is the agent’s psychological value of others perceiving them as type θ . s may be updated by action x if the action is visible.¹⁹ v^s is scaled to units of outcome utility.

Non-conformity utility u^{nc} : In addition to the social signal, agents may suffer a non-conformity penalty from adopting a visible installation proportional to the deviation from the local neighborhood political affiliation. We write expected non-conformity utility, u^{nc} , as:

$$u^{nc} = Vis \cdot V^{nc} \cdot |s(x) - (1 - \bar{A})| \quad (6)$$

where $V^{nc} < 0$ is the penalty from deviating from a local norm (Bernheim, 1994; Akerlof and Kranton, 2000; Austen-Smith and Fryer, 2005; Bursztyn and Jensen, 2015; Brick et al., 2017; Perez-Truglia, 2018; Pryor et al., 2019).²⁰ u^{nc} is expected utility as the agent is

¹⁸We depart by setting $1 - \lambda = 1$, where $1 - \lambda$ controls the importance of past actions in self-diagnosis. While simpler, our version sufficiently motivates the behavior we examine empirically.

¹⁹This implies social signal utility has the potential to “shame” (Butera et al., 2022) in public recognition when non-adoption may result in lower utility relative to pre-decision.

²⁰ V^{nc} may also come from expected costs of deviating from the norm, and is not limited to pure social discomfort. For example, if an individual cares not about social pressure but believes that the future sale price of their home will be impacted by visible solar, the non-conformity penalty may be an internalization of social norms via this expected eventual transaction.

uncertain over how they are perceived by others, but the agent's actions update the perception. Non-conformity utility is zero when $Vis = 0$. V^{nc} scales the deviation in absolute value to units of outcome utility.

Bayesian Updating of ρ and s : An agent holds an initial prior belief regarding their type, $\rho_0(A, \bar{A})$, and an initial prior belief regarding how others perceive their type, $s_0(\bar{A})$, which is a function of the neighborhood political affiliation such that $s_0 = 1 - \bar{A}$. This corresponds to the prior that an individual in a neighborhood, absent new information, is perceived to be the average of the neighborhood. Since type is not explicitly visible, the fraction of the neighborhood registered as a Democrat, $1 - \bar{A}$, establishes the prior s_0 .

Upon taking an action (or not) $x \in \{0, 1\}$, posterior beliefs $\hat{\rho}(x)$ and observer beliefs $\hat{s}(x)$ update via Bayes' Rule:

$$\hat{\rho}(x = 1) = \frac{\rho_0 Pr_H}{\bar{P}}, \quad \hat{\rho}(x = 0) = \frac{\rho_0(1 - Pr_H)}{1 - \bar{P}} \quad (7)$$

$$\hat{s}(x = 1) = \frac{(1 - \bar{A})Pr_H}{\bar{P}^s}, \quad \hat{s}(x = 0) = \frac{(1 - \bar{A})(1 - Pr_H)}{1 - \bar{P}^s} \quad (8)$$

in which we drop the functional arguments for readability. We define $Pr_\theta(\theta, A, \bar{A}, EV, Vis) \equiv Pr(x = 1 | \theta) \in (0, 1)$, which denotes the choice probability of $x = 1$ for type $\theta \in \{\theta^h, \theta^l\}$. $\bar{P}(\theta, A, \bar{A}, EV, Vis) = \rho_0 Pr_H + (1 - \rho_0) Pr_L$ is the unconditional expected choice probability using the prior, and $\bar{P}^s(\theta, A, \bar{A}, EV, Vis) = s_0 Pr_H + (1 - s_0) Pr_L$ is others' expectations about the agent's choice probability using their prior, which differs from $\bar{P}$ because the agent has private information in ρ_0 .

Assumption 1 (Diagnostic Preference Heterogeneity). *Democrats place a higher (or equal) relative valuation on self- and social type diagnosis than Republicans:*

$$\begin{aligned} \tilde{v}^\rho(D) &\equiv v^\rho(\theta^H | D) - v^\rho(\theta^L | D) \geq v^\rho(\theta^H | R) - v^\rho(\theta^L | R) \equiv \tilde{v}^\rho(R) \geq 0 \\ \tilde{v}^s(D) &\equiv v^s(\theta^H | D) - v^s(\theta^L | D) \geq v^s(\theta^H | R) - v^s(\theta^L | R) \equiv \tilde{v}^s(R) \geq 0 \end{aligned}$$

A.2 Net utility of adopting

Let Δ indicate the difference between adopting solar, $x = 1$, and not adopting, $x = 0$. $\Delta U = \Delta u^o + \Delta u^\rho + \Delta u^s + \Delta u^{nc}$ where $\Delta u^o = u^o(1, A, EV, Vis, \theta) - u^o(0, A, EV, Vis, \theta)$. For other components, choosing to adopt updates ρ and s as follows:

Using (4) and (7) we have:

$$\Delta u^p(A, \bar{A}, EV) = \tilde{v}^p(A) \cdot \frac{\rho_0(A, \bar{A}) [1 - \rho_0(A, \bar{A})] [Pr_H - Pr_L]}{\bar{P} [1 - \bar{P}]} \quad (9)$$

Using (5) and (8) yields:

$$\Delta u^s(\bar{A}, EV) = \tilde{v}^s(A) \cdot \frac{(1 - \bar{A})\bar{A} [Pr_H - Pr_L]}{\bar{P}^s [1 - \bar{P}^s]} \quad (10)$$

The resulting net non-conformity differential using (6) is:

$$\begin{aligned} \Delta u^{nc}(\bar{A}, EV) &= V^{nc} \cdot |\hat{s}(1) - (1 - \bar{A})| - |\hat{s}(0) - (1 - \bar{A})| \\ &= V^{nc} \cdot \Delta NC \\ &= V^{nc} \cdot \frac{\bar{A}(1 - \bar{A})(Pr_H - Pr_L)(1 - 2\bar{P}^s)}{\bar{P}^s(1 - \bar{P}^s)} \end{aligned} \quad (11)$$

Assumption 2 (Net outcome utility). *The outcome utility differential $\Delta u^o(A, EV, Vis, \theta)$ satisfies:*

$$\Delta u^o|\theta^h > \Delta u^o|\theta^l, \quad \frac{\partial \Delta u^o}{\partial EV} > 0, \quad \frac{\partial^2 \Delta u^o}{\partial \bar{A} \partial EV} = 0, \quad \text{and} \quad \frac{\partial \Delta u^o}{\partial \bar{A}} = 0. \quad (12)$$

This first assumption is needed to ensure that $Pr_H > Pr_L$ and would follow from high types receiving more warm glow utility than low types. Assuming $\frac{\partial^2 \Delta u^o}{\partial \bar{A} \partial EV} = 0$ guarantees that interaction effects ($\frac{d^2 \Delta U}{d\bar{A} dEV}$) are driven entirely by endogenous signaling and non-conformity mechanisms rather than arbitrary mechanical baseline curvature. $\frac{\partial \Delta u^o}{\partial \bar{A}} = 0$ assumes that the neighborhood political affiliation does not enter the outcome utility directly.

Remark 1. *The last assumption implies that all signaling value from non-visible installations is self-signaling. It is possible that non-visible installations are still socially relevant. If this were the case, some of the self-signal effect that we will estimate also captures some degree of social signal that is not subject to the same conformity disutility that we find as visible solar panels.*

Assumption 3 (The prior). *We assume prior beliefs about one's type are influenced by the political affiliation of neighbors such that $\rho_0(A, \bar{A})$ is decreasing in $\bar{A}$. Social psychology*

research shows that group interactions shape self-perception (*Allport, 1954; Turner and Onorato, 1999; Turner et al., 2008*). Living amongst and interacting with neighbors perceived to be θ^l through their political affiliation leads to a prior ρ_0 that places lower probability on being type θ^h as one wonders “I am surrounded by a lot of θ^l , maybe I’m actually a θ^l ?” Under this assumption, we have that.²¹

$$\frac{\partial \rho_0(A, \bar{A})}{\partial \bar{A}} < 0 \quad \text{for all } A \in \{R, D\} \quad (13)$$

This implies $\partial \Delta u^\rho / \partial \bar{A} > 0$, as the signal carries greater meaning when adopting in a high- $\bar{A}$ (heavily Republican) neighborhood than it does in a low- $\bar{A}$ neighborhood. We refer to this as the self-signal identity restoration effect.

We also assume that one’s own political identity affects one’s initial beliefs such that $\partial \rho_0 / \partial A < 0$.

We normalize the variance of the additional private i.i.d. type-1 extreme value preference shock $\epsilon_{i,x}$ to 1. This yields the standard Logit choice probabilities for type $\theta \in \{H, L\}$:

$$Pr_\theta(A, \bar{A}, EV, Vis) = \frac{1}{1 + \exp(-\Delta U(A, \bar{A}, EV, Vis, \theta))} \quad (14)$$

We make one final assumption bounding the probability of taking action $x = 1$.

Assumption 4 (Choosing $x = 1$ is an infrequent action).

$$Pr_L(A, \bar{A}, EV, Vis) < Pr_H(A, \bar{A}, EV, Vis) < 0.5 \quad (15)$$

We assume prosocial technology adoption is an infrequent action across both types.²²

²¹Kaikati et al. (2017) notes the contrast between differing ideology (e.g., D vs. R) and shared social identity, e.g., neighbors in affecting charitable donations.

²²See Pallottini (2025) for more on the role of prevalence of adoption in social signals.

A.3 Equilibrium

Definition 1 (Choice probabilities). *We define the vector of choice probabilities $\mathbf{Pr}^* = (Pr_H^*, Pr_L^*)$ such that for each type $\theta \in \{H, L\}$:*

$$Pr_\theta^* = \frac{1}{1 + \exp(-\Delta U(A, \bar{A}, EV, Vis, \theta; \mathbf{Pr}^*))} \quad (16)$$

where posterior beliefs $\hat{\rho}(x; \mathbf{Pr}^*)$ and $\hat{s}(x; \mathbf{Pr}^*)$ are evaluated using equilibrium probabilities $\mathbf{Pr}^*$.

Since $\Delta u^o|\theta^h > \Delta u^o|\theta^l$, and this is the only utility component that depends on actual type rather than beliefs about type, we have that $\Delta U(\theta^h) > \Delta U(\theta^l)$. This means that $Pr_H > Pr_L$. This yields the result that, while the agent with high $\tilde{v}^\rho(A)$ gains expected utility from self-diagnosing as a high type with greater probability, their perception of Pr_H , the probability an observationally equivalent individual who is type θ^h , is greater than Pr_L , without knowing who is θ^h and who is θ^l .

Lemma 1 (Existence and uniqueness of equilibrium). *For any environmental state $(A, \bar{A}, EV, Vis)$, an equilibrium fixed point $\mathbf{Pr}^*$ exists by Brouwer's Fixed Point Theorem. Furthermore, the equilibrium is locally unique provided the following feedback condition holds:²³*

$$1 - \Omega^\rho(A, \bar{A}, EV, Vis) - Vis \cdot [\Omega^s(A, \bar{A}, EV) + \Omega^{nc}(\bar{A}, EV)] > 0 \quad (17)$$

Ω^ρ , Ω^s , and Ω^{nc} are derived below.

Remark 2. *By Corollary 1 and Lemma 4, all three feedback terms are non-negative wherever active: $\Omega^\rho > 0$ for $\rho_0 > \rho_0^*$, $\Omega^s > 0$ for $\bar{A} < 1 - \rho_0^*$, and $\Omega^{nc} \geq 0$ unconditionally. This means that self-signaling, social signaling, and non-conformity penalties are all self-reinforcing (rather than self-correcting) feedback channels in this model: none of the three mechanisms dampens the equilibrium response to $\bar{A}$, so the risk of the fixed point losing local uniqueness (and with it, the validity of the Implicit Function Theorem argument underlying Propositions 1–3) is concentrated specifically in the $Vis = 1$ case.*

²³“Locally unique” reflects what the Implicit Function Theorem actually delivers: at any fixed point satisfying the stated condition, that equilibrium is isolated and the fixed-point correspondence is locally differentiable there — exactly what licenses the comparative statics derived throughout this appendix. It does not rule out the existence of other, distant equilibria satisfying the same condition independently.

A.4 Equilibrium Comparative Statics

Our model explains changes in adoption when $\bar{A}$ operates through the self-signal utility, and when visible, the social signal utility and the non-conformity utility as well. To sign the change in adoption expected under an exogenous increase in $\bar{A}$, we derive $d\Delta U/d\bar{A}$. So that we may consider equilibrium effects, we take each utility component's derivative with respect to Pr_θ .

Differentiating Δu^ρ with respect to ρ_0 and choice probabilities Pr_H and Pr_L yields:

$$\Psi_{\rho_0}^\rho(A, \bar{A}, EV, Vis) \equiv \frac{\partial \Delta u^\rho}{\partial \rho_0} = \tilde{v}^\rho(A) \cdot \frac{Pr_H Pr_L (1 - \bar{P})^2 - (1 - Pr_H)(1 - Pr_L) \bar{P}^2}{[\bar{P}(1 - \bar{P})]^2} \quad (18)$$

$$\Psi_{Pr_H}^\rho(A, \bar{A}, EV, Vis) \equiv \tilde{v}^\rho(A) \cdot \rho_0(1 - \rho_0) \left[\frac{Pr_L}{\bar{P}^2} + \frac{1 - Pr_L}{(1 - \bar{P})^2} \right] > 0 \quad (19)$$

$$\Psi_{Pr_L}^\rho(A, \bar{A}, EV, Vis) \equiv -\tilde{v}^\rho(A) \cdot \rho_0(1 - \rho_0) \left[\frac{Pr_H}{\bar{P}^2} + \frac{1 - Pr_H}{(1 - \bar{P})^2} \right] < 0 \quad (20)$$

Differentiating Δu^s with respect to s_0 and choice probabilities:

$$\Psi_{s_0}^s(\bar{A}, EV, Vis) \equiv \frac{\partial \Delta u^s}{\partial s_0} = \tilde{v}^s(A) \cdot \frac{Pr_H Pr_L (1 - \bar{P}^s)^2 - (1 - Pr_H)(1 - Pr_L)(\bar{P}^s)^2}{[\bar{P}^s(1 - \bar{P}^s)]^2}$$

$$\Psi_{Pr_H}^s(\bar{A}, EV, Vis) \equiv \tilde{v}^s(A) \cdot (1 - \bar{A})\bar{A} \left[\frac{Pr_L}{(\bar{P}^s)^2} + \frac{(1 - Pr_L)}{(1 - \bar{P}^s)^2} \right] > 0$$

$$\Psi_{Pr_L}^s(\bar{A}, EV, Vis) \equiv -\tilde{v}^s(A) \cdot (1 - \bar{A})\bar{A} \left[\frac{Pr_H}{(\bar{P}^s)^2} + \frac{(1 - Pr_H)}{(1 - \bar{P}^s)^2} \right] < 0$$

Differentiating Δu^{nc} with respect to $\bar{A}$ and choice probabilities:

$$\Psi^{nc}(\bar{A}, EV, Vis) \equiv \frac{\partial \Delta u^{nc}}{\partial \bar{A}} = V^{nc} \cdot \left[2 - \frac{Pr_H Pr_L}{(\bar{P}^s)^2} - \frac{(1 - Pr_H)(1 - Pr_L)}{(1 - \bar{P}^s)^2} \right]$$

$$\Psi_{Pr_H}^{nc}(\bar{A}, EV, Vis) \equiv V^{nc} \left[\frac{\Delta NC}{Pr_H - Pr_L} - \frac{\bar{A}(1 - \bar{A})^2(Pr_H - Pr_L)(2(\bar{P}^s)^2 - 2\bar{P}^s + 1)}{[\bar{P}^s(1 - \bar{P}^s)]^2} \right]$$

$$\Psi_{Pr_L}^{nc}(\bar{A}, EV, Vis) \equiv V^{nc} \left[-\frac{\Delta NC}{Pr_H - Pr_L} - \frac{\bar{A}^2(1 - \bar{A})(Pr_H - Pr_L)(2(\bar{P}^s)^2 - 2\bar{P}^s + 1)}{[\bar{P}^s(1 - \bar{P}^s)]^2} \right]$$

Applying logit choice variance $\frac{dPr_\theta}{d\Delta U} = Pr_\theta(1 - Pr_\theta)$ defines the self-signal, social signal, and non-conformity equilibrium feedback coefficients $\Omega^\rho, \Omega^s, \Omega^{nc}$.²⁴

²⁴In the proofs, equations (25)–(27) treat the equilibrium feedback through the scalar multiplier $1/(1 - \Omega)$ (and its visible installation analogue), obtained by applying the Implicit Function Theorem as if (Pr_H^*, Pr_L^*) adjusted along a single aggregate channel. A fully rigorous treatment of the fixed point in Definition 1 requires

$$\Omega^\rho(A, \bar{A}, EV, Vis, \theta) \equiv \Psi_{Pr_H}^\rho Pr_H(1 - Pr_H) + \Psi_{Pr_L}^\rho Pr_L(1 - Pr_L) \quad (21)$$

$$\Omega^s(\bar{A}, EV, Vis) \equiv \Psi_{Pr_H}^s Pr_H(1 - Pr_H) + \Psi_{Pr_L}^s Pr_L(1 - Pr_L) \quad (22)$$

$$\begin{aligned} \Omega^{nc}(\bar{A}, EV, Vis) &\equiv \Psi_{Pr_H}^{nc} Pr_H(1 - Pr_H) + \Psi_{Pr_L}^{nc} Pr_L(1 - Pr_L) \\ &= V^{nc}(Pr_H - Pr_L) \bar{A}(1 - \bar{A}) \left[\frac{1 - 2\bar{P}^s}{\bar{P}^s(1 - \bar{P}^s)} (1 - Pr_H - Pr_L) - Q \cdot W \right] \end{aligned} \quad (23)$$

where

$$Q \equiv \frac{2(\bar{P}^s)^2 - 2\bar{P}^s + 1}{[\bar{P}^s(1 - \bar{P}^s)]^2} > 0$$

$$W \equiv (1 - \bar{A}) Pr_H(1 - Pr_H) + \bar{A} Pr_L(1 - Pr_L) > 0$$

We establish the sign of Ω^{nc} in Lemma 4

A.4.1 Sign of $\Psi_{\rho_0}^\rho$, $\Psi_{s_0}^s$, Ω^ρ , and Ω^s

The sign of $\partial \Delta u^\rho / \partial \bar{A}$ in (18) determines the first-order and equilibrium feedback effects of $\bar{A}$ on adoption utility ΔU . We solve for the threshold above which $\Psi_{\rho_0}^\rho < 0$. Setting $\Psi_{\rho_0}^\rho = 0$ and solving for ρ_0 yields the explicit threshold prior ρ_0^* :

$$\rho_0^*(A, \bar{A}, EV, Vis) = \frac{\frac{\sqrt{Pr_H Pr_L}}{\sqrt{Pr_H Pr_L} + \sqrt{(1 - Pr_H)(1 - Pr_L)}} - Pr_L}{Pr_H - Pr_L} \quad (24)$$

Lemma 2 (Most priors $\rho_0 < \rho_0^*$). *Under Assumption 4, the critical threshold belief satisfies $\rho_0^*(A, \bar{A}, EV, Vis) < 0.5$. Consequently, the “diagnostic regime” $\rho_0 > \rho_0^*$, where $\Psi_{\rho_0}^\rho < 0$, expands to encompass the vast majority of prior belief space.*

differentiating *both* equilibrium conditions in Pr_H^* and Pr_L^* jointly with respect to $\bar{A}$, yielding a 2×2 linear system in $(dPr_H^*/d\bar{A}, dPr_L^*/d\bar{A})$ whose solution involves inverting the full Jacobian of cross-type feedback effects, rather than dividing by a scalar. The scalar multiplier used here is exact only in the special case where the two types’ equilibrium responses are proportional (i.e., $dPr_L^*/d\bar{A} = \lambda dPr_H^*/d\bar{A}$ for some fixed λ), which holds, for example, under a representative-agent aggregation or when the Jacobian’s dominant eigenvector aligns with the vector of direct effects $(\partial \Delta U_H / \partial \bar{A}, \partial \Delta U_L / \partial \bar{A})$. Absent this restriction, Ω should be read as a reduced-form approximation to the dominant feedback channel rather than an exact multiplier; it captures the correct sign of the comparative statics under Assumptions 1–4, but not necessarily their precise magnitude. A full 2×2 treatment is left for future work. Allowing Ω to vary by party A via party-specific priors $\rho_0(A, \bar{A})$ and $\tilde{v}(A)$ or party-specific choice variances $Pr_{\theta, A}(1 - Pr_{\theta, A})$ introduces higher-order non-linear feedback effects in the equilibrium denominator. However, because these variance terms affect only the magnitude of multiplier feedback and leave all qualitative comparative static signs unchanged, we evaluate $\Omega^\rho(A, \bar{A}, EV, Vis)$, $\Omega^s(A, \bar{A}, EV)$, and $\Omega^{nc}(\bar{A}, EV)$ at the aggregate level for expositional clarity.

Corollary 1 (Sign of the equilibrium feedback coefficients Ω^ρ and Ω^s). *Under Assumptions 1–4, the coefficients Ω^ρ and Ω^s defined in equations (21)–(22) satisfy*

$$\Omega^\rho(A, \bar{A}, EV, Vis, \theta) = -\rho_0(1 - \rho_0)(Pr_H - Pr_L) \Psi_{\rho_0}^\rho(A, \bar{A}, EV, Vis, \theta),$$

$$\Omega^s(\bar{A}, EV, Vis) = -s_0(1 - s_0)(Pr_H - Pr_L) \Psi_{s_0}^s(\bar{A}, EV, Vis),$$

Consequently, since $\rho_0(1 - \rho_0) > 0$, $s_0(1 - s_0) > 0$, and $Pr_H - Pr_L > 0$ by Assumption 4:

$$\Omega^\rho > 0 \iff \Psi_{\rho_0}^\rho < 0 \iff \rho_0 > \rho_0^*, \quad \Omega^s > 0 \iff \Psi_{s_0}^s < 0 \iff s_0 > \rho_0^*,$$

where ρ_0^* is the threshold of equation (24), which depends only on Pr_H, Pr_L and therefore governs both conditions identically. Substituting $s_0 = 1 - \bar{A}$, the second condition is equivalently

$$\Omega^s > 0 \iff \Psi_{s_0}^s < 0 \iff \bar{A} < 1 - \rho_0^*.$$

By Lemma 2, the regime $\rho_0 > \rho_0^*$ encompasses the vast majority of prior belief space under Assumption 4; hence $\Omega > 0$ holds generically. No analogous claim is available for Ω^s : because $\bar{A}$ is an exogenous environmental parameter rather than an agent belief, Ω^s is positive only for $\bar{A} < 1 - \rho_0^*$ and negative thereafter. This threshold is distinct in general from the non-conformity turning point $\bar{A}^*$ of Lemma 3.

Proof. Substituting equations (19)–(20) into the definition of Ω^ρ and collecting terms over the common denominators $\bar{P}^2$ and $(1 - \bar{P})^2$ yields, after simplification,

$$\Omega^\rho(A, \bar{A}, EV, Vis, \theta) = \tilde{v}^\rho(A) \cdot \rho_0(1 - \rho_0)(Pr_H - Pr_L) \left[\frac{(1 - Pr_H)(1 - Pr_L)}{(1 - \bar{P})^2} - \frac{Pr_H Pr_L}{\bar{P}^2} \right].$$

Comparing the bracketed term to the numerator of $\Psi_{\rho_0}^\rho$ in equation (18) shows the bracket equals $-\Psi_{\rho_0}^\rho$ exactly, giving the first identity. The identical algebra applied to $\Psi_{Pr_H}^s, \Psi_{Pr_L}^s$ yields the second identity, with $\rho_0 \rightarrow s_0$ and $\bar{P} \rightarrow \bar{P}^s$ throughout. The sign equivalences follow immediately from Assumption 4 ($Pr_H > Pr_L$) and the definition of ρ_0^* as the unique root of $\Psi_{\rho_0}^\rho = 0$ (equivalently $\Psi_{s_0}^s = 0$, since the defining equation depends only on Pr_H, Pr_L). $\square$

A.5 Formal Propositions and Proofs

Proposition 1 (Visibility and political affiliation divergence). *Under Assumptions 1–4:*

1. **For Private Goods** ($Vis = 0$): Self-signal identity restoration dominates, yielding a positive marginal expected utility response to neighborhood political affiliation:

$$\left. \frac{d\Delta U}{d\bar{A}} \right|_{Vis=0} = \frac{1}{1 - \Omega^\rho(A, \bar{A}, EV, 0)} \left[\frac{\partial \Delta u}{\partial \bar{A}} + \Psi_{\rho_0}^\rho \frac{\partial \rho_0}{\partial \bar{A}} \right] > 0 \quad (25)$$

Furthermore, Democrats exhibit a stronger marginal identity restoration response than Republicans:

$$\left. \frac{d\Delta U_D}{d\bar{A}} \right|_{Vis=0} \geq \left. \frac{d\Delta U_R}{d\bar{A}} \right|_{Vis=0} > 0 \quad (26)$$

2. **For Visible Goods** ($Vis = 1$): Non-conformity penalties oppose self-signal identity restoration:

$$\left. \frac{d\Delta U}{d\bar{A}} \right|_{Vis=1} = \frac{\frac{\partial \Delta u}{\partial \bar{A}} + \Psi_{\rho_0}^\rho \frac{\partial \rho_0}{\partial \bar{A}} + \Psi_{s_0}^s \frac{\partial s_0}{\partial \bar{A}} + \Psi^{nc}}{1 - \Omega^\rho(A, \bar{A}, EV, 1, \theta) - \Omega^s(\bar{A}, EV, 1) - \Omega^{nc}(\bar{A}, EV, 1)} \quad (27)$$

Provided $\bar{A} < \bar{A}^*$ (the non-conformity turning point of Lemma 3) **and** $\bar{A} < 1 - \rho_0^*$ (equivalently $\Omega^s \geq 0$, per Corollary 1), if political non-conformity penalties $|V^{nc}|$ are sufficiently severe, the comparative static decreases:

$$\left. \frac{d\Delta U}{d\bar{A}} \right|_{Vis=1} < \left. \frac{d\Delta U}{d\bar{A}} \right|_{Vis=0} \quad (28)$$

Proof. For $Vis = 0$, social signaling and non-conformity terms vanish ($\Delta u^s = 0$ and $\Delta u^{nc} = 0$). Applying Lemma 1 and the Implicit Function Theorem across equilibrium fixed points yields equation (25). Under Assumption 4 and Lemma 2, agents operate in the regime where $\Psi_{\rho_0}^\rho < 0$. Combined with Assumption 3 ($\frac{\partial \rho_0}{\partial \bar{A}} < 0$), the self-signal identity restoration term is strictly positive: $\Psi_{\rho_0}^\rho \frac{\partial \rho_0}{\partial \bar{A}} > 0$. Because $\tilde{v}^\rho(D) \geq \tilde{v}^\rho(R)$ by Assumption 1, the positive self-signal identity restoration effect is weakly larger for Democrats.

For $Vis = 1$, applying the Implicit Function Theorem across equilibrium fixed points accounts for the endogenous equilibrium response of choice probabilities (Pr_H, Pr_L) to shifting neighborhood political affiliation $\bar{A}$, yielding equation (27). When the non-conformity penalty weight $|V^{nc}|$ is sufficiently large relative to signaling returns, and provided $\bar{A} < 1 - \rho_0^*$ (equivalently, $\Omega^s \geq 0$ per Corollary 1), (28) holds only if $\Psi^{nc} = \frac{\partial \Delta u^{nc}}{\partial \bar{A}} < 0$ (see Lemma 3).²⁵ $\square$

²⁵An additional condition for this to hold in all cases is that $\Omega^\rho(\cdot, 1) \approx \Omega^\rho(\cdot, 0)$, which we maintain as a simplifying assumption.

Lemma 3 (Non-monotonicity of Ψ^{nc}). *Define the observer-consistent aggregate adoption rate $\bar{P}^s(\bar{A}) \equiv s_0 Pr_H + (1 - s_0) Pr_L$, where $s_0 = 1 - \bar{A}$, so that*

$$\bar{P}^s(\bar{A}) = Pr_H - \bar{A}(Pr_H - Pr_L).$$

Under Assumptions 3–4, the non-conformity utility differential satisfies

$$\Delta u^{nc}(\bar{A}) \leq 0 \quad \text{for all } \bar{A} \in [0, 1],$$

with strict inequality on the open interval $(0, 1)$, and $\Delta u^{nc}(0) = \Delta u^{nc}(1) = 0$. However, $\Delta u^{nc}(\bar{A})$ is not monotone on $(0, 1)$. Its derivative is

$$\Psi^{nc}(\bar{A}) \equiv \frac{\partial \Delta u^{nc}}{\partial \bar{A}} = V^{nc} \left[2 - \frac{Pr_H Pr_L}{(\bar{P}^s)^2} - \frac{(1 - Pr_H)(1 - Pr_L)}{(1 - \bar{P}^s)^2} \right].$$

There exists a unique interior threshold $\bar{A}^* \in (0, 1)$, implicitly defined by

$$\frac{Pr_H Pr_L}{(\bar{P}^{s*})^2} + \frac{(1 - Pr_H)(1 - Pr_L)}{(1 - \bar{P}^{s*})^2} = 2, \quad \bar{P}^{s*} \equiv \bar{P}^s(\bar{A}^*), \quad (29)$$

that is unaffected by the scaling constant V^{nc} , which cancels from the defining condition such that, since $V^{nc} < 0$,

$$\Psi^{nc}(\bar{A}) \leq 0 \quad \text{as } \bar{A} \leq \bar{A}^*.$$

Proof. The boundary values follow directly from the factor $\bar{A}(1 - \bar{A})$ in (11), which is equal to zero at both $\bar{A} = 0$ and $\bar{A} = 1$ and is strictly positive on $(0, 1)$; combined with $\bar{P}^s < \frac{1}{2}$ (so that $1 - 2\bar{P}^s > 0$) under Assumption 4, Δu^{nc} inherits this same zero/negative/zero shape. For the turning point, centering equation (29) at $\bar{P}^s = \frac{1}{2}$ via $e \equiv \bar{P}^s - \frac{1}{2}$ reduces the condition to a depressed quartic in e ,

$$2e^4 - (1 + a + b)e^2 - (b - a)e + \left(\frac{1}{8} - \frac{a+b}{4}\right) = 0, \quad a \equiv Pr_H Pr_L, \quad b \equiv (1 - Pr_H)(1 - Pr_L),$$

which admits a unique real root in the economically relevant range $\bar{P}^s \in (0, \frac{1}{2})$ implied by Assumption 4. Existence and uniqueness of $\bar{A}^*$ then follow from the strict monotonicity of $\bar{P}^s(\bar{A})$ in $\bar{A}$. $\square$

Remark 3. Δu^{nc} starts at 0 (fully Democrat neighborhood, $\bar{A} = 0$), decreases to an interior minimum at $\bar{A}^*$ (the point of maximum non-conformity penalty), then rises back to 0 (fully Republican neighborhood, $\bar{A} = 1$) — a single-troughed shape. At either boundary, the observer's

prior $s_0 \in \{0, 1\}$ leaves no room for either action to move belief away from that prior, so the non-conformity distance collapses regardless of which action is taken.

Lemma 4 (Sign of the non-conformity feedback coefficient Ω^{nc}). *Under Assumptions 1–4, the non-conformity feedback weight satisfies*

$$\Omega^{nc}(A, \bar{A}, EV) = V^{nc}(Pr_H - Pr_L) \bar{A}(1 - \bar{A}) \left[\frac{(1 - 2\bar{P}^s)(1 - Pr_H - Pr_L)}{\bar{P}^s(1 - \bar{P}^s)} - Q \cdot W \right],$$

where $Q \equiv \frac{2(\bar{P}^s)^2 - 2\bar{P}^s + 1}{[\bar{P}^s(1 - \bar{P}^s)]^2} > 0$ and $W \equiv (1 - \bar{A})Pr_H(1 - Pr_H) + \bar{A}Pr_L(1 - Pr_L) > 0$. The bracketed term is strictly negative for all $\bar{A} \in [0, 1]$ under Assumptions 3–4. Consequently,

$$\Omega^{nc} \geq 0 \quad \text{for all } \bar{A} \in [0, 1],$$

with strict inequality for $\bar{A} \in (0, 1)$. Unlike Ω and Ω^p (Corollary 0.1), this sign holds unconditionally, without restriction to a diagnostic-regime subset of the domain.

Proof. Write $Pr_H = \frac{1}{2} - a$, $Pr_L = \frac{1}{2} - b$, so that Assumption 4 is equivalent to $0 < a < b < \frac{1}{2}$. Let

$$N(\bar{A}) \equiv W(2(\bar{P}^s)^2 - 2\bar{P}^s + 1) - (1 - 2\bar{P}^s)(1 - Pr_H - Pr_L)\bar{P}^s(1 - \bar{P}^s)$$

denote the numerator of the bracketed term after clearing its denominator $[\bar{P}^s(1 - \bar{P}^s)]^2$, so that $N(\bar{A}) = -[\text{bracket}] \cdot [\bar{P}^s(1 - \bar{P}^s)]^2$. As a function of $\bar{A}$, N is quadratic, with coefficients

$$c_2 = \frac{(a - b)^2(4ab + 1)}{2}, \quad c_1 = -b(a - b)(2a - 1)(2a + 1), \quad c_0 = \frac{(2a - 1)(2a + 1)(4ab - 1)}{8}.$$

Since $a \neq b$ and $4ab + 1 > 0$, $c_2 > 0$. Since $0 < a < \frac{1}{2}$, $(2a - 1)(2a + 1) = 4a^2 - 1 < 0$; since $a, b < \frac{1}{2}$, $4ab - 1 < 0$; hence $c_0 > 0$. The discriminant is

$$c_1^2 - 4c_2c_0 = -\frac{(a - b)^2(2a - 1)(2a + 1)(2b - 1)(2b + 1)}{4},$$

and since $(2a - 1)(2a + 1) < 0$ and $(2b - 1)(2b + 1) < 0$, their product is positive, making the discriminant strictly negative. A quadratic with positive leading coefficient and negative discriminant is strictly positive for all real $\bar{A}$; hence $N(\bar{A}) > 0$, giving bracket < 0 (since $[\bar{P}^s(1 - \bar{P}^s)]^2 > 0$) and $\Omega^{nc} \geq 0$ throughout $[0, 1]$, with equality only at $\bar{A} \in \{0, 1\}$ where $\bar{A}(1 - \bar{A}) = 0$. $\square$

A.6 Crowding-out by EV

We turn now to examine the crowding-out of the self- and social signals by private payoffs EV .²⁶

Assumption 5 (Sufficient Relative EV -Sensitivity). *Let $Pr'_H \equiv \partial Pr_H / \partial EV$ and $Pr'_L \equiv \partial Pr_L / \partial EV$, both strictly positive by Assumption 2. Define $\Gamma_H \equiv d\Psi_{\rho_0}^\rho / dPr_H$ and $\Gamma_L \equiv d\Psi_{\rho_0}^\rho / dPr_L$ — the total derivatives of $\Psi_{\rho_0}^\rho$ (equation (18)) with respect to Pr_H and Pr_L respectively, accounting for the fact that $\bar{P} = \rho_0 Pr_H + (1 - \rho_0) Pr_L$ itself depends on both. Within the diagnostic regime ($\rho_0 > \rho_0^*$), $\Gamma_L > 0$ always. Define the auxiliary threshold*

$$\Theta(\rho_0, Pr_H, Pr_L) \equiv \max\left(0, -\frac{\Gamma_H}{\Gamma_L}\right),$$

the unique value at which $\Gamma_H + \Gamma_L \Theta = 0$, and note

$$-\Psi_{\rho_0}^\rho + (Pr_H - Pr_L)\Gamma_L > 0$$

(a sum of two positive terms, by Lemma 2 and Assumption 4), so the following threshold is always well-defined:

$$r^*(\rho_0, Pr_H, Pr_L) \equiv \frac{-[\Psi_{\rho_0}^\rho + (Pr_H - Pr_L)\Gamma_H]}{-\Psi_{\rho_0}^\rho + (Pr_H - Pr_L)\Gamma_L}.$$

Within the diagnostic regime, the relative marginal EV -sensitivity of Low versus High types satisfies

$$\frac{Pr'_L}{Pr'_H} > r^*(\rho_0, Pr_H, Pr_L).$$

Remark 4. Γ_L captures how the diagnostic curvature $\Psi_{\rho_0}^\rho$ responds when the Low type's own adoption probability moves, and this channel always erodes the diagnostic gap in the diagnostic regime. Γ_H captures the analogous sensitivity to the High type's own probability, but this channel can go either way. Θ is the threshold at which the diagnostic-gap channel alone changes sign; r^* is the threshold at which the feedback-strength channel changes sign. Because the bracket underlying r^* is provably increasing in Pr'_L / Pr'_H , clearing r^* is both necessary and sufficient for the feedback-strength result, and — since $r^* \geq \Theta$ throughout the diagnostic regime — clearing r^* subsumes the weaker, diagnostic-gap-only requirement $r > \Theta$ as well. Only r^* appears in the statement of Assumption 5; Θ is retained solely as an auxiliary quantity used in the proof of Lemma 5.

²⁶See Dubé et al. (2017) on crowding-out of self-signals, Bénabou and Tirole (2006) and Jee et al. (2025) on crowding-out of reputational social signals, and Bollinger et al. (2022) or LaNauze (2023) on peer effects being stronger for lower- EV solar.

Lemma 5 (Sign of the EV -Sensitivity Derivative and the Feedback-Strength Response).
Under Assumptions 1–4 restricted to the diagnostic regime, and under Assumption 5:

$$(a) \quad \frac{\partial \Psi_{\rho_0}^\rho}{\partial EV} = \Gamma_H \cdot Pr'_H + \Gamma_L \cdot Pr'_L > 0,$$

$$(b) \quad \frac{\partial \Omega^\rho}{\partial EV} = -\rho_0(1 - \rho_0) \left[\Psi_{\rho_0}^\rho (Pr'_H - Pr'_L) + (Pr_H - Pr_L) \cdot \frac{\partial \Psi_{\rho_0}^\rho}{\partial EV} \right] < 0.$$

Proof. Let $r \equiv Pr'_L / Pr'_H$. The expression

$$\Psi_{\rho_0}^\rho (1 - r) + (Pr_H - Pr_L) [\Gamma_H + \Gamma_L r]$$

is affine in r with slope $-\Psi_{\rho_0}^\rho + (Pr_H - Pr_L)\Gamma_L > 0$ (Assumption 5).²⁷ By construction, r^* is the unique value of r at which this expression equals zero, so the expression is strictly positive for all $r > r^*$.

(b) Recall $\Omega^\rho = -\rho_0(1 - \rho_0)(Pr_H - Pr_L)\Psi_{\rho_0}^\rho$ (Corollary 1); differentiating via the product rule on $(Pr_H - Pr_L)$ and $\Psi_{\rho_0}^\rho$ gives

$$\frac{\partial \Omega^\rho}{\partial EV} = -\rho_0(1 - \rho_0) Pr'_H \cdot \left[\Psi_{\rho_0}^\rho (1 - r) + (Pr_H - Pr_L) [\Gamma_H + \Gamma_L r] \right].$$

Since $Pr'_H > 0$ and the term in square brackets is strictly positive for $r > r^*$, $\partial \Omega^\rho / \partial EV < 0$.

(a) Evaluating $\Psi_{\rho_0}^\rho (1 - r) + (Pr_H - Pr_L) [\Gamma_H + \Gamma_L r]$ at $r = \Theta$ gives $\Psi_{\rho_0}^\rho (1 - \Theta)$, since $\Gamma_H + \Gamma_L \Theta = 0$ by definition of Θ . Because $\Psi_{\rho_0}^\rho < 0$ in the diagnostic regime, this quantity has the sign of $-(1 - \Theta)$: negative when $\Theta < 1$, non-negative when $\Theta \geq 1$. In the former case, since the expression is increasing in r , its zero r^* satisfies $r^* > \Theta$; in the latter, $r^* \leq \Theta$. Either way, $r > r^*$ implies $r > \Theta$, so $\Gamma_H + \Gamma_L r > 0$, giving $\partial \Psi_{\rho_0}^\rho / \partial EV = \Gamma_H Pr'_H + \Gamma_L Pr'_L = Pr'_H (\Gamma_H + \Gamma_L r) > 0$. $\square$

²⁷In the corner case of $\tilde{v}(A) = 0$ for some A , this expression degenerates to 0.

Proposition 2 (Self-Signal Substitutability for Private Goods across Parties). *Under Assumptions 1–4, restricted to the diagnostic regime, and under Assumption 5, private payoffs EV and neighborhood political affiliation $\bar{A}$ are strategic substitutes for private goods ($Vis = 0$):*

$$\left. \frac{d^2 \Delta U_A}{d\bar{A} dEV} \right|_{Vis=0} < 0 \quad \text{for } A \in \{R, D\}.$$

The magnitude of self-signal identity-restoration crowding-out is weakly stronger for Democrats than Republicans:

$$\left| \left. \frac{d^2 \Delta U_D}{d\bar{A} dEV} \right|_{Vis=0} \right| \geq \left| \left. \frac{d^2 \Delta U_R}{d\bar{A} dEV} \right|_{Vis=0} \right|.$$

Proof. Differentiating

$$\left. \frac{d\Delta U_A}{d\bar{A}} \right|_{Vis=0} = \frac{\Psi_{\rho_0}^\rho \partial \rho_0 / \partial \bar{A}}{1 - \Omega^\rho}$$

(equation (25)) with respect to EV :

$$\left. \frac{d^2 \Delta U_A}{d\bar{A} dEV} \right|_{Vis=0} = \frac{\partial \rho_0}{\partial \bar{A}} \left[\Psi_{\rho_0}^\rho \cdot \frac{\partial}{\partial EV} \left(\frac{1}{1 - \Omega^\rho} \right) + \frac{1}{1 - \Omega^\rho} \cdot \frac{\partial \Psi_{\rho_0}^\rho}{\partial EV} \right].$$

By Lemma 5(a), $\partial \Psi_{\rho_0}^\rho / \partial EV > 0$, so the second term inside the bracket is strictly positive (since $1/(1 - \Omega^\rho) > 0$ by Lemma 1's stability condition). By Lemma 5(b), $\partial \Omega^\rho / \partial EV < 0$, so

$$\frac{\partial}{\partial EV} \left(\frac{1}{1 - \Omega^\rho} \right) = \frac{\partial \Omega^\rho / \partial EV}{(1 - \Omega^\rho)^2} < 0;$$

combined with $\Psi_{\rho_0}^\rho < 0$ (diagnostic regime), the first term inside the bracket is also strictly positive. The bracket is therefore strictly positive. Since $\partial \rho_0 / \partial \bar{A} < 0$ (Assumption 3), the full expression is strictly negative. Because $\Psi_{\rho_0}^\rho$ and Ω^ρ each scale with $\tilde{v}(A)$, and $\tilde{v}(D) \geq \tilde{v}(R) \geq 0$ (Assumption 1), the magnitude is weakly larger for Democrats. $\square$

Remark 5. *The result rests on two reinforcing mechanisms. The diagnostic-gap channel says rising EV dilutes the informational value of adoption itself. The feedback-strength channel says rising EV also weakens the self-reinforcing equilibrium loop that would otherwise amplify whatever the diagnostic-gap channel does. Assumption 5, the Low type's relative EV -sensitivity clearing the exact threshold r^* , keeps both channels pointed the same direction. Unlike a flat requirement that Low types respond at least as much as High types in absolute terms, r^* is state-dependent and can be well below 1, so the result holds under a strictly larger set of admissible relative sensitivities than a cruder sufficient condition would allow.*

Crowding out of the social signal and non-conformity penalty: Preliminary Definitions

Let $r \equiv Pr'_L/Pr'_H$, and define

$$T \equiv \frac{2Pr_H Pr_L}{(\bar{P}^s)^3} - \frac{2(1 - Pr_H)(1 - Pr_L)}{(1 - \bar{P}^s)^3},$$

$$\Phi_H^{nc} \equiv \left[\frac{1 - Pr_L}{(1 - \bar{P}^s)^2} - \frac{Pr_L}{(\bar{P}^s)^2} \right] + (1 - \bar{A}) T, \quad \Phi_L^{nc} \equiv \left[\frac{1 - Pr_H}{(1 - \bar{P}^s)^2} - \frac{Pr_H}{(\bar{P}^s)^2} \right] + \bar{A} T,$$

so that $\partial^2 \Delta NC / \partial \bar{A} \partial EV = \Phi_H^{nc} Pr'_H + \Phi_L^{nc} Pr'_L$.

With $R^s \equiv 3(\bar{P}^s)^2 - 3\bar{P}^s + 1$, define the total derivatives of $\Psi_{s_0}^s$ with respect to Pr_H, Pr_L :

$$\Gamma_H^s \equiv \frac{d\Psi_{s_0}^s}{dPr_H} = \tilde{v}^s(A) \cdot \frac{\bar{P}^s(1 - \bar{P}^s)[Pr_L(1 - 2\bar{P}^s) + (\bar{P}^s)^2] - 2s_0[Pr_H Pr_L R^s + (\bar{P}^s)^3(1 - Pr_H - Pr_L)]}{(\bar{P}^s)^3(1 - \bar{P}^s)^3},$$

$$\Gamma_L^s \equiv \frac{d\Psi_{s_0}^s}{dPr_L} = \tilde{v}^s(A) \cdot \frac{\bar{P}^s(1 - \bar{P}^s)[Pr_H(1 - 2\bar{P}^s) + (\bar{P}^s)^2] - 2(1 - s_0)[Pr_H Pr_L R^s + (\bar{P}^s)^3(1 - Pr_H - Pr_L)]}{(\bar{P}^s)^3(1 - \bar{P}^s)^3},$$

so that $\frac{\partial^2 \Delta u^s}{\partial \bar{A} \partial EV} = -[\Gamma_H^s Pr'_H + \Gamma_L^s Pr'_L]$.

Define $\bar{A}_{nc} \in (0, 1)$ implicitly by $\Phi_H^{nc}(\bar{A}_{nc}) + \Phi_L^{nc}(\bar{A}_{nc}) = 0$, and $\bar{A}_s \in (0, 1)$ implicitly by $\Gamma_H^s(\bar{A}_s) + \Gamma_L^s(\bar{A}_s) = 0$. Using the threshold r^* from Assumption 5, define

$$\tau(\bar{A}, Pr_H, Pr_L) \equiv \max \left(0, \frac{\Gamma_L^s}{|\Phi_L^{nc}|}, \frac{\Gamma_H^s + \Gamma_L^s r^*}{|\Phi_H^{nc} + \Phi_L^{nc} r^*|} \right).$$

Lemma 6 (Signs of the social signal and non-conformity cross-partial channels). *Under Assumptions 1–4, restricted to the diagnostic regime ($\rho_0 > \rho_0^*$), and under Assumption 5:*

(a) Non-Conformity channel. *For $\bar{A} < \bar{A}_{nc}$:*

$$\frac{\partial^2 \Delta u^{nc}}{\partial \bar{A} \partial EV} > 0.$$

For $\bar{A} > \bar{A}_{nc}$, the sign is indeterminate absent a further assumption bounding r from above.

(b) Social Signal channel. For $\bar{A} > \bar{A}_s$:

$$\frac{\partial^2 \Delta u^s}{\partial \bar{A} \partial EV} > 0.$$

For $\bar{A} < \bar{A}_s$, the term is negative in the overwhelming majority of admissible parameter configurations, but this has not been established as an unconditional result.

Proof. (a) Since $\Delta u^{nc} = V^{nc} \Delta NC$ with $V^{nc} < 0$, and $\partial^2 \Delta NC / \partial \bar{A} \partial EV = Pr'_H [\Phi_H^{nc} + \Phi_L^{nc} r]$: because $\Phi_H^{nc} + \Phi_L^{nc} r$ is decreasing in r (as $\Phi_L^{nc} < 0$ throughout $\bar{A} < \bar{A}_{nc}$), and its value at $r = 1$ is negative by definition of $\bar{A}_{nc}$, it remains negative for all $r \geq 1$ by monotonicity. Hence $\partial^2 \Delta NC / \partial \bar{A} \partial EV < 0$ for $\bar{A} < \bar{A}_{nc}$, and multiplying by $V^{nc} < 0$ gives $\partial^2 \Delta u^{nc} / \partial \bar{A} \partial EV > 0$. For $\bar{A} > \bar{A}_{nc}$, the value at $r = 1$ is positive but, since r is unbounded above under Assumption 5, no sign can be guaranteed for all admissible r .

(b) By identical reasoning applied to $-\Gamma_H^s + \Gamma_L^s r$: for $\bar{A} > \bar{A}_s$, $\Gamma_L^s < 0$ throughout this range, so $\Gamma_H^s + \Gamma_L^s r$ is decreasing in r ; its value at $r = 1$ is negative by definition of $\bar{A}_s$, and remains negative for all $r \geq 1$, giving $-\Gamma_H^s + \Gamma_L^s r > 0$. For $\bar{A} < \bar{A}_s$, the corresponding monotonicity argument does not deliver a guaranteed sign for all admissible r ; the negative sign holds in the large majority of tested cases but is not proven. $\square$

Proposition 3 (Non-conformity dominance over the social signal). *Under Assumptions 1–4, restricted to the diagnostic regime and $\bar{A} < \bar{A}_{nc}$, and under Assumption 5, suppose*

$$\frac{|V^{nc}|}{\tilde{v}^s(\bar{A})} > \tau(\bar{A}, Pr_H, Pr_L).$$

Then the non-conformity channel strictly dominates the social signal channel:

$$\frac{\partial^2 \Delta u^{nc}}{\partial \bar{A} \partial EV} + \frac{\partial^2 \Delta u^s}{\partial \bar{A} \partial EV} > 0,$$

and this condition is tight: the inequality can fail for some admissible Pr'_H, Pr'_L whenever $|V^{nc}| / \tilde{v}^s(\bar{A}) \leq \tau(\bar{A}, Pr_H, Pr_L)$. This dominance is most pronounced for Republicans ($\tilde{v}^s(R) \leq \tilde{v}^s(D)$, Assumption 1), since a smaller $\tilde{v}^s(\bar{A})$ makes the threshold easier to clear for a given V^{nc} .

Proof. Immediate from Lemma 6(b). $\square$

Remark 6. *The two channels comprising this sum push in opposite directions: greater EV attenuates the non-conformity penalty upward toward zero (Lemma 6(a)), while it attenuates*

the social signal downward toward zero over the corresponding range of $\bar{A}$. Proposition 3 does not require the social signal channel to vanish or reverse sign — only that the non-conformity channel's magnitude, scaled by $|V^{nc}|$, exceed the social signal channel's magnitude, scaled by $\tilde{v}^s(A)$, uniformly over the admissible range of relative EV-sensitivity $r \geq r_{min}$.

B Construction of Visibility Measures

Our primary visibility measure quantifies the angular degree of visibility of a potential solar installation from the nearest public roadway. The underlying data are drawn from the confidential Google Sunroof dataset, obtained directly from Google. For each residential structure, Project Sunroof provides the rooftop centroid and a set of roof segments, defined as planar, contiguous roof areas with uniform orientation. For each segment, the data include the azimuth (degrees clockwise from north), pitch (degrees relative to horizontal), and the maximum number of standard-sized solar panels that can be accommodated. The dataset further reports the optimal system size, the distribution of panels across roof segments under this configuration, and the rank ordering of segments by cumulative solar insolation.

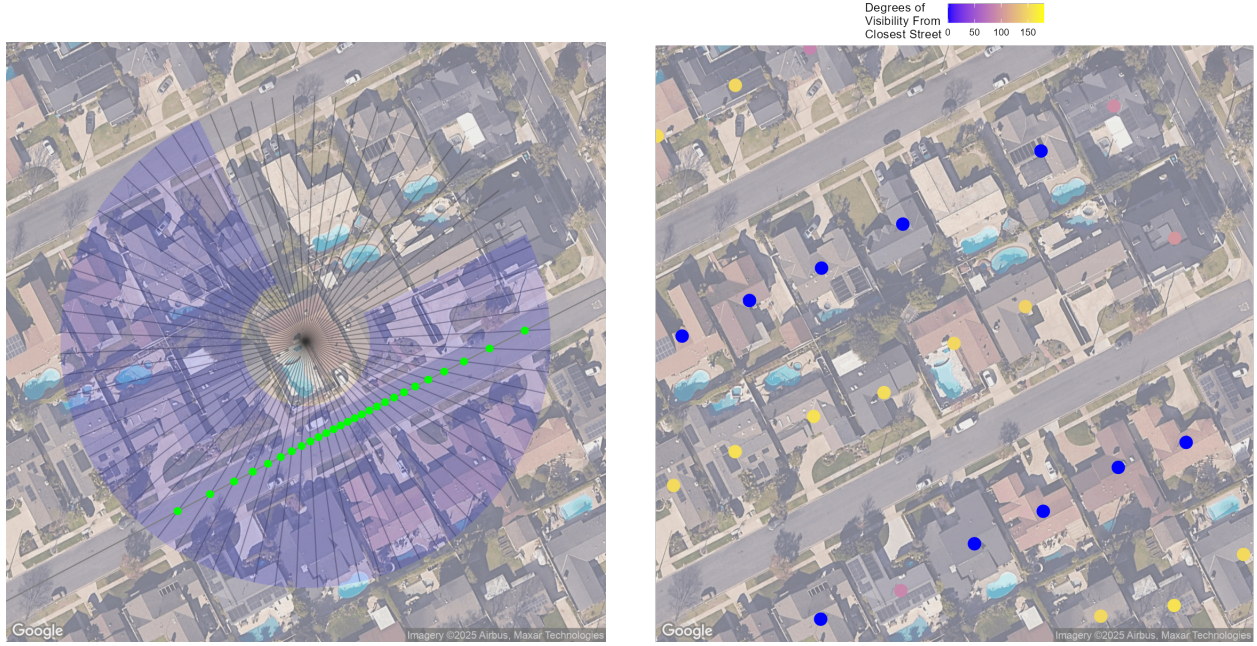
Our algorithmic visibility measure is similar to that of Bollinger et al. (2022), proceeding as follows:

- First, we find the number of roof segments necessary for an optimal installation. We take either the first (sunniest) roof segment or, if the first roof segment does not hold more than 75% of the optimal number of panels, the first two sunniest roof segments.
- Second, from the single reported centroid of the roof, we draw 360 evenly-spaced radii 65m long, each representing a single “degree of visibility”.
- Third, we construct a visibility semi-circle 65m in diameter that represents the 180-degrees of orientation over which a solar panel would be visible. We rotate the semi-circle around the rooftop centroid based on the reported roof segment orientation (azimuth) such that the semi-circle represents the visible range for a panel on that roof segment.
- Fourth, we calculate the distance from the rooftop centroid at which the pitch of the roof segment would render the panel non-visible and remove that area from the semi-circle, making a “semi-donut”.
- Fifth, we find the nearest street to the roof centroid on the CalTrans All Public Road shapefile.²⁸

Each point in the intersection of the the visibility “semi-donut”, the nearest public road, and the 360 drawn radii is counted as a degree of visibility. We show an example in Figure B.1a, which shows the “semi-donut”, the drawn radii, and the intersection points (in green). Because a house may have east-west facing panels, it is possible for the measure to exceed

²⁸Available at <https://caltrans-gis.dot.ca.gov/portal/home/item.html?id=836c3a3e2e7f40b8b36b24addf01cf14>.

180-degrees, especially if the house is located on a sharp bend in the nearest street. This occurs in 0.15% of observations. Therefore, we censor the visibility measure at 180-degrees, and re-scale such that our visibility measure is bounded between 0 and 1 for analysis. 12% of the final sample measures zero visibility and 95% of the sample is below 0.87. A full histogram is shown in Figure B.2



(a) **House visibility measure example.** Each radial line from the focal house represents a potential degree of visibility. The blue circle represents the direction of visibility of the panel(s) oriented in the direction (azimuth) of the optimal installation's roof segment(s). Households with more than one roof segment for an optimal installation, as is the case here, are the union of two semicircles, each oriented in the roof segment's directions. Areas in blue are areas for which the potential solar installation would be visible. The concentric circle in the middle represents areas where rooftop panels would not be visible due to the ground angle and roof pitch. Green points represent the intersection of a degree of visibility (black lines), the nearest street, and the blue visibility area. Each point is a degree of visibility. For simplicity, each line represents 5 degrees of visibility, while in practice, our measure uses one degree each.

(b) **Visibility.** Each home's visibility measure is shown in color. Houses on the south side of the street are largely zero-visibility as the optimal panel orientation is on the back side of the roof. However, one household in the lower left corner is partially visible despite being located on the south side of the street. In this case, the orientation of the roof line is such that one optimal location for a panel is on the street side facing west-southwest, resulting in a visibility measure of approximately 55 degrees due to the street's SW-NE orientation. This household adopted solar, and the placement of panels can be observed to be consistent with the non-zero visibility measure.

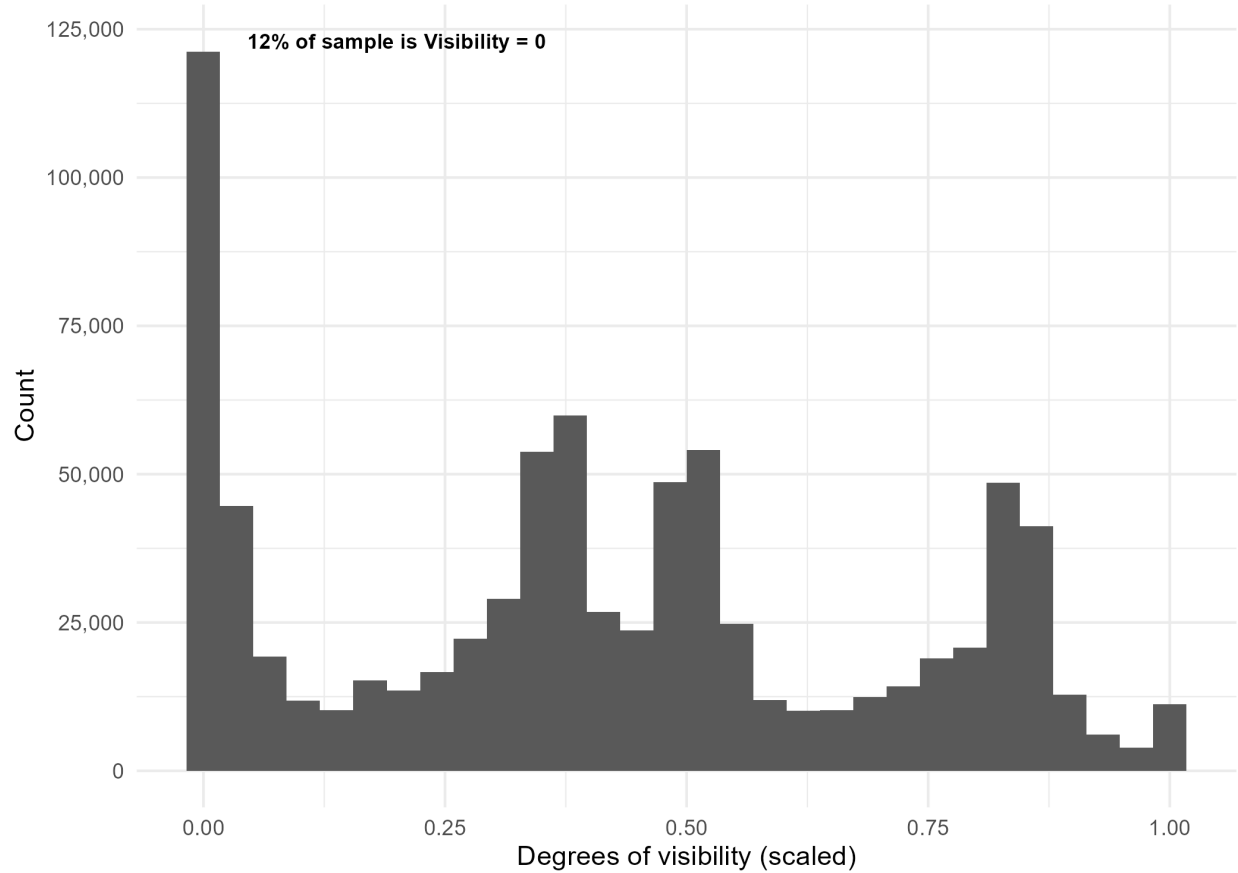


Figure B.2: **Histogram of Visibility.** Histogram of scaled visibility measure as calculated using the process described. 12% of the sample measures 0 visibility, and 1.2% measures the maximum visibility of 1, indicating that censoring does not substantially distort the distribution. Peaks in the scaled value result from the prevalence of north-south / east-west grid oriented neighborhoods.

We also calculate a simpler binary visibility measure based on side-of-street for testing for price differences between sides of the street (see Appendix C.2). We first calculate the angle of a line drawn from a rooftop centroid to the nearest point on the nearest street, where 0 indicates that the home is due north of the nearest street point, and 180 indicates that the home is due south of the nearest street point. Houses located in the northern quadrant 0 ± 45 degrees are designated as “visible.” All others are designated as non-visible. We report results using this measure of visibility in Appendix D.

C Supplemental Identification

A key identifying assumption in our empirical strategy is that, conditional on blockgroup fixed effects, an individual’s characteristics are uncorrelated with the neighborhood’s characteristics, which rules out confounding by selection or sorting. For instance, while a larger neighborhood may have unobserved reasons for attracting households that are registered Republican (and any unobserved characteristics that are correlated with Republican registration), conditional on census blockgroup fixed effects, neighborhood voter affiliation $\bar{A}$ is as-good-as-randomly assigned.²⁹

We address the potential confounding in two ways. First, we follow [Bayer et al. \(2008\)](#) and test this assumption on observable household characteristics. Intuitively, if households select into neighborhoods within a blockgroup based on observed and unobserved characteristics, correlations between observables should persist even when conditioning on blockgroup. We describe our method below, and find that conditioning on blockgroup eliminates within-neighborhood correlations in observable characteristics including voter affiliation and number of children, and reduces within-neighborhood correlations in household income. Second, under the notion that selection on observables tells us something about potential selection on unobservables ([Altonji et al., 2005, 2025](#)), we calculate the breakdown points of [Oster \(2019\)](#) and [Diegert et al. \(2025\)](#) for a key result, where the breakdown point “characterizes the magnitude of selection on unobservables relative to selection on observables needed to overturn one’s baseline findings” ([Diegert et al., 2025](#)). We find that selection on unobservables would have to be of unreasonable size relative to selection on observables to overturn a key portion of our findings.

C.1 As-good-as-random assignment

First, we follow [Bayer et al. \(2008\)](#) to test this assumption on observables. As they note, testing on observables does not directly rule out unobserved correlations that may persist, but selection (or neighborhood sorting) on unobservables is, in some sense, proportional ([Altonji et al., 2005](#)). In essence, if observables show no correlation within neighborhood conditional on neighborhood, it is unlikely that unobservables carry such a condition. Within-neighborhood correlation between neighborhood voter affiliation and a household’s voter affiliation is mechanically induced by the leave-one-out average of the neighborhood and blockgroup means. As such, we expand on [Bayer et al. \(2008\)](#) and test for correlation between a neighborhood’s voter affiliation and an individual’s voter affiliation by first sampling a single pair of households randomly selected from each neighborhood, then estimate the correlation

²⁹As before, we refer to the “hyper-local” measure of voter affiliation $\bar{A}$ as the “neighborhood”.

between them conditional on the blockgroup fixed effects. The blockgroups capture the larger effect of sorting as households are unlikely to be able to sort into neighborhoods due to thinness of the housing market (McCartney and Shah, 2022). If blockgroup fixed effects are sufficient controls, then correlations in household characteristics within neighborhood will not remain after conditioning on neighborhood. We estimate the following regression equations on our two-households-per-neighborhood sample, with and without blockgroup fixed effects, dropping blocks with 2 or fewer observations:

$$\begin{aligned}
Rep_i &= \beta_0 + \beta_1 \bar{A}_{l(i),-i} + \Xi_{n(i)} + \nu_i \\
Rep_i &= \gamma_0 + \gamma_1 Rep_{i,-i} + \Gamma_{n(i)} + \eta_i \\
Children_i &= \delta_0 + \delta_1 Children_{i,-i} + \Delta_{n(i)} + \iota_i \\
Income_i &= \kappa_0 + \kappa_1 Income_{i,-i} + \Omega_{n(i)} + \omega_i
\end{aligned}$$

Where $\bar{A}_{l(i),-i}$ is the randomly sampled household's neighborhood voter affiliation measure measured over all other neighborhood households $-i$, and $Rep_{i,-i}$, $Children_{i,-i}$ and $Income_{i,-i}$ are the randomly-sampled second household voter affiliation, count of children, and income.

Table C.1: **Tests of orthogonality of household characteristics:** Regression results from a regression of household characteristics on a randomly selected second household from within neighborhood (block). (1-2) tests for correlation between household voter affiliation and leave-one-out measure of neighborhood Republican affiliation; (3-4) tests for correlation between household 1 and household 2 voter affiliation, (5-6) tests for correlation between household 1 and household 2 number of children, (7-8) tests for correlation between household 1 income and household 2 income.

Dependent Variables:	Voter affiliation				Children in Hh		Household income	
Model:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
<i>Variables</i>								
Own nbhd ($\bar{A}$)	0.6999*** (0.0150)	0.0173 (0.0204)						
Other Household voter affiliation			0.1172*** (0.0054)	0.0069 (0.0052)				
Children in other Hh					0.0347*** (0.0051)	0.0044 (0.0052)		
Household income in other Hh							0.7278*** (0.0067)	0.0630*** (0.0052)
<i>Fixed-effects</i>								
Blockgroup FE		Yes		Yes		Yes		Yes
<i>Fit statistics</i>								
Observations	41,636	41,396	41,636	41,396	41,636	41,396	41,636	41,396
R ²	0.06039	0.21994	0.01380	0.21996	0.00123	0.14475	0.53472	0.74973

Signif. Codes: ***, 0.001, **, 0.01, *, 0.05, +: 0.1

Results in Table C.1 show that, conditional on the blockgroup fixed effects in (2) and (4), there is no significant correlation between key household voting characteristics within neighborhood. Results in column (2) show that a neighborhood’s leave-one-out measure of voter affiliation and the individual’s affiliation is uncorrelated when conditioning on the blockgroup fixed effects. Results in (4) show that affiliation of household i are uncorrelated with household $-i$, conditional on blockgroup fixed effects. The lack of significant correlation provides support for (but not evidence of) the identifying assumption that no unobservable characteristics correlated with voter affiliation are driving sorting at the neighborhood level.

A similar exercise using count of children in the household yields similar results in (6). However, household income remains correlated, though blockgroup fixed effects dramatically attenuate the relationship (8). This may be due to presence of unobserved amenities (e.g., views, nearby parks) that draw higher-income households. In our main specification, we control for all of the key household characteristics that may drive these decisions, including house size, income, and assessor value (a proxy for wealth), and interact all controls with an indicator for voter affiliation. Below, we discuss the magnitude of additional unobserved characteristics that would threaten identification of the neighborhood effect.

Physical properties of the house are correlated within the neighborhood, conditional on blockgroup, especially in “tract” style housing common in our sample. Furthermore, visibility may be correlated – a neighborhood located along a north-south oriented street is less likely to have visible installations than those located along an east-west street (where half the homes are directly north). Thus, these characteristics are correlated, even conditional on blockgroup, as shown in Table C.2. We include in controls measures of square footage, lot size (2nd degree polynomial), an indicator for year built after 1990, number of stories, assessor value, and number of bedrooms, all of which may be correlated with unobserved preferences for solar. We maintain the assumption that it is infeasible to sort (within a blockgroup) on EV , and test for robustness to this by estimating on a sample of homes sold in 2005 or earlier, well before solar would have been a concern for any adopter, regardless of unobserved tastes for solar. These robustness results are in Table D.4. Results show little change qualitatively or statistically from the main results.

Table C.2: **Tests of orthogonality of house characteristics:** Regression results from a regression of house characteristics on a randomly selected second household from within neighborhood (block). House characteristics are correlated within neighborhood (conditional on blockgroup fixed effects) due to tract nature of housing development. All variables are controlled for, interacted with political affiliation, in all specifications.

Dependent Variables:	Home value		Number of stories		Year built		Home sq. ft.		Lot size		Expected value		Visibility from closest street	
Model:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
<i>Variables</i>														
Other home value	0.5762*** (0.0162)	0.1843*** (0.0054)												
Other home number of stories			0.4382*** (0.0078)	0.1869*** (0.0052)										
Other home year built					0.7842*** (0.0102)	0.5125*** (0.0045)								
Other home sq. ft.							0.6398*** (0.0118)	0.3129*** (0.0050)						
Other lot size									0.5512*** (0.0535)	0.4052*** (0.0046)				
Other expected value											0.6442*** (0.0210)	0.1945*** (0.0164)		
Other home visibility from closest street													0.1469*** (0.0057)	0.1232*** (0.0058)
<i>Fixed-effects</i>														
Blockgroup FE		Yes		Yes		Yes		Yes		Yes		Yes		Yes
<i>Fit statistics</i>														
Observations	41,636	41,396	41,636	41,396	41,636	41,396	41,636	41,396	41,636	41,396	41,636	41,396	41,636	41,396
R ²	0.30388	0.52090	0.19212	0.41161	0.61892	0.71383	0.40451	0.57851	0.31168	0.48133	0.41352	0.62077	0.02158	0.15622

Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1

A final explicit concern is that unobserved energy consumption is greater for Democrats in more Republican neighborhoods relative to Democrats in less-Republican neighborhoods. We test this by taking a subsample of our data consisting of registered voters who live in either the *most* or *least* Republican neighborhoods within a blockgroup. We regress the year built on an indicator for the max-Republican neighborhood and it's interaction with voter affiliation, and include all controls and fixed effects as in the main specification, with the exception of our indicator for year built after 1990.

Results are shown in Table C.3. Year built is not significantly different in the max-Republican neighborhood relative to the min-Republican neighborhood for Democrats/mixed-affiliation (the omitted category) and for Republicans.

Table C.3: Results: Difference in Year Built Between Max and Min Republican Neighborhoods.

Model:	(1)
<i>Variables</i>	
In Max-Rep Block	5.68 (6.06)
In Max-Rep Block $\times$ Republican hh	13.5 (8.29)
In Max-Rep Block $\times$ Unregistered hh	-0.899 (8.08)
<i>Fixed-effects</i>	
Blockgroup-Registered voter FE	Yes
<i>Fit statistics</i>	
Observations	167,616
Pseudo R ²	0.08137
Log-Likelihood	-1,262,974.1
<i>Conley (2km) standard-errors in parentheses</i>	
<i>Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1</i>	

C.2 Side-of-street Sorting: Hedonic regression and correlation with political affiliation

If households with unobserved utility for solar also have unobserved utility for a specific side-of-street, then estimates of the effect of visibility may be confounded. The motivating example for this may be that people who are environmentally-minded may also prefer backyards that lie to the north of their home, and thus prefer the north side of the street. Our setting suggests that yard orientation relative to a home is not particularly predictive of shading or sun exposure – in “tract” style housing which predominates our sample, shading from the home is similar to shading from perimeter trees behind a home, and thus there would be little difference in backyard sunlight or aesthetics between sides of street. Nonetheless, we examine whether or not prices reflect the side-of-street under the assumption that correlated preferences for such an orientation would result in higher sale prices. Approximately half ($N = 485,719$) of sample homes have sold in the last 30 years. We specify the sale price in logs, include all home characteristics used in our main specification, and allow a 4th degree polynomial in sale year to flexibly control for changes in price over time. We find a distinct lack of correlation between side-of-street and sale price:

Table C.4: **Regression of Sale Price on north side of street:** Results from regression of logged sale price on included house characteristics and a polynomial in sale year.

Model:	(1)	(2)	(3)
<i>Variables</i>			
North side of street (visible)	-0.0008 (0.0019)	-0.0021 (0.0013)	-0.0014 (0.0014)
Sq. ft.		1.38×10^{-7} (1.83×10^{-7})	4.5×10^{-8} (9.65×10^{-8})
Lot size (sq. ft.)		0.0011*** (0.0001)	0.0006*** (9.43×10^{-5})
Lot size squared		-6.55×10^{-7} *** (9.09×10^{-8})	-3.97×10^{-7} *** (5.83×10^{-8})
Year built > 1990		0.1119*** (0.0101)	0.0515*** (0.0099)
Number of Bedrooms		0.0676*** (0.0027)	0.0492*** (0.0022)
2+ Stories		0.0677*** (0.0039)	0.0636*** (0.0026)
poly(Saleyear, 4)1		269.7*** (3.470)	271.3*** (3.774)
poly(Saleyear, 4)2		-38.40*** (1.594)	-37.94*** (1.736)
poly(Saleyear, 4)3		-68.59*** (1.218)	-68.43*** (1.314)
poly(Saleyear, 4)4		1.806 (2.279)	1.316 (2.461)
<i>Fixed-effects</i>			
Census blockgroup	Yes	Yes	
Census blockgroup-Street			Yes
<i>Fit statistics</i>			
Observations	485,708	485,708	481,353
Pseudo R ²	0.22912	0.65118	0.75841
Log-Likelihood	-360,621.6	-163,179.0	-111,865.8
<i>Conley (2km) standard-errors in parentheses</i>			
<i>Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1</i>			

Results in column (2) show that side-of-street is uncorrelated with sale price conditional on controls and census blockgroup effects. Column (3) shows the same conditional on blockgroup x street fixed effects. To provide further insight, we also estimate the correlation between side-of-street and political affiliation. A non-significant hedonic price does not rule out sorting if only a small share of the population has a preference for living on the north side of the street. Thus, we check to see if side-of-street and political affiliation are correlated by regressing the binary indicator for north on political affiliation.

Table C.5: **Regression of north side of street on political affiliation:** Results from regression of a binary indicator for side of street on political affiliation.

Model:	(1)	(2)
<i>Variables</i>		
Republican hh	0.0154 (0.0137)	-0.0013 (0.0069)
Unregistered voter hh	0.0032 (0.0087)	0.0005 (0.0070)
Constant	-1.129*** (0.0169)	
<i>Fixed-effects</i>		
Blockgroup		Yes
<i>Fit statistics</i>		
Observations	905,028	901,698
Pseudo R ²	5.78×10^{-6}	0.05580
Log-Likelihood	-503,886.4	-474,782.7
<i>Conley (2km) standard-errors in parentheses</i>		
<i>Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1</i>		

Both with and without blockgroup fixed effects, political affiliation is uninformative about side-of-street. We interpret these results as evidence that side-of-street is likely uncorrelated with unobservable preferences. Furthermore, if solar-minded individuals were able to sort to the north side of the street when purchasing a home, our estimate of visibility would be biased upwards. However, our main results find a negative and significant effect of visibility, including among Democratic/mixed households who would be most likely to sort on visibility.

D Additional Results and Robustness

Table D.1: Main Results including all covariates

Dependent Variable: Model: Neighborhood definition:	(1) Block x Street	Adopt (2) Blockgroup x Street	(3) Block
<i>Variables</i>			
Republican hh	-0.105* (0.053)	-0.105* (0.053)	-0.100+ (0.053)
Expected value	0.188*** (0.017)	0.191*** (0.017)	0.192*** (0.018)
Expected value × Republican hh	0.0006 (0.016)	0.002 (0.016)	0.006 (0.017)
Expected value × Unregistered voter hh	-0.040 (0.026)	-0.041 (0.026)	-0.043 (0.026)
Visible (frac of degrees from closest street)	-0.093*** (0.019)	-0.092*** (0.019)	-0.091*** (0.019)
Visible (frac of degrees from closest street) × Republican hh	-0.006 (0.036)	-0.006 (0.038)	-0.020 (0.040)
Visible (frac of degrees from closest street) × Unregistered voter hh	0.081* (0.039)	0.079* (0.039)	0.084* (0.039)
Visible (frac of degrees from closest street) × Expected value	-0.003 (0.018)	-0.006 (0.018)	-0.004 (0.018)
Republican nbhd	0.188*** (0.050)	0.270*** (0.068)	0.405*** (0.088)
Republican nbhd × Republican hh	-0.136+ (0.080)	-0.178+ (0.101)	-0.230+ (0.127)
Republican nbhd × Unregistered voter hh	-0.025 (0.108)	-0.093 (0.131)	0.019 (0.174)
Republican nbhd × Expected value	-0.170** (0.052)	-0.245*** (0.070)	-0.294*** (0.083)
Republican nbhd × Expected value × Republican hh	0.032 (0.064)	0.039 (0.081)	0.019 (0.097)
Republican nbhd × Expected value × Unregistered voter hh	0.067 (0.078)	0.055 (0.099)	0.134 (0.120)
Visible (frac of degrees from closest street) × Republican nbhd	-0.199* (0.093)	-0.261* (0.113)	-0.328* (0.132)
Visible (frac of degrees from closest street) × Republican nbhd × Republican hh	0.165 (0.161)	0.208 (0.209)	0.416 (0.254)
Visible (frac of degrees from closest street) × Republican nbhd × Unregistered voter hh	0.168 (0.203)	0.266 (0.235)	0.009 (0.278)
Visible (frac of degrees from closest street) × Republican nbhd × Expected value	0.182* (0.079)	0.259* (0.101)	0.205* (0.107)
Home Size	$3.22 \times 10^{-7+}$ (1.92×10^{-7})	$3.22 \times 10^{-7+}$ (1.92×10^{-7})	$3.24 \times 10^{-7+}$ (1.92×10^{-7})
Republican hh × Home Size	0.0003*** (1.82×10^{-5})	0.0003*** (1.82×10^{-5})	0.0003*** (1.82×10^{-5})
Unregistered voter hh × Home Size	0.0003*** (2.96×10^{-5})	0.0003*** (2.95×10^{-5})	0.0003*** (2.95×10^{-5})
Lot size	0.002*** (0.0004)	0.002*** (0.0004)	0.002*** (0.0004)
Republican hh × Lot size	-0.001*** (0.0004)	-0.001*** (0.0004)	-0.001*** (0.0004)
Unregistered voter hh × Lot size	-0.003*** (0.0006)	-0.003*** (0.0006)	-0.003*** (0.0006)
Squared lot size	$-1.34 \times 10^{-6***}$ (3.02×10^{-7})	$-1.33 \times 10^{-6***}$ (3.03×10^{-7})	$-1.33 \times 10^{-6***}$ (3.02×10^{-7})
Republican hh × Squared lot size	$9.17 \times 10^{-7***}$ (2.62×10^{-7})	$9.01 \times 10^{-7***}$ (2.63×10^{-7})	$9.01 \times 10^{-7***}$ (2.64×10^{-7})
Unregistered voter hh × Squared lot size	$1.86 \times 10^{-6***}$ (3.71×10^{-7})	$1.84 \times 10^{-6***}$ (3.72×10^{-7})	$1.86 \times 10^{-6***}$ (3.71×10^{-7})
1(Year Built 1990)	0.041 (0.037)	0.041 (0.037)	0.042 (0.037)
Household income 1,000s	0.005*** (0.0002)	0.005*** (0.0002)	0.005*** (0.0002)
Republican hh × 1(Year Built 1990)	-0.073* (0.036)	-0.073* (0.036)	-0.073* (0.036)
Unregistered voter hh × 1(Year Built 1990)	0.058 (0.054)	0.058 (0.054)	0.058 (0.054)
Republican hh × Household income 1,000s	-0.001*** (0.0002)	-0.001*** (0.0002)	-0.001*** (0.0002)
Unregistered voter hh × Household income 1,000s	0.0007* (0.0003)	0.0007* (0.0003)	0.0007* (0.0003)
Number of bedrooms	0.156*** (0.012)	0.156*** (0.012)	0.156*** (0.012)
Republican hh × Number of bedrooms	-0.052** (0.016)	-0.052** (0.016)	-0.052** (0.016)
Unregistered voter hh × Number of bedrooms	-0.078*** (0.016)	-0.078*** (0.016)	-0.078*** (0.016)
Number of Children	0.014 (0.009)	0.014 (0.009)	0.014 (0.009)
Republican hh × Number of Children	0.084*** (0.014)	0.084*** (0.014)	0.084*** (0.014)
Unregistered voter hh × Number of Children	0.085*** (0.020)	0.085*** (0.020)	0.085*** (0.020)
Assessor value	0.064*** (0.005)	0.064*** (0.005)	0.064*** (0.005)
Republican hh × Assessor value	-0.040*** (0.004)	-0.040*** (0.004)	-0.040*** (0.004)
Unregistered voter hh × Assessor value	-0.020*** (0.005)	-0.019*** (0.005)	-0.019*** (0.005)
2+ stories	0.098*** (0.021)	0.098*** (0.021)	0.098*** (0.021)
Republican hh × 2+ stories	-0.119*** (0.028)	-0.119*** (0.028)	-0.117*** (0.028)
Unregistered voter hh × 2+ stories	-0.035 (0.035)	-0.035 (0.035)	-0.034 (0.035)
<i>Fixed-effects</i>			
Blockgroup-Registered voter	Yes	Yes	Yes
<i>Fit statistics</i>			
Observations	818,086	818,086	818,086
Pseudo R ²	0.10481	0.10482	0.10486
Log-Likelihood	-223,398.9	-223,396.0	-223,385.9

Conley (2km) standard-errors in parentheses
Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1

Table D.1 shows the full results of our main specification, including covariates.

D.1 Linear Probability Model

To ensure robustness to functional form, we estimate our model using OLS in a linear probability model. Results in Table D.2 are unchanged qualitatively and statistically. Around 9% of the sample predictions from the estimates in column (3) are less than zero.

Table D.2: Robustness Results: Estimation using linear probability model.

Dependent Variable:	Adopt		
Model:	(1)	(2)	(3)
Neighborhood definition:	Block x Street	Blockgroup x Street	Block
<i>Variables</i>			
Republican hh	-0.018*** (0.004)	-0.018*** (0.004)	-0.018*** (0.004)
Expected value	0.011*** (0.0010)	0.011*** (0.0010)	0.011*** (0.0010)
Expected value $\times$ Republican hh	0.0006 (0.0009)	0.0006 (0.0009)	0.0005 (0.0009)
Expected value $\times$ Unregistered voter hh	-0.006*** (0.001)	-0.006*** (0.001)	-0.006*** (0.001)
Visible (frac of degrees from closest street)	-0.008*** (0.001)	-0.008*** (0.001)	-0.008*** (0.001)
Visible (frac of degrees from closest street) $\times$ Republican hh	-0.0007 (0.003)	-0.0004 (0.003)	-0.0008 (0.003)
Visible (frac of degrees from closest street) $\times$ Unregistered voter hh	0.006** (0.002)	0.006** (0.002)	0.006** (0.002)
Visible (frac of degrees from closest street) $\times$ Expected value	-0.002+ (0.0009)	-0.001 (0.0009)	-0.001 (0.0009)
Republican nbhd	0.017*** (0.004)	0.025*** (0.006)	0.035*** (0.008)
Republican nbhd $\times$ Republican hh	-0.015* (0.007)	-0.020* (0.009)	-0.025* (0.011)
Republican nbhd $\times$ Unregistered voter hh	-0.009 (0.007)	-0.017+ (0.009)	-0.013 (0.011)
Republican nbhd $\times$ Expected value	0.0003 (0.004)	0.001 (0.005)	0.004 (0.006)
Republican nbhd $\times$ Expected value $\times$ Republican hh	0.004 (0.005)	0.005 (0.006)	0.006 (0.007)
Republican nbhd $\times$ Expected value $\times$ Unregistered voter hh	0.002 (0.004)	-0.0005 (0.005)	0.004 (0.006)
Visible (frac of degrees from closest street) $\times$ Republican nbhd	-0.021** (0.008)	-0.028** (0.009)	-0.036*** (0.011)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Republican hh	0.014 (0.014)	0.018 (0.019)	0.033 (0.021)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Unregistered voter hh	0.020 (0.013)	0.029+ (0.015)	0.017 (0.017)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Expected value	0.005 (0.005)	0.009 (0.007)	0.001 (0.007)
<i>Fixed-effects</i>			
Blockgroup-Registered voter	Yes	Yes	Yes
<i>Fit statistics</i>			
Observations	904,939	904,939	904,939
R ²	0.08137	0.08138	0.08139
RMSE	0.26349	0.26349	0.26348

Conley (2km) standard-errors in parentheses

Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1

Abbreviations used: neighborhood (nbhd), household (hh)

D.2 Heterogeneity of Effects by Income

We further explore heterogeneity in the effect of neighborhood political affiliation and the interaction of neighborhood political affiliation with visibility from the closest street. We test for heterogeneous effects by re-estimating our main results, but allowing our effects consistent with self- and social signaling to vary by income. Results are shown in in Table D.3.

Results indicate no significant difference over income in the effect of visibility interacted with Republican neighborhood for any political affiliation, suggesting that social signaling is not different between low- and high-income households. Results for the interaction of income with effects we interpret as consistent with the self-signal identity restoration effect are negative and significant (see bottom three rows of Table D.3). In our preferred specification in column (3), the effect of income is negative for all affiliations, suggesting that the causal effect of Republican neighborhood (β_1 for D/mixed, $\beta_1 + \beta_{2R}$ for Republican, and $\beta_1 + \beta_{2U}$ for unregistered households) is attenuated downward toward zero in higher-than-average income households, potentially more so for Republican and unregistered households. Our model does not explain, nor does it rule out, income-specific heterogeneity.

Table D.3: Heterogeneity Results: Interacting effects with income

Dependent Variable:	Adopt		
Model:	(1)	(2)	(3)
Neighborhood definition:	Block x Street	Blockgroup x Street	Block
<i>Variables</i>			
Republican hh	-0.232*** (0.053)	-0.231*** (0.054)	-0.226*** (0.054)
Expected value	0.188*** (0.017)	0.191*** (0.017)	0.192*** (0.018)
Expected value $\times$ Republican hh	0.001 (0.016)	0.003 (0.016)	0.007 (0.017)
Expected value $\times$ Unregistered voter hh	-0.041 (0.026)	-0.041 (0.026)	-0.043 (0.026)
Visible (frac of degrees from closest street)	-0.092*** (0.019)	-0.091*** (0.020)	-0.092*** (0.020)
Visible (frac of degrees from closest street) $\times$ Republican hh	-0.007 (0.037)	-0.007 (0.038)	-0.020 (0.040)
Visible (frac of degrees from closest street) $\times$ Unregistered voter hh	0.080* (0.039)	0.080* (0.039)	0.082* (0.039)
Visible (frac of degrees from closest street) $\times$ Expected value	-0.003 (0.018)	-0.006 (0.018)	-0.003 (0.018)
Republican nbhd	0.221*** (0.052)	0.336*** (0.071)	0.481*** (0.090)
Republican nbhd $\times$ Republican hh	-0.099 (0.085)	-0.133 (0.104)	-0.157 (0.134)
Republican nbhd $\times$ Unregistered voter hh	-0.013 (0.109)	-0.109 (0.131)	0.013 (0.175)
Republican nbhd $\times$ Expected value	-0.181*** (0.053)	-0.262*** (0.070)	-0.308*** (0.084)
Republican nbhd $\times$ Expected value $\times$ Republican hh	0.027 (0.065)	0.034 (0.081)	0.007 (0.097)
Republican nbhd $\times$ Expected value $\times$ Unregistered voter hh	0.061 (0.077)	0.055 (0.098)	0.135 (0.119)
Visible (frac of degrees from closest street) $\times$ Republican nbhd	-0.198* (0.094)	-0.255* (0.115)	-0.353** (0.134)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Republican hh	0.107 (0.171)	0.131 (0.218)	0.328 (0.270)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Unregistered voter hh	0.143 (0.203)	0.262 (0.234)	-0.020 (0.280)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Expected value	0.190* (0.080)	0.263** (0.101)	0.215* (0.108)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ D/mixed Hh $\times$ Income	-0.0003 (0.001)	-0.0005 (0.001)	0.0007 (0.002)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Republican hh $\times$ Income	0.001 (0.001)	0.001 (0.002)	0.002 (0.002)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Unregistered voter hh $\times$ Income	0.001 (0.002)	-0.0003 (0.002)	0.003 (0.003)
Income	0.005*** (0.0002)	0.005*** (0.0002)	0.005*** (0.0002)
Republican hh $\times$ Income	-0.001*** (0.0002)	-0.001*** (0.0002)	-0.001*** (0.0002)
Unregistered voter hh $\times$ Income	0.0007* (0.0003)	0.0007* (0.0003)	0.0007* (0.0003)
Republican nbhd $\times$ D/mixed Hh $\times$ Income	-0.0010 ⁺ (0.0005)	-0.002* (0.0007)	-0.002** (0.0008)
Republican nbhd $\times$ Republican hh $\times$ Income	-0.002* (0.0007)	-0.002** (0.0008)	-0.004*** (0.0010)
Republican nbhd $\times$ Unregistered voter hh $\times$ Income	-0.002* (0.001)	-0.002 ⁺ (0.001)	-0.004** (0.002)
<i>Fixed-effects</i>			
Blockgroup-Registered voter	Yes	Yes	Yes
<i>Fit statistics</i>			
Observations	818,086	818,086	818,086
Pseudo R ²	0.10486	0.10488	0.10493
Log-Likelihood	-223,386.5	-223,379.9	-223,366.8

Conley (2km) standard-errors in parentheses

Signif. Codes: ***, 0.001, **, 0.01, *, 0.05, +: 0.1

Income is demeaned for easier interpretation.

D.3 Homes transacted in 2005 or earlier

Results may be biased if households with unobserved taste for solar sort into homes with higher expected value of solar (e.g. better irradiance, etc.) To ensure results are robust to this, we re-estimate our main results using only those homes that were sold in 2005 or earlier. Prior to 2005, solar was extremely rare (1,587 adoptions, 2.1% of all adoptions in our data) and sorting on EV is highly unlikely.

Table D.4: Robustness Results: Sample limited to houses sold in 2005 or earlier only.

Dependent Variable: Model: Neighborhood definition:	(1) Block x Street	Adopt (2) Blockgroup x Street	(3) Block
<i>Variables</i>			
Republican hh	-0.181** (0.062)	-0.179** (0.062)	-0.175** (0.062)
Expected value	0.194*** (0.020)	0.197*** (0.020)	0.198*** (0.020)
Expected value $\times$ Republican hh	0.036+ (0.020)	0.038+ (0.021)	0.035 (0.022)
Expected value $\times$ Unregistered voter hh	-0.039 (0.034)	-0.041 (0.034)	-0.040 (0.034)
Visible (frac of degrees from closest street)	-0.084*** (0.024)	-0.085*** (0.024)	-0.083*** (0.024)
Visible (frac of degrees from closest street) $\times$ Republican hh	-0.031 (0.046)	-0.037 (0.047)	-0.047 (0.050)
Visible (frac of degrees from closest street) $\times$ Unregistered voter hh	0.120* (0.055)	0.121* (0.055)	0.122* (0.055)
Visible (frac of degrees from closest street) $\times$ Expected value	-0.014 (0.021)	-0.017 (0.022)	-0.014 (0.022)
Republican nbhd	0.190** (0.060)	0.268*** (0.077)	0.359*** (0.098)
Republican nbhd $\times$ Republican hh	-0.194* (0.098)	-0.276* (0.120)	-0.278+ (0.158)
Republican nbhd $\times$ Unregistered voter hh	-0.215 (0.144)	-0.231 (0.171)	-0.146 (0.234)
Republican nbhd $\times$ Expected value	-0.200** (0.063)	-0.302*** (0.082)	-0.339*** (0.095)
Republican nbhd $\times$ Expected value $\times$ Republican hh	0.026 (0.075)	0.034 (0.098)	0.097 (0.123)
Republican nbhd $\times$ Expected value $\times$ Unregistered voter hh	0.035 (0.113)	0.093 (0.139)	0.059 (0.177)
Visible (frac of degrees from closest street) $\times$ Republican nbhd	-0.204+ (0.117)	-0.280* (0.137)	-0.274+ (0.161)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Republican hh	0.355+ (0.207)	0.460+ (0.254)	0.591* (0.294)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Unregistered voter hh	0.375 (0.278)	0.301 (0.314)	0.276 (0.378)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Expected value	0.216* (0.100)	0.363** (0.125)	0.174 (0.130)
<i>Fixed-effects</i>			
Blockgroup-Registered voter	Yes	Yes	Yes
<i>Fit statistics</i>			
Observations	512,166	512,166	512,166
Pseudo R ²	0.09583	0.09585	0.09588
Log-Likelihood	-133,659.0	-133,656.7	-133,652.4

Conley (2km) standard-errors in parentheses

Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1

This robustness check also focuses the sample onto households for which our voter affiliation snapshot, which is based on registration and participation in 2014-2016, is accurate. Homes that last sold in 2005 are occupied by the same owners for the vast bulk of years covered by our adoption data (1999 to 2016) as very few adoptions occur prior to 2005. Thus, this robustness check also ensures that the relevant voter registration is represented in the data.

Results in [D.4](#) show little change qualitatively or statistically from the main results. The statistical significance of the interaction between Visibility and Republican neighborhood drops from significant at the 5% level in the main results to the 10% level in (1) and (3) here. The point estimate of the former is only slightly smaller, which we interpret as a case of increased noise, as the sample size drops from over 818k to 512k.

D.4 Stability and Breakdown Point Analysis

We further address the potential for selection into neighborhood on unobservables. Specifically, we may worry that household tastes for more Republican neighborhoods (conditional on census blockgroup and controls including household income and number of children) may be correlated with an unobserved preference for solar. For example, households who value self-reliance and household production of food and/or energy may sort into neighborhoods that are more conducive to household production (larger lots, relaxed rules on animal husbandry, or better soil), which may also be unobserved and correlated with $\bar{A}$.³⁰ This threat to identification can be framed as selection into treatment, where the treatment is defined as a Republican neighborhood, which would bias estimates involving $\bar{A}_i$. [Altonji et al. \(2005\)](#) notes that, though the extent of selection on unobservables cannot be directly observed, assumptions regarding the relative ratio of selection on observed (which is observable) and selection on unobserved can provide intuition for the extent of the selection problem.

D.4.1 Stability

We first present stability results showing the change in estimates as we add controls, a common method in the literature. If adding observed controls does not significantly change results, then adding unobserved controls would likely not do so as well. Results are shown in [Table D.5](#), where column (1) shows our main results specification without census blockgroup-by-registered voter fixed effects nor our controls, (2) shows main results with fixed effects but no controls, and (3) repeats our main results for comparison. Results are wildly different absent the fixed effects, as expected given results in [Table C.1](#), but relatively similar for both main effects and key interaction terms between (2) and (3).

³⁰We control for lot size, but do not observe the latter.

The intuition of Altonji et al. (2025, 2005) focuses on the addition of observable controls *after* conditioning on “variables whose inclusion is a central part of the identification strategy, such as state and time dummy variables in a two-way fixed effect model or the running variable in a regression discontinuity design” (Altonji et al., 2025). Thus, our stability discussion centers on the change in coefficient estimates and R^2 in Table D.5 between columns (2) with blockgroup fixed effects and (3) with blockgroup fixed effects and observable controls. The addition of controls in (3) relative to (2) increases the R^2 from 0.084 to 0.105, indicating a substantial amount of variation explained by the controls. The coefficient on Republican neighborhood moves from 0.627 to 0.405 and remains significant at 5%, while the coefficient on Visibility interacted with Republican neighborhood changes from -0.411 to -0.328 . Both are relatively small changes.

Table D.5: Results showing stability over (1) no fixed effects and no controls; (2) fixed effects and no controls; and (3) Main results

Dependent Variable: Model:	(1)	Adopt (2)	(3)
<i>Variables</i>			
Republican hh	0.068** (0.025)	-0.006 (0.022)	-0.100+ (0.053)
Unregistered voter hh	-0.292*** (0.023)		
Expected value	0.333*** (0.031)	0.208*** (0.019)	0.192*** (0.018)
Expected value $\times$ Republican hh	0.109*** (0.024)	0.039* (0.017)	0.006 (0.017)
Expected value $\times$ Unregistered voter hh	0.027 (0.024)	-0.021 (0.026)	-0.043 (0.026)
Visible (frac of degrees from closest street)	-0.136*** (0.020)	-0.106*** (0.019)	-0.091*** (0.019)
Visible (frac of degrees from closest street) $\times$ Republican hh	-0.038 (0.037)	-0.031 (0.039)	-0.020 (0.040)
Visible (frac of degrees from closest street) $\times$ Unregistered voter hh	0.037 (0.036)	0.066+ (0.038)	0.084* (0.039)
Visible (frac of degrees from closest street) $\times$ Expected value	-0.025 (0.028)	0.0002 (0.019)	-0.004 (0.018)
Republican nbhd	1.92*** (0.147)	0.627*** (0.087)	0.405*** (0.088)
Republican nbhd $\times$ Republican hh	-0.342* (0.157)	-0.043 (0.125)	-0.230+ (0.127)
Republican nbhd $\times$ Expected value	0.155 (0.149)	-0.363*** (0.085)	-0.294*** (0.083)
Republican nbhd $\times$ Expected value $\times$ Republican hh	-0.116 (0.156)	0.019 (0.096)	0.019 (0.097)
Republican nbhd $\times$ Expected value $\times$ Unregistered voter hh	-0.248+ (0.148)	0.150 (0.120)	0.134 (0.120)
Visible (frac of degrees from closest street) $\times$ Republican nbhd	-0.472*** (0.135)	-0.411** (0.131)	-0.328* (0.132)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Republican hh	0.188 (0.233)	0.411+ (0.249)	0.416 (0.254)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Unregistered voter hh	-0.045 (0.245)	0.010 (0.270)	0.009 (0.278)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Expected value	0.164 (0.181)	0.187+ (0.108)	0.205+ (0.107)
Home Size			$3.24 \times 10^{-7+}$ (1.92×10^{-7})
Republican hh $\times$ Home Size			0.0003*** (1.82×10^{-5})
Unregistered voter hh $\times$ Home Size			0.0003*** (2.95×10^{-5})
Lot size			0.002*** (0.0004)
Republican hh $\times$ Lot size			-0.001*** (0.0004)
Unregistered voter hh $\times$ Lot size			-0.003*** (0.0006)
Squared lot size			$-1.33 \times 10^{-6***}$ (3.02×10^{-7})
Republican hh $\times$ Squared lot size			$9.01 \times 10^{-7***}$ (2.64×10^{-7})
Unregistered voter hh $\times$ Squared lot size			$1.86 \times 10^{-6***}$ (3.71×10^{-7})
1(Year Built 1990)			0.042 (0.037)
Houshold income 1,000s			0.005*** (0.0002)
Republican hh $\times$ 1(Year Built 1990)			-0.073* (0.036)
Unregistered voter hh $\times$ 1(Year Built 1990)			0.058 (0.054)
Republican hh $\times$ Houshold income 1,000s			-0.001*** (0.0002)
Unregistered voter hh $\times$ Houshold income 1,000s			0.0007* (0.0003)
Number of bedrooms			0.156*** (0.012)
Republican hh $\times$ Number of bedrooms			-0.052** (0.016)
Unregistered voter hh $\times$ Number of bedrooms			-0.078*** (0.016)
Number of Children			0.014 (0.009)
Republican hh $\times$ Number of Children			0.084*** (0.014)
Unregistered voter hh $\times$ Number of Children			0.085*** (0.020)
Assessor value			0.064*** (0.005)
Republican hh $\times$ Assessor value			-0.040*** (0.004)
Unregistered voter hh $\times$ Assessor value			-0.019*** (0.005)
2+ stories			0.098*** (0.021)
Republican hh $\times$ 2+ stories			-0.117*** (0.028)
Unregistered voter hh $\times$ 2+ stories			-0.034 (0.035)
Constant	-2.29*** (0.028)		
Republican nbhd $\times$ Unregistered voter hh	-0.223 (0.153)	0.090 (0.170)	0.019 (0.174)
<i>Fixed-effects</i>			
Blockgroup-Registered voter		Yes	Yes
<i>Fit statistics</i>			
Observations	818,086	818,086	818,086
Pseudo R ²	0.02208	0.08391	0.10486
Log-Likelihood	-244,043.6	-228,613.0	-223,385.9

Conley (2km) standard-errors in parentheses
Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1

D.4.2 Breakdown Point Analysis

Stability exhibitions do not provide the full information necessary to eliminate concerns regarding selection on observables. To do so, we perform a breakdown point analysis. The breakdown point is a scalar measure of the ratio of selection on unobservables relative to selection on observables at which point our baseline findings are overturned (Diegert et al., 2025). We calculate the breakdown point for each of our models (1)-(3) using four methods. The first two follow Oster (2019), which requires an assumption regarding the R^2 of a full regression, inclusive of the unobservables. We first report δ^O for the assumption that all variation in adoption can be explained by observed and unobserved variables ($R_{max}^2 = 1$). An implication of this assumption is that no further variation in solar adoption occurs after observables and unobservables are determined. In our setting, this is a strong assumption – given that most individuals purchased before solar was widely available, it is unlikely that no further variation in adoption is determined post-purchase.

The second, more common assumption is that the R_{max}^2 is proportional to the R^2 obtained from a regression of adoption on the full suite of controls and variables of interest. Oster (2019) establishes a factor of 1.3 by examining a battery of randomized control trials and calculating the ratio of R^2 , but even this rule-of-thumb is context dependent. Including fewer observable controls would reduce the reference R^2 and, thus, R_{max}^2 , which in turn would increase the breakdown point arbitrarily. In our setting, this regression, after residualizing outcome, treatment, and observed covariates to the blockgroup by registered voter fixed effects, has a relatively low R^2 of approximately around .010 across our specifications.³¹ Even with observed political and demographic controls chosen to maximize explanation of outcome variance, this R^2 indicates very little variation is explained and the remaining “irreducible” error is substantial. This suggests that a lower value of R_{max}^2 is appropriate. We use $R_{max}^2 = 1.3 \times R_{overall}^2 = .013$, which we report in table D.6.

Our preferred breakdown point is that of Diegert et al. (2025), which does not require that unobserved controls be orthogonal to observed controls, and does not require an assumption on R_{max}^2 . In our setting, it is unlikely that the unobserved and observed variables are orthogonal, and the choice of R_{max}^2 is pivotal. Calculating $\bar{r}_x^{DMP}$ requires no assumption on the ratio of outcome (solar adoption) variation explained by the unobserved variables relative to the observed variables ($\bar{r}_y$). Thus, it represents a conservative estimate of the breakdown point. We also calculate the breakdown point under the common assumption that $\bar{r}_X = \bar{r}_y$, which restricts the ratio of the unobserved variables’ impact on outcome relative to the observed variables’ impact on outcome to be equal to the the ratio of the unobserved variables’ impact

³¹The R^2 of the regression including census blockgroup FEs is over .10

on treatment relative to the observed variables' impact on treatment. This breakdown point, $\bar{r}^{DMP}$, is our preferred breakdown point measure, though we report the more conservative $\bar{r}_x^{DMP}$ as well.

For each method, we assess a main effects version of our main results. While methods for assessing selection on unobservables with multiple, interacted treatments (e.g. $Vis \times \bar{A}$) exist, the breakdown point becomes an identified set of dimension equal to the number of main and interactive effects, which is difficult to interpret given our large number of interactions. Furthermore, non-linear models may not have well-defined breakdown points when interactions are present. For tractability and interpretation, we use a linear probability model (LPM) main effects regression for this analysis, dropping all interactions from the main results. The parameter for which we have primary concern for selection on unobservables is the neighborhood political affiliation, $\bar{A}$, which is itself a main effect. We therefore focus our breakdown point analysis on this parameter.³²

The results are shown in Table D.6. In the main effects regression, the point estimate on our preferred measure of neighborhood, Republican block, is similar to the main LPM results at a statistically-significant 0.022 (the main results interact affiliation with neighborhood, producing an estimate of .035 for Democratic/mixed and $0.035 - 0.025 = .010$ for Republican households, see table D.2). δ^O varies a great deal based on the assumption of R_{max}^2 , but the Oster measure indicates that selection into treatment for unobservables would have to be 4.16 times as strong as selection on observables to reduce the coefficient on Republican block to 0. The breakdown point of our preferred breakdown point $\bar{r}^{DMP}$ of 0.720 indicates that unobservables would have to explain at least 72.0% of selection into a Republican block relative to the amount of selection explained by observables in order to overturn the results. This level of selection is highly unlikely. Observables include household income, lot size, home size, assessor value, and number of stories, all of which are highly predictive of adoption in theory and practice. It would be very unlikely for an unobservable to have even a small fraction of explanatory power after conditioning on observables.

Interpretation of a breakdown point is improved with calibration through context. Following Diegert et al. (2025), we explore the relative impact of each observable control in order to better calibrate a reasonable threshold of $\bar{r}_X$. In table D.7, we present $\hat{\rho}_k$, the ratio of selection on control k relative to all other controls, for each control. This provides a range of values of $\bar{r}_X$ for observed controls *as if* each one were unobserved. Only lot size and lot size squared exceed the estimated breakdown point of 72.0%. If this control was, in

³²Note that the stability results in Table D.5 show that the key interaction terms including Republican nbhd $\times$ Expected Value, Visibility $\times$ Republican nbhd, and Visibility $\times$ Republican nbhd $\times$ Expected Value, all show stability when adding observables (2)-(3).

fact, unobserved, it would exceed the breakdown point. However, selection into Republican neighborhoods is driven largely by lot size, assessor value, number of bedrooms, and age of home. Notably, even home size has a value of $\hat{\rho}_k = 3.1\%$ and household income has a value of $\hat{\rho}_k = 52.8$, far lower than the breakdown point threshold of 72.0%. Unobservables of any nature are highly unlikely to match the relative impact of observed lot size and household income controls.

Table D.6: Breakdown Point Analysis Results for Selection on Unobservables

Dependent Variable:	Adopt		
Model:	(1)	(2)	(3)
Nbhd definition:	Block x Street	Block	Blockgroup x Street
<i>Variables</i>			
Republican block-by-street	0.009*** (0.002)		
Republican block		0.022*** (0.005)	
Republican blockgroup-by-street			0.014*** (0.004)
Expected value	0.011*** (0.0007)	0.011*** (0.0007)	0.011*** (0.0007)
Visible (side of street)	-0.005*** (0.0009)	-0.005*** (0.0009)	-0.005*** (0.0009)
Home Size	7.67×10^{-8} (7.85×10^{-8})	7.65×10^{-8} (7.82×10^{-8})	7.66×10^{-8} (7.84×10^{-8})
Lot size	0.0002*** (6.1×10^{-5})	0.0002*** (6.08×10^{-5})	0.0002*** (6.09×10^{-5})
Squared lot size	-1.45×10^{-7} *** (3.87×10^{-8})	-1.44×10^{-7} *** (3.86×10^{-8})	-1.44×10^{-7} *** (3.86×10^{-8})
1(Year Built 1990)	0.008* (0.004)	0.008* (0.004)	0.008* (0.004)
Number of bedrooms	0.014*** (0.0009)	0.014*** (0.0009)	0.014*** (0.0009)
Number of Children	0.004*** (0.0007)	0.004*** (0.0007)	0.004*** (0.0007)
Assessor value	0.005*** (0.0005)	0.005*** (0.0005)	0.005*** (0.0005)
2+ stories	0.006*** (0.001)	0.006*** (0.001)	0.006*** (0.001)
Income	0.0005*** (1.74×10^{-5})	0.0005*** (1.73×10^{-5})	0.0005*** (1.73×10^{-5})
Republican Hh	-0.002+ (0.001)	-0.002+ (0.001)	-0.002+ (0.001)
<i>Fixed-effects</i>			
Blockgroup-Registered voter	Yes	Yes	Yes
Sensitivity:			
$\delta_{r_{max}=1}^O$	0.019	0.019	0.017
$\delta_{r_{max}=1.3 \times \text{overall}}^O$	4.28	4.19	3.76
$\bar{r}_X^{DMP}$	0.123	0.114	0.107
$\bar{r}^{DMP}$	.708	.720	.701
<i>Fit statistics</i>			
Observations	818,086	818,086	818,086
R ²	0.07139	0.07141	0.07139
Within R ²	0.01438	0.01440	0.01438

Conley (2km) standard-errors in parentheses

Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1

Based on the orthogonality tests, stability results, and breakdown points calculated, selection into Republican neighborhood on unobservables is unlikely to overturn the main results.

Table D.7: Calibrating $\bar{r}_X$: Relative Impact of Each Observed Covariate on Republican locale

W_{1k}	$\hat{\rho}_k$ (%)
	Republican neighborhood
Expected value	7.1
Home Size	3.1
Lot size	97.0
Squared lot size	83.5
1(Year Built 1990)	21.0
Number of bedrooms	48.0
Number of Children	10.9
Assessor value	34.8
2+ stories	5.6
Income	52.8
Visible (frac of degrees from closest street)	0.9
Republican Hh	22.0

D.5 Alternative Definition of Neighborhood

We further check robustness to the definition of neighborhood by defining the neighborhood as the four nearest neighbors for each household. We calculate our neighborhood affiliation as the fraction of those households that are registered Republican. To account for rural areas where the nearest neighbors may be entirely unseen or even unknown, we also calculate this measure using only those homes under 200m from the focal home.

Columns (1)-(2) measure neighborhood as the 4 nearest neighbors and include census blockgroup by registered voter fixed effects as in our main specification. Columns (3)-(4) include finer census block fixed effects, possible only due to the very fine 4 nearest neighbor definition of neighborhood. Results in Table D.8 columns (1)-(2) are very similar to our main results that define neighborhood as “block-by-street” in column (1) of Table 3. Only the interaction of Republican neighborhood with Republican voter affiliation changes in significance, even then it remains significant at 10% and changes only when defining neighborhood as the four nearest houses regardless of distance. Columns (3)-(4) are similar in sign and magnitude to (1)-(2) with some exceptions. The interaction of visibility, Republican neighborhood, and expected value is slightly smaller and not significant, while the estimated effect of visibility interacted with Republican neighborhood is now significant only at the 10% level. The former may indicate weaker crowding out than our main estimates. The latter is likely the result of a sizable decrease in sample size, with many census blocks having zero installations which results in being dropped from the estimation sample.

Table D.8: Robustness Results: Neighborhood measured as 4 nearest neighbors.

Dependent Variable:	Adopt			
Model:	(1)	(2)	(3)	(4)
Neighborhood definition:	4NN under 200m	4NN	4NN under 200m	4NN
<i>Variables</i>				
Republican hh	-0.105* (0.053)	-0.101+ (0.053)	-0.077 (0.061)	-0.072 (0.061)
Expected value	0.192*** (0.017)	0.187*** (0.017)	0.197*** (0.020)	0.197*** (0.020)
Expected value $\times$ Republican hh	0.004 (0.016)	0.0004 (0.016)	0.012 (0.018)	0.007 (0.018)
Expected value $\times$ Unregistered voter hh	-0.039 (0.026)	-0.040 (0.026)	-0.048 (0.035)	-0.049 (0.035)
Visible (frac of degrees from closest street)	-0.093*** (0.019)	-0.093*** (0.019)	-0.084*** (0.022)	-0.084*** (0.022)
Visible (frac of degrees from closest street) $\times$ Republican hh	-0.013 (0.036)	-0.011 (0.036)	-0.011 (0.042)	-0.009 (0.042)
Visible (frac of degrees from closest street) $\times$ Unregistered voter hh	0.085* (0.039)	0.083* (0.039)	0.048 (0.053)	0.047 (0.053)
Visible (frac of degrees from closest street) $\times$ Expected value	-0.006 (0.018)	-0.002 (0.018)	-0.017 (0.020)	-0.015 (0.020)
Republican nbhd	0.180*** (0.044)	0.183*** (0.045)	0.147** (0.050)	0.152** (0.050)
Republican nbhd $\times$ Republican hh	-0.108 (0.073)	-0.095 (0.074)	-0.107 (0.085)	-0.099 (0.086)
Republican nbhd $\times$ Unregistered voter hh	-0.077 (0.098)	-0.084 (0.099)	-0.066 (0.128)	-0.079 (0.129)
Republican nbhd $\times$ Expected value	-0.132** (0.043)	-0.144** (0.045)	-0.108* (0.049)	-0.130* (0.051)
Republican nbhd $\times$ Expected value $\times$ Republican hh	-0.019 (0.056)	0.016 (0.058)	-0.026 (0.065)	0.008 (0.066)
Republican nbhd $\times$ Expected value $\times$ Unregistered voter hh	0.062 (0.072)	0.068 (0.073)	-0.039 (0.105)	-0.026 (0.106)
Visible (frac of degrees from closest street) $\times$ Republican nbhd	-0.190* (0.084)	-0.187* (0.084)	-0.179+ (0.096)	-0.176+ (0.097)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Republican hh	0.251+ (0.139)	0.216 (0.141)	0.264 (0.163)	0.233 (0.164)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Unregistered voter hh	0.030 (0.170)	0.042 (0.172)	-0.068 (0.220)	-0.058 (0.222)
Visible (frac of degrees from closest street) $\times$ Republican nbhd $\times$ Expected value	0.183** (0.071)	0.175* (0.072)	0.132 (0.084)	0.137 (0.085)
<i>Fixed-effects</i>				
Blockgroup-Registered voter	Yes	Yes		
Block-Registered voter			Yes	Yes
<i>Fit statistics</i>				
Observations	817,378	818,086	515,353	515,894
Pseudo R ²	0.10479	0.10481	0.11549	0.11553
Log-Likelihood	-223,067.0	-223,397.8	-186,027.0	-186,304.9

Conley (2km) standard-errors in parentheses

Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1

All of our different specifications capture alternative neighborhoods and “audiences” to a potential installation that may influence signaling utility. The models implicitly assume that households of different political affiliations consider the same population to be the relevant audience. It is possible that this would not hold, e.g., Republicans might consider the whole block while Democrats consider only their immediate contiguous neighbors. However, if this were the case, we would expect the relative sizes of the signaling coefficients for households of different political affiliations to vary as we change the definition of the peer group. Since the estimates are fairly similar to our main results with the exception of the coefficient on

Republican nbhd $\times$ Republican hh in (1) which is no longer significant, we find asymmetric audience sizes as a function of political affiliation to be unlikely.

The stability of results on the self-signaling utility (the estimate for “Republican nbhd”) is only somewhat attenuated when including block fixed effects in (3)-(4). We interpret this as suggestive evidence in support of our identification assumption that sorting happens only at the blockgroup level and above. Sorting at the census block level (choosing a home with 4-nearest-neighbors who are more Republican conditional on the census block) is largely infeasible, but controlling for it renders similar results as our main specification.

D.6 Visibility Measure

D.6.1 Alternative definitions of visibility

To check robustness to our definition of “visible”, we substitute our measure of visibility with a binary indicator for the north side of the street, as described in Appendix B.

Results are shown above in Table D.9. Main effects are identical in sign and significance to our primary specification with one exception: Visibility $\times$ Republican nbhd maintains a negative sign, but is significant at only 10% in Column (1), where the neighborhood is defined as a census block. We interpret this as imprecision in the measure of visibility imposed by the continuous measure of visibility. It is unlikely that households consider visibility in such a manner, especially in the context of neighborhood visibility.

Table D.9: Robustness Results: Visibility measured by side-of-street.

Dependent Variable: Model: Neighborhood definition:	(1) Block x Street	Adopt (2) Blockgroup x Street	(3) Block
<i>Variables</i>			
Republican hh	-0.102* (0.051)	-0.104* (0.051)	-0.102* (0.051)
Expected value	0.189*** (0.017)	0.190*** (0.017)	0.193*** (0.017)
Expected value $\times$ Republican hh	0.0008 (0.016)	0.002 (0.016)	0.006 (0.017)
Expected value $\times$ Unregistered voter hh	-0.041 (0.026)	-0.042 (0.026)	-0.044+ (0.026)
Visible (side of street)	-0.068*** (0.015)	-0.068*** (0.015)	-0.067*** (0.015)
Visible (side of street) $\times$ Republican hh	-0.022 (0.023)	-0.014 (0.023)	-0.024 (0.025)
Visible (side of street) $\times$ Unregistered voter hh	0.034 (0.028)	0.035 (0.028)	0.035 (0.028)
Visible (side of street) $\times$ Expected value	-0.005 (0.013)	-0.004 (0.013)	-0.004 (0.014)
Republican nbhd	0.140*** (0.039)	0.217*** (0.058)	0.333*** (0.074)
Republican nbhd $\times$ Republican hh	-0.098 (0.061)	-0.117 (0.077)	-0.120 (0.093)
Republican nbhd $\times$ Unregistered voter hh	0.021 (0.079)	-0.008 (0.104)	0.009 (0.136)
Republican nbhd $\times$ Expected value	-0.111* (0.048)	-0.167* (0.065)	-0.239** (0.077)
Republican nbhd $\times$ Expected value $\times$ Republican hh	0.024 (0.065)	0.030 (0.082)	0.012 (0.098)
Republican nbhd $\times$ Expected value $\times$ Unregistered voter hh	0.064 (0.078)	0.049 (0.098)	0.139 (0.119)
Visible (side of street) $\times$ Republican nbhd	-0.124+ (0.070)	-0.191* (0.087)	-0.254* (0.098)
Visible (side of street) $\times$ Republican nbhd $\times$ Republican hh	0.118 (0.114)	0.083 (0.144)	0.235 (0.164)
Visible (side of street) $\times$ Republican nbhd $\times$ Unregistered voter hh	0.073 (0.128)	0.058 (0.161)	0.055 (0.192)
Visible (side of street) $\times$ Republican nbhd $\times$ Expected value	0.073 (0.068)	0.105 (0.082)	0.132 (0.095)
<i>Fixed-effects</i>			
Blockgroup-Registered voter	Yes	Yes	Yes
<i>Fit statistics</i>			
Observations	818,086	818,086	818,086
Pseudo R ²	0.10483	0.10484	0.10489
Log-Likelihood	-223,392.8	-223,389.1	-223,378.1

Conley (2km) standard-errors in parentheses

Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1

D.6.2 Semi-parametric visibility effects

In case the visibility relationship is not binary, we bin our visibility measure into one of four visibility bins and re-estimate our main specification. The first bin consists of the 12% of the sample for which visibility from the nearest street for a potential installation would be zero. The remaining sample is split into 4 additional bins for a total of five bins. Counts by

adoption are shown in Figure D.1. This specification allows for a more flexible (and even non-monotonic) relationship between visibility and adoption. Notably, if low visibility (e.g. Bin 1 relative to bin 0) has a smaller effect relative to the most visible bin, we can interpret very small amounts of visibility as being negligible.

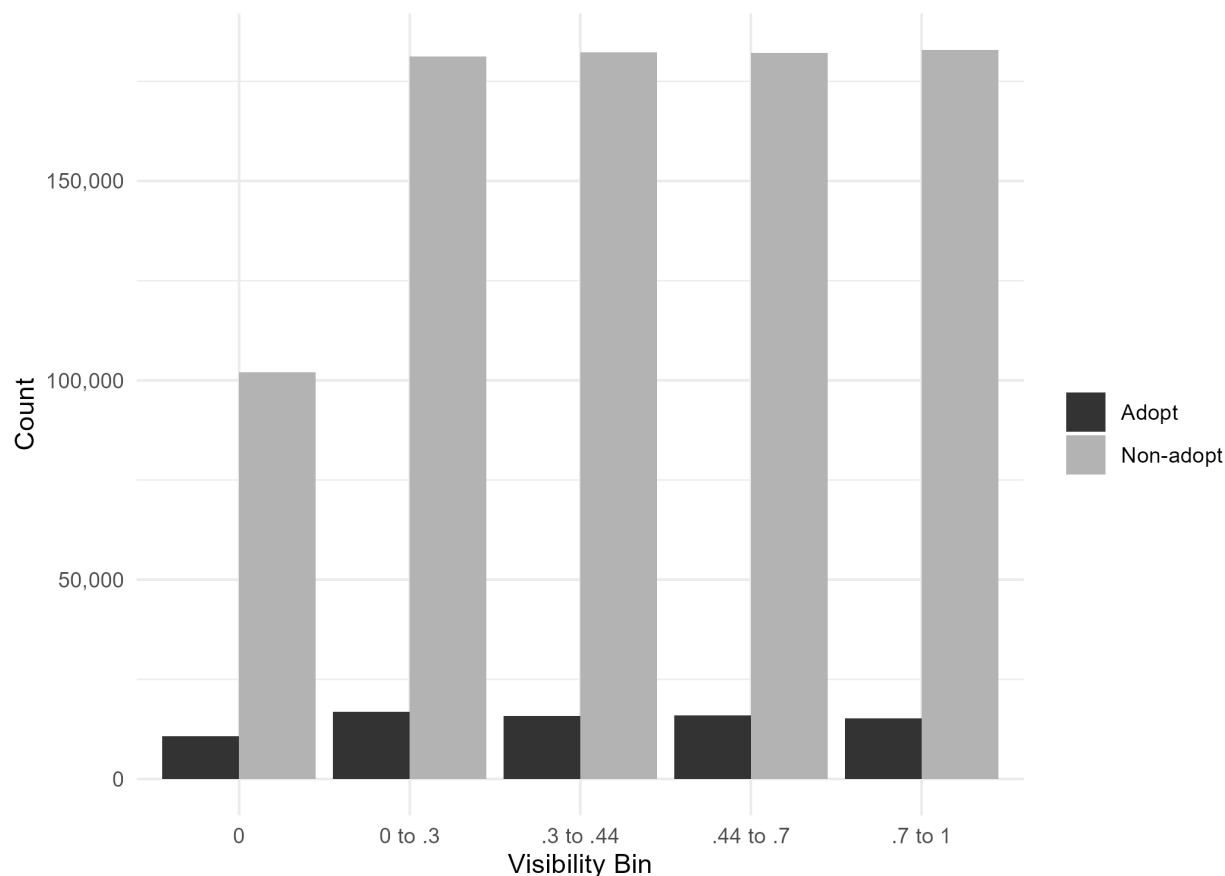


Figure D.1: **Distribution of Adopters and Non-adopters in Visibility Bins.** Count of adopters and non-adopters in each of five visibility bins. The first bin contains all zero-visibility households, comprised of 12% of all households. 14.4% of all adopters lie in the zero-visibility bin, while 12.2% of all non-adopters lie in the zero-visibility bin.

Results in Table D.10 show that the effect of visibility is clearly highest in bin 3 (0.44 to 0.77) and bin 4 (0.77 to 1.00). Small “doses” of visibility are not statistically distinguishable from 0.

Table D.10: Robustness Results: Binning continuous visibility measure.

Dependent Variable: Model: Neighborhood definition:	(1) Block x Street	Adopt (2) Blockgroup x Street	(3) Block
<i>Variables</i>			
Republican hh	-0.140* (0.058)	-0.137* (0.059)	-0.136* (0.059)
Expected value	0.180*** (0.022)	0.183*** (0.022)	0.183*** (0.023)
Expected value × Republican hh	0.001 (0.016)	0.003 (0.016)	0.006 (0.017)
Expected value × Unregistered voter hh	-0.040 (0.026)	-0.040 (0.026)	-0.043 (0.026)
Republican nbhd	0.180* (0.077)	0.242* (0.097)	0.417*** (0.117)
Republican nbhd × Republican hh	-0.087 (0.125)	-0.144 (0.152)	-0.174 (0.191)
Republican nbhd × Unregistered voter hh	0.083 (0.163)	0.094 (0.189)	0.186 (0.244)
Republican nbhd × Expected value	-0.204** (0.066)	-0.268** (0.086)	-0.282** (0.100)
Republican nbhd × Expected value × Republican hh	0.032 (0.064)	0.037 (0.081)	0.016 (0.097)
Republican nbhd × Expected value × Unregistered voter hh	0.066 (0.077)	0.052 (0.099)	0.134 (0.120)
VisBin1	-0.033 (0.020)	-0.034 (0.021)	-0.030 (0.021)
VisBin2	-0.037+ (0.021)	-0.037+ (0.021)	-0.034 (0.021)
VisBin3	-0.044* (0.020)	-0.045* (0.020)	-0.040* (0.020)
VisBin4	-0.094*** (0.021)	-0.094*** (0.021)	-0.090*** (0.022)
VisBin1 × Republican nbhd	0.014 (0.094)	0.027 (0.115)	0.007 (0.131)
VisBin2 × Republican nbhd	-0.124 (0.100)	-0.119 (0.125)	-0.226+ (0.137)
VisBin3 × Republican nbhd	-0.028 (0.098)	-0.064 (0.119)	-0.146 (0.135)
VisBin4 × Republican nbhd	-0.216* (0.100)	-0.226+ (0.120)	-0.327* (0.136)
VisBin1 × Republican hh	0.053 (0.036)	0.047 (0.038)	0.049 (0.039)
VisBin2 × Republican hh	0.045 (0.037)	0.048 (0.038)	0.049 (0.041)
VisBin3 × Republican hh	0.021 (0.036)	0.011 (0.038)	0.008 (0.040)
VisBin4 × Republican hh	0.028 (0.039)	0.026 (0.040)	0.018 (0.043)
VisBin1 × Unregistered voter hh	0.046 (0.038)	0.049 (0.038)	0.053 (0.039)
VisBin2 × Unregistered voter hh	0.037 (0.038)	0.039 (0.038)	0.044 (0.040)
VisBin3 × Unregistered voter hh	0.030 (0.039)	0.032 (0.039)	0.039 (0.040)
VisBin4 × Unregistered voter hh	0.082* (0.039)	0.083* (0.039)	0.089* (0.040)
VisBin1 × Expected value	0.004 (0.021)	0.002 (0.021)	0.005 (0.022)
VisBin2 × Expected value	0.011 (0.020)	0.012 (0.020)	0.012 (0.021)
VisBin3 × Expected value	0.011 (0.020)	0.010 (0.020)	0.014 (0.021)
VisBin4 × Expected value	0.003 (0.019)	0.001 (0.019)	0.004 (0.020)
VisBin1 × Republican nbhd × Republican hh	-0.067 (0.152)	-0.024 (0.185)	-0.035 (0.230)
VisBin2 × Republican nbhd × Republican hh	0.093 (0.169)	0.053 (0.203)	0.101 (0.249)
VisBin3 × Republican nbhd × Republican hh	-0.078 (0.159)	0.040 (0.203)	0.145 (0.232)
VisBin4 × Republican nbhd × Republican hh	0.141 (0.166)	0.162 (0.207)	0.331 (0.262)
VisBin1 × Republican nbhd × Unregistered voter hh	-0.131 (0.189)	-0.256 (0.217)	-0.301 (0.259)
VisBin2 × Republican nbhd × Unregistered voter hh	-0.079 (0.208)	-0.120 (0.238)	-0.087 (0.258)
VisBin3 × Republican nbhd × Unregistered voter hh	-0.094 (0.193)	-0.149 (0.232)	-0.380 (0.278)
VisBin4 × Republican nbhd × Unregistered voter hh	0.113 (0.206)	0.146 (0.238)	-0.002 (0.282)
VisBin1 × Republican nbhd × Expected value	0.104 (0.082)	0.145 (0.095)	0.035 (0.116)
VisBin2 × Republican nbhd × Expected value	0.037 (0.078)	-0.038 (0.105)	0.033 (0.125)
VisBin3 × Republican nbhd × Expected value	0.229** (0.082)	0.321** (0.099)	0.174 (0.114)
VisBin4 × Republican nbhd × Expected value	0.135+ (0.075)	0.157 (0.098)	0.101 (0.106)
<i>Fixed-effects</i>			
Blockgroup-Registered voter	Yes	Yes	Yes
<i>Fit statistics</i>			
Observations	818,086	818,086	818,086
Pseudo R ²	0.10485	0.10487	0.10489
Log-Likelihood	-223,388.2	-223,382.9	-223,377.8

Conley (2km) standard-errors in parentheses
Signif. Codes: ***: 0.001, **: 0.01, *: 0.05, +: 0.1

This finding is important to an interpretation of the neighborhood effect as the self-signal. Since we cannot rule out some neighbor or adjacent property visibility of a potential panel, even for those with zero visibility from the nearest street, showing a negligible role for small amounts of visibility supports the argument that the neighborhood effect is capturing self-signaling, and not other socially-moderated effects.

D.7 Control for pre-2012 adopters

“Early” adopters of new technology may differ from general adopters in unobserved ways that are important when considering technology diffusion (Langer and Lemoine, 2022). We re-estimate our main specification with the aim of avoiding drawing conclusions from early adopters by first adding up all early-adopters in a neighborhood, then dropping these adopters from the analysis but including the total number of early adoptions in the neighborhood, as well as the share of early adoptions that were undertaken by Republican households.

We do so for two reasons. First, early adopters may differ in unobserved ways that are important to consider for external validity of our findings (Langer and Lemoine, 2022). Second, early adopters may influence subsequent adoptions. If households obtain information about adopting solar through networks correlated with political affiliation (e.g. Republican households influence other Republicans through word of mouth, visibility, etc. (Bollinger and Gillingham, 2012; Bollinger et al., 2022), then the estimated effect of the neighborhood political affiliation may absorb this effect. This effect cannot be disentangled from unobserved factors common in the peer effects literature (Manski, 1993), thus we do not place any interpretation on the resulting coefficient. In this application, it serves as a control for early-adopter effects and influences. Finally, controlling for early adopters focuses our identification on the post-2012 time period, where our voter affiliation snapshot (based on voter activity in 2014 and 2016) is taken.³³

³³Our data structure precludes a fully dynamic estimation due to the local nature of our neighborhood definition. The median number of total adoptions in a neighborhood defined at the block level is 1. Identifying dynamic or heterogeneous effects of prior installations under any finer timescale is not feasible.

Table D.11: Robustness Results: Control for pre-2012 early-adopters.

Dependent Variable:	Post-2012 Adopt		
Model:	(1)	(2)	(3)
Neighborhood definition:	Block x Street	Blockgroup x Street	Block
<i>Variables</i>			
Republican hh	-0.092 ⁺ (0.056)	-0.095 ⁺ (0.056)	-0.083 (0.056)
Expected value	0.182*** (0.017)	0.186*** (0.017)	0.188*** (0.018)
Republican hh × Expected value	0.011 (0.017)	0.011 (0.017)	0.019 (0.018)
Unregistered voter hh × Expected value	-0.029 (0.028)	-0.029 (0.028)	-0.032 (0.028)
Visible (frac of degrees from closest street)	-0.080*** (0.021)	-0.083*** (0.021)	-0.076*** (0.021)
Visible (frac of degrees from closest street) × Republican hh	0.0008 (0.039)	0.004 (0.040)	-0.020 (0.043)
Visible (frac of degrees from closest street) × Unregistered voter hh	0.069 (0.042)	0.069 (0.042)	0.071 ⁺ (0.043)
Visible (frac of degrees from closest street) × Expected value	-0.013 (0.019)	-0.017 (0.019)	-0.017 (0.020)
Republican nbhd	0.210*** (0.054)	0.291*** (0.073)	0.422*** (0.095)
Republican nbhd × Republican hh	-0.114 (0.089)	-0.151 (0.113)	-0.205 (0.138)
Republican nbhd × Unregistered voter hh	-0.052 (0.117)	-0.170 (0.140)	-0.030 (0.191)
Republican nbhd × Expected value	-0.142* (0.056)	-0.226** (0.072)	-0.276** (0.087)
Visible (frac of degrees from closest street) × Republican nbhd	-0.199 ⁺ (0.102)	-0.248* (0.126)	-0.349* (0.144)
Visible (frac of degrees from closest street) × Republican nbhd × Republican hh	0.145 (0.173)	0.143 (0.227)	0.464 ⁺ (0.273)
Visible (frac of degrees from closest street) × Republican nbhd × Unregistered voter hh	0.167 (0.220)	0.246 (0.259)	0.048 (0.304)
Visible (frac of degrees from closest street) × Republican nbhd × Expected value	0.104 (0.089)	0.178 (0.111)	0.202 ⁺ (0.114)
N.adoptOld.blockstreet	0.093*** (0.016)		
Rshare.adoptOld.blockstreet	0.021 (0.040)		
N.adoptOld.blockstreet × Republican hh	-0.003 (0.024)		
N.adoptOld.blockstreet × Unregistered voter hh	-0.007 (0.033)		
Republican hh × Rshare.adoptOld.blockstreet	-0.015 (0.060)		
Unregistered voter hh × Rshare.adoptOld.blockstreet	-0.008 (0.082)		
Republican nbhd × Republican hh × Expected value	0.024 (0.069)	0.052 (0.085)	-0.027 (0.103)
Republican nbhd × Unregistered voter hh × Expected value	0.074 (0.081)	0.076 (0.102)	0.160 (0.124)
N.adoptOld.BGS		0.036*** (0.007)	
Rshare.adoptOld.BGS		0.031 (0.028)	
N.adoptOld.BGS × Republican hh		-0.013 (0.008)	
N.adoptOld.BGS × Unregistered voter hh		-0.014 (0.014)	
Republican hh × Rshare.adoptOld.BGS		0.048 (0.041)	
Unregistered voter hh × Rshare.adoptOld.BGS		0.009 (0.059)	
N.adoptOld.block			0.058*** (0.010)
Rshare.adoptOld.block			0.024 (0.031)
N.adoptOld.block × Republican hh			-0.016 (0.013)
N.adoptOld.block × Unregistered voter hh			-0.013 (0.020)
Republican hh × Rshare.adoptOld.block			-0.032 (0.042)
Unregistered voter hh × Rshare.adoptOld.block			0.024 (0.058)
<i>Fixed-effects</i>			
Blockgroup-Registered voter	Yes	Yes	Yes
<i>Fit statistics</i>			
Observations	796,838	796,838	796,838
Pseudo R ²	0.10039	0.10033	0.10043
Log-Likelihood	-198,489.3	-198,502.5	-198,480.3

Conley (2km) standard-errors in parentheses
Signif. Codes: ***, 0.001, **, 0.01, *, 0.05, +: 0.1
BGS is blockgroup x street, blockstreet is block x street

Results in Table D.11 are robust to the inclusion of both the count of early adopters in a neighborhood, the political affiliation of the early adopters, and the interaction of own-affiliation with these measures. The main results for Republican nbhd are of larger magnitude in all three specifications and significant. The visibility measure interacted with Republican nbhd in column (1) is significant only at the 10% level. Results are otherwise robust. The coefficient on the count of early adopters ($N.adoptOld$) is positive and significant, consistent with both targeted marketing (those targeted before 2012 are likely targeted after as well) and information diffusion. None of the heterogeneity by political affiliation parameters show a significant difference across political affiliation.

Because our main results persist (and in many cases become stronger), even when controlling for early adopters, and because political affiliation of early adopters appears to have no correlation with adoption, we interpret our results as robust to early adoption dynamics and unaffected by the voter affiliation “snapshot” used to construct our data.

D.8 Split Samples

We split the sample between household political affiliation and blockgroup density (splitting at median density). Qualitatively, voter affiliation results in Table D.12 are similar to the pooled sample, unregistered households generally fall between Democratic/mixed and Republican, with the exception of the main effect of visibility. On density results, effects of Republican neighborhood affiliation (self-signaling) and the interaction of visibility with Republican neighborhood affiliation (social signaling) are strongest in the high-density subsample. While point estimates are larger in the high-density subsample, an inter-sample test for difference in these estimates between densities fails to reject for any of the neighborhood definitions. Point estimates are consistent with our model — the internalization of neighborhood affiliation as a prior belief for own-type should be stronger when houses are closer together and interactions are more frequent and salient.

Table D.12: Robustness Results: Sample split by registration.

Nblid definition: Sample split: Model:	Block x Street					Blockgroup x Street					Block				
	Dem (1)	Rep (2)	Unreg (3)	All: Low Dens (4)	All: Hi Dens (5)	Dem (6)	Rep (7)	Unreg (8)	All: Low Dens (9)	All: Hi Dens (10)	Dem (11)	Rep (12)	Unreg (13)	All: Low Dens (14)	All: Hi Dens (15)
<i>Variables</i>															
Expected value	0.201*** (0.020)	0.162*** (0.025)	0.135*** (0.029)	0.156*** (0.026)	0.214*** (0.028)	0.203*** (0.020)	0.164*** (0.025)	0.135*** (0.029)	0.162*** (0.020)	0.209*** (0.028)	0.204*** (0.020)	0.180*** (0.026)	0.132*** (0.030)	0.171*** (0.022)	0.214*** (0.028)
Republican nbhd	0.164** (0.050)	0.064 (0.075)	0.196* (0.098)	0.085 (0.064)	0.208* (0.087)	0.235*** (0.070)	0.105 (0.101)	0.204* (0.118)	0.098 (0.085)	0.315** (0.125)	0.329*** (0.092)	0.247* (0.131)	0.439** (0.163)	0.251* (0.119)	0.305* (0.145)
Republican nbhd × Expected value	-0.175** (0.059)	-0.011 (0.083)	-0.241* (0.095)	-0.134* (0.062)	-0.186* (0.100)	-0.293*** (0.079)	-0.028 (0.118)	-0.287* (0.117)	-0.206** (0.080)	-0.334*** (0.130)	-0.334*** (0.095)	-0.182 (0.141)	-0.191 (0.171)	-0.314** (0.098)	-0.127 (0.147)
Republican nbhd × Republican hh				0.017 (0.094)	-0.181 (0.153)				0.072 (0.123)	-0.292 (0.189)				-0.001 (0.224)	-0.080 (0.224)
Republican nbhd × Unregistered voter hh				0.035 (0.135)	-0.004 (0.181)				0.070 (0.163)	-0.197 (0.231)				0.251 (0.217)	-0.056 (0.280)
Republican nbhd × Expected value × Republican hh				0.030 (0.073)	0.099 (0.142)				0.039 (0.095)	0.180 (0.176)				0.090 (0.113)	0.031 (0.190)
Republican nbhd × Expected value × Unregistered voter hh				0.049 (0.088)	0.114 (0.166)				0.036 (0.112)	0.074 (0.201)				-0.021 (0.139)	0.421* (0.214)
Visible (frac of degrees from closest street)	-0.086*** (0.019)	-0.103** (0.037)	-0.023 (0.035)	-0.103*** (0.029)	-0.079*** (0.028)	-0.087*** (0.019)	-0.104** (0.038)	-0.024 (0.035)	-0.103*** (0.029)	-0.084** (0.028)	-0.085*** (0.019)	-0.110** (0.040)	-0.016 (0.035)	-0.095** (0.030)	-0.080*** (0.028)
Visible (frac of degrees from closest street) × Republican nbhd	-0.180* (0.096)	-0.030 (0.149)	-0.118 (0.190)	-0.082 (0.122)	-0.324* (0.160)	-0.242* (0.116)	-0.020 (0.098)	-0.070 (0.217)	-0.095 (0.146)	-0.471* (0.205)	-0.284* (0.135)	0.042 (0.222)	-0.358 (0.259)	-0.204 (0.181)	-0.395* (0.219)
Visible (frac of degrees from closest street) × Expected value	-0.013 (0.022)	0.007 (0.040)	0.035 (0.039)	0.028 (0.029)	-0.039 (0.027)	-0.014 (0.022)	0.008 (0.040)	0.034 (0.039)	0.018 (0.030)	-0.040 (0.028)	-0.014 (0.023)	-0.016 (0.042)	0.040 (0.040)	0.021 (0.031)	-0.044 (0.028)
Visible (frac of degrees from closest street) × Republican nbhd × Expected value	0.161 (0.112)	-0.068 (0.159)	0.556** (0.212)	0.103 (0.110)	0.215 (0.147)	0.324* (0.131)	-0.080 (0.221)	0.536* (0.232)	0.225 (0.138)	0.157 (0.209)	0.212 (0.146)	0.167 (0.249)	0.297 (0.293)	0.162 (0.146)	0.010 (0.196)
Visible (frac of degrees from closest street) × Republican hh				0.016 (0.048)	-0.039 (0.057)				0.029 (0.050)	-0.042 (0.058)				-0.011 (0.056)	-0.034 (0.058)
Visible (frac of degrees from closest street) × Unregistered voter hh				0.066 (0.056)	0.093 (0.058)				0.067 (0.056)	0.104* (0.059)				0.071 (0.059)	0.082 (0.059)
Visible (frac of degrees from closest street) × Republican nbhd × Republican hh				-0.114 (0.193)	0.672* (0.303)				-0.219 (0.252)	1.03** (0.373)				0.165 (0.331)	0.681* (0.396)
Visible (frac of degrees from closest street) × Republican nbhd × Unregistered voter hh				0.083 (0.276)	0.316 (0.336)				0.071 (0.314)	0.657 (0.402)				0.014 (0.377)	-0.026 (0.460)
Republican hh				-0.005 (0.060)	-0.212** (0.082)				-0.011 (0.061)	-0.210* (0.082)				0.0005 (0.082)	-0.208* (0.082)
Expected value × Republican hh				0.015 (0.020)	0.022 (0.024)				0.016 (0.021)	0.024 (0.024)				0.012 (0.023)	0.024 (0.024)
Expected value × Unregistered voter hh				-0.019 (0.030)	-0.050 (0.045)				-0.019 (0.030)	-0.061 (0.045)				-0.013 (0.031)	-0.041 (0.046)
<i>Fixed-effects</i>															
Blockgroup-Registered voter	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
<i>Fit statistics</i>															
Observations	457,138	152,609	186,976	352,332	465,754	457,138	152,609	186,976	352,332	465,754	457,138	152,609	186,976	352,332	465,754
Pseudo R ²	0.0702	0.10784	0.11825	0.09597	0.10077	0.09705	0.10785	0.11820	0.09599	0.10079	0.09707	0.10792	0.11821	0.09607	0.10077
Log-Likelihood	-127,058.4	-50,926.4	-42,596.2	-114,335.7	-108,357.0	-127,055.0	-50,925.9	-42,598.3	-114,334.2	-108,354.7	-127,051.2	-50,922.2	-42,598.0	-114,323.4	-108,357.5
<i>Conley (2km) standard-errors in parentheses</i>															
Signif. Codes: ***, 0.001, **, 0.01, *, 0.05, +: 0.1															