ORIGINAL ARTICLE
Journal Section

Simple rules to guide expert classifications

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Funding information
N/A

Judges, doctors, and managers are among those decision makers who must often choose a course of action under limited time, with limited knowledge, and without the aid of a computer. Because data-driven methods typically outperform unaided judgments, resource-constrained practitioners can benefit from simple, statistically derived rules that can be applied mentally. In this work, we formalize longstanding observations about the efficacy of improper linear models to construct accurate yet easily applied rules. To test the performance of this approach, we conduct a large-scale evaluation in 22 domains and focus in depth on one: judicial decisions to release or detain defendants while they await trial. In these domains, we find that simple rules rival the accuracy of complex prediction models that base decisions on considerably more information. Further, comparing to unaided judicial decisions, we find that simple rules substantially outperform the human experts. To conclude, we present an analytical framework that sheds light on why simple rules perform as well as they do.

KEYWORDS
Heuristics, judgment and decision making, policy evaluation, sensitivity analysis

Abbreviations: FTA, failure to appear; RoR, release on recognizance.

All authors contributed equally.

1 | INTRODUCTION

In field settings, decision makers often choose a course of action based on experience and intuition rather than on statistical analysis (Klein, 2017). This includes doctors classifying patients based on their symptoms (McDonald, 1996), judges setting bail amounts (Dhami, 2003) or making parole decisions (Danziger et al., 2011), and managers determining which ventures will succeed (Åstebro and Elhedhli, 2006) or which customers to target (Wübben and Wangenheim, 2008). Despite the prevalence of this approach, a large body of work shows that in many domains intuitive inferences are inferior to those based on statistical models (Meehl, 1954; Dawes, 1979; Dawes et al., 1989; Camerer and Johnson, 1997; Tetlock, 2005; Kleinberg et al., 2015, 2017).

In this work, we generalize from research on improper linear models (Einhorn and Hogarth, 1975; Green, 1977; Dawes, 1979; Gigerenzer and Goldstein, 1996; Waller and Jones, 2011) to suggest a straightforward method for constructing simple yet accurate decision rules that experts can apply mentally. This *select-regress-and-round* method results in rules that are fast, frugal, and clear: fast in that decisions can be made quickly in one's mind, without the aid of a computer; frugal in that they require very little information to reach a decision; and clear in that they expose the grounds on which classifications are made.

Decision rules satisfying these criteria have many benefits. Fast rules that can be applied mentally reduce transaction costs, encouraging persistent use. In medicine, frugal rules require fewer tests, which saves time, money, and, in the case of triage situations, lives (Marewski and Gigerenzer, 2012). Frugal decision rules incorporating predictors that are broadly related to outcomes of interest are well suited for settings in which a model highly customized for one population may not generalize to other populations (Wyatt and Altman, 1995). The clarity of simple rules provides insight into how systems work and exposes where models may be improved (Gleicher, 2016; Sull and Eisenhardt, 2015), which may encourage adoption of such tools in clinical settings (Wyatt and Altman, 1995). Clarity can even become a legal requirement when society demands to know how algorithmic decisions are being made (Goodman and Flaxman, 2016; Corbett-Davies et al., 2017).

After describing the select-regress-and-round method, we evaluate its efficacy on 21 datasets from the UCI Machine Learning repository, and show that in many cases simple rules are competitive with state-of-the-art machine learning algorithms. To illustrate in detail the value of simple rules, we present a case study of judicial decisions for pretrial release. Based on an analysis of over 100,000 cases, we show that simple rules substantially improve upon the efficiency and equity of unaided judicial decisions. In particular, we estimate that judges can detain one-third fewer defendants while simultaneously increasing the number that appear at their court dates.[1] In the judicial context, as in many policy settings, it is statistically challenging to evaluate decision rules based solely on historical data. The key difficulty is that one cannot observe what would have happened under an alternative course of action. What would have happened, for example, if one released a defendant who in reality was detained? We address this issue by first estimating the relevant counterfactual outcomes, and then assessing the sensitivity of our estimates to unobserved confounding, generalizing the technique of Rosenbaum and Rubin (1983a).

Our results add to a growing literature on *interpretable machine learning*. In addition to methods for better understanding complex machine learning models and data structures (Kim et al., 2015; Ribeiro et al., 2016), several methods have been introduced to construct interpretable decision rules, similar to the simple decision rules we discuss here. For example, Van Belle et al. (2012) use convex optimization to build interval coded scoring models for binary outcomes. More general methods for constructing interpretable decision rules have been recently proposed, including supersparse linear integer models (SLIM) (Ustun and Rudin, 2016), Bayesian rule lists (Wang and Rudin, 2015), and interpretable decision sets (Lakkaraju et al., 2016). These methods all produce rules that are easy to interpret and to

[1] Kleinberg et al. (2017) recently and independently proposed using machine learning models to assist judicial decisions, but they do not consider simple rules.

apply but the methods differ considerably on the ease of rule creation. As an important practical consideration, the method we investigate here can be carried out by a practitioner without extensive training in statistics, using popular open-source software—though it bears emphasis that appropriate application of all statistical methods requires both domain knowledge and familiarity with the relevant data.

2 | SELECT-REGRESS-AND-ROUND: A SIMPLE METHOD FOR CREATING SIMPLE RULES

We begin by presenting a simple method—which we call *select-regress-and-round*—for constructing simple decision rules. This procedure generalizes ideas that appear throughout the judgment and decision-making literature on improper linear scoring rules, and formalizes heuristics used by practitioners in creating decision aids.

The rules we construct are designed to aid classification or ranking decisions by assigning each item in consideration a score z, computed as a linear combination of a subset S of the item features:

$$z = \sum_{j \in S} w_j x_j,$$

where the weights w_j are integers. In the cases we consider, the features themselves are typically 0-1 indicator variables (indicating, for example, whether a person is male, or whether an individual is 26–30 years old), and so the rule reduces to a weighted checklist, in which one simply sums up the (integer) weights of the applicable attributes.[2] Often, one seeks to make dichotomous decisions (*e.g.*, whether to detain or to release an individual pending trial), which amounts to setting a threshold and then taking a particular course of action if and only if the score is above that threshold.

This class of rules has two natural dimensions of complexity: the number of features included in the subset S, and the magnitude of the weights. Given integers $k \geq 1$ and $M \geq 1$, we apply the following three-step procedure to construct rules with at most k features and integer weights bounded by M (*i.e.*, $|S| \leq k$ and $-M \leq w_j \leq M$).

1. **Select.** From the full set of features, select k features via forward stepwise selection. This is done by iteratively adding the feature that minimizes AIC. For fixed k, we note that standard selection metrics (*e.g.*, AIC or BIC) are theoretically guaranteed to yield the same set of features.
2. **Regress.** Using only these k selected features, train an L^1-regularized (lasso) logistic regression model to the data, which yields (real-valued) fitted coefficients $\beta_1, \ldots, \beta_k$.
3. **Round.** Rescale the coefficients to be in the range $[-M, M]$, and then round the rescaled coefficients to the nearest integer. Specifically, set

$$w_j = \text{Round}\left(\frac{M\beta_j}{\max_i |\beta_i|}\right).$$

This *select-regress-and-round* method for rule construction extends research on unit-weighted linear models by incorporating feature selection and by adopting more general integer weights to generate a richer family of rules, the accuracy of which we examine in the next section. In practice, we recommend that developers of such rules apply the

[2] It is possible to directly apply select-regress-and-round to continuous features but, in the spirit of simplicity and interpretability, we recommend discretizing continuous covariates, using, for example, three equal-sized bins, as proposed in Gelman and Park (2009). But in practice, as always, domain knowledge and technical considerations play an important role in determining appropriate transformations or discretization schemes. For example, rather than simply partitioning an age covariate into three bins, one might use 10-year buckets. Similarly, one might collapse categorical features with several levels into a smaller number of more semantically meaningful groups.

procedure for a range of small values of k and M that are appropriate to their domain, and then pick the values that perform best on context-specific metrics, balancing simplicity with performance, an approach we illustrate below.

We note that rules constructed in this way may have fewer than k features, since the lasso regression in Step 2 may result in coefficients that are identically zero, and rescaling and rounding coefficients in Step 3 may zero-out additional terms. For Step 2, the regularization parameter, λ, is chosen via cross-validation.[3] In our applications, following Friedman et al. (2010), we explore a regularization path with 1,000 values of λ spaced evenly on the log-scale in the range $(\lambda_{\text{min}}, \lambda_{\text{max}})$, where $\lambda_{\text{min}} = 10^{-4}$ and λ_{max} is selected as the minimum value such that all coefficients are regularized to zero.

3 | EVALUATING THE EFFICACY OF SIMPLE RULES

We apply the select-regress-and-round procedure to 21 publicly available datasets to examine the tradeoff between rule complexity and performance. These datasets all come from the UCI Machine Learning Repository (see Table 1), and were selected according to four criteria: (1) the dataset involves binary classification (as opposed to a regression problem), where we set the majority class as the target of prediction for those datasets whose outcome variable takes more than two values; (2) the dataset is provided in a standard and complete form; (3) the dataset involves more than 10 (binarized) features; and (4) the classification problem is one that a human could plausibly learn to solve with the given features. For example, we included a dataset in which the task was to determine whether cells were malignant or benign based on various biological attributes of the cells, but we excluded image recognition tasks in which the features were represented as pixel values. This fourth requirement limits the scope of our analysis and conclusions to domains in which human decision makers typically act without the aid of a computer.

3.1 | Benchmarking to complex prediction models

We benchmark the performance of our simple rules against three standard statistical models: logistic regression, L^1-regularized logistic regression, and random forest. Random forest, in particular, is considered to be one of the best off-the-shelf classification algorithms in machine learning (Fernández-Delgado et al., 2014; Kleinberg et al., 2017). These models were fit in `R` with the `glm`, `glmnet`, and `randomForest` packages, respectively. For the L^1-regularized logistic regression models, the `cv.glmnet` method was used to determine the best value of the regularization parameter λ with nested 10-fold cross-validation and 1,000 values of λ. We used 1,000 trees for the random forest models.

Across the 21 UCI datasets, variables are documented as *categorical* (discrete and unordered), *ordinal* (discrete and ordered), or *continuous*. For our simple rules, we represent discrete covariates—both categorical and ordinal—as a series of binary indicator variables, with one indicator per category. In particular, for simplicity, we ignore the category ranking in ordinal variables. Further, all continuous features are discretized into three approximately equal-sized bins representing (categorical) low, medium, and high values of the feature, following Gelman and Park (2009). For the three complex models, we include the above feature representations, as well as the original (non-discretized) values of continuous variables. Also, for ordinal variables—in addition to their unordered categorical representation—we include a feature representation that preserves the order of categories.[4]

[3] Unlike λ, the parameters k and M cannot be selected via an automated procedure unless one formally quantifies the tradeoff between performance and simplicity, since both performance and complexity increase with larger values of k and M. However, in practice, as we show below, one might achieve approximately the same performance as a traditional logistic regression model with relatively small values of k and M, meaning the tradeoff may be negligible.

[4] As is common, the categories of an ordinal variable are represented as sequential integers, with our complex models fitting orthogonal polynomials to these integer values (Chambers et al., 1992).

TABLE 1 Summary of UCI datasets. For each domain, we report: the name of the dataset; the number of rows and features (columns excluding the class label) in the original dataset; the number of complete rows with no missing data; the number of continuous features, as well as the number of features after discretizing continuous variables and expanding categorical variables to binary indicators; and the proportion of instances in the most common class (proportion majority). The context of each domain is presented in detail in Appendix A.

Domain	Instances	Features	Complete instances	Continuous features	Binarized features	Proportion majority
1. adult	32,561	14	30,162	4	96	25
2. annealing	798	38	798	7	54	76
3. audiology-std	200	41	190	0	55	24
4. bank	41,188	20	41,188	9	62	11
5. bankruptcy	250	6	250	0	13	43
6. car	1,728	6	1,728	0	16	70
7. chess-krvk	28,056	6	28,056	0	35	10
8. chess-krvkp	3,196	36	3,196	0	37	52
9. congress-voting	435	16	232	0	17	53
10. contrac	1,473	9	1,473	2	20	43
11. credit-approval	690	15	653	6	44	45
12. ctg	2,126	38	2,126	33	67	78
13. cylinder-bands	541	39	279	19	65	65
14. dermatology	366	34	358	34	69	31
15. german_credit	1,000	20	1,000	7	56	70
16. heart-cleveland	303	13	299	6	26	46
17. ilpd	583	10	579	9	20	72
18. mammo	961	5	830	1	18	49
19. mushroom	8,124	22	5,644	0	76	38
20. wine	178	13	178	13	27	40
21. wine_qual	6,497	12	6,468	11	24	63

On each of the UCI datasets we analyze here, we construct a family of simple rules having $k \in \{1, \ldots, 10\}$ features, with feature weights bounded by $M \in \{1, 2, 3\}$. We count the number of features k prior to binarization. For example, a categorical covariate with five possible values—and hence converted to five binary variables—counts as one of the k features in the simple rule, not five. The head-to-head comparison with complex models provides a difficult test for the simple rules in part because the simple rules can only base their predictions on 1 to 10 features. The complex models, in contrast, can train and predict with all the features in a domain, which number between 5 and 41 with a mean of 20.[5] We provide the complex models with an additional advantage over the simple rules by including continuous and ordinal features in their native representation as well as their unordered, discretized versions. In contrast, the simple rules only include the unordered, discretized versions.

Figure 1 shows model performance, measured in terms of mean cross-validated AUC (i.e., area under the receiver operating characteristic curve) across the 21 datasets, as a function of model size and coefficient range. The AUC for

[5] In theory, the out-of-sample performance of logistic regression could be improved by incorporating a variable selection step. However, we find no qualitative difference in performance when adding a variable selection step in our specific case.

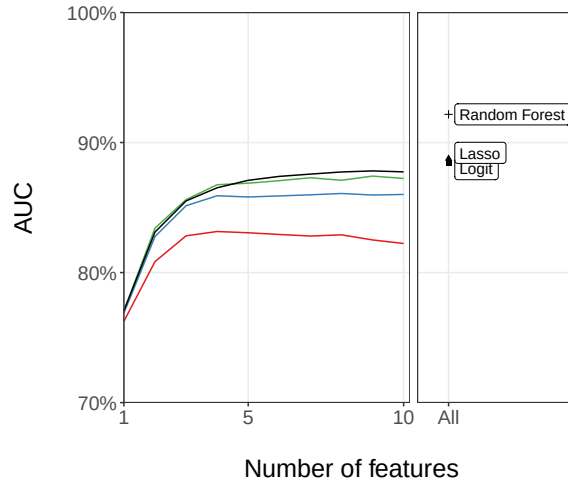

FIGURE 1 Performance of simple and complex rules. Performance is measured in terms of mean cross-validated AUC over all 21 datasets. The black line represents simple models with no rounding, and the green, blue, and red lines represent simple models rounding coefficients to $[-3, 3]$, $[-2, 2]$, and $[-1, 1]$, respectively. The simple models can predict with up to 10 features. The number of "all" features used by random forest, lasso, and logistic regression varied by domain, with an average of 20.

each model on each dataset is computed via 10-fold cross-validation. We find that simple rules with only five features and integer coefficients between -3 and 3 perform on par with logistic regression and L^1-regularized logistic regression trained on the full set of features. For 1 to 10 features, the [-3, 3] model (green line) differs from the unrounded lasso model (black line) by less than 1 percentage point. The performance of the random forest model—which is designed to capture non-linear structure—is somewhat better: trained on all features, random forest achieves a mean AUC of 92%; the mean AUC is 87% for simple rules with at most five features and integer coefficients between -3 and 3.

In Appendix B.2, we examine the performance of select-regress-and-round for each of the 21 UCI datasets separately. As Figures B2 and B3 demonstrate, across almost all datasets, simple rules have comparable AUC to logistic regression (with or without regularization), and have slightly lower AUC than a random forest model. As these results indicate, complex prediction methods certainly have their advantages, but the gap in performance between simple rules and fully optimized prediction methods is not as large as one might have thought.

3.2 | Benchmarking to integer programming

The simple rules we construct take the form of a linear scoring rule with integer weights. To produce such rules, mixed-integer programming is a natural alternative to our select-regress-and-round method, and supersparse linear integer models—abbreviated SLIM (Ustun and Rudin, 2016)—is the leading instantiation of that approach, to which we now compare. Integer programming is an NP-hard problem, and so following Ustun and Rudin (2016) we set a time limit for SLIM: a 10-minute limit is set in the original paper, but we allow up to 6 hours of computation per model. For 7 of the 21 datasets, SLIM found an integer-optimal solution within the time limit, and it returned approximate solutions in the remaining 14 cases.

Figure 2 compares the binary classification accuracy of SLIM and select-regress-and-round on the 21 UCI datasets,

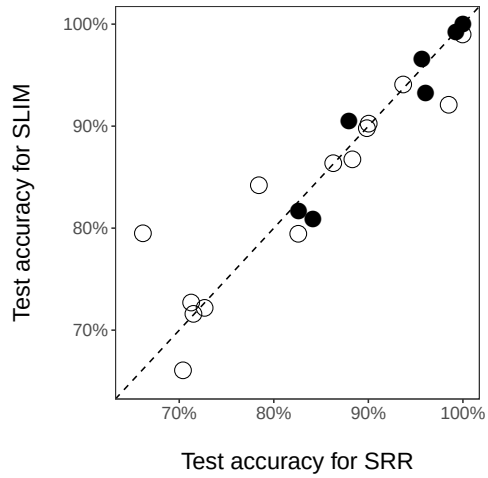

FIGURE 2 Comparing binary classification accuracy for select-regress-and-round (SRR) and SLIM on 21 UCI datasets. Solid dots are cases in which SLIM successfully found an optimal integer solution, while the open circles are cases in which the time limit of 6 hours was exceeded.

where each point corresponds to a dataset. Both methods are constrained to produce rules with at most five features and integer coefficients between -3 and 3.[6] We show 0-1 accuracy as opposed to AUC, since SLIM only produces optimized binary decisions, for which AUC is not applicable. In computing 0-1 accuracy for select-regress-and-round, we select a cut point that corresponds to approximately 0.5 on the probability scale. Accuracy is computed out-of-sample via 10-fold cross-validation. Both methods for producing simple rules yield comparable results: averaged across all 21 datasets, SLIM and select-regress-and-round both achieve a mean accuracy of 86%. Even in the 7 cases where SLIM found integer-optimal solutions, performance is nearly identical to the simple select-regress-and-round method.

In terms of classification accuracy, select-regress-and-round generates rules that are on par with those obtained by solving mixed-integer programs. We note, however, two advantages of our approach. First, whereas select-regress-and-round yields results almost instantaneously, integer programs can be computationally expensive to solve. Second, our approach is relatively simple, both conceptually and technically, accordingly easing adoption for practitioners.

4 | CASE STUDY: PRETRIAL RELEASE DECISIONS

To illustrate the value—and challenges—of applying simple decision rules in practice, we now turn to the domain of pretrial release determinations and present an extended case study. In the United States, defendants are typically *arraigned* shortly after arrest in a court appearance where they are provided with written notice of the charges alleged by the prosecutor. At this time, a judge must decide whether a defendant, while awaiting trial, should be *released on one's own recognizance* (RoR), or alternatively, subject to monetary bail. In practice, if the judge rules that bail be set,

[6] In comparing with SLIM, we now define the number of features k to be the number of *binarized* variables—for both SLIM and select-regress-and-round—since this method of accounting is what is used by SLIM. For example, while a single categorical variable with five possible values would have been considered as one feature in the previous section, each possible value is counted as a feature here, and hence including the entire categorical variable would result in a model with five features.

defendants often await trial in jail since many of them do not have the financial resources to post bail. Moreover, when defendants are able to post bail, they often do so by contracting with a bail bondsman and in turn incur hefty fees. The judge, however, has a legal obligation to consider taking measures necessary to secure the defendant's appearance at required court proceedings. Pretrial release decisions must thus balance flight risk against the high burden that bail requirements place on defendants. In practice, judges may consider other factors—e.g., a defendant's threat to public safety or ability to afford bail—but flight risk is the only legally relevant factor for the specific jurisdiction we analyze below.

A key statistical challenge in this setting is that one cannot directly observe the effects of hypothetical decision rules. Unlike the class of prediction problems discussed in Section 3, outcomes in this domain are affected by a judge's decisions, and one only observes the outcomes that result from those decisions. For example, if a proposed policy recommends releasing some defendants who in reality were detained by the judge, one does not observe what would have happened had the rule been followed. This counterfactual estimation problem—also known as offline policy evaluation (Dudík et al., 2011)—is common in many domains. We address it here by adapting tools from causal inference to the policy setting, including the method of Rosenbaum and Rubin (1983a) for assessing the sensitivity of estimated causal effects to unobserved confounding.

Our analysis is based on 165,000 adult cases involving nonviolent offenses charged by a large urban prosecutor's office and arraigned in criminal court between 2010 and 2015. This set was obtained by starting with a random sample of 200,000 cases provided to us by the prosecutor's office, and then restricting to those cases involving nonviolent offenses and for which the records were complete and accurate. Our initial sample of 200,000 cases does not include instances where defendants accepted a plea deal at arraignment, obviating the need for a pretrial release decision. For each case, we have a rich set of attributes: 49 features describe characteristics of the current charges (*e.g.*, theft, gun-related), and 15 describe characteristics of the defendant (*e.g.*, gender, age, prior arrests). We also observe whether the defendant was RoR'd, and whether the defendant failed to appear (FTA) at any of the subsequent court dates. We note that even if bail is set, a defendant may still fail to appear since one can post bail and then miss a court date. Overall, 69% of defendants are RoR'd, and 15% of RoR'd defendants fail to appear. Of the remaining 31% of defendants for whom bail is set, 45% are eventually released and 9% fail to appear. As a result, the overall FTA rate is 13%.

In our analysis below, we randomly divide the full set of 165,000 cases into three approximately equal subsets; we use the first fold to construct decision rules (both simple and complex), and the second and third to evaluate these rules, as described next.

4.1 | Rule construction

We start by constructing traditional, complex statistical decision rules for balancing flight risk against the burdens of bail. These rules serve as a benchmark for evaluating the simple rules we create below. On the first fold of the data, we restrict to cases in which the judge RoR'd the defendant, and then fit an L^1-regularized logistic (lasso) regression and random forest, using the procedures described in Section 3.1, to estimate the likelihood an individual fails to appear at any of their subsequent court dates. We fit these models on all available information about the case and the defendant, excluding race.[7] The fitted models let us compute risk scores (*i.e.*, estimated flight risk if RoR'd) for any defendant. These risk scores can in turn be converted to a binary decision rule by selecting a threshold for releasing individuals. One might, for example, RoR a defendant if and only if their estimated flight risk is below 20%.

To start, we construct a family of simple rules by applying select-regress-and-round as described in Section 2, using

[7] We excluded race from the presented results due to legal and policy concerns with basing decisions on protected attributes (Corbett-Davies et al., 2017). We note, however, that including race does not significantly affect performance.

TABLE 2 A simple rule for estimating flight risk with 5 features: age, prior FTAs, major charge category, housing instability, and defense attorney type. A defendant's flight risk is obtained by summing the corresponding scores for the features that apply to the case.

Feature	Score	Feature	Score
$18 \leq$ age < 26	2	1 prior FTA	2
$26 \leq$ age < 31	1	2 prior FTAs	3
Major charge group A	-2	3 or more prior FTAs	3
Major charge group B	-1	Unstable housing	3
Major charge group C	1	Defense attorney type A	2
Major charge group D	2	Defense attorney type B	-1
Major charge group E	2	Defense attorney type C	-3

all available features.[8] The resulting rule using 5 features with integer coefficients between -3 and 3 is presented in Table 2. Unsurprisingly, missing court appearances in the past is a strong indicator of flight risk, and an individual's risk also declines with age, in line with conventional wisdom. The rule in Table 2, however, may be inappropriate for implementation given that some features and their associated scores could be challenged as undesirable. For example, defendants with unstable housing are rated higher risk, which may be statistically true but which could lead to adverse outcomes for poorer defendants. Particularly in policy domains, feature selection often requires careful thought.

In practice, we recommend that variable selection incorporates domain expertise. For example, starting from a list of predictive features, as in Table 2, one might exclude problematic variables. Based on discussions with experts in our partner prosecutor's office, we ultimately used only two features—age and prior history of failing to appear—which are generally viewed as acceptable considerations in pretrial decision making. In this case, we can think of the "select" step in the select-regress-and-round strategy as incorporating both human and machine judgment. Specifically, we fit the following model:

$$\Pr(Y_i = 1) = \text{logit}^{-1}\left(\beta_0 + \beta_1^{\text{priors}} H_i^1 + \beta_2^{\text{priors}} H_i^2 + \beta_3^{\text{priors}} H_i^3 + \beta_{4+}^{\text{priors}} H_i^{4+} + \beta_{18-20}^{\text{age}} A_i^{18-20} + \cdots + \beta_{46-50}^{\text{age}} A_i^{46-50}\right),$$

where $Y_i \in \{0, 1\}$ indicates whether the i-th defendant failed to appear; $H_i^* \in \{0, 1\}$ indicates the defendant's number of prior failures to appear (exactly one, two, three, or at least four); and $A_i^* \in \{0, 1\}$ indicates the binned age of the defendant (18–20, 21–25, 26–30, 31–35, 36–40, 41–45, or 46–50). The parameters β_*^{priors} and β_*^{age} are the coefficients corresponding to each binary indicator variable. For identifiability, indicator variables for no prior FTAs and age 51-and-older are omitted. As before, this model is fit on the subset of cases in the first fold of data for which the judge released the defendant. Next, we rescale the age and prior FTA coefficients so that they lie in the interval $[-3, 3]$; specifically, we multiply each coefficient by the constant

$$\frac{3}{\max\left(|\beta_1^{\text{prior}}|, \ldots, |\beta_{4+}^{\text{priors}}|, |\beta_{18-20}^{\text{age}}|, \ldots, |\beta_{46-50}^{\text{age}}|\right)}.$$

Finally, we round the rescaled coefficients to the nearest integer.

[8] The exact discretization scheme used for numerical features—such as age and a defendant's number of prior failures to appear—was determined in consulta-

Feature	Score	Feature	Score
18 ≤ age < 21	3	no prior FTAs	0
21 ≤ age < 31	2	1 prior FTA	2
31 ≤ age < 51	1	2 or more prior FTAs	3
51 ≤ age	0		

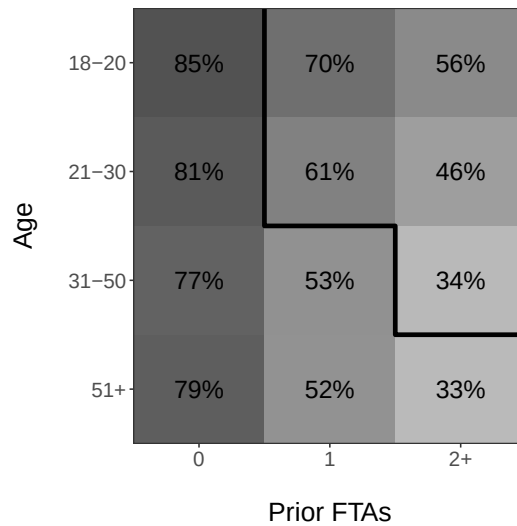

FIGURE 3 A simple rule for estimating flight risk, where a defendant's risk is obtained by summing the appropriate scores for age and prior history of failing to appear (FTA). Also shown is a graphical representation of this rule, based on setting a release threshold of 3.5. Groups to the left of the solid black line in the grid are those that would be released under the rule. For comparison, the shading and numbers in the grid show the proportion of defendants that were actually RoR'd by judges in each group.

Figure 3 shows the result of this procedure. For any defendant, a risk score can be computed by summing the relevant terms in the table. These risk scores can be converted to a binary decision rule by selecting a threshold for releasing individuals. For example, one might RoR a defendant if and only if their risk score is below 3.5; a graphical representation of such a binary decision rule is also shown in the figure.

We note that the application of a simple rule derived from the select-regress-and-round procedure yields an integer score for each defendant. However, in practice it may be useful to also have a probabilistic estimate of each defendant's risk (*i.e.*, the probability that a defendant will fail to appear if released). A given integer score can be converted to a probabilistic risk estimate by considering all released defendants in the training set with that score, and then computing the empirical frequency that those defendants failed to appear. Figure 4 shows the empirical frequency of FTA for each risk score based on the simple rule shown in Figure 3. These empirical frequencies represent probabilistic risk estimates associated with each integer score; for example, the threshold score of 3.5 in Figure 3 corresponds to a risk estimate of 20%. In Appendix B.1, we examine the robustness of these probability estimates, and find that they are comparable to estimates from more complex prediction models.

tion with domain experts in the prosecutor's office with which we worked.

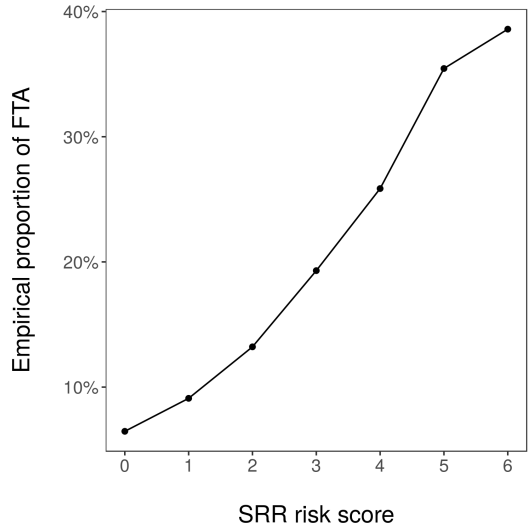

FIGURE 4 The empirical frequency of FTA for each risk score, based on the simple rule shown in Figure 3.

4.2 | Policy evaluation

The AUC is a useful general measure of performance, and hence the metric we consider when evaluating the 21 UCI datasets in Section 3. But in applied settings it is often necessary to directly measure the costs and benefits of any given rule. We do that here by assessing decision rules for pretrial release on two key dimensions: (1) the proportion of defendants who are released under the rule; and (2) the resulting proportion who fail to appear at their court proceedings. It is straightforward to estimate the former, since one need only apply the rule to historical data to see what actions would have been recommended.[9] For example, if defendants are released if and only if their risk score is below 3.5, 79% would be RoR'd; under this rule, bail would be required of only two-thirds as many defendants relative to the status quo. Forecasting the proportion who would fail to appear, however, is generally much more difficult. The key problem is that for any particular defendant, we only observe the outcome (*i.e.*, whether or not the defendant failed to appear) conditional on the action the judge ultimately decided to take (*i.e.*, RoR or bail). Since the action taken by the judge may differ from that prescribed by the decision rule, we do not always observe what would have happened under the rule. This problem of *offline policy evaluation* (Dudík et al., 2011) is a specific instance of the fundamental problem of causal inference.

To rigorously describe the estimation problem and our approach, we introduce some notation. For concreteness, we frame our methodology in terms of the pretrial release example, but the ideas presented here are common to many policy decisions. We denote the observed set of cases by $\Omega = \{(x_i, a_i, r_i)\}$, where x_i is a case, $a_i \in \{\text{RoR}, \text{bail}\}$ is the action taken by the judge, and $r_i \in \{0, 1\}$ indicates whether the defendant failed to appear at a scheduled court date. We write $r_i(\text{RoR})$ and $r_i(\text{bail})$ to mean the *potential outcomes*: what would have happened under the two possible judicial actions. For any policy π, our goal is to estimate the FTA rate under the policy:

$$V^{\pi} = \frac{1}{|\Omega|} \sum_i r_i(\pi(x_i)),$$

[9] In theory, implementing a decision rule could alter the equilibrium distribution of defendants. We do not consider such possible effects, and assume the distribution of defendants is not affected by the rule itself.

TABLE 3 An illustrative example of response surface modeling for offline policy evaluation. For each defendant, $\hat{r}$(RoR) and $\hat{r}$(bail) are model-based estimates of the likelihood of FTA under each potential action. In cases where the observed action equals the proposed action, the observed outcome (FTA or not) is used to estimate the policy's effect; otherwise, the model-based estimates are used. The gray shading indicates which values are used in each instance. The overall FTA rate under the policy is estimated by averaging the shaded values over all cases.

Proposed action (π)	Observed action (a)	Observed outcome ($r(a)$)	$\hat{r}$(RoR)	$\hat{r}$(bail)
RoR	RoR	0	20%	10%
Bail	Bail	1	80%	30%
Bail	RoR	1	90%	70%
RoR	Bail	0	30%	25%
RoR	RoR	0	20%	15%

where $\pi(x)$ denotes the action prescribed under the rule. The key statistical challenge is that only one of the two potential outcomes, $r_i = r_i(a_i)$, is observed. We note that policy evaluation is a generalization of estimating average treatment effects. Namely, the average treatment effect can be expressed as $V^{\pi_{\text{RoR}}} - V^{\pi_{\text{bail}}}$, where π_{RoR} is the policy under which everyone is released and π_{bail} is defined analogously.

Here we take a straightforward and popular statistical approach to estimating V^{π}: response surface modeling (Hill, 2012).[10] With response surface modeling, the idea is to use a standard prediction model (*e.g.*, logistic regression or random forest) to estimate the effect on each defendant of each potential judicial action. The model estimates of these potential outcomes are denoted by $\hat{r}_i(t)$, for $t \in \{\text{RoR}, \text{bail}\}$. Our estimate of V^{π} is then given by

$$\hat{V}^{\pi} = \frac{1}{|\Omega|} \sum_i \left[r_i \mathbf{I}(\pi(x_i) = a_i) + \hat{r}_i(\pi(x_i)) \mathbf{I}(\pi(x_i) \neq a_i) \right],$$

where $\mathbf{I}(\cdot)$ is an indicator function evaluating to 1 if its argument is true and to 0 otherwise. If the prescribed action is in fact taken by the judge, then $r_i = r_i(\pi(x_i))$ is directly observed and can be used; otherwise we approximate the potential outcome with $\hat{r}_i(\pi(x_i))$. Table 3 illustrates this method for a hypothetical example.

Response surface modeling implicitly assumes that a judge's action is *ignorable* given the observed covariates (*i.e.*, that conditional on the observed covariates, those who are RoR'd are similar to those who are not). Formally, ignorability means that

$$(r(\text{RoR}), r(\text{bail})) \perp\!\!\!\perp a \mid x.$$

This ignorability assumption is typically unavoidable, and is similarly required for methods based on propensity scores (Rosenbaum and Rubin, 1983b, 1984; Cassel et al., 1976; Robins et al., 1994; Robins and Rotnitzky, 1995; Kang and Schafer, 2007; Dudík et al., 2011). We examine this assumption in detail in Section 4.3, and find that our conclusions are robust under a common model of unobserved heterogeneity.

To carry out this approach, we derive estimates $\hat{r}_i(t)$ via an L^1-regularized logistic regression (lasso) model trained on the second fold of our data. For each individual, the model estimates likelihood of FTA given all the observed features and the action taken by the judge. In contrast to the rule construction described above, this time we train the model on all cases (not just those for which the judge RoR'd the defendant) and include as a predictor the judge's action (RoR or bail);

[10] We investigated two alternative approaches—inverse propensity weighting (Rosenbaum and Rubin, 1983b, 1984) and doubly robust estimation (Cassel et al., 1976; Robins et al., 1994; Robins and Rotnitzky, 1995; Kang and Schafer, 2007; Dudík et al., 2011)—and found qualitatively similar results.

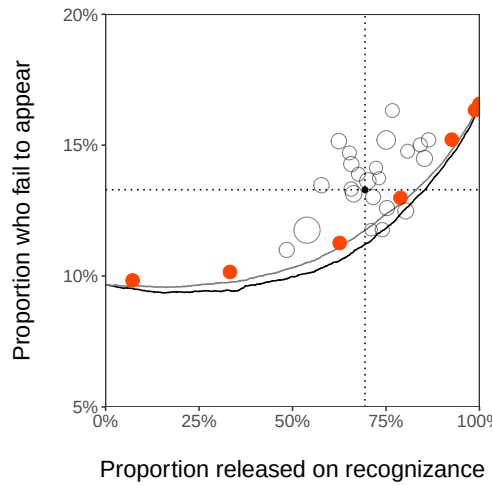

FIGURE 5 Evaluation of simple and complex decision rules. Each point on the solid lines corresponds to decision rules derived from a random forest (gray) or lasso (black) risk model using all 64 features with varying thresholds for release. The red points correspond to policies based on the simple risk score using just two features for all possible release thresholds. The simple rules perform nearly identically to the random forest models, and comparably to the lasso models. The open circles show the observed RoR and FTA rates for each judge in our data who presided over at least 1,000 cases, sized in proportion to their case load. In nearly every instance, the statistical decision rules outperform the human decision maker.

we also include the defendant's race.[11] Then, on the third fold of the data, we use the observed and model-estimated outcomes to approximate the overall FTA rate for any decision rule.

Figure 5 shows estimated RoR and FTA rates for a variety of pretrial release rules. Points on the solid lines correspond to rules constructed via the lasso (black line) and random forest (gray line) models that use all 64 available features, as described above, for various decision thresholds. The red points correspond to rules based on the simple scoring procedure in Figure 3, using just age and prior FTA, again corresponding to various decision thresholds. For each rule, the horizontal axis shows the estimated proportion of defendants ROR'd under the rule, and the vertical axis shows the estimated proportion of defendants who would fail to appear at their court dates. The solid black dot shows the status quo: 69% of defendants RoR'd and a 13% FTA rate. Finally, the open circles show the observed RoR and FTA rates for each of the 23 judges in our data who have presided over at least 1,000 cases, sized in proportion to their case load.

The plot illustrates three key points. First, simple rules that consider only two features—age and prior FTAs—perform nearly identically to state-of-the-art machine learning models (random forest and lasso regression) that incorporate all 64 available features. Second, the statistically informed policies in the lower right quadrant all achieve higher rates of RoR and, simultaneously, lower rates of FTA than the status quo. In particular, by releasing defendants if and only if their risk score is below 3.5, we expect to release 79% of defendants while achieving an FTA rate of 13%. Relative to the existing policy, following this rule would result in detaining one-third fewer defendants while also slightly

[11] Although it is legally problematic to use race when *making* decisions, its use is acceptable—and indeed often required—when *evaluating* decisions. The model was fit with the `glmnet` package in `R`. The `cv.glmnet` method was used to determine the best value for the regularization parameter λ with 10-fold cross-validation and 1,000 values of λ. The model includes all pairwise interactions between the judge's decision and defendant's features. We opt for lasso instead of random forest for this prediction task because we empirically found lasso to yield better predictions in this case.

decreasing the overall FTA rate—from 13.3% to 13.0%. Finally, for nearly every judge, there is a statistical decision rule that simultaneously yields both a higher rate of release and a lower rate of FTA than the judge currently achieves. The statistical decision rules consistently outperform the human decision makers.

Why do these statistical decision rules outperform the experts? Figure 3 sheds light on this phenomenon. Each cell in the grid corresponds to defendants binned by their age and prior number of FTAs. Under a rule that releases defendants if and only if their risk score is below 3.5, one would release everyone to the left of the solid black line, and set bail for everyone to the right of the line. The number in each cell shows the proportion of defendants in each bin who were actually released, and the cell shading graphically indicates this proportion. Aside from the lowest risk defendants, who have no prior FTAs, the likelihood of being released does not correlate strongly with estimated flight risk. For example, the high risk group of young defendants with two or more prior FTAs is released at about the same rate as the low risk group of older defendants with one prior FTA. This low correlation between flight risk and release decision is in part attributable to extreme differences in release rates across judges, with some releasing more than 90% of defendants and others releasing just 50%.[12] Whereas defendants experience dramatically different outcomes based on the judge they happened to appear in front of, statistical decision rules improve efficiency in part by ensuring consistency.

4.3 | Sensitivity to unobserved heterogeneity

As noted above, our estimation strategy assumes that the judicial action taken is ignorable given the observed covariates. Under this ignorability assumption, one can accurately estimate the potential outcomes. Judges, however, might base their decisions in part on information that is not recorded in the data, which could in turn bias our estimates. For example, a judge, upon meeting a defendant, might surmise that his flight risk is higher than one would expect based on the recorded covariates alone, and may accordingly require the defendant to post bail. In this case, since our estimates are based only on the recorded data, we may underestimate the defendant's counterfactual likelihood of failing to appear if released.

We take two approaches to gauge the robustness of our results to such hidden heterogeneity. First, on each subset of cases handled by a single judge, we use response surface modeling to estimate V^{π}. Each judge has idiosyncratic criteria for releasing defendants, as evidenced by the dramatically different release rates across judges; accordingly, the types and proportion of cases for which the policy π coincides with the observed action differ from judge to judge. This variation allows us to assess the sensitivity of our estimates to the observed actions $\{a_i\}$. In particular, if unobserved heterogeneity were significant, we would expect our estimates to systematically vary depending on the proportion of observed judicial actions that agree with the policy π. Figure 6 shows the results of this analysis for the simple decision rule described in Figure 3, where each point corresponds to a judge. We find that the FTA rate of the decision rule is consistently estimated to be approximately 12–14%. Moreover, some judges act in concordance with the decision rule in nearly 80% of cases; for this subset of judges, where our estimates are largely based on directly observed outcomes, we again find FTA is estimated at around 12–14%.

As a second robustness check, we adapt the method of Rosenbaum and Rubin (1983a) for assessing the sensitivity of estimated causal effects to an unobserved binary covariate. We specifically tailor their approach to offline policy evaluation. At a high level, we assume there is an unobserved covariate $u \in \{0, 1\}$ that affects both a judge's decision (RoR or bail) and also the outcome conditional on that action. For example, u might indicate that a defendant is sympathetic, and sympathetic defendants may be more likely to be RoR'd and also more likely to appear at their court proceedings. Our key assumption is that a judge's action is ignorable given the observed covariates x and the unobserved

[12] Defendants are not perfectly randomly assigned to judges for arraignment, but in practice judges see a similar distribution of defendants.

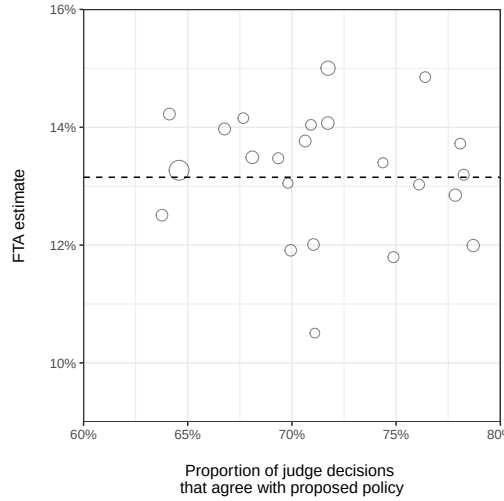

FIGURE 6 Robustness of estimated FTA rate for the simple decision rule. FTA rate is estimated by applying response surface modeling to each judge's cases, where each point corresponds to a judge; the dashed horizontal line indicates the FTA rate of the decision rule estimated on the full set of cases. Though judges have different criteria for releasing defendants—and the corresponding response models may thus differ—the FTA rate of the decision rule is consistently estimated to be approximately 12–14%.

covariate u:

$$(r(\text{RoR}), r(\text{bail})) \perp\!\!\!\perp a \,\big|\, x, u. \tag{1}$$

There are four key parameters in this framework: (1) the probability that $u = 1$; (2) the effect of u on the judge's decision; (3) the effect of u on the defendant's likelihood of FTA if RoR'd; and (4) the effect of u on the defendant's likelihood of FTA if bail is set. Our goal is to quantify the extent to which our estimate of V^{π} changes as a function of these parameters.

Without loss of generality, we can write

$$\Pr(a = \text{RoR}|u, x) = \text{logit}^{-1}\left(\gamma_x + u\alpha_x\right) \tag{2}$$

for appropriately chosen parameters γ_x and α_x that depend on the observed covariates x. We note that randomness in judicial decisions may arise from a multitude of factors, including idiosyncrasies in how judges are assigned to cases. Here α_x is the change in log-odds of being RoR'd when $u = 0$ versus when $u = 1$. For $t \in \{\text{RoR}, \text{bail}\}$, we can similarly write

$$\Pr(r(t)|u, x) = \text{logit}^{-1}\left(\beta_x^t + u\delta_x^t\right) \tag{3}$$

for parameters β_x^t and δ_x^t. In this case, δ_x^{RoR} is the change in log-odds of failing to appear if RoR'd when $u = 0$ versus when $u = 1$, and δ_x^{bail} is the corresponding change if bail is set.

Now, for any posited values of $\Pr(u = 1|x)$, α_x, δ_x^{RoR} and δ_x^{bail}, we use the observed data to estimate γ_x, β_x^{RoR} and

β_x^{bail}. We do this in three steps. First, by (2),

$$\Pr(a = \text{RoR}|x) = \Pr(u = 0|x) \cdot \text{logit}^{-1}(\gamma_x) + \Pr(u = 1|x) \cdot \text{logit}^{-1}(\gamma_x + \alpha_x).$$

The left-hand side of the equation can be estimated with a regression model fit to the data. For fixed values of $\Pr(u = 1|x)$ and α_x, the right-hand side is a continuous, increasing function of γ_x that takes on values from 0 to 1 as γ_x goes from $-\infty$ to $+\infty$. There is thus a unique value $\hat{\gamma}_x$ such that the right-hand side equals $\hat{\Pr}(a = \text{RoR}|x)$. Rosenbaum and Rubin (1983a) derive a simple closed form solution for $\hat{\gamma}_x$, facilitating fast computation on large datasets, which we omit for space.

Second, we use the fitted values of γ_x to estimate the distribution of u given the observed covariates and judicial action. By Bayes' rule,

$$\begin{aligned}\Pr(u = 1|a = t, x) &= \frac{\Pr(a = t|u = 1, x)\Pr(u = 1|x)}{\Pr(a = t|x)} \\ &= \frac{\Pr(a = t|u = 1, x)\Pr(u = 1|x)}{\Pr(a = t|u = 1, x)\Pr(u = 1|x) + \Pr(a = t|u = 0, x)\Pr(u = 0|x)}.\end{aligned}$$

With $\hat{\gamma}_x$, the $\Pr(a = t|u, x)$ terms on the right-hand side can be estimated from (2), and we can thus approximate the left-hand side.

Third, we have

$$\begin{aligned}\Pr(r(t) = 1|a = t, x) &= \Pr(u = 0|a = t, x)\Pr(r(t) = 1|a = t, x, u = 0) + \Pr(u = 1|a = t, x)\Pr(r(t) = 1|a = t, x, u = 1) \\ &= \Pr(u = 0|a = t, x)\Pr(r(t) = 1|x, u = 0) + \Pr(u = 1|a = t, x)\Pr(r(t) = 1|x, u = 1) \\ &= \Pr(u = 0|a = t, x) \cdot \text{logit}^{-1}\left(\beta_x^t\right) + \Pr(u = 1|a = t, x) \cdot \text{logit}^{-1}\left(\beta_x^t + \delta_x^t\right).\end{aligned}$$

The second equality above follows from the ignorability assumption stated in (1), and the third equality follows from (3). The left-hand side can be approximated by the quantity $\hat{r}_x(t)$ that we obtain via response surface modeling. Importantly, $\hat{r}_x(t)$ is a reasonable estimate of $\Pr(r(t) = 1|a = t, x)$ even though it may not be a good estimate of $r_x(t)$. This distinction is indeed the rationale of our sensitivity analysis. Given our above estimate of $\Pr(u = 1|a = t, x)$ and our assumed value of δ_x^t, the only unknown on the right-hand side is β_x^t. As before, there is a unique value $\hat{\beta}_x^t$ that satisfies the constraint.

With $\hat{\beta}_x^t$ in hand, we can now approximate the potential outcome for the action *not* taken:

$$\Pr(r(\bar{t}) = 1|a = t, x)$$

where $\bar{t} = \text{RoR}$ if $t = \text{bail}$, and vice versa. Specifically, we have

$$\hat{\Pr}(r(\bar{t}) = 1|a = t, x) = \hat{\Pr}(u = 0|a = t, x) \cdot \text{logit}^{-1}\left(\hat{\beta}_x^{\bar{t}}\right) + \hat{\Pr}(u = 1|a = t, x) \cdot \text{logit}^{-1}\left(\hat{\beta}_x^{\bar{t}} + \delta_x^{\bar{t}}\right). \tag{4}$$

Finally, the Rosenbaum and Rubin estimator adapted to policy evaluation is

$$\hat{V}_{\text{RR}}^{\pi} = \frac{1}{|\Omega|}\sum_i \left[r_i \mathbf{I}(\pi(x_i) = a_i) + \hat{r}_i(\bar{a}_i)\mathbf{I}(\pi(x_i) \neq a_i)\right],$$

where $\hat{r}_i(\bar{a}_i) = \hat{\Pr}(r(\bar{a}_i) = 1|a_i, x_i)$ is computed via (4).

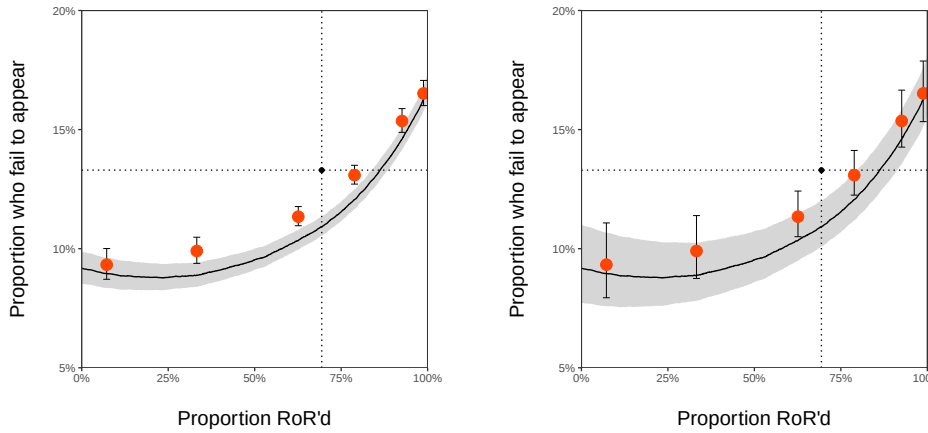

FIGURE 7 Sensitivity of FTA estimates to unobserved heterogeneity. The gray bands (for the complex rules using lasso) and the error bars (for the simple rules) indicate minimum and maximum FTA estimates for a variety of parameter settings. In the left-hand plot, we assume $\alpha = \log 2$ and consider all combinations of $p(u = 1) \in \{0.1, 0.2, \ldots, 0.9\}$, $\delta^{\text{RoR}} \in \{-\log 2, 0, \log 2\}$, and $\delta^{\text{bail}} \in \{-\log 2, 0, \log 2\}$, where all parameters are constant independent of x. In the right-hand plot, we consider a more extreme situation, with $\alpha = \log 3$, $\delta^{\text{RoR}} \in \{-\log 3, 0, \log 3\}$, and $\delta^{\text{bail}} \in \{-\log 3, 0, \log 3\}$. The results are relatively stable in these parameter regimes.

Figure 7 shows the results of computing $\hat{V}^{\pi}_{\text{RR}}$ on our data in two parameter regimes. In the first (left-hand plot), we assume $\alpha = \log 2$ and consider all combinations of $p(u = 1) \in \{0.1, 0.2, \ldots, 0.9\}$, $\delta^{\text{RoR}} \in \{-\log 2, 0, \log 2\}$, and $\delta^{\text{bail}} \in \{-\log 2, 0, \log 2\}$. All parameters are constant independent of x. We thus assume that holding the observed covariates fixed, a defendant with $u = 1$ has twice the odds of being RoR'd as one with $u = 0$, and that u can double or halve the odds a defendant fails to appear. For each complex policy (*i.e.*, one based on lasso), the gray bands show the minimum and maximum value of $\hat{V}^{\pi}_{\text{RR}}$ across all parameters in this set; the error bars on the red points show the analogous quantity for the simple rules. In the right-hand plot, we consider a more extreme situation, with $\alpha = \log 3$, $\delta^{\text{RoR}} \in \{-\log 3, 0, \log 3\}$, and $\delta^{\text{bail}} \in \{-\log 3, 0, \log 3\}$. We find that our estimates are relatively stable in these parameter regimes. In the first case ($\alpha = \log 2$) the estimated FTA rate for a given policy typically varies by only half a percentage point. Even in the more extreme setting ($\alpha = \log 3$), policies are typically stable to about one percentage point. It thus seems our conclusions are robust to potentially unobserved heterogeneity across defendants.

5 | THE ROBUSTNESS OF BINARY CLASSIFICATION

Why is it that simple rules often perform as well as the most sophisticated statistical methods? In part, it is because binary classification accuracy is relatively robust to error in the underlying predictive model, an observation that we formalize in Proposition 1 below.

To establish this result, we start by considering the prediction scores generated via a standard statistical method—such as logistic regression trained on the full set of available features—which we call the "true" scores. As in linear discriminant analysis, we assume that the true scores for positive and negative instances are normally distributed with equal variance: $\text{N}(\mu_p, \sigma^2)$ and $\text{N}(\mu_n, \sigma^2)$, respectively. The homoscedasticity assumption guarantees the Bayes optimal classifier is a threshold rule on the scores. For scores estimated via logistic regression, the normality assumption

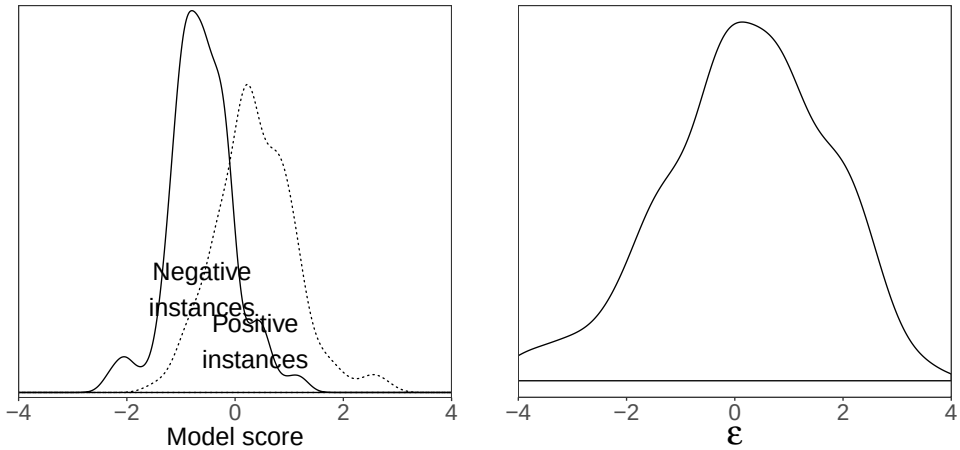

FIGURE 8 Empirical estimation of noise added by simple rules. Left panel: empirical distribution of prediction scores, on the logit scale, for positive and negative instances of one UCI dataset (`heart-cleveland`), generated via an L^1-regularized logistic regression model. Right panel: empirical distribution of ϵ for select-regress-and-round applied to the same dataset.

is reasonable if we consider the scores on the logit scale rather than on the probability scale. Figure 8 (left panel) shows such scores for one of the UCI datasets, `heart-cleveland`. We further assume that the process of generating simple rules—both limiting the number of features and also restricting the possible values of the weights—can be viewed as adding normal, mean-zero noise $\mathrm{N}(0, \sigma_\epsilon^2)$ to the true scores; Figure 8 (right panel) plots the distribution of this noise for the same `heart-cleveland` dataset considered in the left panel.[13] Thus, with simple rules, instead of making classification decisions based on the true scores, we assume decisions are made in terms of a noisy approximation. Under this analytic framework, Proposition 1 shows that the drop in classification performance (as measured by AUC) can be expressed in terms of the "true AUC" (*i.e.*, the AUC under the true scores) and $\gamma = \sigma_\epsilon^2/\sigma^2$, the ratio of the noise to the within-class variance of the true scores. In particular, we find that when the magnitude of the noise is on par with (or smaller than) the score variance (*i.e.*, $\gamma \lesssim 1$), then the AUC of the noisy approximation is comparable to the true AUC.

Proposition 1 *For a binary classification task, let Y be a continuous random variable that denotes the prediction score of a random instance, and let Y_p and Y_n denote the conditional distributions of Y for positive and negative instances, respectively. Suppose $Y_p \sim N(\mu_p, \sigma^2)$ and $Y_n \sim N(\mu_n, \sigma^2)$. Then, for $\epsilon \sim N(0, \sigma_\epsilon^2)$ and $\hat{Y} = Y + \epsilon$,*

$$\mathrm{AUC}_{\hat{Y}} = \Phi\left(\frac{\Phi^{-1}(\mathrm{AUC}_Y)}{\sqrt{1+\gamma}}\right), \tag{5}$$

where $\gamma = \sigma_\epsilon^2/\sigma^2$, and Φ is the CDF for the standard normal.

Proof In general, AUC is equal to the probability that a randomly selected positive instance has a higher prediction score than a randomly selected negative instance, and so $\mathrm{AUC}_Y = \Pr(Y_p - Y_n > 0)$ (Su and Liu, 1993). Since $Y_p - Y_n$ is

[13] We estimate the noise distribution by taking the difference between the simple and true scores. Before taking the difference, we convert the simple scores to the scale of true scores by dividing the simple scores by M, the scaling factor used when generating the rule.

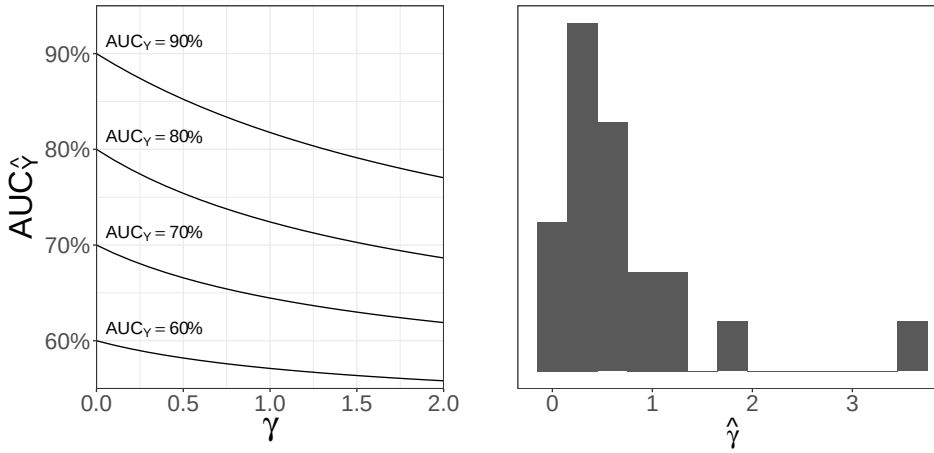

FIGURE 9 Theoretical analysis of simple rules. Left panel: the theoretical change in AUC, as a function of γ. Right panel: distribution of $\hat{\gamma}$, estimated across all simple rules for 21 datasets with $k = 5$ and $M = 3$.

normally distributed with mean $\mu_p - \mu_n$ and variance $2\sigma^2$,

$$\frac{Y_p - Y_n - (\mu_p - \mu_n)}{\sqrt{2}\sigma} \sim \mathrm{N}(0, 1).$$

Hence,

$$\begin{aligned}\mathrm{AUC}_Y &= \Pr\left(\frac{Y_p - Y_n - (\mu_p - \mu_n)}{\sqrt{2}\sigma} > -\frac{\mu_p - \mu_n}{\sqrt{2}\sigma}\right)\\ &= \Phi\left(\frac{\mu_p - \mu_n}{\sqrt{2}\sigma}\right),\end{aligned}$$

where the last equality follows from symmetry of the normal distribution.

Now define $\hat{Y}_p = Y_p + \epsilon$, so $\hat{Y}_p \sim \mathrm{N}(\mu_p, \sigma^2 + \sigma_\epsilon^2)$, with $\hat{Y}_n$ defined similarly. A short computation shows that

$$\begin{aligned}\mathrm{AUC}_{\hat{Y}} &= \Pr(\hat{Y}_p > \hat{Y}_n)\\ &= \Phi\left(\frac{\mu_p - \mu_n}{\sqrt{2\sigma^2 + 2\sigma_\epsilon^2}}\right)\\ &= \Phi\left(\frac{\Phi^{-1}(\mathrm{AUC}_Y)}{\sqrt{1+\gamma}}\right).\end{aligned}$$

Proposition 1 establishes a direct theoretical link between performance and noise in model specification. To give a better sense of how the analytic expression for $\mathrm{AUC}_{\hat{Y}}$ varies with AUC_Y and γ, Figure 9 (left panel) shows this expression for various parameter values. For example, the figure shows that for $\mathrm{AUC}_Y = 90\%$ and $\gamma = 0.5$, we have $\mathrm{AUC}_{\hat{Y}} = 85\%$. That is, if the amount of noise is equal to half the within-class variance of the true scores, then the drop in performance is relatively small.

While connecting model performance to model noise, Proposition 1 leaves unanswered how much noise simple

rules add to the underlying scores. This question seems difficult to answer theoretically. We can, however, empirically estimate how much noise simple rules add in the datasets we analyze.[14] The right panel of Figure 9 shows the distribution of $\hat{\gamma}$ across the 21 UCI datasets we consider, when using rules with five features and a coefficient range of -3 to 3, with an average value of $\hat{\gamma} = 0.22$. This low empirically observed noise is in line with our finding that such simple rules perform well on these datasets.[15]

6 | CONCLUSION

Our work extends past research on improper linear models by formalizing and evaluating a simple method for constructing simple rules—rules that experts can apply mentally to guide classification decisions. These simple rules take the form of a short checklist whose factors have small integer weights. In 22 domains of varying size and complexity, the rules produced by the select-regress-and-round method rivaled the accuracy of regularized logistic regression models, while using only a fraction of the information. In a detailed analysis of pretrial release decisions, the simple rules outperformed human judges and matched machine learning models that incorporated 64 features. (In Appendix C, we provide another detailed demonstration of select-regress-and-round to assess credit risk, and reach similar conclusions.)

Although our focus has been on the comparison between simple, statistically informed decision rules and more complex machine learning methods, our results are also in accordance with an extensive literature comparing predictions by human experts to those based on statistical models. Over 60 years ago, Meehl contrasted *clinical* methods for predicting behavior, which rely on professional judgment, with *actuarial* methods, which rely on statistically derived patterns in data (Meehl, 1954). Subsequently, large meta-analyses have consistently demonstrated that actuarial methods outperform clinical approaches, including in the context of predicting criminal activity (Ægisdóttir et al., 2006), and even for judgments by the most experienced professionals. These results hold in the judicial context as well, where clinical assessments of risk by judges are generally worse at predicting recidivism than actuarial formulas (Gottfredson, 1999). Our analysis of a large dataset of judicial decisions provides further evidence that simple statistical models can outperform experts in a high-stakes domain.

Statistically informed rules, and simple checklists in particular, may result in improved accuracy and consistency compared to unaided human decisions, but a number of open questions remain. First, in many contexts, allowing human overrides of algorithmic decision aids may be legally mandated, but such overrides can reduce accuracy (Krauss, 2004). In the criminal justice setting, past work indeed suggests that judges may not apply the recommendations of risk assessments in a consistent manner (Christin, 2017; DeMichele et al., 2018). It is important to strike an appropriate balance, allowing for human overrides in exceptional instances while not degrading overall performance. Second, in contexts where an outcome variable has a nonlinear relationship with a set of predictors, the simple rules produced by select-regress-and-round may not be flexible enough to make useful predictions (e.g., in Figure B3, all linear model show poor performance on the `chess-krvk` dataset). One solution may be to allow additional model flexibility in select-regress-and-round, though that approach could be at odds with the goals of transparency and interpretability. Finally, it is unclear how well simple rules would work in domains with little training data, but we also note that prediction tasks using small sample sizes remain challenging for more complex methods.

Our results complement a growing body of work in statistics and computer science on interpretable machine learning, in which sophisticated algorithms are used to create simple scoring systems and rule sets (Ustun and Rudin,

[14] To estimate $\gamma = \sigma_\epsilon^2/\sigma^2$ for a specific simple rule on a given dataset, we first compute the average within-class variance of the true scores, where these scores are generated via an L^1-regularized logistic regression model. We estimate σ_ϵ^2 by taking the variance of the noise, as described in Footnote 13.

[15] In Appendix B.1, we further test the empirical robustness of probabilistic risk predictions, in addition to binary classification, using simple rules. We find that probabilistic estimates from our simple rules are comparable to those from more complex statistical models.

2016; Wang and Rudin, 2015; Lakkaraju et al., 2016). While many of these rule construction methods offer great flexibility, they in turn require considerable computational resources and expertise to carry out. In contrast, the method we propose can easily be carried out by ordinary practitioners using popular open-source software. It has long been noted that statistical models tend to outperform unaided human judgment (Einhorn and Hogarth, 1975; Green, 1977; Dawes, 1979; Gigerenzer and Goldstein, 1996; Waller and Jones, 2011). We hope that providing practitioners with models that are both easy to apply and easy to construct will increase their adoption and, ultimately, the quality of decisions.

ACKNOWLEDGEMENTS

We thank Avi Feller, Andrew Gelman, Gerd Gigerenzer, Art Owen, and Berk Ustun for helpful conversations. Code to replicate our results is publicly available at: https://github.com/stanford-policylab/simple-rules.

REFERENCES

Ægisdóttir, S., White, M. J., Spengler, P. M., Maugherman, A. S., Anderson, L. A., Cook, R. S., Nichols, C. N., Lampropoulos, G. K., Walker, B. S., Cohen, G. et al. (2006) The meta-analysis of clinical judgment project: Fifty-six years of accumulated research on clinical versus statistical prediction. *The Counseling Psychologist*, **34**, 341–382.

Åstebro, T. and Elhedhli, S. (2006) The effectiveness of simple decision heuristics: Forecasting commercial success for early-stage ventures. *Management Science*, **52**, 395–409.

Bareiss, E. R., Porter, B. W. and Wier, C. C. (1988) Protos: an exemplar-based learning apprentice. *International Journal of Man-Machine Studies*, **29**, 549 – 561. URL: `http://www.sciencedirect.com/science/article/pii/S0020737388800129`.

Bohanec, M. and Rajkovic, V. (1988) Knowledge acquisition and explanation for multi-attribute decision making. In *8th Intl Workshop on Expert Systems and their Applications*, 59–78.

Camerer, C. F. and Johnson, E. J. (1997) The process-performance paradox in expert judgment. In *Research on judgment and decision making: Currents, connections, and controversies* (eds. W. M. Goldstein and R. M. Hogarth). Cambridge University Press.

Cassel, C. M., Särndal, C. E. and Wretman, J. H. (1976) Some results on generalized difference estimation and generalized regression estimation for finite populations. *Biometrika*, **63**, 615–620.

Chambers, J. M., Hastie, T. J. et al. (1992) *Statistical models in S*, vol. 251. Wadsworth & Brooks/Cole Advanced Books & Software Pacific Grove, CA.

Christin, A. (2017) Algorithms in practice: Comparing web journalism and criminal justice. *Big Data & Society*, **4**, 2053951717718855.

Corbett-Davies, S., Pierson, E., Feller, A., Goel, S. and Huq, A. (2017) Algorithmic decision making and the cost of fairness. In *Proceedings of the 23rd ACM SIGKDD International Conference on Knowledge Discovery and Data Mining*, 797–806. ACM.

Danziger, S., Levav, J. and Avnaim-Pesso, L. (2011) Extraneous factors in judicial decisions. *Proceedings of the National Academy of Sciences*, **108**, 6889–6892.

Dawes, R. M. (1979) The robust beauty of improper linear models in decision making. *American Psychologist*, **34**, 571.

Dawes, R. M., Faust, D. and Meehl, P. E. (1989) Clinical versus actuarial judgment. *Science*, **243**, 1668–1674.

DeMichele, M., Baumgartner, P., Barrick, K., Comfort, M., Scaggs, S. and Misra, S. (2018) What do criminal justice professionals think about risk assessment at pretrial? URL: `https://ssrn.com/abstract=3168490`.

Dhami, M. K. (2003) Psychological models of professional decision making. *Psychological Science*, **14**, 175–180.

Dudík, M., Langford, J. and Li, L. (2011) Doubly robust policy evaluation and learning. *ICML*. URL: `http://arxiv.org/abs/1103.4601`.

Einhorn, H. J. and Hogarth, R. M. (1975) Unit weighting schemes for decision making. *Organizational Behavior and Human Performance*, **13**, 171–192.

Fernández-Delgado, M., Cernadas, E., Barro, S. and Amorim, D. (2014) Do we need hundreds of classifiers to solve real world classification problems? *J. Mach. Learn. Res*, **15**, 3133–3181.

Friedman, J., Hastie, T. and Tibshirani, R. (2010) Regularization paths for generalized linear models via coordinate descent. *Journal of statistical software*, **33**, 1.

Gelman, A. and Park, D. K. (2009) Splitting a predictor at the upper quarter or third and the lower quarter or third. *The American Statistician*, **63**, 1–8.

Gigerenzer, G. and Goldstein, D. G. (1996) Reasoning the fast and frugal way: models of bounded rationality. *Psychological review*, **103**, 650.

Gleicher, M. (2016) A framework for considering comprehensibility in modeling. *Big Data*, **4**, 75–88.

Goodman, B. and Flaxman, S. (2016) EU regulations on algorithmic decision-making and a "right to explanation". In *ICML workshop on human interpretability in machine learning (WHI 2016), New York, NY. http://arxiv. org/abs/1606.08813 v1*.

Gottfredson, D. M. (1999) *Effects of judges' sentencing decisions on criminal careers*. US Department of Justice, Office of Justice Programs, National Institute of Justice.

Green, B. F. (1977) Parameter sensitivity in multivariate methods. *Multivariate Behavioral Research*, **12**, 263–287.

Hill, J. L. (2012) Bayesian nonparametric modeling for causal inference. *Journal of Computational and Graphical Statistics*.

Kang, J. D. and Schafer, J. L. (2007) Demystifying double robustness: A comparison of alternative strategies for estimating a population mean from incomplete data. *Statistical science*, 523–539.

Kim, B., Shah, J. A. and Doshi-Velez, F. (2015) Mind the gap: A generative approach to interpretable feature selection and extraction. In *Advances in Neural Information Processing Systems*, 2260–2268.

Klein, G. A. (2017) *Sources of power: How people make decisions*. MIT press.

Kleinberg, J., Lakkaraju, H., Leskovec, J., Ludwig, J. and Mullainathan, S. (2017) Human decisions and machine predictions. *The Quarterly Journal of Economics*, **133**, 237–293.

Kleinberg, J., Ludwig, J., Mullainathan, S. and Obermeyer, Z. (2015) Prediction policy problems. *The American Economic Review*, **105**.

Krauss, D. A. (2004) Adjusting risk of recidivism: Do judicial departures worsen or improve recidivism prediction under the federal sentencing guidelines? *Behavioral sciences & the law*, **22**, 731–750.

Lakkaraju, H., Bach, S. H. and Leskovec, J. (2016) Interpretable decision sets: A joint framework for description and prediction. In *Proceedings of the 22nd International Conference on Knowledge Discovery and Data Mining*.

Marewski, J. N. and Gigerenzer, G. (2012) Heuristic decision making in medicine. *Dialogues Clin Neurosci*, **14**, 77–89.

McDonald, C. J. (1996) Medical heuristics: The silent adjudicators of clinical practice. *Annals of Internal Medicine*, **124**, 56–62.

Meehl, P. E. (1954) *Clinical vs. statistical prediction*. Minneapolis: University of Minnesota Press.

Ribeiro, M. T., Singh, S. and Guestrin, C. (2016) "Why Should I Trust You?": Explaining the Predictions of Any Classifier. *CoRR*, 21–27.

Robins, J. M. and Rotnitzky, A. (1995) Semiparametric efficiency in multivariate regression models with missing data. *Journal of the American Statistical Association*, **90**, 122–129.

Robins, J. M., Rotnitzky, A. and Zhao, L. P. (1994) Estimation of regression coefficients when some regressors are not always observed. *Journal of the American statistical Association*, **89**, 846–866.

Rosenbaum, P. R. and Rubin, D. B. (1983a) Assessing sensitivity to an unobserved binary covariate in an observational study with binary outcome. *Journal of the Royal Statistical Society. Series B (Methodological)*, 212–218.

— (1983b) The central role of the propensity score in observational studies for causal effects. *Biometrika*, **70**, 41–55.

— (1984) Reducing bias in observational studies using subclassification on the propensity score. *Journal of the American statistical Association*, **79**, 516–524.

Su, J. Q. and Liu, J. S. (1993) Linear combinations of multiple diagnostic markers. *Journal of the American Statistical Association*, **88**, 1350–1355.

Sull, D. and Eisenhardt, K. M. (2015) *Simple rules: How to thrive in a complex world*. Houghton Mifflin Harcourt.

Tetlock, P. (2005) *Expert political judgment: How good is it? How can we know?* Princeton University Press.

Ustun, B. and Rudin, C. (2016) Supersparse linear integer models for optimized medical scoring systems. *Machine Learning*, **102**, 349–391.

Van Belle, V. M., Van Calster, B., Timmerman, D., Bourne, T., Bottomley, C., Valentin, L., Neven, P., Van Huffel, S., Suykens, J. A. and Boyd, S. (2012) A mathematical model for interpretable clinical decision support with applications in gynecology. *PloS one*, **7**, e34312.

Waller, N. and Jones, J. (2011) Investigating the performance of alternate regression weights by studying all possible criteria in regression models with a fixed set of predictors. *Psychometrika*, **76**, 410.

Wang, F. and Rudin, C. (2015) Falling rule lists. In *Artificial Intelligence and Statistics*, 1013–1022.

Wübben, M. and Wangenheim, F. V. (2008) Instant customer base analysis: Managerial heuristics often get it right. *Journal of Marketing*, **72**, 82–93.

Wyatt, J. C. and Altman, D. G. (1995) Commentary: Prognostic models: clinically useful or quickly forgotten? *BMJ*, **311**, 1539–1541.

Appendix A | DESCRIPTION OF UCI DATASETS

Here we provide a short description of the classification task associated with each of the 21 UCI datasets that we consider in Section 3.

- `adult`: Predict whether income exceeds $50K/yr based on census data. Also known as "Census Income" dataset.
- `annealing`: Classify steel types based on various annealing properties.
- `audiology-std`: Standardized version of the original audiology database first presented in Bareiss et al. (1988).
- `bank`: The data is related with direct marketing campaigns (phone calls) of a Portuguese banking institution. The classification goal is to predict if the client will subscribe a term deposit.
- `bankruptcy`: Predict bankruptcy based on qualitative parameters measured by experts.
- `car`: Determine whether a car is "acceptable" or not, based on quantitative attributes. Originally presented in Bohanec and Rajkovic (1988).
- `chess-krvk`: Chess end-game data for white king and rook against black king (KRK). Classification task is to determine whether white can win or not.
- `chess-krvkp`: Chess end-game data for king and rook versus king and pawn on `A7` (usually abbreviated `KRKPA7`). The pawn on `A7` means it is one square away from queening. It is the king and rook's turn (white) to move. The goal is to classify whether white can win or not.
- `congress-voting`: 1984 United Stated congressional voting records. The task is to classify votes as republican or democrat.
- `contrac`: A subset of the 1987 National Indonesia Contraceptive Prevalence Survey. The samples are married women who were either not pregnant or do not know if they were at the time of interview. The problem is to predict the current contraceptive method choice (no use, long-term methods, or short-term methods) of a woman based on her demographic and socio-economic characteristics.
- `credit-approval`: A collection of credit card applications. The task is to determine whether the application was approved or not.
- `ctg`: Measurements of fetal heart rate (FHR) and uterine contraction (UC) features on cardiotocograms classified by expert obstetricians. The task is to classify the fetal state as normal, suspect, or pathologic.
- `cylinder-bands`: Predict process delays known as "cylinder bands" in rotogravure printing.
- `dermatology`: The aim of this dataset is to determine the type of Eryhemato-Squamous Disease.
- `german_credit`: This dataset classifies people described by a set of attributes as good or bad credit risks.
- `heart-cleveland`: The goal is to determine the presence of heart disease in the patients. The outcome is integer valued from 0 (no presence) to 4. Experiments with the Cleveland database have concentrated on simply attempting to distinguish presence (values 1,2,3,4) from absence (value 0).
- `ilpd`: This dataset contains 416 liver patient records and 167 non liver patient records. The data were collected from the northeast of Andhra Pradesh, India.
- `mammo`: Discrimination of benign and malignant mammographic masses based on BI-RADS attributes and the patient's age.
- `mushroom`: From the Audobon Society Field Guide. Mushrooms are described in terms of physical characteristics. The task is to classify them as either poisonous or edible.
- `wine`: Using chemical analysis, determine the origin of wines.
- `wine_qual`: Two datasets are included, related to red and white vinho verde wine samples, from the north of Portugal. The goal is to model wine quality based on physicochemical tests.

Appendix B | ADDITIONAL RESULTS FOR UCI DATA

B.1 | Robustness of probability estimates with simple rules

To gauge the robustness of probability estimates derived from simple rules, we compare the mean absolute deviation of those estimates to the predictions from a lasso model that uses all available features. For each integer score produced by a select-regress-and-round model, we compute the corresponding probability estimate for the simple rule by considering all cases in the training set with that score, and then computing the empirical frequency of the outcome of interest, as detailed in Section 4.1. As shown in Figure B1, using five features and rounding coefficients to the interval $[-3, 3]$, probability estimates using select-regress-and-round deviate from the lasso predictions by about 6 percentage points on average across the UCI datasets.

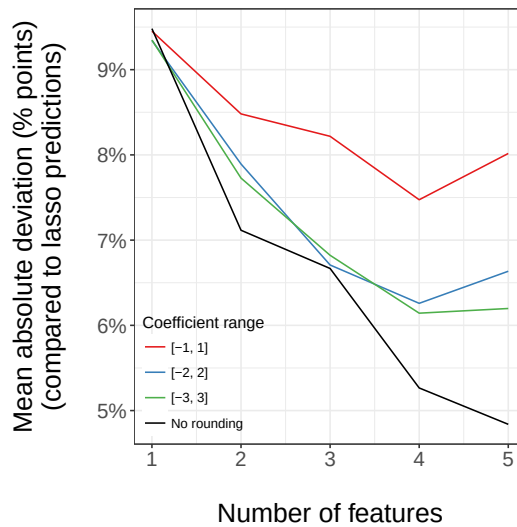

FIGURE B1 Mean absolute deviation of probability estimates based on simple rules compared to those from a complex lasso model, averaged over all the UCI datasets.

B.2 | Detailed results for individual datasets

Here, we disaggregate the results in Section 3.1 to compare simple rules and complex models on each of the 21 UCI datasets. First, Figure B2 compares the performance of select-regress-and-round using up to 5 features and rounding to the nearest integer in the range [-3, 3] (i.e., $k = 5$, $M = 3$) against each of the three benchmark models for each individual dataset. In Figure B2, each point represents a dataset, and the corresponding horizontal and vertical position shows the cross-validated AUC of the complex models and simple rules, respectively. For the logistic regression and lasso comparisons, all points are very close to the diagonal, indicating that select-regress-and-round performs on par with these complex models for each individual dataset. On the other hand, we see that a random forest model outperforms simple rules in many situations. Next, Figure B3 provides a more detailed comparison by replicating Figure 1 for each individual dataset. We observe that in general, model comparisons on individual datasets are similar to those that average over all datasets. Finally, in Figure B4 we similarly plot performance, but replace the horizontal axis with

the number of non-zero coefficients instead of features. A model can have more non-zero coefficients than features, because a categorical variable with more than two categories will yield more than two non-zero coefficients after each category has been binarized. For example, the rule presented in Figure 3 has two features: age and prior FTAs, but five non-zero coefficients.

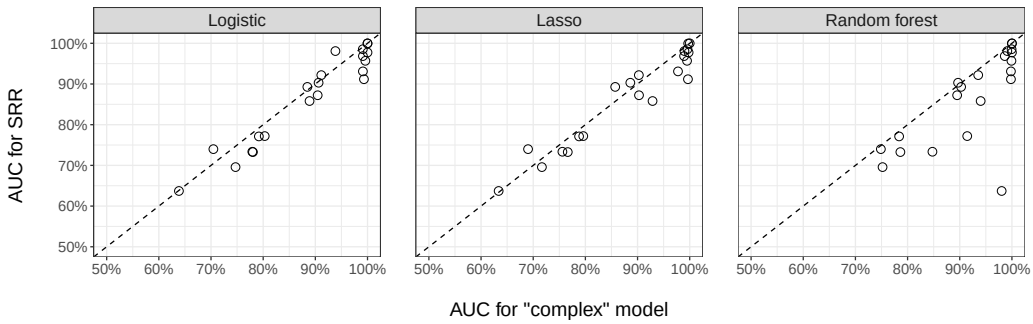

FIGURE B2 Performance comparison across each of the 21 UCI datasets, for simple rules using up to 5 features, and rounding to the nearest integer within the range [-3, 3] (i.e., $k = 5$, $M = 3$). While random forest generally outperforms select-regress-and-round, simple rules are comparable in performance with logistic and lasso regression across most datasets.

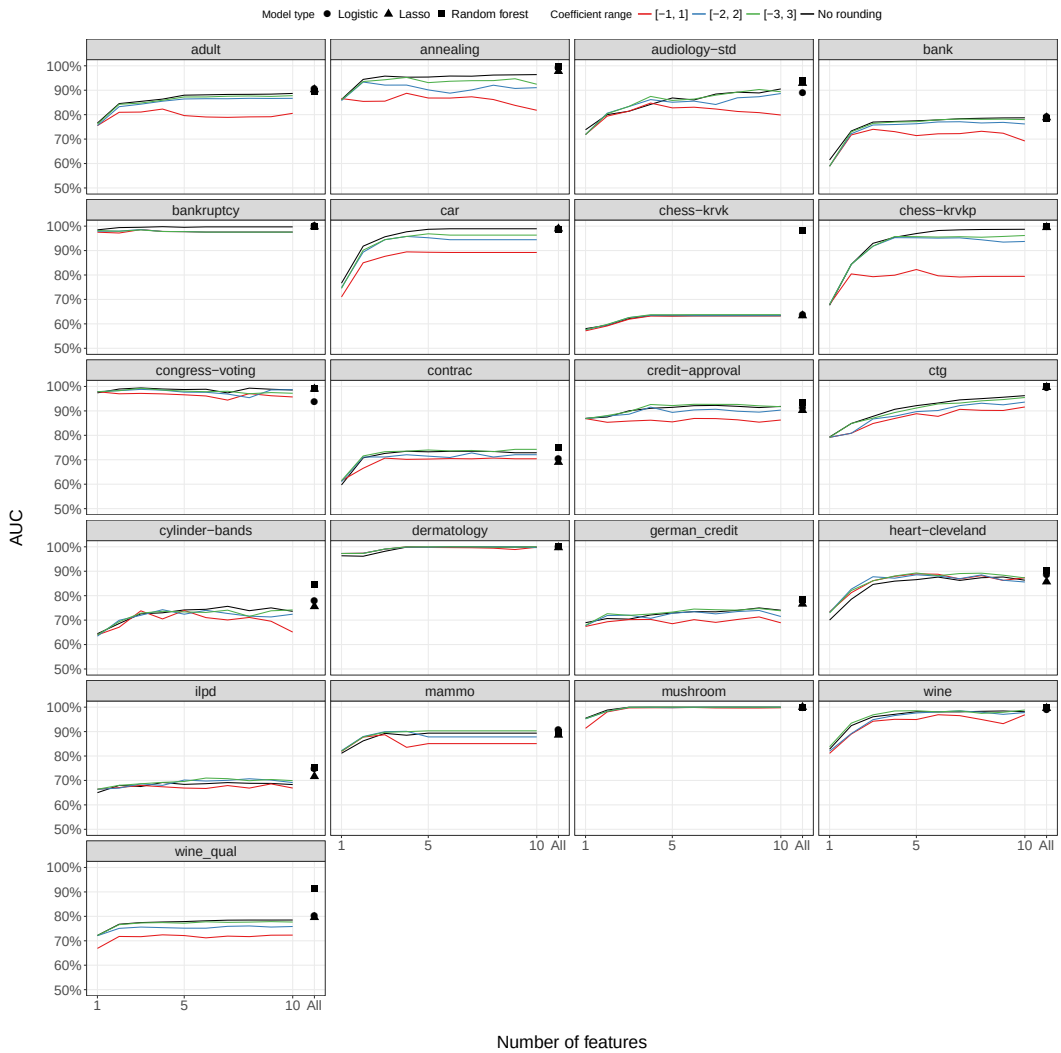

FIGURE B3 Performance comparison of AUC for each of the 21 UCI datasets, for simple rules with different k and M parameters.

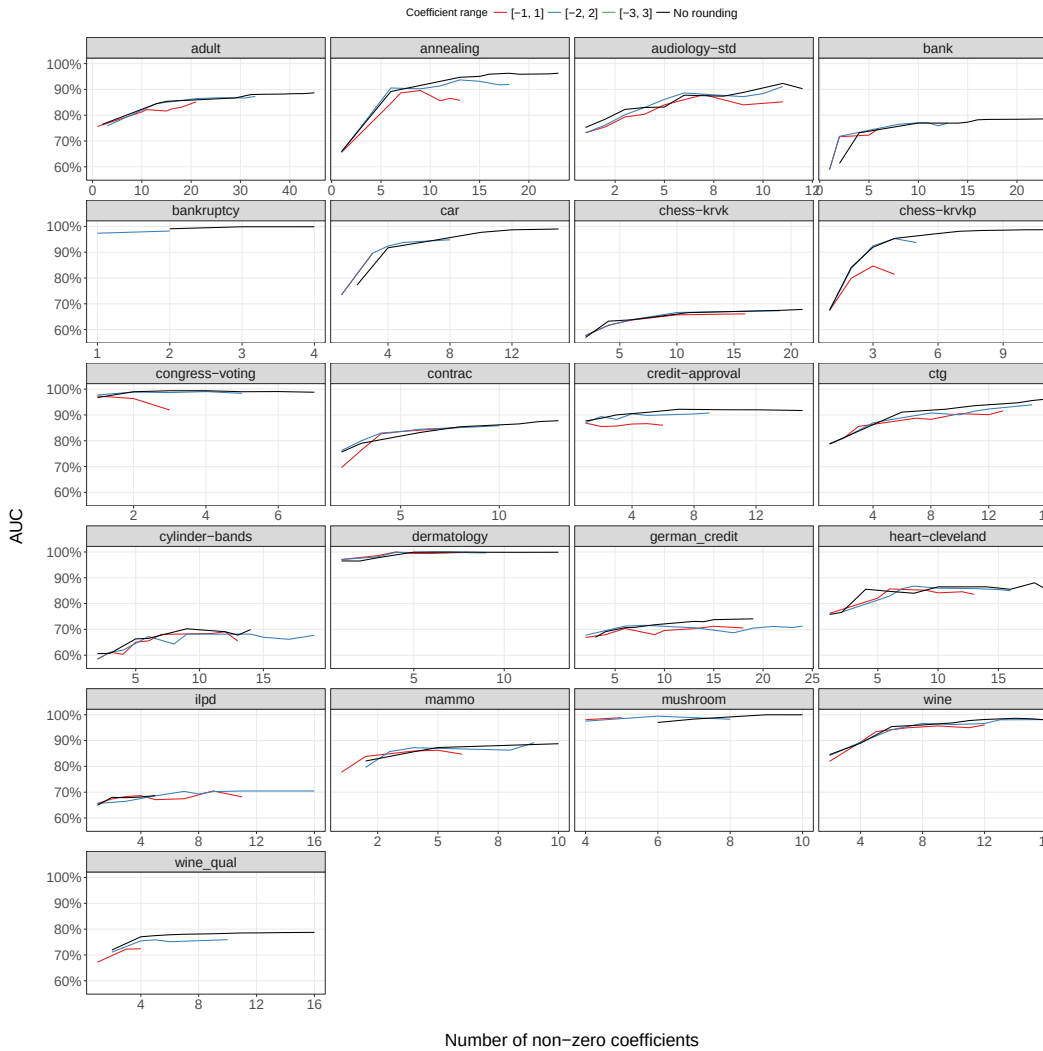

FIGURE B4 Performance comparison of AUC for each of the 21 UCI datasets, for simple rules with different k and M parameters. Compared to Figure B3, this figure shows the number of binarized features with non-zero coefficients on the horizontal axis.

Appendix C | ALTERNATIVE CASE STUDY WITH `GERMAN_CREDIT` DATA

Here, we illustrate select-regress-and-round on the `german_credit` data. This dataset consists of 1,000 cases labeled as having either good or bad credit. Each row is described by 20 features; 7 continuous and 13 categorical.[16] The goal is to estimate the risk of default (labeled as "bad" credit) for each case. We split the data randomly into 900 cases for training and 100 cases for evaluation.[17]

As a benchmark, we first fit a random forest model with 1,000 trees on the training data using all 20 available features. The result is a complex model that achieves 0.80 AUC on the test set. A full ROC curve for the complex model is presented as a blue line in Figure C6. According to the steps presented in Section 2, we build a simple rule to score the risk of default for a given case. Note that as described previously, we discretize continuous features into three bins of approximately equal sizes to prioritize simplicity. This is achieved by discretizing each continuous feature at the 33rd and 67th percentile in the training data, and applying the same cut-offs to the test data. For example, Figure C5 shows the distribution of the `duration in months` feature in the training data, which represents the months an applicant has lived at their current address. The 33rd and 67th percentile of this feature are 12 and 24, respectively, hence the feature is discretized at these points for both the train and test data. In detail, we perform the following steps:

1. **Select.** From the full set of 20 features, we select $k = 5$ features via forward stepwise selection. The features that are selected, in order, are: `checking account status`, `months lived at current address`, `credit history`, `savings account/bonds`, and `guarantors`.
2. **Regress.** Using the five selected features, we train an L^1-regularized (lasso) logistic regression model to predict whether the credit is good (0) or bad (1) for each case. The regularization parameter, λ, is chosen via 10-fold cross-validation. Following Friedman et al. (2010), we explore a regularization path with 1,000 values of λ spaced evenly on log-scale in the range $(\lambda_{\text{min}}, \lambda_{\text{max}})$, where $\lambda_{\text{min}} = 10^{-4}$ and λ_{max} is set to 0.141, the minimum value such that all coefficients are regularized to zero. We find that $\lambda^\star$, the value of λ that maximizes cross-validated performance is about 0.004. The second column of Table C1 shows the fitted lasso model coefficients.
3. **Round.** We rescale the coefficients of the model from Step 2 to be in the range $[-3, 3]$ (e.g., $M = 3$), and then round the rescaled coefficients to the nearest integer. The final scores corresponding to each variable are listed in the third column of Table C1.

Figure C6 shows the comparison of ROC curve performance on the held-out test set, between the random forest model (blue) and our simple rule in Table C1 (red). In practice, a decision maker would typically select a threshold based on various costs to determine which loan applications to approve or reject. However, our results demonstrate that across all threshold values, a simple, transparent rule achieves nearly identical performance when compared with a considerably more complex model.

This case study and all results can be replicated by running the `case_study.R` script provided in our public code repository: https://github.com/stanford-policylab/simple-rules. In addition, we have made it easy to generate simple rules for any combination of parameters for each of the 21 UCI datasets, by providing an `R` markdown file that can be run using freely available software.

[16] A full description of the dataset can be found at the UCI repository: https://archive.ics.uci.edu/ml/datasets/Statlog+%28German+Credit+Data%29

[17] We use a single 9:1 split for simplicity here, but note that our main results reported in Figure 1 were obtained via 10-fold cross validation.

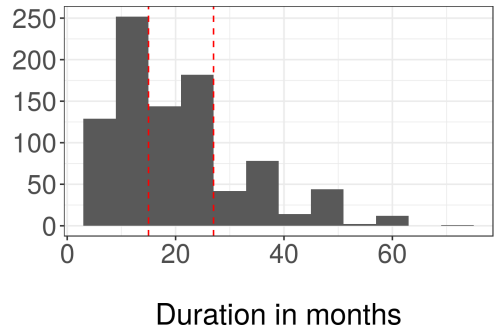

FIGURE C5 Distribution of the continuous feature `duration in months` in the training data. Red dashed lines indicate values at the 33rd and 67th percentiles, which are used to discretize the continuous feature into three bins of approximately equal size.

TABLE C1 Simple rule for determining risk of default, derived using select-regress-and-round on the `german_credit` dataset with $k = 5$ and $M = 3$. Variables with a zero Lasso coefficient have been omitted.

Selected feature	Lasso coefficient	SRR score
Checking account status		
Less than 200 DM	0.3	1
200 DM or above	0.92	2
No checking account	1.56	3
Months lived at current address		
Between 12 and 24 months	-0.39	-1
24 months or more	-1.02	-3
Credit history		
All credits at this bank paid back	-0.36	-1
Delayed payments in the past	0.56	1
Unpaid credits existing (not at this bank)	1.23	3
Savings account/bonds		
$100 \leq$ value < 500 DM	0.1	0
$500 \leq$ value < 1000 DM	0.67	2
1000 DM $\leq$ value	1.002	2
No known savings account	0.99	2
Guarantors		
Co-applicant	-0.34	-1
Guarantor	1.35	3

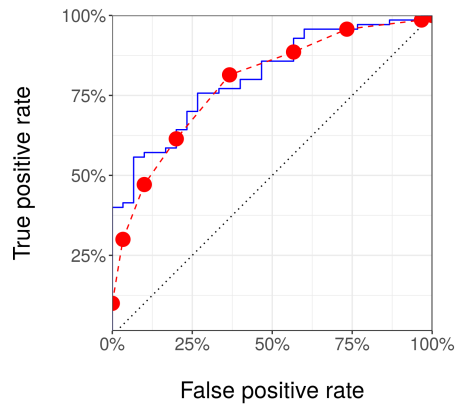

FIGURE C6 ROC curve comparing performance of a complex (random forest, in blue) model that uses all 20 available features to predict whether an individual has "good" or "bad" credit, versus the simple rule derived using SRR, as shown in Table C1 (red). The ROC curve for simple rules is shown as points, since a simple rule results in discrete cut-offs. The complex model and simple rule achieve AUCs of 0.80 and 0.78, respectively. Note that while a decision maker would typically select a threshold based on various costs given a risk score, our results show that a simple rule achieves almost identical performance compared to a complex model for all possible threshold values, with the additional benefit of being transparent and interpretable.