

RESEARCH ARTICLE 

Model Averaging and Double Machine Learning

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ABSTRACT

This paper discusses pairing double/debiased machine learning (DDML) with *stacking*, a model averaging method for combining multiple candidate learners, to estimate structural parameters. In addition to conventional stacking, we consider two stacking variants available for DDML: *Short-stacking* exploits the cross-fitting step of DDML to substantially reduce the computational burden, and *pooled stacking* enforces common stacking weights over cross-fitting folds. Using calibrated simulation studies and two applications estimating gender gaps in citations and wages, we show that DDML with stacking is more robust to partially unknown functional forms than common alternative approaches based on single pre-selected learners. We provide Stata and R software implementing our proposals.

JEL Classification: C21, C26, C52, C55, J01, J08

1 | Introduction

Motivated by their robustness to partially unknown functional forms, supervised machine learning estimators are increasingly leveraged for causal inference. For example, lasso-based approaches such as the post-double-selection lasso (PDS lasso) of Belloni, Chernozhukov, and Hansen (2014) have become popular estimators of causal effects under conditional unconfoundedness in applied economics (e.g., Gilchrist and Sands 2016; Dhar, Jain, and Jayachandran 2022). Yet, a recent literature also raises practical concerns about the use of machine learning for causal inference. Wüthrich and Zhu (2023) find that lasso often fails to select relevant confounders in small samples while inference based on linear regression performs relatively well. Giannone, Lenza, and Primiceri (2021) and Kolesár, Müller, and Roelsgaard (2023) argue that the sparsity assumption, on which the lasso fundamentally relies, is frequently not plausible in economic data sets. Angrist and Frandsen (2022) show that conditioning on confounders using random forests may yield spurious results in IV

regressions.¹ In an application to the evaluation of active labor market programs, Goller et al. (2020) find that random forests are not suitable for the estimation of propensity scores. A key characteristic shared by many of these studies using machine learning for causal inference is the focus on a single pre-selected machine learner.

This paper revisits the application of machine learning for causal inference in light of this recent literature. In particular, we highlight the benefits of pairing double/debiased machine learning (DDML) estimators of Chernozhukov et al. (2018) with stacking (Breiman 1996; Wolpert 1996; van der Laan, Polley, and Hubbard 2007). DDML can leverage generic machine learners meeting mild convergence rate requirements for the estimation of common (causal) parameters. Stacking is a form of model averaging that allows selecting among or combining multiple candidate machine learners relying on different regularization assumptions rather than requiring an ad hoc choice between them. Based on a diverse set of applications and calibrated simulation stud-

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ies, we show that the synthesis of stacking and DDML improves the robustness of estimates of target parameters to the underlying structure of the data, and illustrate the finite sample performance of stacking-based DDML estimators. The results suggest that stacking with a rich set of candidate estimators can address some of the shortcomings highlighted in the recent literature on causal inference with single pre-selected machine learners.

We further consider two alternate ways of combining stacking and DDML aimed at improving practical feasibility and stability in finite samples: *Short-stacking* leverages the cross-fitting step of DDML to reduce the computational burden of stacking substantially. *Pooled stacking* decreases the variance of stacking-based learners across the DDML cross-fitting folds. Both approaches facilitate interpretability compared to conventional stacking by enforcing common stacking weights. We complement the paper with software packages for Stata and R that implement the proposed approaches (Ahrens et al. 2024; Wiemann et al. 2024).

Model averaging techniques have a long tradition in economics and statistics. In the time-series literature, the idea of using “optimal” weights to combine forecasts goes back to the 1960s (Crane and Crotty 1967; Bates and Granger 1969). Loss-minimizing combinations of a pre-specified set of estimators were introduced under the term stacking to the statistics literature by Wolpert (1992) and Breiman (1996) and generalized by van der Laan, Polley, and Hubbard (2007). Stacking fits a final (typically parametric) learner on a set of cross-validated (CV) predicted values derived from distinct candidate learners. A popular choice is to constrain the weights attached to each candidate learner to be non-negative and sum to one. The benefits of combining multiple estimators into a “super learner” via stacking to improve robustness to the structure of the underlying data-generating process are well known in the econometrics and statistics literature. Under appropriate restrictions on the data generating process and loss-function, van Laan and Dudoit (2003) show asymptotic

equivalence between stacking and the best-performing candidate learner.²

Model averaging methods and stacking are widely used in time-series forecasting and macro-econometrics (for recent reviews, see Steel 2020; Wang et al. 2023). Yet, despite its theoretical appeal, stacking has hitherto been rarely used for the estimation of causal effects in economics or other social sciences. An important exception is van der Laan and Rose (2011), who recommend stacking in the context of targeted maximum likelihood. Instead, estimators are often based on parametric (frequently linear) specifications or single pre-selected machine learners. This can have severe consequences for the properties of causal effect estimators if the given choice is ill-suited for the application at hand. A simple example is shown in Figure 1, which compares the performance of DDML using either CV lasso or a feed-forward neural network to estimate a partially linear model across two different data-generating processes. The results show that the bias associated with each learner strongly depends on the structure of the data. Since true functional forms are often unknown in the social sciences, indiscriminate choices of machine learners in practice can thus result in poor estimates. DDML with stacking is a practical solution to this problem. As the example showcases, DDML using stacking is associated with low bias when considering a rich set of candidate learners that are individually most suitable to different structures of the data.

We conduct simulation studies calibrated to real economic datasets to demonstrate that stacking approaches can safeguard against ill-chosen or poorly tuned estimators in practical settings. Throughout, stacking estimators are associated with relatively low bias regardless of the simulated data-generating process, in sharp contrast to the data-dependent performance of the causal effect estimators based on single pre-selected learners. The proposed stacking approaches thus appear relevant in the ubiquitous

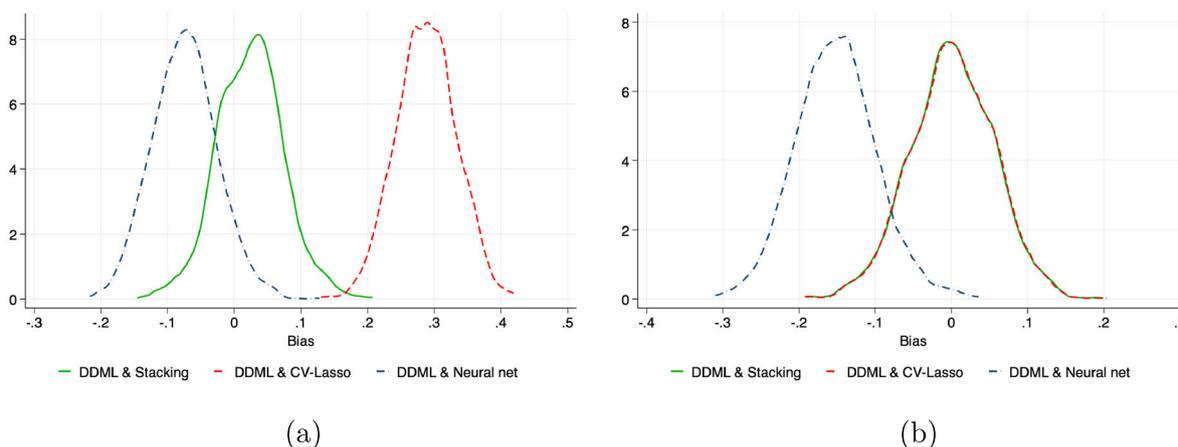


FIGURE 1 | Estimation bias of DDML with cross-validated lasso, feed-forward neural net and stacking. The figures show kernel density plots comparing the bias of DDML paired with either cross-validated lasso, a feed-forward neural net (with two hidden layers of size 20), or a stacking learner combining 13 candidate learners (including cross-validated lasso and ridge, random forests, gradient-boosted trees, and feed-forward neural nets). See Ahrens et al. (2024), where this example is taken from, for details on the specification of each learner. With respect to the data-generating processes, we generate 1000 samples of size $n = 1000$ using the partially linear model $Y_i = \theta_0 D_i + c_Y g(\mathbf{X}_i) + \varepsilon_i$, $D_i = c_D g(\mathbf{X}_i) + u_i$ where X_j are drawn from $\mathcal{N}(0, \Sigma)$ with $\Sigma_{i,k} = 0.5^{|j-k|}$. ε_i and u_i are drawn from standard normal distributions. In (a), the nuisance function is $g(\mathbf{X}_i) = X_{i,1}X_{i,2} + X_{i,3}^2 + X_{i,4}X_{i,5} + X_{i,6}X_{i,7} + X_{i,8}X_{i,9} + X_{i,10} + X_{i,11}^2 + X_{i,12}X_{i,13}$. In (b), the nuisance function is $g(\mathbf{X}_i) = \sum_j 0.9^j X_{i,j}$. c_Y and c_D are two constants chosen to ensure that the R^2 of the regression of Y onto X is approximately 0.5. (a) Nonlinear DGP. (b) Approximately sparse linear DGP.

scenario where there is uncertainty about the set of control variables, correct functional form or the appropriate regularization assumption.

By revisiting the simulation design of Wüthrich and Zhu (2023), we further show that stacking can outperform linear regression for even small sample sizes. We argue that the poor small sample performance of lasso-based approaches is partially driven by the choice of covariate transformations and illustrate how stacking can accommodate a richer set of specifications, including competing parametric models. We also find that short-stacking and pooled stacking may outperform DDML paired with conventional stacking in small to moderate sample sizes. Paired with its lower computational cost, this finding suggests that short-stacking may be an attractive baseline approach to select and combine competing reduced form specifications.

Finally, we demonstrate the value of pairing DDML with stacking with two applications. First, we examine gender gaps in citations of articles published in top-30 economic journals from 1983 to 2020, and assess how the difference in citations change when conditioning on content and quality proxied by the abstract text. Estimating these conditional differences is a challenging statistical problem due to the non-standard nature of text data, which is increasingly encountered in economic applications (see also, e.g., Ash and Hansen 2023; Ash, Chen, and Ornaghi 2024; Eberhardt, Facchini, and Rueda 2023). Second, we revisit a UK sample of the OECD Skill Survey to estimate semiparametric Kitagawa-Oaxaca-Binder estimates of the unexplained gender wage gap. Both applications highlight that estimators of structural parameters based on single learners can be highly sensitive to the underlying structure of the data and/or poor tuning. The applications further demonstrate that DDML with stacking is a simple and practical solution to resolve the difficult problem of choosing a particular candidate learner in practice. Further, we observe that the optimal stacking weights often vary across reduced-form equations—meaning that different conditional expectation functions in the same data set are best estimated using different learners. This behavior sharply contrasts with common estimation approaches, such as OLS and PDS lasso, that impose the same form for each conditional expectation function.

The remainder of the paper is organized as follows: Section 2 provides a brief review of DDML. Section 3 discusses DDML with stacking, short-stacking, and pooled stacking. Section 4 presents our calibrated simulation studies. Section 5 discusses the applications, and Section 6 concludes.

2 | DDML

This section outlines DDML as discussed in Chernozhukov et al. (2018). Throughout, we focus on the partially linear model as a natural extension of commonly applied linear regression methods. Despite its simplicity, the partially linear model illustrates practical challenges in the application of DDML that can be addressed by stacking. We highlight, however, that our discussion also applies to the wide range of models outlined in Chernozhukov et al. (2018) and more generally to estima-

tion of low-dimensional structural parameters in the presence of high-dimensional nuisance functions.³ Stacking could also be applied in settings where DDML is applied with nonparametric targets as in Colangelo and Lee (2023).

The partially linear model is defined by a random vector (Y, D, X^\top, U) with joint distribution characterized by

$$Y = \theta_0 D + g_0(X) + U, \quad (1)$$

where Y is the outcome, D is the scalar variable of interest, and X is a vector of control variables. The parameter of interest θ_0 and the unknown nuisance function g_0 are such that the corresponding residual U satisfies the conditional orthogonality property $E[Cov(U, D|X)] = 0$. These properties are analogous to the orthogonality properties of residuals in multiple linear regression with the key difference here being that g_0 need not be linear in the controls.

Albeit a seemingly small change in specification, the partially linear model has several important advantages over linear regression. For discrete D , for example, results in Angrist and Krueger (1999) imply that θ_0 can be interpreted as a positively weighted average of incremental changes in the conditional expectation function $E[Y|D = d, X]$. Under appropriate conditional unconfoundedness assumptions, θ_0 thus corresponds to a convex combination of conditional average treatment effects.⁴ Importantly, these interpretations remain valid even if the additive separability assumption of the partially linear model fails. Linear regression coefficients, in contrast, do not correspond to positively weighted averages of causal effects without imposing strong linearity assumptions that are questionable in real applications.⁵

The advantages of the partially linear model in the interpretation of its parameter of interest come at the cost of a more challenging estimation problem relative to estimating a model that is linear in a pre-specified set of variables. Estimators for θ_0 are based on the solution to the moment equation

$$E[(Y - \ell_0(X) - \theta_0(D - m_0(X)))(D - m_0(X))] = 0,$$

given by

$$\theta_0 = \frac{E[(Y - \ell_0(X))(D - m_0(X))]}{E[(D - m_0(X))^2]},$$

where $\ell_0(X) \equiv E[Y|X]$ and $m_0(X) \equiv E[D|X]$ are the conditional expectations of the outcome and variable of interest given the controls, respectively. Since conditional expectation functions are high-dimensional in the absence of strong functional form assumptions, a sample analog estimator for θ_0 requires nonparametric first-step estimators for the nuisance parameters ℓ_0 and m_0 . While nonparametric estimation generally reduces bias compared to linear regression alternatives, the increased variance associated with more flexible functional form estimation introduces additional statistical challenges: To allow for statistical inference on θ_0 , the nonparametric estimators need to converge sufficiently quickly to the true conditional expectation functions as the sample size increases.

DDML defines a class of estimators that allows for statistical inference on the parameter of interest θ_0 while only imposing

relatively mild convergence requirements on the nonparametric estimators. These mild requirements are central to the wide applicability of DDML as they permit the use of a large variety of machine learners.⁶

Two key devices permit the mild convergence requirements of DDML: identification of the parameter of interest based on Neyman-orthogonal moment conditions and estimation using cross-fitting. Neyman-orthogonal moment conditions are insensitive to local perturbations around the true nuisance parameter.⁷

Cross-fitting is a sample-splitting approach that addresses the *own-observation bias* that arises when the nuisance parameter estimation and the estimation of θ_0 are applied to the same observation. In practice, cross-fitting is implemented by randomly splitting a sample $\{(Y_i, D_i, X_i^\top)\}_{i \in I}$ indexed by $I = \{1, \dots, n\}$ into K evenly-sized folds, denoted as I_1, \dots, I_K . For each fold k , the conditional expectations ℓ_0 and m_0 are estimated using only observations not in the k th fold—that is, in $I_k^c \equiv I \setminus I_k$ —resulting in $\hat{\ell}_{I_k^c}$ and $\hat{m}_{I_k^c}$, respectively, where the subscript I_k^c indicates the subsample used for estimation. The out-of-sample predictions for an observation i in the k th fold are then computed via $\hat{\ell}_{I_k^c}(X_i)$ and $\hat{m}_{I_k^c}(X_i)$. Repeating this procedure for all K folds then allows for computation of the DDML estimator for θ_0 :

$$\hat{\theta}_n = \frac{\frac{1}{n} \sum_{i=1}^n \left(Y_i - \hat{\ell}_{I_{k_i}^c}(X_i) \right) \left(D_i - \hat{m}_{I_{k_i}^c}(X_i) \right)}{\frac{1}{n} \sum_{i=1}^n \left(D_i - \hat{m}_{I_{k_i}^c}(X_i) \right)^2},$$

where k_i denotes the fold of the i th observation.

Since the cross-fitting algorithm depends on the randomized fold split, and since some machine learners rely on randomization too, DDML estimates vary with the underlying random-number generator and seed. To reduce dependence on randomization, it is thus worthwhile to repeat the cross-fitting procedure and apply mean or median aggregation over DDML estimates (see Remark 2 in Ahrens et al. 2024). We show in Section 5 that repeating the cross-fitting procedure is a useful diagnostic tool, allowing one to gauge the stability of DDML estimators.

Under the conditions of Chernozhukov et al. (2018)—including, in particular, the convergence requirements on the nonparametric estimators— $\hat{\theta}_n$ is root- n asymptotically normal around θ_0 . As already highlighted by the example in Figure 1, however, a poorly chosen or poorly tuned machine learner for the estimation of nuisance parameters $\hat{\ell}$ and \hat{m} can have detrimental effects on the properties of $\hat{\theta}_n$. Since no machine learner can be best across all settings, this raises the difficult question of which learner to apply in a particular setting. In the next section, we discuss how DDML can be paired with stacking to provide a practical solution to the choice of learner. We also illustrate how the cross-fitting structure naturally arising in DDML estimators can be leveraged to substantially reduce the computational burden otherwise associated with stacking.

3 | Pairing DDML with Stacking Approaches

This section discusses the estimation of structural parameters by pairing DDML with stacking approaches. After the discussion of

DDML with conventional stacking, we introduce two stacking variants that leverage the cross-fitting structure of DDML estimators: short-stacking and pooled stacking. To fix ideas, we focus on the nuisance parameter $\ell_0(X) = E[Y|X]$ arising in the partially linear model where we consider an i.i.d. sample $\{(Y_i, X_i)\}_{i \in I}$. Further, we consider a rich set of J pre-selected base or candidate learners. The set of learners could include distinct parametric and nonparametric estimators—for example, linear or logistic regression, regularized regression such as the lasso, or tree-based methods such as random forests—as well as the same algorithm with varying (hyper-)tuning parameters or different (basis) expansions of the control variables. It is important to note that the set of candidate learners for stacking can readily incorporate commonly used unregularized learners such as linear or logistic regression; in practice, sometimes, the best-performing candidate learner may be one such learner.

3.1 | DDML with Conventional Stacking

Combining DDML with conventional stacking involves two layers of re-sampling, as we illustrate in Figure 2. The *cross-fitting layer* divides the sample into K cross-fitting folds, denoted by I_1, \dots, I_K . In each cross-fitting step $k \in \{1, \dots, K\}$, the stacking learner is trained on the training sample which excludes fold I_k and which we label $T_k \equiv I \setminus I_k$. Fitting the stacking learner, in turn, requires subdividing the training sample T_k further into V cross-validation folds. This second sample split constitutes the *cross-validation layer*. We denote the cross-validation folds in cross-fitting step k by $T_{k,1}, \dots, T_{k,V}$. Each candidate learner $j \in \{1, \dots, J\}$ is CV across these folds, yielding CV predicted values for each learner.

The final learner fits the outcome Y_i against the CV predicted values of each candidate learner. The most common choice is to construct a convex combination via constrained least squares (CLS), with weights restricted to be non-negative and summing to one. Specifically, for each k , candidate learners are combined to solve

$$\begin{aligned} \min_{w_{k,1}, \dots, w_{k,J}} \sum_{i \in T_k} \left(Y_i - \sum_{j=1}^J w_{k,j} \hat{\ell}_{T_{k,v(i)}}^{(j)}(X_i) \right)^2 \\ \text{s.t. } w_{k,j} \geq 0, \sum_{j=1}^J |w_{k,j}| = 1. \end{aligned}$$

Here, $\hat{\ell}_{T_{k,v(i)}}^{(j)}(X_i)$ denotes the out-of-sample predicted value for observation i , which is calculated from training candidate learner j on sub-sample $T_{k,v(i)}^c \equiv T_k \setminus T_{k,v(i)}$, that is, all step- k cross-validation folds but fold $(k, v(i))$ which is the fold of the i th observation. We call the resulting $\hat{w}_{k,j}$ the *stacking weights*. The stacking predictions are obtained as $\sum_j \hat{w}_{k,j} \hat{\ell}_{T_k}^{(j)}(X_i)$ where each learner j is re-fit on T_k .

Although various options for combining candidate learners are available, CLS facilitates the interpretation of stacking as a weighted average of candidate learners (Hastie et al. 2009). Due to this constraint, CLS tends to set some stacking weights to exactly zero. The constraint also regularizes the final estimator, which is important to mitigate issues arising from potential multicollinearity of the candidate learners. An alternative to CLS,

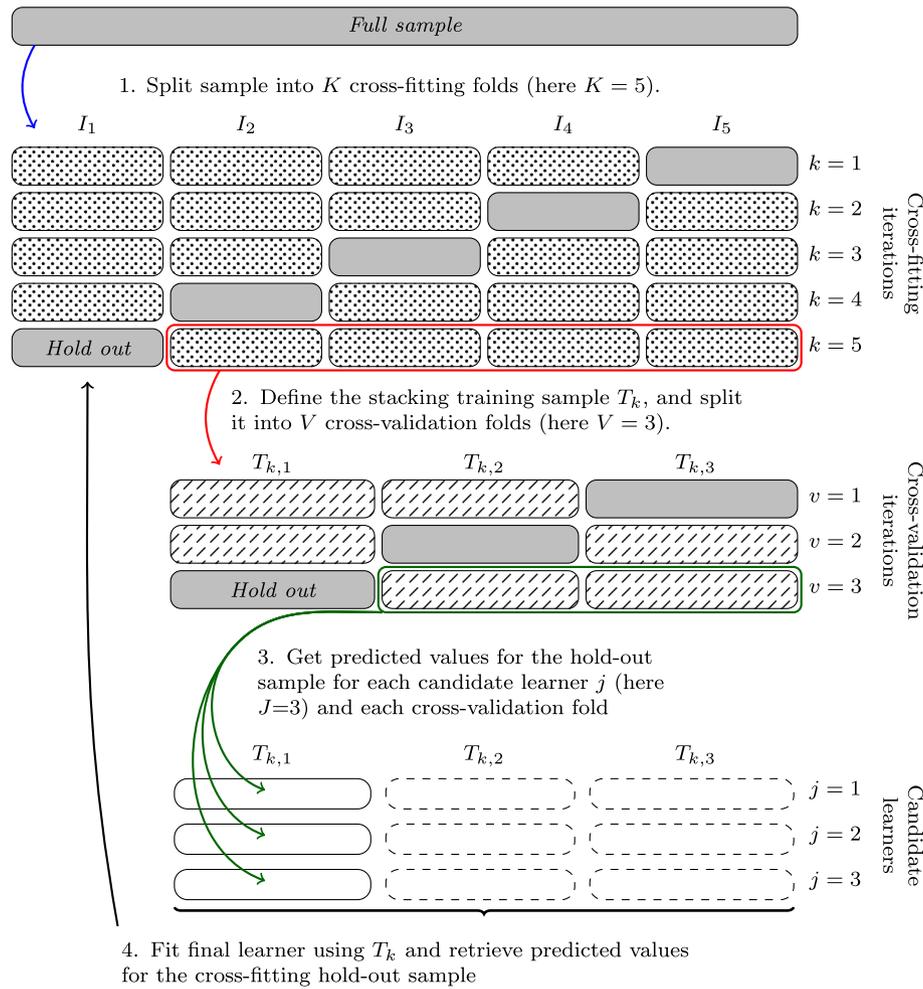


FIGURE 2 | Cross-fitting with conventional stacking. The diagram illustrates cross-fitting with conventional stacking. The diagram uses $K = 5$ cross-fitting folds, $V = 3$ cross-validation folds and $J = 3$ candidate learners. Step 1: The sample randomly is split into cross-fitting folds I_1, \dots, I_K . Step 2: In each step $k \in \{1, \dots, K\}$ of the cross-fitting process, we define the training data as $T_k \equiv I \setminus I_k$. The k -step training data is then split into V sub-partitions, which we denote by $T_{k,1}, \dots, T_{k,V}$. Step 3: For each cross-fit step k and cross-validation step $v \in \{1, \dots, V\}$, fit each base learner $j \in \{1, \dots, J\}$ on $T_k \setminus T_{k,v}$ and obtain predicted values for the cross-validation hold-out sample. Step 4: Fit the final learner on sample T_k to obtain predicted values for the cross-fitting hold-out sample I_k .

which we refer to as *single-best learner*, is to impose the constraint that $w_{k,j} \in \{0,1\}$ and $\sum_j w_{k,j} = 1$, implying that only the candidate learner with lowest CV loss is used as the final estimator. Under appropriate restrictions on the data-generating process and loss function, van der Laan and Dudoit (2003) show asymptotic equivalence between stacking and the best-performing candidate learner.⁸

A drawback of DDML with stacking is its computational complexity. Considering the estimation of a single candidate learner as the unit of complexity (and ignoring the cost of fitting the final learner), DDML with stacking heuristically has a computational cost proportional to $K \times V \times J$. For example, when considering DDML with $K = 5$ cross-fitting folds and $J = 10$ candidate learners that are combined based on $V = 5$ fold cross-validation, more than 250 candidate learners need to be individually estimated. Although DDML with stacking is “embarrassingly parallel” and can thus be expected to decrease in computational time nearly linearly in the number of available computing processes, the

increased complexity limits its application to moderately complex applications. Another potential concern (which we investigate in Section 4.2) is that DDML with stacking might not perform well in small samples, given that candidate learners are effectively trained on approximately $\frac{(K-1)(V-1)}{KV}$ % of the full sample (see Figure 2). These two concerns motivate *short-stacking*.

3.2 | DDML with Short-Stacking

In the context of DDML, we propose to take a short-cut: Instead of fitting the final learner on the CV fitted values in each step k of the cross-fitting process, we can directly train the final learner on the cross-fitted values using the full sample; see Figure 3. Formally, candidate learners are combined to solve

$$\min_{w_1, \dots, w_J} \sum_{i=1}^n \left(Y_i - \sum_{j=1}^J w_j \hat{\ell}_{I_{k(i)}}^{(j)}(X_i) \right)^2 \quad \text{s.t. } w_j \geq 0, \sum_{j=1}^J |w_j| = 1$$

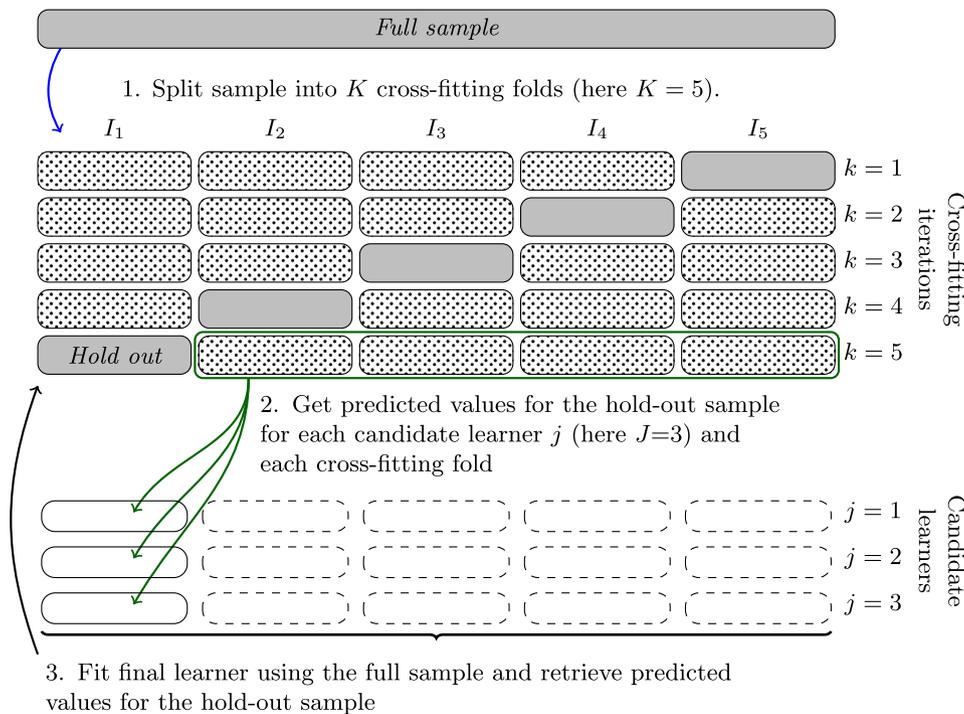


FIGURE 3 | Cross-fitting with *short-stacking*. The diagram illustrates cross-fitting with short-stacking. The diagram uses $K = 5$ cross-fitting folds and $J = 3$ candidate learners. Step 1: The sample randomly is split into cross-fitting folds I_1, \dots, I_K . Step 2: In each step $k \in \{1, \dots, K\}$ of the cross-fitting process, we define the training data as $I_k^c = I \setminus I_k$. Fit each base learner $j \in \{1, \dots, J\}$ on the training data and obtain predicted values for the cross-fitting hold-out sample I_k . Step 3: Fit the final learner on the full sample to obtain predicted values for the cross-fitting hold-out sample.

where w_j are the short-stacking weights. Cross-fitting thus serves a double purpose: First, it avoids the own-observation bias by avoiding overlap between the samples used for estimating high-dimensional nuisance functions and the samples used for estimating structural parameters. Second, it yields out-of-sample predicted values which we leverage for constructing the final stacking learner. As a consequence, the computational cost of DDML with short stacking is heuristically only proportional to $K \times J$ in units of estimated candidate learners. In the example from the previous paragraph, short-stacking thus requires estimating about 200 fewer candidate learners.

We recommend DDML with short-stacking in settings where the number of candidate learners is small relative to the sample size, that is, $J \ll n$. We believe this setting provides a good approximation to current applications of machine learning in economics and other social sciences where it is rare to consider more than a few candidate learners. If instead the number of considered learners is very large relative to the sample size—that is, settings in which inference for standard linear regression on J variables is invalid—pairing DDML with short-stacking may introduce bias.⁹

3.3 | DDML with Pooled Stacking

While DDML with conventional stacking has one vector of weights per cross-fitting fold, short-stacking yields a single weight for each learner. A single weight for each learner decreases the variance of the final estimator and facilitates the interpretation of the stacking weights. Another way of achieving

common stacking weights is DDML with pooled stacking. Pooled stacking relies on the same two-layer re-sampling strategy as conventional stacking but combines candidate learners to solve

$$\min_{w_1, \dots, w_J} \sum_{i \in I} \sum_{k \neq k(i)} \left(Y_i - \sum_{j=1}^J w_j \hat{\ell}_{T_{k,e(i)}^c}^{(j)}(\mathbf{X}_i) \right)^2$$

s.t. $w_j \geq 0, \sum_{j=1}^J |w_j| = 1.$

That is, pooled stacking collects the CV predicted values that are calculated in each step k of the cross-fitting process for each learner j and estimates the stacking weights based on the pooled data set. We note that the computational costs are approximately the same as for DDML with conventional stacking.

4 | The Practical Benefits of DDML with Stacking: Two Simulation Studies

In this section, we discuss two simulation studies illustrating the advantages of pairing DDML with stacking over alternative approaches based on single pre-selected learners. We begin with a simulation calibrated to household data on wealth and 401k eligibility from the 1991 wave of the Survey of Income and Program Participation (SIPP) in Section 4.1. In Section 4.2, we revisit the simulation of Wüthrich and Zhu (2023) to assess the robustness of DDML with stacking approaches in very small samples.

4.1 | Simulation Calibrated to the SIPP 1991 Household Data

To assess the performance of DDML with conventional stacking, short-stacking and pooled stacking in a realistic setting, we consider the analysis of 401(k) eligibility and total financial assets in Poterba, Venti, and Wise (1995) as the basis for an empirically calibrated Monte Carlo simulation. The application has recently been revisited by Belloni et al. (2017), Chernozhukov et al. (2018), and Wüthrich and Zhu (2023) to approximate high-dimensional confounding factors using machine learning. We focus on estimating the partially linear model discussed in the previous section. The outcome is measured as net financial assets, the treatment variable is an indicator for eligibility to the 401(k) pension scheme, and the set of controls includes age, income, education in years, family size, as well as indicators for two-earner status, home ownership, and participation in two alternative pension schemes.

The simulation involves three steps. In the calibration step, we fit two generative models to the $n = 9915$ households from the 1991 wave of the Survey of Income and Program Participation. The first generative model is fully linear while the second is partially linear, allowing controls to enter nonlinearly through gradient-boosted trees fitted to the real data. This approach is aimed at extracting and magnifying the linear or nonlinear structures in the empirical conditional distributions, respectively, enabling us to compare the performance of estimators across favorable and unfavorable structures of the data. The generative step then simulates datasets of size $n_b = \{9915, 99150\}$ from the respective fully linear model and the partially linear model. Throughout, we set the effect of 401(k) eligibility on total financial wealth to $\theta_0 = 6000$. Finally, in the estimation step, we fit various estimators to bootstrapped samples of the generated datasets and assess their statistical properties. We outline the steps used for constructing the two generative models in more detail in Algorithm 1.

For each bootstrap sample, we calculate estimates of the effect of 401(k) eligibility on simulated net financial assets. The estimators we consider are linear regression, the post-double selection (PDS) lasso estimator proposed by Belloni, Chernozhukov, and Hansen (2014), as well as DDML estimators with and without stacking. The candidate learners of the DDML estimators are linear regression, CV lasso and ridge regression with interactions and second-order polynomial expansions of the controls, CV lasso and ridge with no interactions but 10th-order polynomial expansions of the controls, two versions of random forests, two versions of gradient-boosted trees, and feed-forward neural nets with three hidden layers of size five (see notes in Table 1 for details). We estimate DDML paired with conventional stacking, short-stacking and pooled stacking, and consider different methods to construct the final conditional expectation function estimator: CLS, unconstrained linear regression (OLS), selecting the single best estimator, and an unweighted average.

Algorithm 1 Algorithm for the calibrated Monte Carlo simulation.

Let $\{(y_i, d_i, x_i)\}_{i=1, \dots, n}$ denote the observed sample, where i is a household in the 1991 SIPP and y_i , d_i , and x_i respectively denote net financial assets, an indicator for 401(k) eligibility, and the vector of control variables.

1. Using the full sample, obtain the slope coefficient $\hat{\theta}_{OLS} \approx 5896$ from linear regression of y_i against d_i and x_i in the original data. Construct the partial residuals $y_i^{(r)} = y_i - \hat{\theta}_{OLS}d_i$, $\forall i$.
2. Fit a supervised learning estimator (either linear regression or gradient boosting) to predict $y_i^{(r)}$ with the controls x_i . Denote the fitted estimator by \tilde{g} . Similarly, fit a supervised learning estimator to predict d_i with x_i and denote the fitted estimator by \tilde{h} .
3. Repeat to generate simulated samples of size n_b :
 - (a) Sample from the empirical distribution of x_i by bootstrapping n_b observations from the original data. Denote the bootstrapped sample by \mathcal{D}_b .
 - (b) Draw $v_i \stackrel{iid}{\sim} \mathcal{N}(0, \kappa_1)$ and $\varepsilon_i \stackrel{iid}{\sim} \mathcal{N}(0, \kappa_2)$, where κ_1 and κ_2 are simulation hyperparameters. Define

$$\begin{aligned} \tilde{d}_i^{(b)} &= \mathbb{1}\{\tilde{h}(x_i) + v_i \geq 0\} \\ \tilde{y}_i^{(b)} &= \theta_0 \tilde{d}_i^{(b)} + \tilde{g}(x_i) + \varepsilon_i \quad \forall i \in \mathcal{D}_b \end{aligned}$$

where we set $\theta_0 = 6000$ to roughly resemble the magnitude of the regression coefficient of 401(k) eligibility in the full data.

Note: We set the hyper-parameter κ_1 and κ_2 to approximately match the variance of 401(k) eligibility and wealth in the data. The values of the simulation hyperparameters (κ_1, κ_2) differ slightly depending on the supervised learning estimator used to fit the reduced form equations in the data. We take $\kappa_1 = 0.35$ in both scenarios but take $\kappa_2 = 55,500$ when using linear regression and $\kappa_2 = 54,000$ when using gradient boosting. Differences arise because gradient boosting reduces residual variance in the true data.

Table 1 presents the mean bias, median absolute bias (MAB) and coverage rates of a 95% confidence interval associated with estimates of the effect of 401(k) eligibility on net financial assets. The left and right panels correspond to results based on data simulated from the linear (Panel A) and nonlinear (Panel B) generative models, respectively. The CLS weights associated with each candidate learner are shown in Table 2.¹⁰

Given the construction of the generative models, we would expect that linear regression performs best in the fully linear setting and that DDML with gradient boosting performs best in the nonlinear setting where the nuisance function is generated by gradient boosting. The simulation results confirm this intuition, showing that the two procedures achieve among the lowest bias and median absolute bias in the data-generating processes that are based on them. Researchers are rarely certain of the functional structure in economic applications,

TABLE 1 | Bias and coverage rates in the linear and nonlinear DGP.

	Panel (A): Linear DGP						Panel (B): Nonlinear DGP					
	$n_b = 9915$			$n_b = 99,150$			$n_b = 9915$			$n_b = 99,150$		
	Bias	MAB	Rate	Bias	MAB	Rate	Bias	MAB	Rate	Bias	MAB	Rate
Panel A. Full sample methods												
OLS	49.9	793.8	0.95	-6.8	281.2	0.95	-2588.9	2576.5	0.58	-2632.3	2611.5	0
PDS-lasso	48.4	787.1	0.95	-4.2	280.8	0.95	-2598.7	2590.1	0.58	-2631.6	2609.5	0
Panel B. DDML methods												
Candidate learners												
OLS	46.2	818.1	0.94	-6.9	283.1	0.95	-2613.0	2634.2	0.58	-2635.4	2615.9	0
Lasso with CV (2nd-order poly)	50.9	806.6	0.95	-6.2	284.8	0.95	703.7	1052.3	0.91	718.5	712.8	0.60
Ridge with CV (2nd-order poly)	48.2	806.9	0.94	-6.9	283.7	0.96	767.4	1080.8	0.90	729.3	724.0	0.60
Lasso with CV (10th-order poly)	248.1	1034.5	0.94	55.9	285.9	0.95	-4109.0	1799.9	0.90	7.4	306.5	0.94
Ridge with CV (10th-order poly)	1230.1	1321.9	0.91	31.6	283.0	0.96	-5126.2	2215.7	0.89	9.6	307.8	0.94
Random forest (low regularization)	-74.7	1031.3	0.89	-25.2	344.0	0.88	-96.1	1037.1	0.90	-37.5	328.0	0.87
Random forest (high regularization)	69.1	891.2	0.94	-23.5	287.6	0.93	-159.7	904.4	0.94	-4.2	280.4	0.95
Gradient boosting (low regularization)	12.1	817.0	0.94	-24.2	285.1	0.96	8.5	866.0	0.94	30.9	275.1	0.96
Gradient boosting (high regularization)	114.8	823.8	0.94	66.9	285.6	0.95	162.0	857.2	0.94	200.1	314.6	0.93
Neural net	394.2	943.6	0.93	9.1	287.5	0.94	-601.3	1063.9	0.93	-131.9	310.0	0.93
Stacking approaches												
Stacking: CLS	42.8	813.4	0.94	-7.5	282.9	0.96	133.9	1049.5	0.94	37.8	271.0	0.95
Stacking: average	107.7	821.9	0.94	-6.5	273.6	0.95	94.0	1035.6	0.94	72.3	289.3	0.96
Stacking: OLS	-129.3	878.7	0.94	-9.4	283.8	0.95	-204.8	1184.9	0.94	17.5	268.3	0.96
Stacking: single-best	43.7	819.8	0.94	-8.6	281.4	0.95	-121.9	976.2	0.94	30.9	275.1	0.96
Short-stacking: CLS	45.0	794.9	0.94	-7.0	282.6	0.95	162.7	865.1	0.94	33.6	266.3	0.95
Short-stacking: average	107.7	821.9	0.94	-6.5	273.6	0.95	94.0	1035.6	0.94	72.3	289.3	0.96
Short-stacking: OLS	37.6	803.7	0.94	-7.8	282.0	0.95	123.6	863.5	0.94	29.5	265.3	0.96
Short-stacking: single-best	44.4	817.8	0.94	-8.3	281.9	0.95	71.7	868.4	0.94	30.9	275.1	0.96
Pooled stacking: CLS	58.6	819.5	0.95	-7.1	283.9	0.96	209.8	921.0	0.94	37.5	269.8	0.95
Pooled stacking: average	107.7	821.9	0.94	-6.5	273.6	0.95	94.0	1035.6	0.94	72.3	289.3	0.96
Pooled stacking: OLS	46.5	805.2	0.94	-7.8	284.0	0.96	234.5	1003.2	0.94	30.6	266.3	0.96
Pooled stacking: single-best	46.9	823.5	0.94	-8.3	282.6	0.95	103.3	904.6	0.94	30.9	275.1	0.96

Note: The table reports mean bias, median absolute bias (MAB), and coverage rate of a 95% confidence interval for the listed estimators. We consider DDML with $K = 2$ cross-fit folds and the following individual learners: OLS with elementary covariates, CV lasso and CV ridge with second-order polynomials and interactions, CV lasso and CV ridge with 10th-order polynomials but no interactions, random forest with low regularization (8 predictors considered at each leaf split, no limit on the number of observations per node, bootstrap sample size of 70%), highly regularized random forest (5 predictors considered at each leaf split, at least 10 observation per node, bootstrap sample size of 70%), gradient-boosted trees with low regularization (500 trees, maximum depth of 3 and a learning rate of 0.01), gradient-boosted trees with high regularization (250 trees, maximum depth of 3 and a learning rate of 0.01), and feed-forward neural nets with three hidden layers of size five. For reference, we report two estimators using the full sample: OLS and PDS lasso. Finally, we report results for DDML paired with conventional stacking, short-stacking and pooled stacking where the final estimator is either CLS, OLS, the unweighted average of candidate learners or the single-best candidate learner. Results are based on 1000 replications.

however, so that it is more interesting to consider their respective performance in the non-favorable setting. In the nonlinear data-generating process, linear regression is among the estimators with the worst performance across all three measures. Similarly, gradient boosting-based DDML is non-optimal in the linear data-generating process. It is outperformed by linear regression and CV lasso, both of which enforce a linear functional form on the control variables, in terms of MAB.

The simulation results are consequences of the “no free lunch” theorem in machine learning (Wolpert 1996). Informally, the

theorem states that there exists no estimator that performs best across all empirical settings. Researchers must, therefore, carefully match estimators to their application. However, with limited knowledge about underlying data-generating processes and few functional form restrictions implied by economic theory, the number of plausibly suitable estimators is typically large.

The bottom section of Table 1 reports results for DDML combined with the three stacking approaches outlined in Section 3. For each stacking approach, we consider stacking weights estimated by (CLS) as outlined in Section 3, set equal to $1/J$

TABLE 2 | Average stacking weights with CLS.

	Stacking		Pooled stacking		Short-stacking	
	$E[Y X]$	$E[D X]$	$E[Y X]$	$E[D X]$	$E[Y X]$	$E[D X]$
Panel A. Linear DGP and $n_b = 9915$						
OLS	0.668	0.501	0.738	0.573	0.692	0.492
Lasso with CV (2nd-order poly)	0.105	0.144	0.093	0.131	0.118	0.130
Ridge with CV (2nd-order poly)	0.068	0.054	0.050	0.040	0.068	0.063
Lasso with CV (10th-order poly)	0.027	0.073	0.021	0.059	0.020	0.085
Ridge with CV (10th-order poly)	0.033	0.043	0.018	0.027	0.024	0.057
Random forest (low regularization)	0.013	0.011	0.009	0.007	0.009	0.008
Random forest (high regularization)	0.017	0.024	0.012	0.018	0.013	0.024
Gradient boosting (low regularization)	0.030	0.043	0.024	0.034	0.020	0.040
Gradient boosting (high regularization)	0.020	0.060	0.020	0.070	0.018	0.060
Neural net	0.019	0.049	0.014	0.041	0.018	0.043
Panel B. Linear DGP and $n_b = 99,150$						
OLS	0.770	0.325	0.824	0.348	0.769	0.291
Lasso with CV (2nd-order poly)	0.066	0.026	0.055	0.012	0.068	0.013
Ridge with CV (2nd-order poly)	0.074	0.041	0.057	0.028	0.094	0.049
Lasso with CV (10th-order poly)	0.015	0.207	0.011	0.225	0.011	0.198
Ridge with CV (10th-order poly)	0.023	0.111	0.018	0.095	0.019	0.129
Random forest (low regularization)	0.003	0.003	0.002	0.002	0.002	0.002
Random forest (high regularization)	0.006	0.009	0.005	0.006	0.005	0.006
Gradient boosting (low regularization)	0.019	0.160	0.015	0.175	0.013	0.186
Gradient boosting (high regularization)	0.004	0.008	0.003	0.004	0.003	0.003
Neural net	0.019	0.110	0.009	0.106	0.016	0.122
Panel C. nonlinear DGP and $n_b = 9915$						
OLS	0.011	0.015	0.007	0.012	0.004	0.007
Lasso with CV (2nd-order poly)	0.035	0.057	0.024	0.051	0.019	0.039
Ridge with CV (2nd-order poly)	0.161	0.229	0.192	0.268	0.114	0.237
Lasso with CV (10th-order poly)	0.053	0.080	0.047	0.071	0.048	0.062
Ridge with CV (10th-order poly)	0.071	0.064	0.060	0.029	0.059	0.056
Random forest (low regularization)	0.045	0.011	0.043	0.006	0.043	0.005
Random forest (high regularization)	0.019	0.069	0.010	0.065	0.012	0.065
Gradient boosting (low regularization)	0.521	0.233	0.548	0.251	0.632	0.339
Gradient boosting (high regularization)	0.014	0.191	0.005	0.203	0.004	0.139
Neural net	0.071	0.051	0.065	0.043	0.064	0.049
Panel D. Nonlinear DGP and $n_b = 99,150$						
OLS	0	0	0	0	0	0
Lasso with CV (2nd-order poly)	0	0	0	0	0	0
Ridge with CV (2nd-order poly)	0	0.036	0	0.037	0	0.026
Lasso with CV (10th-order poly)	0	0.001	0	0	0	0
Ridge with CV (10th-order poly)	0	0.036	0	0.034	0	0.024
Random forest (low regularization)	0.153	0.003	0.154	0.001	0.180	0.001
Random forest (high regularization)	0	0.060	0	0.063	0	0.071
Gradient boosting (low regularization)	0.845	0.853	0.846	0.858	0.819	0.871
Gradient boosting (high regularization)	0	0	0	0	0	0
Neural net	0.001	0.012	0	0.006	0.001	0.007

Note: The table shows the average stacking weights associated with the candidate learner for DDML with conventional stacking, pooled stacking and short-stacking. The final learner is CLS. The bootstrap sample size is denoted by n_b . The number of cross-fitting folds is $K = 2$. Results are based on 1000 replications. See Table 1 for more information. The final learner weights using OLS and single best are reported in Tables S1.3 and S1.4.

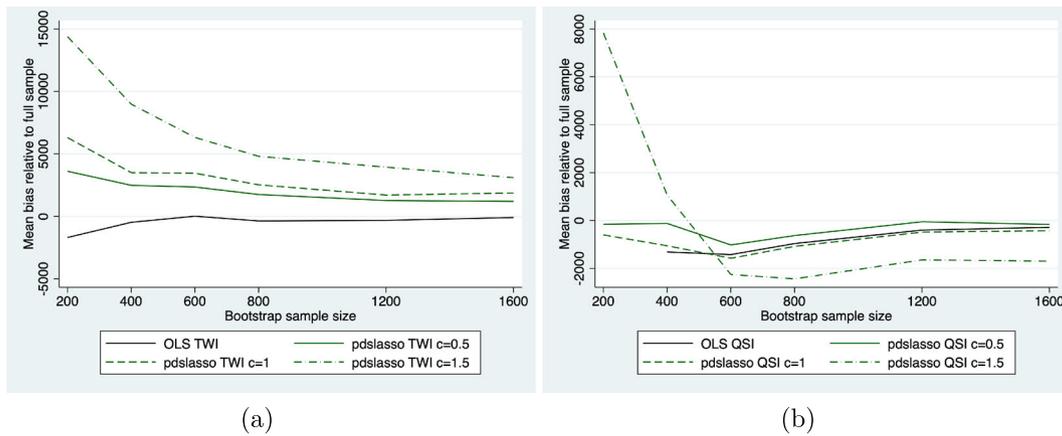


FIGURE 4 | Replication of Figure 8 in Wüthrich and Zhu (2023). The figures report the mean bias calculated as the mean difference to the full sample estimates. Full sample estimates reported in Table S2.1. Following WZ, we draw 1000 bootstrap samples of size $n_b = \{200, 400, 600, 800, 1200, 1600\}$. “TWI” indicates that the predictors have been expanded by two-way interactions. “QSI” refers to the quadratic spline and interactions specification of Belloni et al. (2017). (a) Bias (TWI). (b) Bias (QSI).

(average), estimated without constraint by OLS (OLS), and by selecting only the single best candidate learner (single-best). We find that short-stacking performs similarly to, and sometimes better than, conventional and pooled stacking, while being computationally much cheaper (as shown in Table S1.5). For example, at $K = 10$ and $V = 5$, DDML combined with short-stacking ran around 4.3 times faster on the full sample than DDML with conventional or pooled stacking, which is roughly in line with a speed improvement by a factor of $1/V$.¹¹

The simulation set-up is favorable to using single-best as the final learner because there is one “true” candidate learner. However, single-best does not visibly outperform CLS, although single-best always selects the correct learner in the nonlinear DGP if the sample size is sufficiently large (see Table S1.4). We believe that CLS is, in practice, a good default choice. In a setting where there is a single learner that does a distinctly better job approximating the target conditional expectation function, CLS should assign a very large weight to that learner and thus approximate single-best. Otherwise, when there are several learners with similar performance or learners that perform differentially well for different observations, there are potential gains from combining the different learners.

The bias of the OLS final learner is overall similar to CLS, except when employing conventional stacking under the nonlinear DGP for $n_b = 9915$ where the average bias is almost four times as large.¹² The unweighted average appears sub-optimal for $n_b = 99150$ under the nonlinear DGP. Poor performance of unweighted averaging is to be expected in settings where, as in our setup, the candidate set includes learners that are not well-matched to the DGP.¹³ We also note that the computational advantage of short-stacking with the unweighted average over short-stacking with estimated weights amounts to one (constrained) regression per conditional expectation function and is thus minimal.

The CLS weights in Table 2 indicate that stacking approaches successfully assign the highest weights to the estimators aligning with the data-generating process (i.e., either OLS or gradient

boosting) among the ten included candidate learners, illustrating the ability to adapt to different data structures. Specifically, the stacking methods applied to the linear data-generating process assign the largest weight to linear models while they assign the largest weights to the gradient-boosting estimators and the lowest weights to estimators that impose a linear functional form on the control variables in the nonlinear data-generating process.¹⁴ We conclude that DDML paired with stacking approaches reduces the burden of choice researchers face when selecting between candidate learners and specifications by allowing for the simultaneous consideration of multiple options, thus implying attractive robustness properties across a variety of data-generating processes.

4.2 | DDML and Stacking in Very Small Samples

A possible concern for estimators relying on machine learning is that they might not perform well for very small samples, given that their flexibility comes at the cost of increased variance compared to parametric estimators. Wüthrich and Zhu (2023), henceforth WZ) use two simulations to demonstrate that PDS lasso tends to underselect controls, which may result in a substantial small-sample bias. They also show that the bias heavily depends on the exact lasso penalty chosen (i.e., whether the plugin penalty of Belloni, Chernozhukov, and Hansen 2014, is scaled by 0.5 or 1.5) and argue in favor of OLS with appropriately chosen standard errors over PDS lasso in high-dimensional settings.

We revisit the 401(k) simulation set-up in WZ to assess if DDML with stacking suffers from similar issues in small samples and to compare the performance of DDML paired with stacking with PDS lasso and OLS. Following WZ, we run simulations on bootstrap samples of the data for $n_b = \{200, 400, 800, 1600\}$ and approximate the bias as the mean difference relative to the full-sample estimates ($n = 9915$).¹⁵ WZ consider two sets of controls: two-way interactions (TWI) and quadratic splines with interactions (QSI) (as in Belloni et al. 2017). The number of predictors is 167 and 272, respectively. Figure 4 replicates the main results of WZ (Figure 8 in their paper). Figure 4a,b shows the bias relative to

the full sample estimate for the TWI and QSI specification based on OLS and PDS lasso with tuning parameter equal to the plug-in penalty of Belloni, Chernozhukov, and Hansen (2014) scaled by c for $c \in \{0.5, 1, 1.5\}$. It is noteworthy that the speed at which the bootstrapped estimates converge to the full-sample estimate depends on the set of controls for the PDS lasso, but less so for OLS. While PDS lasso with $c = \{0.5, 1\}$ and OLS perform similarly if QSI controls are used, PDS lasso converges much more slowly to the full-sample estimate with TWI controls.

The DDML-stacking framework allows us to choose between, and combine, OLS and lasso with both the TWI and QSI set of controls. Another advantage of DDML over PDS lasso is that we can leverage lasso with CV penalization for a fully data-driven penalization approach. Table 3 compares the performance of the full-sample estimators OLS and PDS lasso (shown in Panel A) to DDML-stacking estimators only relying on OLS and CV lasso with TWI and QSI controls as candidate learners (Panel B). We again consider conventional stacking, short-stacking and pooled stacking together with either CLS or single-best as the final learner. We set the number of cross-fitting folds to $K = 10$ (but also consider $K = 2$ below for comparison in Panel E).

Across all sample sizes, the DDML-stacking estimators strictly outperform both OLS specifications, as well as PDS lasso with TWI, and exhibit overall similar performance to PDS lasso utilizing QSI controls and $c = \{0.5, 1\}$. The differences across DDML-stacking estimators are relatively minor. The CLS short-stacking weights reported in Table 4, Panels A and B reveal that CV-lasso with QSI controls receives the largest weights, while both OLS specifications contribute jointly between nearly zero (at $n_b = 200$) and only up to 15% (for $n_b = 1600$ and the estimation of $E[D|X]$). When selecting only a single candidate learner, CV-lasso with QSI is chosen in more than three-fourths of bootstrap iterations for the estimation of $E[Y|X]$ and $E[D|X]$ (Panels C and D in Table 4), suggesting that CV-lasso with QSI controls is strictly preferable over OLS and lasso with TWI controls in this application. This simulation exercise again highlights that relying on poorly chosen specifications that are not validated against other choices might be suboptimal. In practice, the researcher does not know whether TWI or QSI controls perform better and whether to use OLS or lasso. Crucially, DDML paired with stacking allows for simultaneous consideration of OLS and lasso with both TWI and QSI controls and thus resolves the choice between learners and control specifications in a data-driven manner.

In the next step, we expand the set of candidate learners by two types of random forests, two types of gradient-boosted trees and a feed-forward neural net. In principle, widening the set of candidate learners increases robustness to a larger class of unknown confounding structures. We show the results in Panel C of Table 3. When measuring performance based on the difference to the full-sample estimates, we find there are benefits of extending the set of candidate learners for bootstrap sample sizes of $n_b = 800$ or larger. The results are generally comparable across conventional, short and pooled stacking. However, single-best exhibits a lower bias for small bootstrap sample sizes versus CLS, while pooled stacking with CLS appears to perform worse. The CLS weights reported in Table S2.2 illustrate how DDML-stacking

estimators adapt to the sample size. For example, for smaller sample sizes, a larger weight is put on OLS in the estimation of $E[Y|X]$. In Panel D of Table 3, we report results for each candidate learner individually. DDML-stacking approaches perform better than most individual candidate learners and similar to the best-performing individual learner, which is DDML with CV lasso and QSI controls. In Panel E, we also show results if we reduce the number of folds to $K = 2$. The performance deteriorates drastically for smaller sample sizes, indicating that—while DDML stacking appears competitive for small sample sizes—it is important to increase the number of folds to ensure larger training samples for the CEF estimators.

A drawback of measuring the bias as the difference to the full-sample estimate is that we do not gain insights about convergence to the true parameter. We thus revisit the calibrated simulation exercise from Section 4.1, which allows us to measure the bias as the difference to the true parameter. When the DGP is linear (see Figure 5a), DDML with short-stacking or pooled stacking and using CLS performs overall similarly to OLS. DDML with conventional stacking exhibits relatively large bias with $n_b = 200$. If the true DGP is nonlinear, see Figure 5b, OLS and PDS-Lasso are unable to recover the true effect, while DDML with short and pooled stacking and CLS yield reasonably close approximations of the true parameter even for small sample sizes. DDML with conventional stacking is competitive only for larger samples. We provide extensive results for mean bias and coverage rates in Tables S2.3–S2.5.

To conclude, the results highlight the risks of relying on inappropriate functional form assumptions. DDML paired with stacking approaches—when combined with a diverse set of candidate learners—imposes weaker conditions on the underlying data-generating process compared to relying on a single pre-selected learner. Short-stacking and pooled stacking outperform conventional stacking in small samples. We conjecture the improvement is due to short and pooled stacking imposing common weights across cross-fitting folds. The use of a common set of weights imposes regularization that is consistent with learner performance being stable across subsamples, which seems like a natural benchmark. For pooled stacking, this additional regularization should reduce weight variance while coming with relatively little cost in terms of bias. Because short-stacking does not make use of the additional train and test splits from conventional stacking, it has more potential to suffer from an additional over-fitting bias. When a small number of learners is considered, this additional bias should be small relative to the variance reduction obtained from not needing to estimate fold-specific stacking weights. The simulation results provide support in favor of this conjecture.

5 | Applications

In this section, we use two applications to illustrate how pairing DDML and stacking can increase the robustness of structural parameter estimates to the underlying structure of the data. In the first application, we estimate gaps in citations of articles in top economics journals across different gender compositions among the authors. We condition on the abstract to

TABLE 3 | Mean bias relative to full-sample estimates.

	Bootstrap sample size n_b					
	200	400	600	800	1200	1600
Panel A. Full-sample estimators						
OLS QSI	-2083.5	-910.2	-806.4	-809.9	-677.2	-626.5
OLS TWI	-1694.5	-475.4	13.2	-366	-320.3	-91.3
Post double lasso QSI $c = 0.5$	409.2	-308.9	-204	-503.1	-571.6	-354.1
Post double lasso QSI $c = 1$	-179.1	-1113.5	-639.4	-1063.2	-1000.5	-523.5
Post double lasso QSI $c = 1.5$	8021.3	739.9	-1526.2	-2434.4	-2255.4	-1863.5
Post double lasso TWI $c = 0.5$	3611.2	2484.4	2347.2	1748.3	1270.4	1197.5
Post double lasso TWI $c = 1$	6303.3	3501.1	3453.1	2523.9	1702.4	1871.8
Post double lasso TWI $c = 1.5$	14,386.1	8981.9	6317.9	4802.2	3939	3094.5
Panel B. DDML-stacking with only OLS and CV lasso ($K = 10$)						
Short-stacking: CLS	1020	-113.8	-181.1	-538.2	-575.6	-292.4
Short-stacking: single-best	1002.3	-122.2	-270.1	-499.7	-550.3	-197.7
Pooled stacking: CLS	925.7	-237.3	-319.1	-628.1	-711	-370.5
Pooled stacking: single-best	782.3	-200.5	-358.9	-541.2	-580.2	-237.5
Stacking: CLS	1155.8	-254.7	-266.9	-645	-633	-315.1
Stacking: single-best	999.5	-23.6	-184.9	-503.9	-571.1	-248.2
Panel C. DDML-stacking with all candidate learners ($K = 10$)						
Short-stacking: CLS	1355.1	342.2	403.3	34.2	-103.9	43.8
Short-stacking: single-best	669.2	113.5	144.6	-182.3	-272.6	48.9
Pooled stacking: CLS	2849.3	1345.7	1197	383.8	-102.3	-10.6
Pooled stacking: single-best	724.1	-69.4	45	-250.7	-309	-19.4
Stacking: CLS	1394.1	296.9	344.5	2.8	-168.5	56.9
Stacking: single-best	718.4	-47	104.3	-141.5	-318.6	42.5
Panel D. DDML with candidate learners ($K = 10$)						
OLS	963	-150.8	210	-161.7	-235.5	31.8
Lasso with CV (TWI)	5948.6	3223.1	2589.1	1706.2	872.2	734.1
Ridge with CV (TWI)	4137.3	1853.8	1617.5	951.8	657.5	879.2
Lasso with CV (QSI)	297.5	-343.9	-311.9	-551.8	-597.1	-239.8
Ridge with CV (QSI)	426.1	-111	85.3	-240.8	-294.4	-7.8
Random forest (low regularization)	1852.8	618.3	709.6	259.7	7.7	95.5
Random forest (high regularization)	9987.4	4270.1	2940.2	1919.5	1037.8	925
Gradient boosting (low regularization)	772.3	-25	306.3	70.7	-127.2	113
Gradient boosting (high regularization)	1060.8	94.3	564.6	292.5	44.2	228.6
Neural net	8892.3	7481.2	6915.4	5653.2	3716.5	2224.2
Panel E. DDML-stacking with all candidate learners ($K = 2$)						
Short-stacking: CLS	1842.3	1078.3	-144.4	61.2	446.7	282.9
Short-stacking: single-best	1303.5	582.3	-436.4	-248.8	194	111.1
Pooled stacking: CLS	2799	1471.3	159.5	209.8	572.7	508.8
Pooled stacking: single-best	1791.9	622.9	-542.3	-296.3	144.7	84.8
Stacking: CLS	1924.6	1196.1	-191.2	59.4	390.3	310.9
Stacking: single-best	1173.4	549.6	-604.2	-285	181.8	138.3

Note: The table reports the mean bias calculated as the mean difference to the full sample estimates. Following WZ, we draw 1000 bootstrap samples of size n_b . In Panel A, we show results for the full-sample estimators OLS and PDS lasso using either two-way interactions as controls (denoted TWI) or the quadratic spline and interactions specification of Belloni et al. (2017), denoted as QSI. We scale the PDS lasso penalty by $c = 0.5, 1, \text{ or } 1.5$. In Panel B, we report results for DDML with stacking approaches and only relying on OLS and CV lasso. In Panel C, we consider a larger set of candidate learners. These are: OLS, CV lasso and CV ridge with either TWI or QSI controls, random forest with low regularization (8 predictors considered at each leaf split, no limit on the number of observations per node, bootstrap sample size of 70%) or high regularization (5 splitting predictors, at least 10 observation per node, bootstrap sample size of 70%), gradient-boosted tree with either low (500 trees, learning rate of 0.01, maximum depth of 3) or high (250 trees, learning rate of 0.01, maximum depth of 3) regularization, and a neural net with three hidden layers of size 5. Panel D shows results for these individual candidate learners. In Panels B–D, we use $K = 10$ cross-fitting folds and $R = 5$ cross-fitting repetitions. Panel D uses the same specifications as Panel C but uses $K = 2$.

TABLE 4 | Short-stacking weights.

Estimator	Observations						
	200	400	600	800	1200	1600	9915
Panel A. Constrained least squares. $E[Y X]$, $K = 10$							
OLS (TWI)	0.01	0.042	0.062	0.078	0.098	0.113	0.013
OLS (QSI)	0	0	0.002	0.008	0.023	0.032	0.128
Lasso with CV (TWI)	0.249	0.2	0.196	0.171	0.158	0.14	0.214
Lasso with CV (QSI)	0.74	0.758	0.74	0.742	0.721	0.716	0.645
Panel B. Constrained least squares. $E[D X]$, $K = 10$							
OLS (TWI)	0.005	0.037	0.055	0.074	0.1	0.127	0.13
OLS (QSI)	0	0	0.001	0.003	0.011	0.022	0.134
Lasso with CV (TWI)	0.264	0.163	0.137	0.119	0.114	0.111	0.232
Lasso with CV (QSI)	0.731	0.8	0.807	0.803	0.775	0.74	0.504
Panel C. Single-best. $E[Y X]$, $K = 10$							
OLS (TWI)	0	0	0	0	0.001	0	0
OLS (QSI)	0	0	0	0	0	0	0
Lasso with CV (TWI)	0.186	0.141	0.128	0.112	0.09	0.081	0
Lasso with CV (QSI)	0.814	0.859	0.872	0.888	0.909	0.919	1
Panel D. Single-best. $E[D X]$, $K = 10$							
OLS (TWI)	0	0	0	0	0	0	0
OLS (QSI)	0	0	0	0	0	0	0
Lasso with CV (TWI)	0.239	0.126	0.098	0.079	0.06	0.068	0.003
Lasso with CV (QSI)	0.761	0.874	0.902	0.921	0.94	0.932	0.997

Note: The table reports the stacking weights corresponding to the DDML short-stacking estimators in Figure 3. Panels A and B use constrained least squares. Panels C and D rely on the single-best final learner. Panels A and C refer to the estimation of $E[Y|X]$; Panels C and D refer to the estimation of $E[D|X]$. See notes below Table 3 for more information.

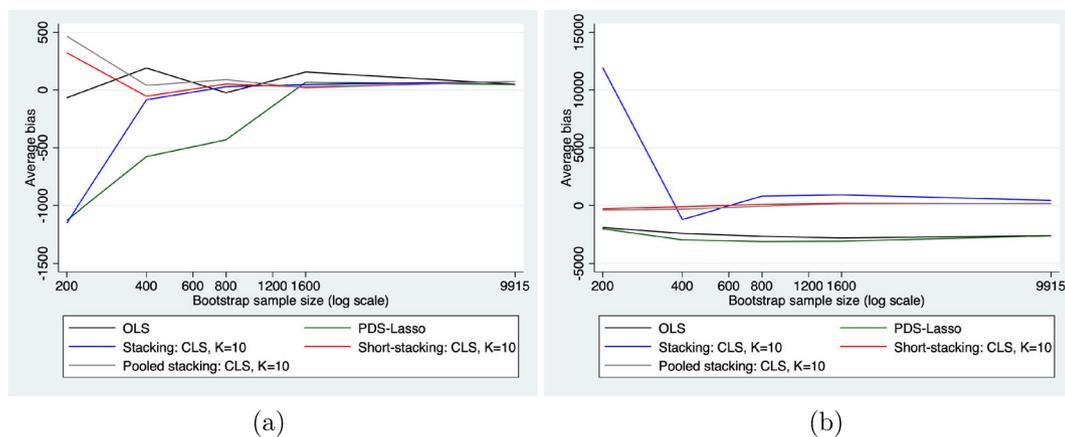


FIGURE 5 | Mean bias for very small sample sizes. The figure shows results from the calibrated simulation in Table 1, but with smaller bootstrap sample sizes. See table notes in Table 1 for more information. Full results for bias and coverage in small samples can be found in Tables S2.3, S2.4, and S2.5. (a) Linear DGP. (b) Nonlinear DGP.

proxy for the content and quality of the paper and demonstrate that stacking-based DDML is a practical solution to challenging estimation problems using text data. In the second application, we revisit the UK sample of the OECD Skills Survey for Kitagawa-Oaxaca-Blinder-type estimates of the unexplained gender wage gap where we condition on a large set of individual characteristics. Both applications pertain to the

literature on gender gaps in various domains, for example, entry to STEM programs (Card and Payne 2021), ICT literacy (Siddiq and Scherer 2019), or wages (Strittmatter and Wunsch 2021; Bonaccolto-Töpfer and Briel 2022) and are methodologically also closely related to the broader literature on discriminatory attitudes towards minority groups (e.g., Hangartner, Kopp, and Siegenthaler 2021).

5.1 | Gender Gap in Citations

This section uses DDML with stacking to estimate a partially linear model applied to average differences in citations of articles published in top-30 economic journals from 1983 to 2020 by the gender composition of the authors. Following Card et al. (2020), we distinguish between papers with (imputed) all-male, all-female, and mixed-gender authorship.¹⁶ Instead of conditioning on hand-coded characteristics such as JEL codes, we leverage the abstract text as a proxy for the topic and quality of the article. Estimating these conditional differences is a challenging statistical problem due to the non-standard nature of text data, and researchers are faced with two key decisions when operationalizing an estimator using text data: how to encode the text data into numerical features, and how to select a suitable learner given the encoded data. Both decisions are ex-ante challenging, but also practically highly relevant as text data is becoming increasingly encountered in economic applications (e.g., Gentzkow and Shapiro 2010; Ash, Chen, and Ornaghi 2024; Widmer, Galletta, and Ash 2023). We show that these decisions can be consequential and that by simultaneously considering different encoding procedures and multiple learners, DDML with stacking provides a simple practical solution to both problems.

In documenting average differences in citations, the analysis presented also contributes to the broader literature on gender biases in academia (e.g., Lundberg and Stearns 2019; Card et al. 2020; Hengel 2022). It is well documented that women are under-represented in academia, especially in senior positions (Ceci et al. 2014; Lundberg and Stearns 2019). A possible reason for the persistent gap in representation include is that scholarly work produced by women faces more skeptical scrutiny compared to work produced by their male counterparts (Hengel 2022; Krawczyk and Smyk 2016). Higher scrutiny could be, for example, reflected at the refereeing stage when a publication decision is made and, as we examine here, after publication when scholarly work is attributed by other scholars through citations (Card et al. 2020; Roberts, Stewart, and Nielsen 2020; Grossbard, Yilmazer, and Zhang 2021).

Throughout our analysis, we focus on a descriptive characterization of the average gaps in citations across different gender compositions of the authors as given by θ_0 in the partially linear model of Equation (1) where Y denotes log-citations and D is a two-dimensional vector whose first component is an indicator for all-female authorship and whose second component is an indicator for mixed-gender authorship. The vector X collects the content of the abstract and a set of year-of-publication indicators. The two components of θ_0 may thus be interpreted as summarizing the average relative difference in total citations between all-male and all-female authorship, and all-male and mixed-gender authorship, respectively, conditional on the article's year of publication and abstract. Throughout, we make no conditional unconfoundedness assumptions that would be necessary for causal interpretations.

We consider a sample of 27,599 articles that have been published between 1983 and 2020. The data were sourced from Scopus and is a subsample of the data analyzed in Advani et al. (2021), who kindly shared their data with us. For each article, we have a record

of the citation count and the authors' names, which we use to infer the authors' gender.¹⁷ In the sample, 6.3% of articles are authored by only female authors, and 22.9% have authors from both genders.

Before turning to estimation, the text of the abstract needs to be transformed into a numerical vector. To admit estimation conditional on the content of the abstract, it is necessary to find a representation (referred to as an embedding) of the text that is lower-dimensional but captures its core meaning. An active literature in statistics and computer science provides solutions to this problem, suggesting a large variety of algorithms to construct text embeddings (see the overview in Ash and Hansen 2023). Thus, in addition to the choice of candidate learner, researchers intent on using text data for their analysis are faced with the additional choice of embedding algorithm. To illustrate how stacking-based DDML can help support this choice, we consider two procedures for encoding the text of the abstract into numerical features: First, we consider a bag-of-words model summarizing the text as (stemmed) word counts (as used in, e.g., Enke 2020; Esposito et al. 2023). In our data, this results in a 211-dimensional vector of word counts for each abstract. Second, since the bag-of-words approach disregards the word order and context, we construct word embeddings generated by a pre-trained BERT model, a transformer-based large-language model (Devlin et al. 2018). In particular, for each abstract, we extract the 768-dimensional vector of weights from the last hidden layer of the BERT model that was pre-trained on a large corpus of (uncased) English text data.¹⁸ Instead of embedding individual words, BERT attempts to reconstruct both whole sentences and the context of these sentences, making it particularly suitable to characterize the content of the abstracts. Recently, Bajari et al. (2023) use BERT to construct embeddings of product descriptions on Amazon.com.

The numerical abstract embeddings are then used in several base learners. We consider OLS, PDS lasso and DDML with CV lasso, CV ridge, `XGBOOST` (Chen and Guestrin 2016), random forests and a feed-forward neural net (see table notes for details).¹⁹ The base learners are aggregated by pairing DDML with either conventional stacking or short-stacking, and with either CLS or single-best.²⁰ The final estimator thus simultaneously combines both text embedding algorithms and machine learning algorithms.

Figure 6 shows estimates of the average relative difference in total citations between all-male and all-female authorship (top-panel) and all-male and mixed-gender authorship (bottom-panel), respectively, for different control specifications and estimators. When we only condition on the publication year, the citation penalty for all-female authorship is close to zero, while there is a large positive effect of +22.6% ($s.e. = 1.7$) for mixed-gender authorship. We next employ PDS lasso to add the abstract text either in the form of word counts or as BERT features. Using the latter, the citation gap increases to -7.4% (2.9) for articles with all-female authorship, while the average relative difference of articles with mixed-gender authorship reduces to +9.2% (1.7). The estimates are qualitatively similar when using word counts instead of BERT features.

In the figure, we also show five cross-fitting repetitions of pairing DDML with each candidate learner (in green) and the median

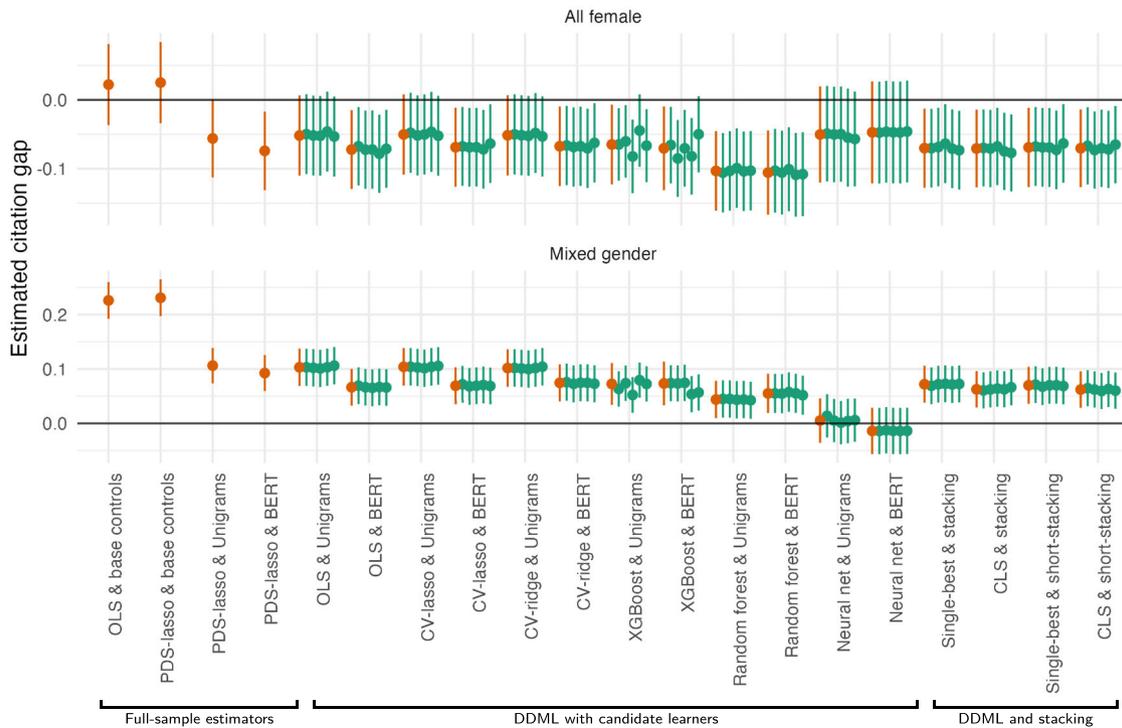


FIGURE 6 | The citation gap by authors' gender composition. The figure shows estimates of θ_0 summarizing average relative difference in total citations between all-male and all-female authorship, and all-male and mixed-gender authorship, respectively, conditional on the article's year of publication and abstract. Error-bars show heteroskedasticity-robust 95% confidence intervals. The sample mean (and standard deviation) of citation counts, all female and mixed gender are 99.8 (254.0), 0.062 (0.242), and 0.229 (0.420), respectively. We consider the following estimators: OLS, PDS lasso, and DDML with the following candidate learners: OLS, CV ridge, CV lasso, XGBoost (using 500 trees, learning rate of 0.3, maximum depth of 6), random forest (using 500 trees), and feed-forward neural net (early stopping with 15 rounds, 0.5 dropout, 0.1 learning rate, 0.1 validation split, 50 epochs, 500 batch size, and 3 hidden layers of size 10). Finally, we pair DDML with either conventional stacking or short-stacking and with either CLS or single-best as the final learner based on the above candidate learners. Throughout, we use five cross-fitting repetitions, five cross-validation folds, and five cross-fitting folds. Results from each cross-fitting replication are illustrated in green, and median aggregates across the cross-fitting replications are shown in orange. The sample includes 29,185 articles published between 1983 and 2020 in top-30 economics journals. A tabular version is provided in Table S3.1. Authors that could not be assigned a gender with less than 70% probability are excluded from the analysis, but we show results based on thresholds of 60% and 90% in Figure S3.1.

aggregates over these repetitions (in orange). There are considerable differences across DDML estimators, with the median estimates of the citation gap ranging between -4.7% (3.8) and -10.6 (3.1) for articles with all-female authorship and between -1.4 (2.2) and $+10.4$ (1.8) for articles with mixed-gender authorship, highlighting that different candidate learner specifications can yield vastly different effect sizes. These stark differences emphasize the need to choose and tune CEF estimators carefully. Without thoroughly validating each candidate learner, judging which results are more credible is difficult. Furthermore, it is noteworthy that some candidate learners, specifically those based on XGBoost in this example, exhibit substantial instability across cross-fitting repetitions.

We show results from pairing DDML and stacking approaches on the right-hand side of the same figure. Relative to the DDML estimates based on the individual candidate learners, the stacking approaches yield lower variability over cross-fitting repetitions. All four stacking-based approaches agree on an average relative difference in citations of around -7.0% (2.9) for articles with all-female authorship and suggest a citation advantage of

between $+6.2$ (1.7) and $+7.0$ (1.7) for articles with mixed-gender authorship.

Table 5 shows the stacking weights of conventional and short-stacking with CLS as the final learner along with mean-squared prediction errors. The stacking estimators assign small weights to learners exhibiting a relatively large MSPE and large variability over cross-fitting repetitions. For example, in the CEF estimation of log citations, the neural nets have an MSPE that is around 50% larger than that of other learners. Stacking assigns, as desired, zero weights to the neural nets, whereas OLS leveraging BERT as one of the best-performing learners receives the largest weights. It is noteworthy that the stacking weights often vary markedly across CEFs, highlighting that there is no reason to assume that the same learner is best suited for estimating both $E[Y|X]$ and $E[D|X]$. This insight is especially important since most estimation approaches (including OLS and PDS lasso) impose the same structure for each CEF.

The results on the citation gaps in top economic journals conditional on the content of the abstract are consistent with a citation penalty for all-female authored articles, possibly due to a higher

TABLE 5 | Stacking weights in the gender citation gap application.

	Citations		All female		Mixed gender	
	Conv.	Short	Conv.	Short	Conv.	Short
Panel A. Stacking and short-stacking weights						
OLS and unigrams	0.1	0.069	0.113	0.155	0.135	0.114
OLS and BERT	0.35	0.368	0.09	0.106	0.174	0.196
CV-lasso and unigrams	0	0	0.049	0.013	0	0
CV-lasso and BERT	0.119	0.102	0.363	0.396	0.129	0.166
CV-ridge and unigrams	0	0	0	0	0	0
CV-ridge and BERT	0	0	0.361	0.31	0.383	0.329
XGBoost and unigrams	0.217	0.236	0.008	0.008	0.017	0.028
XGBoost and BERT	0.171	0.174	0.016	0.01	0.033	0.035
Random forest and unigrams	0.047	0.055	0.02	0.026	0.149	0.151
Random forest and BERT	0	0	0.001	0	0	0
Neural net and unigrams	0	0	0	0	0	0
Neural net and BERT	0	0	0	0	0	0
Panel B. Mean-squared prediction error						
OLS and unigrams		1.341		0.058		0.166
OLS and BERT		1.294		0.059		0.166
CV-lasso and unigrams		1.339		0.058		0.165
CV-lasso and BERT		1.274		0.057		0.161
CV-ridge and unigrams		1.341		0.058		0.165
CV-ridge and BERT		1.286		0.057		0.161
XGBoost and unigrams		1.439		0.075		0.196
XGBoost and BERT		1.472		0.068		0.191
Random forest and unigrams		1.358		0.059		0.167
Random forest and BERT		1.52		0.059		0.169
Neural net and unigrams		2.082		0.059		0.177
Neural net and BERT		2.207		0.059		0.177

Note: Panel A shows stacking weights for conventional stacking (labelled “Conv.”) and short-stacking (labelled “Short”) by candidate learners and by variable. Panel B reports the mean-squared prediction error. The final learner is constrained least squares. The stacking weights and mean-squared prediction errors are averaged over cross-fitting repetitions. Treatment variables are an indicator for all-female authors and mixed-gender authors.

degree of skepticism towards all-female author teams compared to all-male author teams. However, similar to Card et al. (2020) and Maddi and Gingras (2021), the estimates also suggest a conditional citation advantage of articles with mixed-gender authorship.

5.2 | Gender Gap in Wages

The gap in wages between men and women is a central measure of economic gender equality and has been the focus of an extensive empirical literature (see, e.g., the review in Blau and Kahn 2017). The classic approach to estimating the unexplained gender wage gap relies on a linear version of the Kitagawa-Oaxaca-Blinder decomposition (Kitagawa 1955; Oaxaca 1973; Blinder 1973; for an overview, see Fortin, Lemieux, and Firpo 2011). Several recent articles by Bonaccolto-Töpfer and Briel (2022), Strittmatter and Wunsch (2021), Böheim and Stöllinger (2021), and Bach, Chernozhukov, and Spindler (2023),

among others, focus instead on semi-parametric decompositions of the wage gap leveraging more flexible machine learning algorithms. Much of this literature focuses, however, on lasso-based approaches, even though there is no apparent reason to favor sparsity-based approaches over learners relying on other regularization assumptions. In contrast to the recent literature that primarily focuses on lasso-based approaches, we consider a diverse set of candidate learners and aggregate them via stacking.

The parameter of interest in this application is the unexplained gender wage gap, which is the expected difference in wages after conditioning on observed characteristics. Formally,

$$\theta_0 \equiv E[E[Y|D = 1, X] - E[Y|D = 0, X]|D = 1],$$

where Y denotes the logarithm of wages, D is an indicator equal to one for women, and X is a vector of potentially many individual characteristics. The parameter is well-defined if $P(D = 1|X) > 0$ with probability 1.²¹

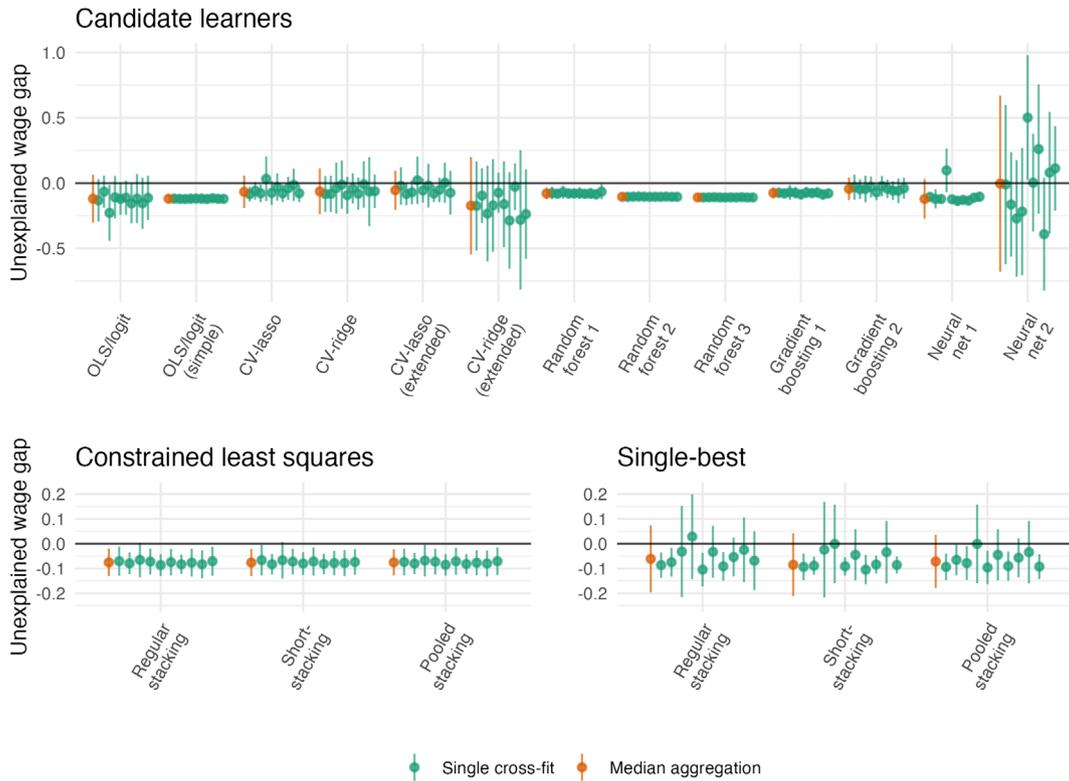


FIGURE 7 | Unexplained gender wage gap. The figure reports DDML estimates of the unexplained gender wage gap based on several different learners. 95% heteroskedasticity-robust confidence intervals are shown. The sample mean (and standard deviation) of log earnings and gender are 2.768 (0.579) and 0.583 (0.493), respectively. The candidate learners are OLS (for the outcome equation) and logit (for the propensity scores) with the base and simple set of controls; CV-lasso and CV-ridge with the base and extended set of controls; three random forests with 500 trees and minimum leaf sizes of 1, 50, and 100; two gradient-boosted trees with and without early stopping (500 trees, with and without early stopping after 10 iterations, maximum depth of 3); two feed-forward neural nets with hidden layer sizes of (40,20, 1,20, 50) and (30,30, 30) and early stopping. We report results for the individual base learners in the top panel. In the bottom panels, we show DDML paired with conventional, short and pooled stacking based on the base learners and with either CLS (left panel) or single-best (right panel) as the final learner. We use 10 cross-fitting folds and 10 cross-fitting repetitions. Results from each cross-fitting replication are illustrated in green, and median aggregates across the cross-fitting replications are shown in orange. A tabular version is provided in Tables S4.4 and S4.5.

In the absence of functional form assumptions, estimation of θ_0 is a challenging statistical problem due to its dependence on unknown conditional expectation functions that need to be non-parametrically estimated. Analogous to the DDML estimator for the partially linear model outlined in Section 2, we consider estimation of θ_0 via the split-sample analog of the efficient score function for θ_0 ²²—that is,

$$\hat{\theta}_n = \frac{1}{n} \sum_{i=1}^n \left(\frac{D_i(Y_i - \hat{g}_{k_i}^c(0, X_i))}{\hat{p}_{k_i}^c} - \frac{\hat{m}_{k_i}^c(X_i)(1 - D_i)(Y_i - \hat{g}_{k_i}^c(0, X_i))}{\hat{p}_{k_i}^c(1 - \hat{m}_{k_i}^c(X_i))} \right) / \frac{1}{n} \sum_{i=1}^n \frac{D_i}{\hat{p}_{k_i}^c},$$

where $\hat{g}_{k_i}^c$ and $\hat{m}_{k_i}^c$ are cross-fitted estimators for $g_0(D, X) \equiv E[Y|D, X]$ and $m_0(X) \equiv E[D|X]$ and $\hat{p}_{k_i}^c$ is a cross-fitted estimator of $P(D = 1)$.

Following Forshaw et al. (2024), we take the data for this application from the UK sample of the OECD Skills Survey, which was collected in 2011-12 and comes with a rich set of covariates, including age, experience, education, occupation, and industry. The final data includes 4836 British respondents. We specify three sets of control variables. The *simple* set of controls only includes a selection of essential covariates: age (in levels and squared), years of education, a literacy and numeracy test score, years of

tenure in the current job (in levels and squared), education level, hours worked per week, and number of children. The *base* set of controls adds, among others, management level, age of children, and parents' education level.²³ Furthermore, we interact age and tenure with all categorical covariates. The *extended* set of controls comprises all variables and interacts each continuous covariate with each categorical covariate.

We include a diverse set of candidate learners to allow for a high level of flexibility. We employ regression approaches for the estimation of the CEF of log wages (i.e., $E[Y|D, X]$) and classification approaches for the CEF estimation of gender (i.e., $E[D|X]$). Our candidate learners are linear (or logistic) regression with the simple and base set of controls; linear (or logistic) CV-lasso and CV-ridge with the base and extended set of controls; three random forests with 500 regression (or classification) trees and minimum leaf sizes of 1, 50 and 100; two types of gradient-boosted regression (or classification) trees with and without early stopping (500 trees, maximum depth of 3); two feed-forward neural nets with hidden layer sizes of (40, 20, 1, 20, 50) and (30, 30, 30) and early stopping. Finally, we aggregate the candidate learners via conventional, short and pooled stacking, using either CLS or single-best as the final learner.

Figure 7 reports results for individual candidate learners (on the top) and stacking approaches (on the bottom). We show results from 10 cross-fitting repetitions (in green) and the median aggregates (in orange). We again find that some candidate learners exhibit substantial variability over cross-fitting repetitions, which is also reflected in the large median-aggregate standard errors. The variability is especially large for CV-ridge with the extended set of controls and the neural nets, which are the candidate learners exhibiting the largest MSE (see Table S4.1). The stacking results are, in contrast, relatively stable over cross-fit repetitions when using CLS as the final learner. Interestingly, the results when relying on single-best as the final learner are noticeably more variable than when using CLS, indicating that a combination of candidate learners seems to better fit the data than a single learner. The stacking weights and MSE in Table S4.1 confirm that there is no single candidate learner dominating the others. The instability of the single-best final learner is reflected in the stacking standard errors. Given this potential for instability of choosing a single candidate learner, we recommend favoring CLS over single-best if one is not confident that one of the chosen learners will be significantly better than the rest, and thus stably selected, which seems likely to be the most common setting in practice.

6 | Conclusion

This article assesses the performance of DDML estimators in realistic settings using applications and simulation studies calibrated to real economic data. We highlight that estimators of structural parameters based on single pre-selected (machine) learners can be highly sensitive to the underlying structure of the data and/or poor tuning, and we show that pairing DDML with stacking can help alleviate these concerns, provided that a sufficiently diverse set of candidate learners is considered.

We discuss pairing DDML with conventional stacking but also suggest two alternative stacking approaches: Short-stacking, which substantially reduces the computational burden by leveraging the cross-fitting naturally arising in the computation of DDML estimates, and pooled stacking, which decreases the variance of the stacking estimator by imposing common stacking weights over cross-fitting folds. In our simulations, both strategies are competitive with conventional stacking in settings with large and moderate sample sizes and are better in small samples. The advantages of short-stacking are particularly worth highlighting, given its substantially lower computational cost.

A key advantage of the DDML-stacking approach is that it accommodates both traditional parametric and nonparametric specifications by allowing simultaneous consideration of, for example, OLS with several sets of controls, sparsity-based learners, tree-based ensembles and neural networks. In this sense, researchers are not forcibly deviating from standard (often linear) specifications unless the data suggests there is reason to. While machine-learning-based causal methods may yield fundamentally different results from linear regression only in specific examples, the additional robustness to unexpected structures in the data thus seems to come at relatively little cost.

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Conflicts of Interest

The authors declare no conflicts of interest.

Open Research Badges



This article has been awarded Open Data Badge for making publicly available the digitally-shareable data necessary to reproduce the reported results. Data is available at <https://doi.org/10.15456/jae.2024286.1440194347>.

Data Availability Statement

The authors provide replication code through the *Journal of Applied Econometrics* Data Archive and share data for all examples with the exception of the application in Section 5.1. The data that support the findings of this application are available on request from the corresponding author. The data are not publicly available due to privacy or ethical restrictions.

Endnotes

- ¹ See also Angrist (2022) for additional discussion.
- ² Hastie, Tibshirani, and Friedman (2009) and van der Laan and Rose (2011) provide textbook treatments of stacking and super learning. See also Hansen and Racine (2012) for discussion of jackknife (leave-one-out) stacking. Clyde and Iversen (2013) and Le and Clarke (2017) provide a Bayesian interpretation of stacking in the setting where the true model is not among the candidates.
- ³ A key example not explicitly discussed in Chernozhukov et al. (2018) is doubly-robust estimation of difference-in-difference parameters with staggered treatment assignment as in Callaway and Sant’Anna (2021) and Chang (2020). In settings with conditional parallel trends assumptions, high-dimensional nuisance functions arise in the estimation of group-time specific average treatment effect on the treated. The pairing of DDML and stacking, as proposed in this paper, also directly applies to the estimator of Callaway and Sant’Anna (2021) under a conditional unconfoundedness assumption.
- ⁴ Similarly, for continuous D , θ_0 corresponds to a positively weighted average of derivatives of the conditional expectation function $E[Y|D = d, X]$ with respect to d . Under a conditional unconfoundedness assumption, θ_0 is thus a convex combination of derivatives of the causal response function.
- ⁵ In the context of IV estimation where instrument validity relies on controlling for observed confounders, Blandhol et al. (2022) emphasize that, in the absence of strong functional form assumptions, two stage least squares does not generally correspond to a convex combination of local average treatment effects (LATE). We note that the IV analog to the partially linear model does admit a causal interpretation under the

LATE assumptions, just as the partially linear model admits a weakly causal interpretation under conditional unconfoundedness.

⁶The exact convergence rate requirement for nonparametric estimators depends on the parameter of interest. Chernozhukov et al. (2018) name the crude rate requirement of $o(n^{-1/4})$ but provide examples where the rate requirement is considerably weaker. Recent contributions show that these requirements are satisfied by specific instances of machine learners; see, for example, results for lasso (Bickel, Ritov, and Tsybakov 2009; Belloni et al. 2012), random forests (Wager and Walther 2016; Wager and Athey 2018; Athey, Tibshirani, and Wager 2019), neural networks (Schmidt-Hieber 2020; Farrell, Liang, and Misra 2021), and boosting (Luo, Spindler, and Kück 2022). The exact asymptotic properties of many other machine learners remain an active research area.

⁷In the context of the partially linear model, the formal Neyman-orthogonality requirement is

$$0 = \frac{\partial}{\partial \lambda} E[(Y - \{\ell_0(X) + \lambda(\ell(X) - \ell_0(X))\}) - \tau_0(D - \{m_0(X) + \lambda(m(X) - m_0(X))\})] \Big|_{\lambda=0} \\ \times (D - \{m_0(X) + \lambda(m(X) - m_0(X))\}) \Big|_{\lambda=0}$$

for arbitrary measurable functions ℓ and m , which can easily be verified using properties of the residuals.

⁸The *scikit-learn* (Buitinck et al. 2013) routines `StackingRegressor` and `StackingClassifier` implement stacking for Python. In Stata, stacking regression and classification are available via `pystacked`, which is a Stata front-end for these Python routines (Ahrens, Hansen, and Schaffer 2023).

⁹Suppose, for simplicity, we consider ordinary (unconstrained) least squares as the final learner. Heuristically, the regression of Y_i against J sets of cross-fitted predicted values is akin to a conventional least squares regression of Y_i against J observed regressors where good performance would require $J/n \rightarrow 0$, ignoring that the cross-fitted predicted values are estimated. The additional regularization by CLS should further weaken this rate requirement.

¹⁰Further results are provided in Appendix S1. Table S1.1 gives the mean-squared prediction errors (MSPE) for each candidate learner for comparison. Table S1.2 reports standard errors of the bias. Tables S1.3 and S1.4 show the stacking weights when using single-best and OLS as the final learner, respectively.

¹¹The computations were performed on the high-performance cluster of the ETH Zurich. Each instance used a single core of an AMD EPYC processor with 2.25–2.6 GHz (nominal)/3.3–3.5 GHz (peak) and 4 GB RAM. The run time of DDML with conventional stacking was 2393 s on the full sample, while short-stacking ran in only 540 s.

¹²Table S1.3 shows that the OLS stacking weights are often outside the unit interval. The weights associated with the neural net are particularly large (in absolute value), suggesting that OLS might be more sensitive to outliers than CLS.

¹³By contrast, the unweighted average is known to often perform well in time-series settings (e.g., Clemen 1989; Timmermann 2006) and, in particular, when the optimal weights are close to equal (see Wang et al. 2023, Section 2.6, for a summary and wider discussion). Such a scenario is ruled out in our simulation setup, which captures the situation where the researcher does not know whether a linear or nonlinear learner would be more appropriate.

¹⁴The rates at which each candidate learner is selected by the single-best final learner are shown in Table S1.4 and provide similar insights.

¹⁵The full-sample estimates are reported in Table S2.1.

¹⁶As we explain below, we impute the gender mix of authors from the authors' names.

¹⁷We use the software *Namsor*, which frequently ranks among the best-performing algorithms for gender classification using names

(Sebo 2021; Krstovski, Lu, and Xu 2023) and is widely used in academic studies (e.g., Bursztyjn et al. 2021; Sebo and Schwarz 2023). In the main specification, we exclude articles of authors whose gender could not be classified with a probability of less than 70%, but we show that results are similar when we apply thresholds of 60% or 90%; see Figure S3.1. Our sample includes 586 articles for which no citation is recorded. These were excluded from the analysis. We also provide results using the number of citations (instead of log-citations) in Table S3.1.

¹⁸The model `bert-base-uncased` is freely available from, among others, the Python library `huggingface`.

¹⁹To reduce the run time, we use regression approaches both for the estimation of $E[Y|X]$ and $E[D|X]$.

²⁰We omit pooled stacking from this application since the R package `ddml`, which was used for this application, does not currently support pooled stacking.

²¹As in the previous section, we focus our analysis on a descriptive parameter of interest and do not make conditional unconfoundedness assumptions that would be necessary for causal interpretations.

²²See, for example, equation (5.4) in Chernozhukov et al. (2018).

²³The base set adds the following variables to the reduced set: area of study, part of larger organization, management position, type of contract, job satisfaction, health status, living with a partner, age of youngest child, immigration age, mother's and father's highest level of education, immigration status of parents, informal job-related education in last 12 months, and informal non-job-related education in last 12 months.

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Supporting Information

Additional supporting information can be found online in the Supporting Information section.