

# Unbiased Instrumental Variables Estimation Under Known First-Stage Sign

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## Abstract

We derive mean-unbiased estimators for the structural parameter in instrumental variables models with a single endogenous regressor where the sign of one or more first stage coefficients is known. In the case with a single instrument, the unbiased estimator is unique. For cases with multiple instruments we propose a class of unbiased estimators and show that an estimator within this class is efficient when the instruments are strong. We show numerically that unbiasedness does not come at a cost of increased dispersion in models with a single instrument: in this case the unbiased estimator is less dispersed than the 2SLS estimator. Our finite-sample results apply to normal models with known variance for the reduced-form errors, and imply analogous results under weak instrument asymptotics with an unknown error distribution.

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# 1 Introduction

Researchers often have strong prior beliefs about the sign of the first stage coefficient in instrumental variables models, to the point where the sign can reasonably be treated as known. This paper shows that knowledge of the sign of the first stage coefficient allows us to construct an estimator for the coefficient on the endogenous regressor which is unbiased in finite samples when the reduced form errors are normal with known variance. When the distribution of the reduced form errors is unknown, our results lead to estimators that are asymptotically unbiased under weak IV sequences as defined in Staiger & Stock (1997).

As is well known, the conventional two-stage least squares (2SLS) estimator may be severely biased in overidentified models with weak instruments. While the 2SLS estimator performs better in the just-identified case according to some measures of central tendency, in this case it has no first moment.<sup>1</sup> A number of papers have proposed alternative estimators to reduce particular measures of bias, e.g. Angrist & Krueger (1995), Imbens et al. (1999), Donald & Newey (2001), and Akerberg & Devereux (2009), but none of the resulting estimators is unbiased either in finite samples or under weak instrument asymptotics. Indeed, Hirano & Porter (2015) show that mean, median, and quantile unbiased estimation are all impossible in the linear IV model with an unrestricted parameter space for the first stage.

We show that by exploiting information about the sign of the first stage we can circumvent this impossibility result and construct an unbiased estimator. Moreover, the resulting estimators have a number of properties which make them appealing for applications. In models with a single instrumental variable, which include many empirical applications, we show that there is a unique unbiased estimator. Moreover, we show that this estimator is substantially less dispersed than the usual 2SLS estimator in finite samples. Under standard (“strong instrument”) asymptotics, the unbiased esti-

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<sup>1</sup>If we instead consider median bias, 2SLS exhibits median bias when the instruments are weak, though this bias decreases rapidly with the strength of the instruments.

mator has the same asymptotic distribution as 2SLS, and so is asymptotically efficient in the usual sense. In over-identified models many unbiased estimators exist, and we propose unbiased estimators which are asymptotically efficient when the instruments are strong. Further, we show that in over-identified models we can construct unbiased estimators which are robust to small violations of the first stage sign restriction. We also derive a lower bound on the risk of unbiased estimators in finite samples, and show that this bound is attained in some models.

The rest of this section discusses the assumption of known first stage sign, introduces the setting and notation, and briefly reviews the related literature. Section 2 introduces the unbiased estimator for models with a single excluded instrument. Section 3 treats models with multiple instruments and introduces unbiased estimators which are robust to small violations of the first stage sign restriction. Section 4 presents simulation results on the performance of our unbiased estimators. Section 5 discusses illustrative applications using data from Hornung (2014) and Angrist & Krueger (1991). Proofs and auxiliary results are given in a separate appendix.<sup>2</sup>

## 1.1 Knowledge of the First-Stage Sign

The results in this paper rely on knowledge of the first stage sign. This is reasonable in many economic contexts. In their study of schooling and earnings, for instance, Angrist & Krueger (1991) note that compulsory schooling laws in the United States allow those born earlier in the year to drop out after completing fewer years of school than those born later in the year. Arguing that quarter of birth can reasonably be excluded from a wage equation, they use this fact to motivate quarter of birth as an instrument for schooling. In this context, a sign restriction on the first stage amounts to an assumption that the mechanism claimed by Angrist & Krueger works in the expected direction: those born earlier in the year tend to drop out earlier. More generally, empirical researchers often have some mechanism in mind for why a model

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<sup>2</sup>The appendix is available online at <https://sites.google.com/site/isaiahandrews/working-papers>

is identified at all (i.e. why the first stage coefficient is nonzero) that leads to a known sign for the direction of this mechanism (i.e. the sign of the first stage coefficient).

In settings with heterogeneous treatment effects, a first stage monotonicity assumption is often used to interpret instrumental variables estimates (see Imbens & Angrist 1994, Heckman et al. 2006). In the language of Imbens & Angrist (1994), the monotonicity assumption requires that either the entire population affected by the treatment be composed of “compliers,” or that the entire population affected by the treatment be composed of “defiers.” Once this assumption is made, our assumption that the sign of the first stage coefficient is known amounts to assuming the researcher knows which of these possibilities (compliers or defiers) holds. Indeed, in the examples where they argue that monotonicity is plausible (involving draft lottery numbers in one case and intention to treat in another), Imbens & Angrist (1994) argue that all individuals affected by the treatment are “compliers” for a certain definition of the instrument.

It is important to note, however, that knowledge of the first stage sign is not always a reasonable assumption, and thus that the results of this paper are not always applicable. In settings where the instrumental variables are indicators for groups without a natural ordering, for instance, one typically does not have prior information about signs of the first stage coefficients. To give one example, Aizer & Doyle Jr. (2013) use the fact that judges are randomly assigned to study the effects of prison sentences on recidivism. In this setting, knowledge of the first stage sign would require knowing a priori which judges are more strict.

## 1.2 Setting

For the remainder of the paper, we suppose that we observe a sample of  $T$  observations  $(Y_t, X_t, Z_t')$ ,  $t = 1, \dots, T$  where  $Y_t$  is an outcome variable,  $X_t$  is a scalar endogenous regressor, and  $Z_t$  is a  $k \times 1$  vector of instruments. Let  $Y$  and  $X$  be  $T \times 1$  vectors with row  $t$  equal to  $Y_t$  and  $X_t$  respectively, and let  $Z$  be a  $T \times k$  matrix with row  $t$  equal to  $Z_t'$ . The usual linear IV model, written in reduced-form, is

$$\begin{aligned} Y &= Z\pi\beta + U \\ X &= Z\pi + V \end{aligned} \quad (1)$$

To derive finite-sample results, we treat the instruments  $Z$  as fixed and assume that the errors  $(U, V)$  are jointly normal with mean zero and known variance-covariance matrix  $Var((U', V)')$ .<sup>3</sup> As is standard (see, for example, D. Andrews et al. (2006)), in contexts with additional exogenous regressors  $W$  (for example an intercept), we define  $Y, X, Z$  as the residuals after projecting out these exogenous regressors. If we denote the reduced-form and first-stage regression coefficients by  $\xi_1$  and  $\xi_2$ , respectively, we can see that

$$\begin{pmatrix} \xi_1 \\ \xi_2 \end{pmatrix} = \begin{pmatrix} (Z'Z)^{-1} Z'Y \\ (Z'Z)^{-1} Z'X \end{pmatrix} \sim N \left( \begin{pmatrix} \pi\beta \\ \pi \end{pmatrix}, \begin{pmatrix} \Sigma_{11} & \Sigma_{12} \\ \Sigma_{21} & \Sigma_{22} \end{pmatrix} \right) \quad (2)$$

for

$$\Sigma = \begin{pmatrix} \Sigma_{11} & \Sigma_{12} \\ \Sigma_{21} & \Sigma_{22} \end{pmatrix} = \left( I_2 \otimes (Z'Z)^{-1} Z' \right) Var((U', V)') \left( I_2 \otimes (Z'Z)^{-1} Z' \right)' \quad (3)$$

We assume throughout that  $\Sigma$  is positive definite. Following the literature (e.g. Moreira & Moreira 2013), we consider estimation based solely on  $(\xi_1, \xi_2)$ , which are sufficient for  $(\pi, \beta)$  in the special case where the errors  $(U_t, V_t)$  are iid over  $t$ . All uniqueness and efficiency statements therefore restrict attention to the class of procedures which depend on the data through only these statistics. The conventional generalized method of moments (GMM) estimators belong to this class, so this restriction still allows efficient estimation under strong instruments. We assume that the sign of each component  $\pi_i$  of  $\pi$  is known, and in particular assume that the parameter space for  $(\pi, \beta)$  is

$$\Theta = \left\{ (\pi, \beta) : \pi \in \Pi \subseteq (0, \infty)^k, \beta \in B \right\} \quad (4)$$

for some sets  $\Pi$  and  $B$ . Note that once we take the sign of  $\pi_i$  to be known, assuming  $\pi_i > 0$  is without loss of generality since this can always be ensured by redefining  $Z$ .

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<sup>3</sup>Note that we assume a homogenous  $\beta$ , which will generally rule out heterogenous treatment effect models with multiple instruments.

In this paper we focus on models with fixed instruments, normal errors, and known error covariance, which allows us to obtain finite-sample results. As usual, these finite-sample results will imply asymptotic results under mild regularity conditions. Even in models with random instruments, non-normal errors, serial correlation, heteroskedasticity, clustering, or any combination of these, the reduced-form and first stage estimators will be jointly asymptotically normal with consistently estimable covariance matrix  $\Sigma$  under mild regularity conditions. Consequently, the finite-sample results we develop here will imply asymptotic results under both weak and strong instrument asymptotics, where we simply define  $(\xi_1, \xi_2)$  as above and replace  $\Sigma$  by an estimator for the variance of  $\xi$  to obtain feasible statistics. Appendix B provides the details of these results.<sup>4</sup> In the main text, we focus on what we view as the most novel component of the paper: finite-sample mean-unbiased estimation of  $\beta$  in the normal problem (2).

### 1.3 Related Literature

Our unbiased IV estimators build on results for unbiased estimation of the inverse of a normal mean discussed in Voinov & Nikulin (1993). More broadly, the literature has considered unbiased estimators in numerous other contexts, and we refer the reader to Voinov & Nikulin for details and references. To our knowledge the only other paper to treat finite sample mean-unbiased estimation in IV models is Hirano & Porter (2015), who find that unbiased estimators do not exist when the parameter space is unrestricted. The nonexistence of unbiased estimators has been noted in other nonstandard econometric contexts by Hirano & Porter (2012).

The broader literature on the finite sample properties of IV estimators is huge: see Phillips (1983) and Hillier (2006) for references. While this literature does not study

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<sup>4</sup>The feasible analogs of the finite-sample unbiased estimators discussed here are asymptotically unbiased in general models in the sense of converging in distribution to random variables with mean  $\beta$ . Note, however, that outside the exact normal case it will not in general be true that means of the feasible estimators themselves will converge to  $\beta$  as the sample size increases, since convergence in distribution does not suffice for convergence of moments.

unbiased estimation in finite samples, there has been substantial research on higher order asymptotic bias properties: see the references given in the first section of the introduction, as well as Hahn et al. (2004) and the references therein.

Our interest in finite sample results for a normal model with known reduced form variance is motivated by the weak IV literature, where this model arises asymptotically under weak IV sequences as in Staiger & Stock (1997) (see also Appendix B). In contrast to Staiger & Stock, however, our results allow for heteroskedastic, clustered, or serially correlated errors as in Kleibergen (2007). The primary focus of recent work on weak instruments has, however, been on inference rather than estimation. See Andrews (2014) for additional references.

Sign restrictions have been used in other settings in the econometrics literature, although the focus is often on inference or on using sign restrictions to improve population bounds, rather than estimation. Recent examples include Moon et al. (2013) and several papers cited therein, which use sign restrictions to partially identify vector autoregression models. Inference for sign restricted parameters has been treated by D. Andrews (2001) and Gouriéroux et al. (1982), among others.

## 2 Unbiased Estimation with a Single Instrument

To introduce our unbiased estimators, we first focus on the just-identified model with a single instrument,  $k = 1$ . We show that unbiased estimation of  $\beta$  in this context is linked to unbiased estimation of the inverse of a normal mean. Using this fact we construct an unbiased estimator for  $\beta$ , show that it is unique, and discuss some of its finite-sample properties. We note the key role played by the first stage sign restriction, and show that our estimator is equivalent to 2SLS (and thus efficient) when the instruments are strong.

In the just-identified context  $\xi_1$  and  $\xi_2$  are scalars and we write

$$\Sigma = \begin{pmatrix} \Sigma_{11} & \Sigma_{12} \\ \Sigma_{21} & \Sigma_{22} \end{pmatrix} = \begin{pmatrix} \sigma_1^2 & \sigma_{12} \\ \sigma_{12} & \sigma_2^2 \end{pmatrix}.$$

The problem of estimating  $\beta$  therefore reduces to that of estimating

$$\beta = \frac{\pi\beta}{\pi} = \frac{E[\xi_1]}{E[\xi_2]}. \quad (5)$$

The conventional IV estimate  $\hat{\beta}_{2SLS} = \frac{\xi_1}{\xi_2}$  is the natural sample-analog of (5). As is well-known, however, this estimator has no integer moments. This lack of unbiasedness reflects the fact that the expectation of the ratio of two random variables is not in general equal to the ratio of their expectations.

The form of (5) nonetheless suggests an approach to deriving an unbiased estimator. Suppose we can construct an estimator  $\hat{\tau}$  which (a) is unbiased for  $1/\pi$  and (b) depends on the data only through  $\xi_2$ . If we then define

$$\hat{\delta}(\xi, \Sigma) = \left( \xi_1 - \frac{\sigma_{12}}{\sigma_2^2} \xi_2 \right), \quad (6)$$

we have that  $E[\hat{\delta}] = \pi\beta - \frac{\sigma_{12}}{\sigma_2^2}\pi$ , and  $\hat{\delta}$  is independent of  $\hat{\tau}$ . Thus,  $E[\hat{\tau}\hat{\delta}] = E[\hat{\tau}]E[\hat{\delta}] = \beta - \frac{\sigma_{12}}{\sigma_2^2}$ , and  $\hat{\tau}\hat{\delta} + \frac{\sigma_{12}}{\sigma_2^2}$  will be an unbiased estimator of  $\beta$ . Thus, the problem of unbiased estimation of  $\beta$  reduces to that of unbiased estimation of the inverse of a normal mean.

## 2.1 Unbiased Estimation of the Inverse of a Normal Mean

A result from Voinov & Nikulin (1993) shows that unbiased estimation of  $1/\pi$  is possible if we assume its sign is known. Let  $\Phi$  and  $\phi$  denote the standard normal cdf and pdf respectively.

**Lemma 2.1.** *Define*

$$\hat{\tau}(\xi_2, \sigma_2^2) = \frac{1}{\sigma_2} \frac{1 - \Phi(\xi_2/\sigma_2)}{\phi(\xi_2/\sigma_2)}.$$

*For all  $\pi > 0$ ,  $E_\pi[\hat{\tau}(\xi_2, \sigma_2^2)] = \frac{1}{\pi}$ .*

The derivation of  $\hat{\tau}(\xi_2, \sigma_2^2)$  in Voinov & Nikulin (1993) relies on the theory of bilateral Laplace transforms, and offers little by way of intuition. Verifying unbiasedness is a straightforward calculus exercise, however: for the interested reader, we work through the necessary derivations in the proof of Lemma 2.1.



From the formula for  $\hat{\tau}$ , we can see that this estimator has two properties which are arguably desirable for a restricted estimate of  $1/\pi$ . First, it is positive by definition, thereby incorporating the restriction that  $\pi > 0$ . Second, in the case where positivity of  $\pi$  is obvious from the data ( $\xi_2$  is very large relative to its standard deviation), it is close to the natural plug-in estimator  $1/\xi_2$ . The second property is an immediate consequence of a well-known approximation to the tail of the normal cdf, which is used extensively in the literature on extreme value limit theorems for normal sequences and processes (see Equation 1.5.4 in Leadbetter et al. 1983, and the remainder of that book for applications). We discuss this further in Section 2.5.

## 2.2 Unbiased Estimation of $\beta$

Given an unbiased estimator of  $1/\pi$  which depends only on  $\xi_2$ , we can construct an unbiased estimator of  $\beta$  as suggested above. Moreover, this estimator is unique.

**Theorem 2.1.** *Define*

$$\begin{aligned}\hat{\beta}_U(\xi, \Sigma) &= \hat{\tau}(\xi_2, \sigma_2^2) \hat{\delta}(\xi, \Sigma) + \frac{\sigma_{12}}{\sigma_2^2} \\ &= \frac{1}{\sigma_2} \frac{1 - \Phi(\xi_2/\sigma_2)}{\phi(\xi_2/\sigma_2)} \left( \xi_1 - \frac{\sigma_{12}}{\sigma_2^2} \xi_2 \right) + \frac{\sigma_{12}}{\sigma_2^2}.\end{aligned}$$

The estimator  $\hat{\beta}_U(\xi, \Sigma)$  is unbiased for  $\beta$  provided  $\pi > 0$ .

Moreover, if the parameter space (4) contains an open set then  $\hat{\beta}_U(\xi, \Sigma)$  is the unique non-randomized unbiased estimator for  $\beta$ , in the sense that any other estimator  $\hat{\beta}(\xi, \Sigma)$  satisfying

$$E_{\pi, \beta} \left[ \hat{\beta}(\xi, \Sigma) \right] = \beta \quad \forall \pi \in \Pi, \beta \in B$$

also satisfies

$$\hat{\beta}(\xi, \Sigma) = \hat{\beta}_U(\xi, \Sigma) \quad a.s. \quad \forall \pi \in \Pi, \beta \in B.$$

Note that the conventional IV estimator can be written as

$$\hat{\beta}_{2SLS} = \frac{\xi_1}{\xi_2} = \frac{1}{\xi_2} \left( \xi_1 - \frac{\sigma_{12}}{\sigma_2^2} \xi_2 \right) + \frac{\sigma_{12}}{\sigma_2^2}.$$

Thus,  $\hat{\beta}_U$  differs from the conventional IV estimator only in that it replaces the plug-in estimate  $1/\xi_2$  for  $1/\pi$  by the unbiased estimate  $\hat{\tau}$ . From results in e.g. Baricz (2008), we have that  $\hat{\tau} < 1/\xi_2$  for  $\xi_2 > 0$ , so when  $\xi_2$  is positive  $\hat{\beta}_U$  shrinks the conventional IV estimator towards  $\sigma_{12}/\sigma_2^2$ .<sup>5</sup> By contrast, when  $\xi_2 < 0$ ,  $\hat{\beta}_U$  lies on the opposite side of  $\sigma_{12}/\sigma_2^2$  from the conventional IV estimator. Interestingly, one can show that the unbiased estimator is uniformly more likely to be of the correct sign than is the conventional estimator, in the sense that

$$Pr_{\pi,\beta}\{\text{sign}(\hat{\beta}_U) = \text{sign}(\beta)\} \geq Pr_{\pi,\beta}\{\text{sign}(\hat{\beta}_{2LS}) = \text{sign}(\beta)\},$$

with strict inequality at some points.<sup>6</sup>

### 2.3 Risk and Moments of the Unbiased Estimator

The uniqueness of  $\hat{\beta}_U$  among nonrandomized estimators implies that  $\hat{\beta}_U$  minimizes the risk  $E_{\pi,\beta}\ell\left(\tilde{\beta}(\xi, \Sigma) - \beta\right)$  uniformly over  $\pi, \beta$  and over the class of unbiased estimators  $\tilde{\beta}$  for any loss function  $\ell$  such that randomization cannot reduce risk. In particular, by Jensen's inequality  $\hat{\beta}_U$  is uniformly minimum risk for any convex loss function  $\ell$ . This includes absolute value loss as well as squared error loss or  $L^p$  loss for any  $p \geq 1$ . However, elementary calculations show that  $|\hat{\beta}_U|$  has an infinite  $p$ th moment for  $p > 1$ . Thus the fact that  $\hat{\beta}_U$  has uniformly minimal risk implies that any unbiased estimator must have an infinite  $p$ th moment for any  $p > 1$ . In particular, while  $\hat{\beta}_U$  is the uniform minimum mean absolute deviation unbiased estimator of  $\beta$ , it is minimum variance unbiased only in the sense that all unbiased estimators have infinite variance. We record this result in the following theorem.

**Theorem 2.2.** *For  $\varepsilon > 0$ , the expectation of  $|\hat{\beta}_U(\xi, \Sigma)|^{1+\varepsilon}$  is infinite for all  $\pi, \beta$ . Moreover, if the parameter space (4) contains an open set then any unbiased estimator of  $\beta$*

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<sup>5</sup>Under weak instrument asymptotics as in Staiger & Stock (1997) and homoskedastic errors,  $\sigma_{12}/\sigma_2^2$  is the probability limit of the OLS estimator, though this does not in general hold under weaker assumptions on the error structure.

<sup>6</sup>This property is far from unique to the unbiased estimator, however.

has an infinite  $1 + \varepsilon$  moment.

## 2.4 The Role of the Sign Restriction

In the introduction we argued that it is frequently reasonable to assume that the sign of the first-stage relationship is known, and Theorem 2.1 shows that this restriction suffices to allow mean-unbiased estimation of  $\beta$  in the just-identified model. In fact, a restriction on the parameter space is necessary for an unbiased estimator to exist.

In the just-identified linear IV model with parameter space  $\{(\pi, \beta) \in \mathbb{R}^2\}$ , Theorem 2.5 of Hirano & Porter (2015) implies that no mean, median, or quantile unbiased estimator can exist. Given this negative result, the positive conclusion of Theorem 2.1 may seem surprising. The key point is that by restricting the sign of  $\pi$  to be strictly positive, the parameter space  $\Theta$  as defined in (4) violates Assumption 2.4 of Hirano & Porter (2015), and so renders their negative result inapplicable. Intuitively, assuming the sign of  $\pi$  is known provides just enough information to allow mean-unbiased estimation of  $\beta$ . For further discussion of this point we refer the interested reader to Appendix C.

## 2.5 Behavior of $\hat{\beta}_U$ When $\pi$ is Large

While the finite-sample unbiasedness of  $\hat{\beta}_U$  is appealing, it is also natural to consider performance when the instruments are highly informative. This situation, which we will model by taking  $\pi$  to be large, corresponds to the conventional strong-instrument asymptotics where one fixes the data generating process and takes the sample size to infinity.<sup>7</sup>

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<sup>7</sup>Formally, in the finite-sample normal IV model (1), strong-instrument asymptotics will correspond to fixing  $\pi$  and taking  $T \rightarrow \infty$ , which under mild conditions on  $Z$  and  $Var((U', V)')$  will result in  $\Sigma \rightarrow 0$  in (2). However, it is straightforward to show that the behavior of  $\hat{\beta}_U$ ,  $\hat{\beta}_{2SLS}$ , and many other estimators in this case will be the same as the behavior obtained by holding  $\Sigma$  fixed and taking  $\pi$  to infinity. We focus on the latter case here to simplify the exposition. See Appendix B, which provides asymptotic results with an unknown error distribution, for asymptotic results under  $T \rightarrow \infty$ .

As we discussed above, the unbiased and conventional IV estimators differ only in that the former substitutes  $\hat{\tau}(\xi_2, \sigma_2^2)$  for  $1/\xi_2$ . These two estimators for  $1/\pi$  coincide to a high order of approximation for large values of  $\xi_2$ . Specifically, as noted in Small (2010) (Section 2.3.4), for  $\xi_2 > 0$  we have

$$\sigma_2 \left| \hat{\tau}(\xi_2, \sigma_2^2) - \frac{1}{\xi_2} \right| \leq \left| \frac{\sigma_2^3}{\xi_2^3} \right|.$$

Thus, since  $\xi_2 \xrightarrow{p} \infty$  as  $\pi \rightarrow \infty$ , the difference between  $\hat{\tau}(\xi_2, \sigma_2^2)$  and  $1/\xi_2$  converges rapidly to zero (in probability) as  $\pi$  grows. Consequently, the unbiased estimator  $\hat{\beta}_U$  (appropriately normalized) has the same limiting distribution as the conventional IV estimator  $\hat{\beta}_{2SLS}$  as we take  $\pi \rightarrow \infty$ .

**Theorem 2.3.** *As  $\pi \rightarrow \infty$ , holding  $\beta$  and  $\Sigma$  fixed,*

$$\pi \left( \hat{\beta}_U - \hat{\beta}_{2SLS} \right) \xrightarrow{p} 0.$$

*Consequently,  $\hat{\beta}_U \xrightarrow{p} \beta$  and*

$$\pi \left( \hat{\beta}_U - \beta \right) \xrightarrow{d} N \left( 0, \sigma_1^2 - 2\beta\sigma_{12} + \beta^2\sigma_2^2 \right).$$

Thus, the unbiased estimator  $\hat{\beta}_U$  behaves as the standard IV estimator for large values of  $\pi$ . Consequently, one can show that using this estimator along with conventional standard errors will yield asymptotically valid inference under strong-instrument asymptotics. See Appendix B for details.

### 3 Unbiased Estimation with Multiple Instruments

We now consider the case with multiple instruments, where the model is given by (1) and (2) with  $k$  (the dimension of  $Z_t$ ,  $\pi$ ,  $\xi_1$  and  $\xi_2$ ) greater than 1. As in Section 1.2, we assume that the sign of each element  $\pi_i$  of the first stage vector is known, and we normalize this sign to be positive, giving the parameter space (4).

Using the results in Section 2 one can construct an unbiased estimator for  $\beta$  in many different ways. For any index  $i \in \{1, \dots, k\}$ , the unbiased estimator based on  $(\xi_{1,i}, \xi_{2,i})$

will, of course, still be unbiased for  $\beta$  when  $k > 1$ . One can also take non-random weighted averages of the unbiased estimators based on different instruments. Using the unbiased estimator based on a fixed linear combination of instruments is another possibility, so long as the linear combination preserves the sign restriction. However, such approaches will not adapt to information from the data about the relative strength of instruments and so will typically be inefficient when the instruments are strong.

By contrast, the usual 2SLS estimator achieves asymptotic efficiency in the strongly identified case (modeled here by taking  $\|\pi\| \rightarrow \infty$ ) when errors are homoskedastic. In fact, in this case 2SLS is asymptotically equivalent to an infeasible estimator that uses knowledge of  $\pi$  to choose the optimal combination of instruments. Thus, a reasonable goal is to construct an estimator that (1) is unbiased for fixed  $\pi$  and (2) is asymptotically equivalent to 2SLS as  $\|\pi\| \rightarrow \infty$ .<sup>8</sup> In the remainder of this section we first introduce a class of unbiased estimators and then show that a (feasible) estimator in this class attains the desired strong IV efficiency property. Further, we show that in the over-identified case it is possible to construct unbiased estimators which are robust to small violations of the first stage sign restriction. Finally, we derive bounds on the attainable risk of any estimator for finite  $\|\pi\|$  and show that, while the unbiased estimators described above achieve optimality in an asymptotic sense as  $\|\pi\| \rightarrow \infty$  regardless of the direction of  $\pi$ , the optimal unbiased estimator for finite  $\pi$  will depend on the direction of  $\pi$ .

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<sup>8</sup>In the heteroskedastic case, the 2SLS estimator will no longer be asymptotically efficient, and a two-step GMM estimator can be used to achieve the efficiency bound. Because it leads to simpler exposition, and because the 2SLS estimator is common in practice, we consider asymptotic equivalence with 2SLS, rather than asymptotic efficiency in the heteroskedastic case, as our goal. As discussed in Section 3.3 below, however, our approach generalizes directly to efficient estimators in non-homoskedastic settings.

### 3.1 A General Class of Unbiased Estimators

Let

$$\xi(i) = \begin{pmatrix} \xi_{1,i} \\ \xi_{2,i} \end{pmatrix} \text{ and } \Sigma(i) = \begin{pmatrix} \Sigma_{11,ii} & \Sigma_{12,ii} \\ \Sigma_{21,ii} & \Sigma_{22,ii} \end{pmatrix}$$

be the reduced form and first stage estimators based on the  $i$ th instrument and their variance matrix, respectively, so that  $\hat{\beta}_U(\xi(i), \Sigma(i))$  is the unbiased estimator based on the  $i$ th instrument. Given a weight vector  $w \in \mathbb{R}^k$  with  $\sum_{i=1}^k w_i = 1$ , let

$$\hat{\beta}_w(\xi, \Sigma; w) = \sum_{i=1}^k w_i \hat{\beta}_U(\xi(i), \Sigma(i)).$$

Clearly,  $\hat{\beta}_w$  is unbiased so long as  $w$  is nonrandom. Allowing  $w$  to depend on the data  $\xi$ , however, may introduce bias through the dependence between the weights and the estimators  $\hat{\beta}_U(\xi(i), \Sigma(i))$ .

To avoid this bias we first consider a randomized unbiased estimator and then take its conditional expectation given the sufficient statistic  $\xi$  to eliminate the randomization. Let  $\zeta \sim N(0, \Sigma)$  be independent of  $\xi$ , and let  $\xi^{(a)} = \xi + \zeta$  and  $\xi^{(b)} = \xi - \zeta$ . Then  $\xi^{(a)}$  and  $\xi^{(b)}$  are (unconditionally) independent draws with the same marginal distribution as  $\xi$ , save that  $\Sigma$  is replaced by  $2\Sigma$ . If  $T$  is even,  $Z'Z$  is the same across the first and second halves of the sample, and the errors are iid, then  $\xi^{(a)}$  and  $\xi^{(b)}$  have the same joint distribution as the reduced form estimators based on the first and second half of the sample. Thus, we can think of these as split-sample reduced-form estimates.

Let  $\hat{w} = \hat{w}(\xi^{(b)})$  be a vector of data dependent weights with  $\sum_{i=1}^k \hat{w}_i = 1$ . By the independence of  $\xi^{(a)}$  and  $\xi^{(b)}$ ,

$$E \left[ \hat{\beta}_w(\xi^{(a)}, 2\Sigma; \hat{w}(\xi^{(b)})) \right] = \sum_{i=1}^k E [\hat{w}_i(\xi^{(b)})] \cdot E \left[ \hat{\beta}_U(\xi^{(a)}(i), 2\Sigma(i)) \right] = \beta. \quad (7)$$

To eliminate the noise introduced by  $\zeta$ , define the ‘‘Rao-Blackwellized’’ estimator

$$\hat{\beta}_{RB} = \hat{\beta}_{RB}(\xi, \Sigma; \hat{w}) = E \left[ \hat{\beta}_w(\xi^{(a)}, 2\Sigma; \hat{w}(\xi^{(b)})) \middle| \xi \right].$$

Unbiasedness of  $\hat{\beta}_{RB}$  follows immediately from (7) and the law of iterated expectations. While  $\hat{\beta}_{RB}$  does not, to our knowledge, have a simple closed form, it can be computed by integrating over the distribution of  $\zeta$ . This can easily be done by simulation, taking the sample average of  $\hat{\beta}_w$  over simulated draws of  $\xi^{(a)}$  and  $\xi^{(b)}$  while holding  $\xi$  at its observed value.

### 3.2 Equivalence with 2SLS under Strong IV Asymptotics

We now propose a set of weights  $\hat{w}$  which yield an unbiased estimator asymptotically equivalent to 2SLS. To motivate these weights, note that for  $W = Z'Z$  and  $e_i$  the  $i$ th standard basis vector, the 2SLS estimator can be written as

$$\hat{\beta}_{2SLS} = \frac{\xi_2' W \xi_1}{\xi_2' W \xi_2} = \sum_{i=1}^k \frac{\xi_2' W e_i e_i' \xi_2}{\xi_2' W \xi_2} \frac{\xi_{1,i}}{\xi_{2,i}},$$

which is the GMM estimator with weight matrix  $W = Z'Z$ . Thus, the 2SLS estimator is a weighted average of the 2SLS estimates based on single instruments, where the weight for estimate  $\xi_{1,i}/\xi_{2,i}$  based on instrument  $i$  is equal to  $\frac{\xi_2' W e_i e_i' \xi_2}{\xi_2' W \xi_2}$ . This suggests the unbiased Rao-Blackwellized estimator with weights  $\hat{w}_i^*(\xi^{(b)}) = \frac{\xi_2^{(b)'} W e_i e_i' \xi_2^{(b)}}{\xi_2^{(b)'} W \xi_2^{(b)}}$ :

$$\hat{\beta}_{RB}^* = \hat{\beta}_{RB}(\xi, \Sigma; \hat{w}) = E \left[ \hat{\beta}_w(\xi^{(a)}, 2\Sigma; \hat{w}^*(\xi^{(b)})) \middle| \xi \right]. \quad (8)$$

The following theorem shows that  $\hat{\beta}_{RB}^*$  is asymptotically equivalent to  $\hat{\beta}_{2SLS}$  in the strongly identified case, and is therefore asymptotically efficient if the errors are iid.

**Theorem 3.1.** *Let  $\pi \rightarrow \infty$  with  $\|\pi\|/\min_i \pi_i = \mathcal{O}(1)$ . Then  $\|\pi\|(\hat{\beta}_{RB}^* - \hat{\beta}_{2SLS}) \xrightarrow{P} 0$ .*

The condition that  $\|\pi\|/\min_i \pi_i = \mathcal{O}(1)$  amounts to an assumption that the “strength” of all instruments is of the same order. As discussed below in Section 3.4, this assumption can be relaxed by redefining the instruments.

To understand why Theorem 3.1 holds, consider the “oracle” weights  $w_i^* = \frac{\pi' W e_i e_i' \pi}{\pi' W \pi}$ . It is easy to see that  $\hat{w}_i^* - w_i^* \xrightarrow{P} 0$  as  $\|\pi\| \rightarrow \infty$ . Consider the oracle unbiased estimator  $\hat{\beta}_{RB}^o = \hat{\beta}_{RB}(\xi, \Sigma; w^*)$ , and the oracle combination of individual 2SLS estimators

$\hat{\beta}_{2SLS}^o = \sum_{i=1}^k w_i^* \frac{\xi_{1,i}}{\xi_{2,i}}$ . By arguments similar to those used to show that statistical noise in the first stage estimates does not affect the 2SLS asymptotic distribution under strong instrument asymptotics, it can be seen that  $\|\pi\|(\hat{\beta}_{2SLS}^o - \hat{\beta}_{2SLS}) \xrightarrow{p} 0$  as  $\|\pi\| \rightarrow \infty$ . Further, one can show that  $\hat{\beta}_{RB}^o = \hat{\beta}_w(\xi, \Sigma; w^*) = \sum_{i=1}^k w_i^* \hat{\beta}_U(\xi(i), \Sigma(i))$ . Since this is just  $\hat{\beta}_{2SLS}^o$  with  $\hat{\beta}_U(\xi(i), \Sigma(i))$  replacing  $\xi_{i,1}/\xi_{i,2}$ , it follows by Theorem 2.3 that  $\|\pi\|(\hat{\beta}_{RB}^o - \hat{\beta}_{2SLS}^o) \xrightarrow{p} 0$ . Theorem 3.1 then follows by showing that  $\|\pi\|(\hat{\beta}_{RB} - \hat{\beta}_{RB}^o) \xrightarrow{p} 0$ , which follows for essentially the same reasons that first stage noise does not affect the asymptotic distribution of the 2SLS estimator but requires some additional argument. We refer the interested reader to the proof of Theorem 3.1 in Appendix A for details.

### 3.3 Efficient Estimation with Non-Homoskedastic Errors

The estimator  $\hat{\beta}_{RB}^*$  proposed above may be viewed as deficient because it is asymptotically efficient only under homoskedastic errors. We now discuss an extension of the results above to efficient estimation without a homoskedasticity assumption. In models with non-homoskedastic errors the two step GMM estimator given by

$$\hat{\beta}_{GMM, \hat{W}} = \frac{\xi_2' \hat{W} \xi_1}{\xi_2' \hat{W} \xi_2} \text{ where } \hat{W} = \left( \Sigma_{11} - \hat{\beta}_{2SLS}(\Sigma_{12} + \Sigma_{21}) + \hat{\beta}_{2SLS}^2 \Sigma_{22} \right)^{-1}, \quad (9)$$

is asymptotically efficient under strong instruments. Here,  $\hat{W}$  is an estimate of the inverse of the variance matrix of the moments  $\xi_1 - \beta \xi_2$ , which the GMM estimator sets close to zero. Let

$$\hat{w}_{GMM,i}^*(\xi^{(b)}) = \frac{\xi_2^{(b)'} \hat{W}(\xi^{(b)}) e_i e_i' \xi_2^{(b)}}{\xi_2^{(b)'} \hat{W}(\xi^{(b)}) \xi_2^{(b)}} \quad (10)$$

where

$$\hat{W}(\xi^{(b)}) = \left( \Sigma_{11} - \hat{\beta}(\xi^{(b)})(\Sigma_{12} + \Sigma_{21}) + \hat{\beta}(\xi^{(b)})^2 \Sigma_{22} \right)^{-1}$$

for a preliminary estimator  $\hat{\beta}(\xi^{(b)})$  of  $\beta$  based on  $\xi^{(b)}$ . The Rao-Blackwellized estimator formed by replacing  $\hat{w}^*$  with  $\hat{w}_{GMM}^*$  in the definition of  $\hat{\beta}_{RB}^*$  gives an unbiased estimator that is asymptotically efficient under strong instrument asymptotics with non-homoskedastic errors. We refer the reader to Appendix A for details.



### 3.4 Robust Unbiased Estimation

So far, all the unbiased estimators we have discussed required  $\pi_i > 0$  for all  $i$ . Even when the first stage sign is dictated by theory, however, we may be concerned that this restriction may fail to hold exactly in a given empirical context. To address such concerns, in this section we show that in over-identified models we can construct estimators which are robust to small violations of the sign restriction. Our approach has the further benefit of ensuring asymptotic efficiency when, while  $\|\pi\| \rightarrow \infty$ , the elements  $\pi_i$  may increase at different rates.

Let  $M$  be a  $k \times k$  invertible matrix such that all elements are strictly positive, and

$$\tilde{\xi} = (I_2 \otimes M)\xi, \quad \tilde{\Sigma} = (I_2 \otimes M)\Sigma(I_2 \otimes M)', \quad \tilde{W} = M^{-1}'WM^{-1}.$$

The GMM estimator based on  $\tilde{\xi}$  and  $\tilde{W}$  is numerically equivalent to the GMM estimator based on  $\xi$  and  $W$ . In particular, for many choices of  $W$ , including all those discussed above, estimation based on  $(\tilde{\xi}, \tilde{W}, \tilde{\Sigma})$  is equivalent to estimation based on instruments  $ZM^{-1}$  rather than  $Z$ .

Note that for  $\tilde{\pi} = M\pi$ ,  $\tilde{\xi}$  is normally distributed with mean  $(\tilde{\pi}'\beta, \tilde{\pi}')$  and variance  $\tilde{\Sigma}$ . Thus, if we construct the estimator  $\hat{\beta}_{RB}^*$  from  $(\tilde{\xi}, \tilde{W}, \tilde{\Sigma})$  instead of  $(\xi, W, \Sigma)$ , we obtain an unbiased estimator provided  $\tilde{\pi}_i > 0$  for all  $i$ . Since all elements of  $M$  are strictly positive this is a strictly weaker condition than  $\pi_i > 0$  for all  $i$ . By Theorem 3.1,  $\hat{\beta}_{RB}^*$  constructed from  $\tilde{\xi}$  and  $\tilde{W}$  will be asymptotically efficient as  $\|\tilde{\pi}\| \rightarrow \infty$  so long as  $\tilde{\pi} = M\pi$  is nonnegative and satisfies  $\|\tilde{\pi}\|/\min_i \tilde{\pi}_i = \mathcal{O}(1)$ . Note, however, that

$$\min_i \tilde{\pi}_i \geq (\min_{i,j} M_{ij})\|\pi\| = (\min_{i,j} M_{ij}) \frac{\|\pi\|}{\|M\pi\|} \|\tilde{\pi}\| \geq (\min_{i,j} M_{ij}) \left( \inf_{\|u\|=1} \frac{\|u\|}{\|Mu\|} \right) \|\tilde{\pi}\|$$

so  $\|\tilde{\pi}\|/\min_i \tilde{\pi}_i = \mathcal{O}(1)$  now follows automatically from  $\|\pi\| \rightarrow \infty$ .

Conducting estimation based on  $\tilde{\xi}$  and  $\tilde{W}$  offers a number of advantages for many

different choices of  $M$ . One natural class of transformations  $M$  is

$$M = \begin{bmatrix} 1 & c & c & \cdots & c \\ c & 1 & c & \cdots & c \\ c & c & 1 & \cdots & c \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ c & c & c & \cdots & 1 \end{bmatrix} \text{Diag}(\Sigma_{22})^{-\frac{1}{2}} \quad (11)$$

for  $c \in [0, 1)$  and  $\text{Diag}(\Sigma_{22})$  the matrix with the same diagonal as  $\Sigma_{22}$  and zeros elsewhere. For a given  $c$ , denote the estimator  $\hat{\beta}_{RB}^*$  based on the corresponding  $(\tilde{\xi}, \tilde{W}, \tilde{\Sigma})$  by  $\hat{\beta}_{RB,c}^*$ . One can show that  $\hat{\beta}_{RB,0}^* = \hat{\beta}_{RB}^*$  based on  $(\xi, W, \Sigma)$ , and going forward we let  $\hat{\beta}_{RB}^*$  denote  $\hat{\beta}_{RB,0}^*$ .

We can interpret  $c$  as specifying a level of robustness to violations on the sign restriction for  $\pi_i$ . In particular, for a given choice of  $c$ ,  $\tilde{\pi}$  will satisfy the sign restriction provided that for each  $i$ ,

$$-\pi_i / \sqrt{\Sigma_{22,ii}} < c \cdot \sum_{j \neq i} \pi_j / \sqrt{\Sigma_{22,jj}},$$

that is, provided the expected z-statistic for testing that each wrong-signed  $\pi_i$  is equal to zero is less than  $c$  times the sum of the expected z-statistics for  $j \neq i$ . Larger values of  $c$  provide a greater degree of robustness to violations of the sign restriction, while all choices of  $c \in (0, 1)$  yield asymptotically equivalent estimators as  $\|\pi\| \rightarrow \infty$ . For finite values of  $\pi$  however, different choices of  $c$  yield different estimators, so we explore the effects of different choices below using the Angrist & Krueger (1991) dataset. Determining the optimal choice of  $c$  for finite values of  $\pi$  is an interesting topic for future research.

### 3.5 Bounds on the Attainable Risk

While the class of estimators given above has the desirable property of asymptotic efficiency as  $\|\pi\| \rightarrow \infty$ , it is useful to have a benchmark for the performance for finite  $\pi$ . In Appendix D, we derive a lower bound for the risk of any unbiased estimator at

a given  $\pi^*, \beta^*$ . The bound is based on the risk in a submodel with a single instrument and, as in the single instrument case, shows that any unbiased estimator must have an infinite  $1 + \varepsilon$  absolute moment for  $\varepsilon > 0$ . In certain cases, which include large parts of the parameter space under homoskedastic errors  $(U_t, V_t)$ , the bound can be attained. The estimator that attains the bound turns out to depend on the value  $\pi^*$ , which shows that no uniform minimum risk unbiased estimator exists. See Appendix D for details.

## 4 Simulations

In this section we present simulation results on the performance of our unbiased estimators. We first consider models with a single instrument and then turn to over-identified models. Since the parameter space in the single-instrument model is small, we are able to obtain comprehensive simulation results in this case, studying performance over a wide range of parameter values. In the over-identified case, by contrast, the parameter space is too large to comprehensively explore by simulation so we instead calibrate our simulations to the Staiger & Stock (1997) specifications for the Angrist & Krueger (1991) dataset.

### 4.1 Performance with a Single Instrument

The estimator  $\hat{\beta}_U$  based on a single instrument plays a central role in all of our results, so in this section we examine the performance of this estimator in simulation. For purposes of comparison we also discuss results for the two-stage least squares estimator  $\hat{\beta}_{2SLS}$ . The lack of moments for  $\hat{\beta}_{2SLS}$  in the just-identified context renders some comparisons with  $\hat{\beta}_U$  infeasible, however, so we also consider the performance of the Fuller (1977) estimator with constant one,

$$\hat{\beta}_{FULL} = \frac{\xi_2 \xi_1 + \sigma_{12}}{\xi_2^2 + \sigma_2^2}$$

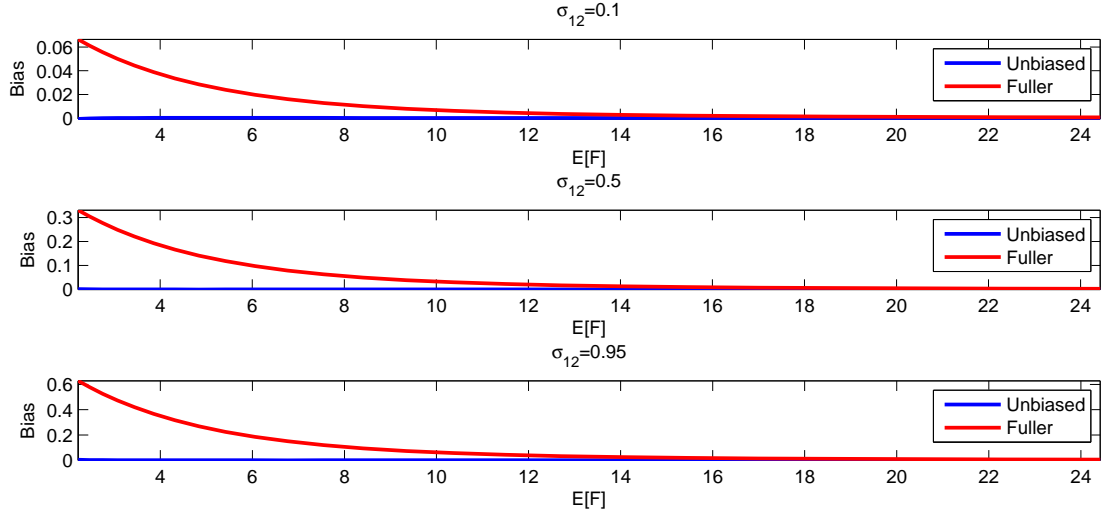


Figure 1: Bias of single-instrument estimators, plotted against mean  $E[F]$  of first-stage F-statistic, based on 10 million simulations.

which we define as in Mills et al. (2014).<sup>9</sup> Note that in the just-identified case considered here  $\hat{\beta}_{FULL}$  also coincides with the bias-corrected 2SLS estimator (again, see Mills et al.).

While the model (2) has five parameters in the single-instrument case,  $(\beta, \pi, \sigma_1^2, \sigma_{12}, \sigma_2^2)$ , an equivariance argument implies that for our purposes it suffices to fix  $\beta = 0$ ,  $\sigma_1 = \sigma_2 = 1$  and consider the parameter space  $(\pi, \sigma_{12}) \in (0, \infty) \times [0, 1)$ . See Appendix E for details. Since this parameter space is just two-dimensional, we can fully explore it via simulation.

#### 4.1.1 Estimator Location

We first compare the bias of  $\hat{\beta}_U$  and  $\hat{\beta}_{FULL}$  (we omit  $\hat{\beta}_{2SLS}$  from this comparison, as it does not have a mean in the just-identified case). We consider  $\sigma_{12} \in \{0.1, 0.5, 0.95\}$  and examine a wide range of values for  $\pi > 0$ .<sup>10</sup>

<sup>9</sup>In the case where  $U_t$  and  $V_t$  are correlated or heteroskedastic across  $t$ , the definition of  $\hat{\beta}_{FULL}$  above is the natural extension of the definition considered in Mills et al. (2014).

<sup>10</sup>We restrict attention to  $\pi \geq 1$  in the bias plots. Since the first stage F-statistic is  $F = \xi_2^2$  in the present context, this corresponds to  $E[F] \geq 2$ . The expectation of  $\hat{\beta}_U$  ceases to exist at  $\pi = 0$ , and

If, rather than considering mean bias, we instead consider median bias, we find that  $\hat{\beta}_U$  and  $\hat{\beta}_{2SLS}$  generally exhibit smaller median bias than  $\hat{\beta}_{FULL}$ . There is no ordering between  $\hat{\beta}_U$  and  $\hat{\beta}_{2SLS}$  in terms of median bias, however, as the median bias of  $\hat{\beta}_U$  is smaller than that of  $\hat{\beta}_{2SLS}$  for very small values of  $\pi$ , while the median bias of  $\hat{\beta}_{2SLS}$  is smaller for larger values  $\pi$ .

#### 4.1.2 Estimator Dispersion

The lack of moments for  $\hat{\beta}_{2SLS}$  complicates comparisons of dispersion, since we cannot consider mean squared error or mean absolute deviation, and also cannot recenter  $\hat{\beta}_{2SLS}$  around its mean. As an alternative, we instead consider the full distribution of the absolute deviation of each estimator from its median. In particular, for the estimators  $(\hat{\beta}_U, \hat{\beta}_{2SLS}, \hat{\beta}_{FULL})$  we calculate the zero-median residuals

$$(\varepsilon_U, \varepsilon_{2SLS}, \varepsilon_{FULL}) = \left( \hat{\beta}_U - \text{med} \left( \hat{\beta}_U \right), \hat{\beta}_{2SLS} - \text{med} \left( \hat{\beta}_{2SLS} \right), \hat{\beta}_{FULL} - \text{med} \left( \hat{\beta}_{FULL} \right) \right).$$

Our simulation results suggest a strong stochastic ordering between these residuals (in absolute value). In particular we find that  $|\varepsilon_{2SLS}|$  approximately dominates  $|\varepsilon_U|$ , which in turn approximately dominates  $|\varepsilon_{FULL}|$ , both in the sense of first order stochastic dominance. In particular, for  $\tau \in \{0.001, 0.002, \dots, 0.999\}$  the  $\tau$ -th quantile of  $|\varepsilon_{2SLS}|$  in simulation is never more than  $10^{-4}$  smaller than the  $\tau$ -th quantile of  $|\varepsilon_U|$ , and the  $\tau$ -th quantile of  $|\varepsilon_U|$  is never more than  $10^{-3}$  smaller than the  $\tau$ -th quantile of  $|\varepsilon_{FULL}|$ , both uniformly over  $\tau$  and  $(\pi, \sigma_{12})$ .<sup>11</sup> Thus, our simulations demonstrate that  $\hat{\beta}_{2SLS}$  is more dispersed around its median than is  $\hat{\beta}_U$ , which is in turn more dispersed around its median than  $\hat{\beta}_{FULL}$ . To illustrate this finding, Figure 2 plots the median of  $|\varepsilon|$  for  $\pi$  close to zero the heavy tails of  $\hat{\beta}_U$  mean that the sample average of  $\hat{\beta}_U$  in simulation can differ substantially from zero even with a large number of simulation replications.

<sup>11</sup>By contrast, the  $\tau$ -th quantile of  $|\varepsilon_{2SLS}|$  may exceed corresponding quantile of  $|\varepsilon_U|$  by as much as 483, or (in proportional terms) by as much as a factor of 32, while the  $\tau$ -th quantile of  $|\varepsilon_U|$  may exceed the corresponding quantile of  $|\varepsilon_{FULL}|$  by as much as 37, or (in proportional terms) by as much as a factor of 170.

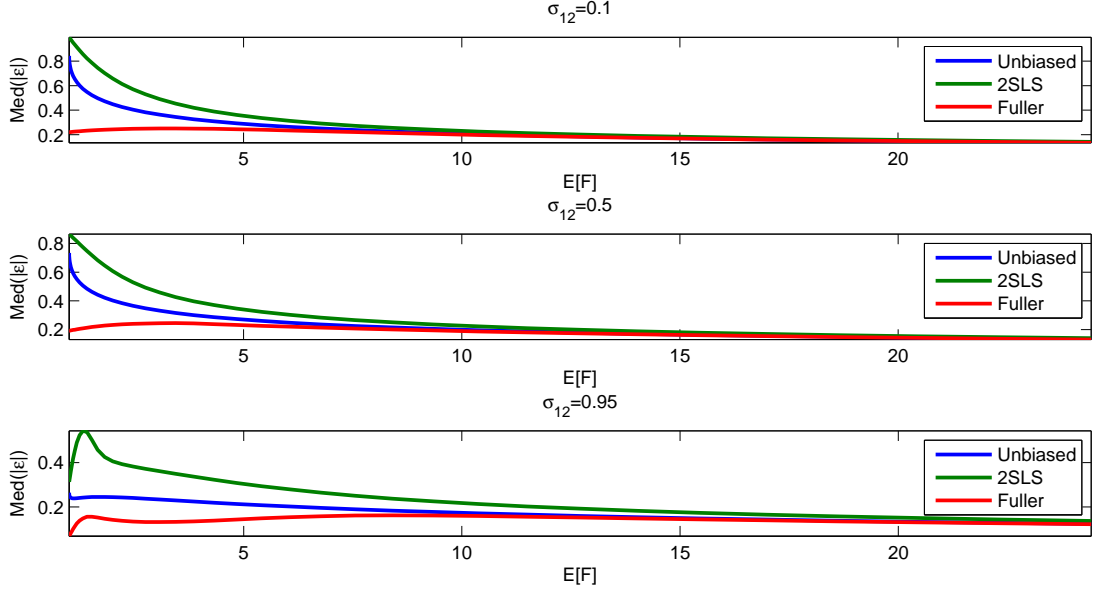


Figure 2: Median of  $|\varepsilon| = \left| \hat{\beta} - \text{med}(\hat{\beta}) \right|$  for single-instrument IV estimators, plotted against mean  $E_{\pi}[F]$  of first-stage F-statistic, based on 10 million simulations.

the different estimators. While Figure 2 considers only one quantile and three values of  $\sigma_{12}$ , a more extensive discussion of our simulations results is given in Appendix F.

This numerical result is consistent with analytical results on the tail behavior of the estimators. In particular,  $\hat{\beta}_{2SLS}$  has no moments, reflecting thick tails in its sampling distribution, while  $\hat{\beta}_{FULL}$  has all moments, reflecting thin tails. As noted in Section 2.3, the unbiased estimator  $\hat{\beta}_U$  has a first moment but no more, and so falls between these two extremes.

## 4.2 Performance with Multiple Instruments

In models with multiple instruments, if we assume that errors are homoskedastic an equivariance argument closely related to that in just-identified case again allows us to reduce the dimension of the parameter space. Unlike in the just-identified case, however, the matrix  $Z'Z$  and the direction of the first stage,  $\pi/\|\pi\|$ , continue to matter (see Appendix E for details). As a result, the parameter space is too large to fully explore by

simulation, so we instead calibrate our simulations to the Staiger & Stock (1997) specifications for the 1930-1939 cohort in the Angrist & Krueger (1991) data. While there is statistically significant heteroskedasticity in this data, this significance appears to be the result of the large sample size rather than substantively important deviations from homoskedasticity. In particular, procedures which assume homoskedasticity produce very similar answers to heteroskedasticity-robust procedures when applied to this data. Thus, given that homoskedasticity leads to a reduction of the parameter space as discussed above, we impose homoskedasticity in our simulations.

In each of the four Staiger & Stock (1997) specifications we estimate  $\pi/\|\pi\|$  and  $Z'Z$  from the data (ensuring, as discussed in Appendix G, that  $\pi/\|\pi\|$  satisfies the sign restriction). After reducing the parameter space by equivariance and calibrating  $Z'Z$  and  $\pi/\|\pi\|$  to the data, the model has two remaining free parameters: the norm of the first stage,  $\|\pi\|$ , and the correlation  $\sigma_{UV}$  between the reduced-form and first-stage errors. We examine behavior for a range of values for  $\|\pi\|$  and for  $\sigma_{UV} \in \{0.1, 0.5, 0.95\}$ . Further details on the simulation design are given in Appendix G.

For each parameter value we simulate the performance of  $\hat{\beta}_{2SLS}$ ,  $\hat{\beta}_{FULL}$  (which is again the Fuller estimator with constant equal to one), and  $\hat{\beta}_{RB}^*$  as defined in Section 3.2. We also consider the robust estimators  $\hat{\beta}_{RB,c}^*$  discussed in Section 3.4 for  $c \in \{0.1, 0.5, 0.9\}$ , but find that all three choices produce very similar results and so focus on  $c = 0.5$  to simplify the graphs.<sup>12</sup> Even with a million simulation replications, simulation estimates of the bias for the unbiased estimators (which we know to be zero from the results of Section 3) remain noisy relative to e.g. the bias in 2SLS in some calibrations, so we do not plot the bias estimates and instead focus on the mean absolute deviation (MAD)  $E_{\pi,\beta} \left[ \left| \hat{\beta} - \beta \right| \right]$  since, unlike in the just-identified case, the MAD for 2SLS is now finite. We also plot the lower bound on the mean absolute deviation of unbiased estimators discussed in Section 3.5.

Several features become clear from these results. As expected, the performance of 2SLS is typically worse for models with more instruments or with a higher degree of cor-

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<sup>12</sup>All results for the *RB* estimators are based on 1,000 draws of  $\zeta$ .

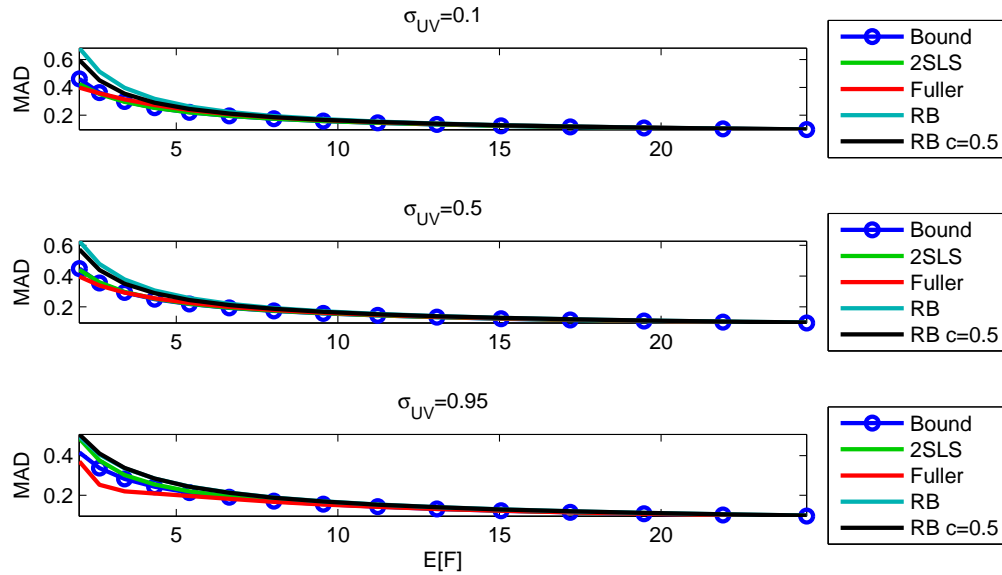


Figure 3: Mean absolute deviation of estimators in simulations calibrated to specification I of Staiger & Stock (1997), which has  $k=3$ , based on 1 million simulations.

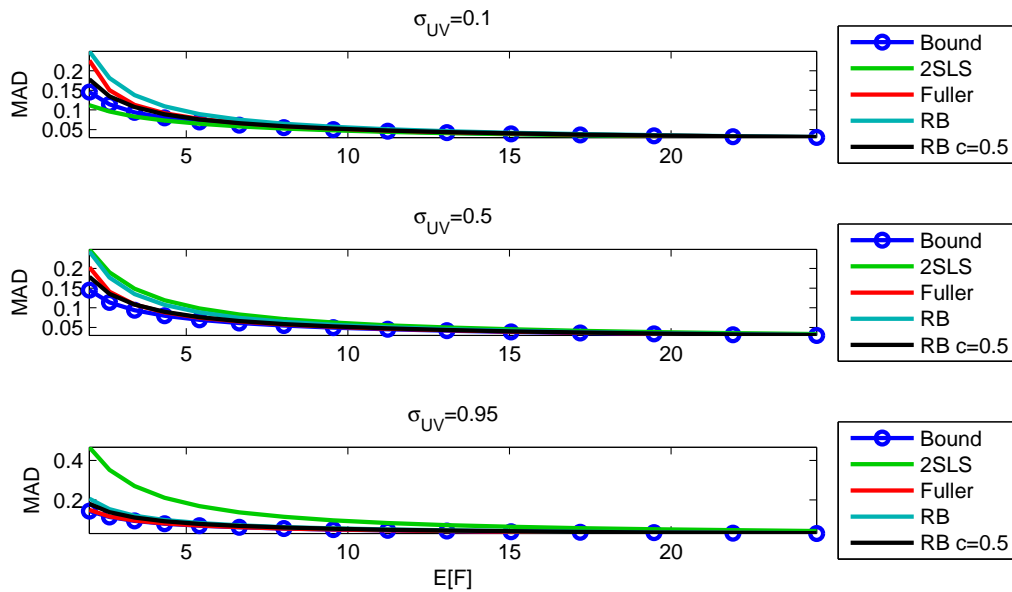


Figure 4: Mean absolute deviation of estimators in simulations calibrated to specification II of Staiger & Stock (1997), which has  $k=30$ , based on 1 million simulations.



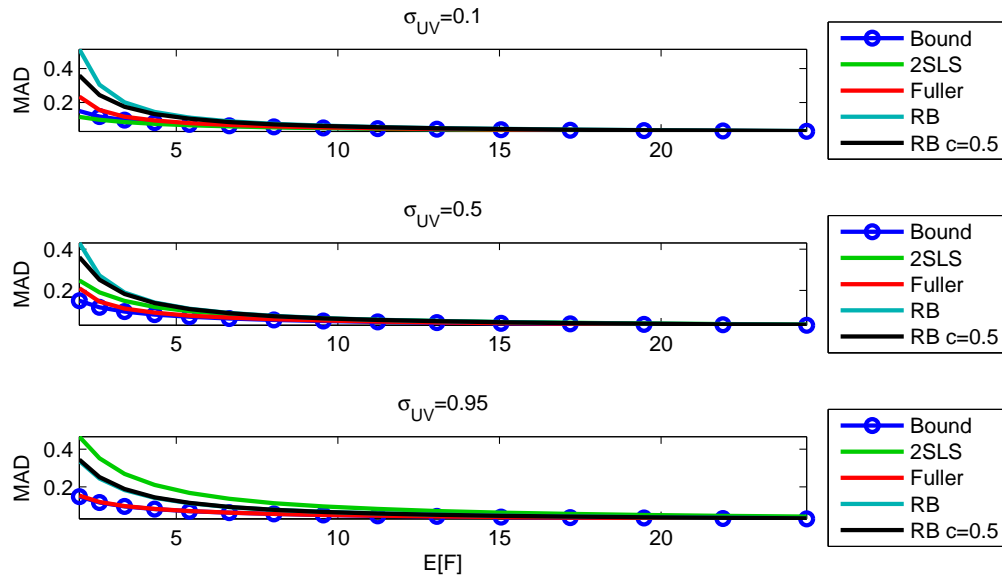


Figure 5: Mean absolute deviation of estimators in simulations calibrated to specification III of Staiger & Stock (1997), which has  $k=28$ , based on 1 million simulations.

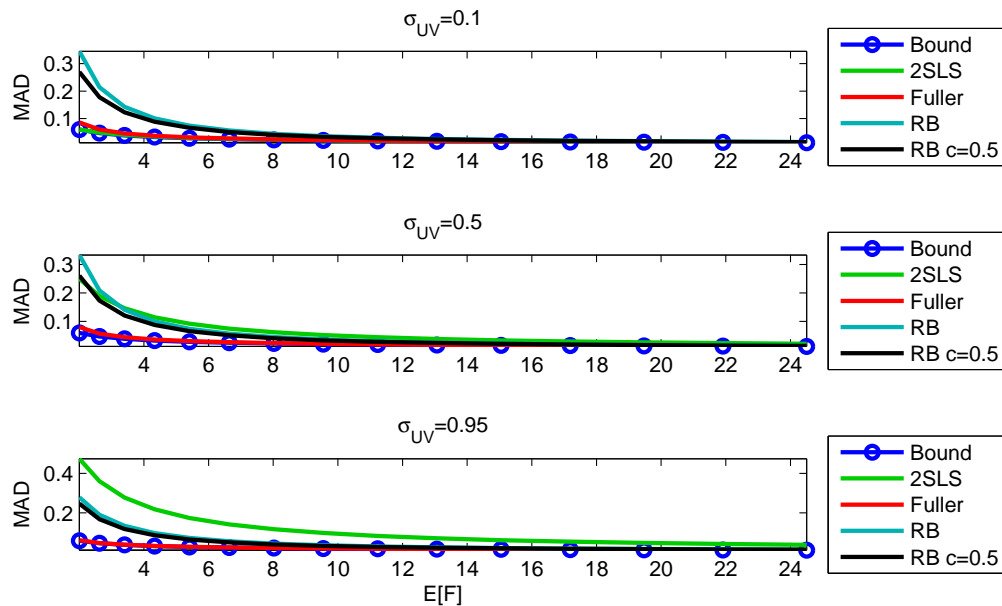


Figure 6: Mean absolute deviation of estimators in simulations calibrated to specification IV of Staiger & Stock (1997), which has  $k=178$ , based on 100,000 simulations.

relation between the reduced-form and first-stage errors (i.e. higher  $\sigma_{UV}$ ). The robust unbiased estimator  $\hat{\beta}_{RB,0.5}$  generally outperforms  $\hat{\beta}_{RB}^* = \hat{\beta}_{RB,0}^*$ . Since the estimators with  $c = 0.1$  and  $c = 0.9$  perform very similarly to that with  $c = 0.5$ , they outperform  $\hat{\beta}_{RB}^*$  as well. The gap in performance between the RB estimators and the lower bound on MAD over the class of all unbiased estimators is typically larger in specifications with more instruments. Interestingly, we see that the Fuller estimator often performs quite well, and has MAD close to or below the lower bound for the class of unbiased estimators in most designs. While this estimator is biased, its bias decreases quickly in  $\|\pi\|$  in the designs considered. Thus, at least in the homoskedastic case, this estimator seems a potentially appealing choice if we are willing to accept bias for small values of  $\pi$ .

## 5 Empirical Applications

We calculate our proposed estimators in two empirical applications. First, we consider the data and specifications used in Hornung (2014) to examine the effect of seventeenth century migrations on productivity. For our second application, we study the Staiger & Stock (1997) specifications for the Angrist & Krueger (1991) dataset on the relationship between education and labor market earnings.

### 5.1 Hornung (2014)

Hornung (2014) studies the long term impact of the flight of skilled Huguenot refugees from France to Prussia in the seventeenth century. He finds that regions of Prussia which received more Huguenot refugees during the late seventeenth century had a higher level of productivity in textile manufacturing at the start of the nineteenth century. To address concerns over endogeneity in Huguenot settlement patterns and obtain an estimate for the causal effect of skilled immigration on productivity, Hornung (2014) considers specifications which instrument Huguenot immigration to a given re-

gion using population losses due to plague at the end of the Thirty Years' War. For more information on the data and motivation of the instrument, see Hornung (2014).

Hornung's argument for the validity of his instrument clearly implies that the first-stage effect should be positive, but the relationship between the instrument and the endogenous regressors appears to be fairly weak. In particular, the four IV specifications reported in Tables 4 and 5 of Hornung (2014) have first-stage F-statistics of 3.67, 4.79, 5.74, and 15.35, respectively. Thus, it seems that the conventional normal approximation to the distribution of IV estimates may be unreliable in this context. In each of the four main IV specifications considered by Hornung, we compare 2SLS and Fuller (again with constant equal to one) to our estimator. Since there is only a single instrument in this context, the model is just-identified and the unbiased estimator is unique. In each specification we also compute and report an identification-robust Anderson-Rubin confidence set for the coefficient on the endogenous regressor. The results are reported in Table 1.

As we can see from Table 1, our unbiased estimates are systematically smaller than the 2SLS estimates computed in Hornung (2014). Fuller estimates are, in turn, smaller than our unbiased estimates. Despite this, only in specification II is the unbiased estimate excluded from a 90% 2SLS Wald confidence set computed using the standard errors reported by Hornung. As the identification-robust Anderson Rubin confidence sets we report make clear, a wide range of parameter values are consistent with the restrictions of the IV model. Indeed, we see that in all specifications considered, 95% AR confidence sets contain both positive and negative values of arbitrarily large magnitude.

## **5.2 Angrist & Krueger (1991)**

Angrist & Krueger (1991) are interested in the relationship between education and labor market earnings. They argue that students born later in the calendar year face a longer period of compulsory schooling than those born earlier in the calendar year, and that quarter of birth is a valid instrument for years of schooling. As we note above their

Specification	Estimator	I	II	III	IV
$X$ : Percent Huguenots in 1700	2SLS	3.48	3.38	1.67	
	Fuller	1.00	0.97	1.01	
	Unbiased	1.52	1.46	1.15	
$X$ : log Huguenots in 1700	2SLS				0.07
	Fuller				0.06
	Unbiased				0.06
95% AR Confidence Set		$(-\infty, 0.56] \cup [1.03, \infty)$	$(-\infty, 0.64] \cup [0.95, \infty)$	$(-\infty, \infty)$	$(-\infty, \infty)$
Other controls		Yes	Yes	Yes	Yes
Observations		150	150	186	186
Number of Towns		57	57	71	71
First Stage F-Statistic		3.67	4.79	5.74	15.35

Table 1: Results in Hornung (2014) data. Specifications in columns I and II correspond to Table 4 columns (3) and (5) in Hornung (2014), respectively, while columns III and IV correspond to Table 5 columns (3) and (6) in Hornung (2014).  $Y$  =log output,  $X$  as indicated, and  $Z$  =unadjusted population losses in I, interpolated population losses in II, and population losses averaged over several data sources in III and IV. See Hornung (2014). The 2SLS and Fuller rows report two stage least squares and Fuller estimates, respectively, while Unbiased reports  $\hat{\beta}_U$ . Other controls include a constant, a dummy for whether a town had relevant textile production in 1685, measurable inputs to the production process, and others as in Hornung (2014). As in Hornung (2014), all covariance estimates are clustered at the town level.

argument implies that the sign of the first-stage effect is known. A substantial literature, beginning with Bound et al. (1995), notes that the relationship between the instruments and the endogenous regressor appears to be quite weak in some specifications considered in Angrist & Krueger (1991). Here we consider four specifications from Staiger & Stock (1997), based on the 1930-1939 cohort. See Angrist & Krueger (1991) and Staiger & Stock (1997) for more on the data and specification.

We calculate unbiased estimators  $\hat{\beta}_{RB}^*$ ,  $\hat{\beta}_{RB,0.1}^*$ ,  $\hat{\beta}_{RB,0.5}^*$ , and  $\hat{\beta}_{RB,0.9}^*$ . In all cases we take  $W = Z'Z$ . To calculate confidence sets we use the quasi-CLR (or GMM-M) test of Kleibergen (2005), which simplifies to the CLR test of Moreira (2003) under homoskedasticity and so delivers nearly-optimal confidence sets in that case (see Mikusheva 2010). Thus, since as discussed above the data in this application appears reasonably close to homoskedasticity, we may reasonably expect the quasi-CLR confidence set to perform well. All results are reported in Table 2.

A few points are notable from these results. First, we see that in specifications I and II, which have the largest first stage F-statistics, the unbiased estimates are quite close to the other point estimates. Moreover, in these specifications the choice of  $c$  makes little difference. By contrast, in specification III, where the instruments appear to be quite weak, the unbiased estimates differ substantially, with  $\hat{\beta}_{RB}^*$  yielding a negative point estimate and  $\hat{\beta}_{RB,c}^*$  for  $c \in \{0.1, 0.5, 0.9\}$  yielding positive estimates substantially larger than the other estimators considered.<sup>13</sup> A similar, though less pronounced, version of this phenomenon arises in specification IV, where unbiased estimates are smaller than those based on conventional methods and  $\hat{\beta}_{RB}^*$  is almost 20% smaller than estimates based on other choices of  $c$ .

As in the simulations there is very little difference between the estimates for  $c \in \{0.1, 0.5, 0.9\}$ . In particular, while not exactly the same, the estimates coincide once

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<sup>13</sup>All unbiased estimates are calculated by averaging over 100,000 draws of  $\zeta$ . For all estimates except  $\hat{\beta}_{RB}^*$  in specification III, the residual randomness is small. For  $\hat{\beta}_{RB}^*$  in specification III, however, redrawing  $\zeta$  yields substantially different point estimates. This issue persists even if we increase the number of  $\zeta$  draws to 1,000,000.

Specification	I	II	III	IV
	$\hat{\beta}$	$\hat{\beta}$	$\hat{\beta}$	$\hat{\beta}$
2SLS	0.099	0.081	0.060	0.081
Fuller	0.100	0.084	0.058	0.100
LIML	0.100	0.084	0.057	0.098
$\hat{\beta}_{RB}^*$ ,	0.097	0.085	-0.041	0.056
$\hat{\beta}_{RB}, c = 0.1$	0.098	0.083	0.135	0.066
$\hat{\beta}_{RB}, c = 0.5$	0.098	0.083	0.135	0.066
$\hat{\beta}_{RB}, c = 0.9$	0.098	0.083	0.135	0.066
First Stage F	30.582	4.625	1.579	1.823
QCLR CS	[0.059,0.144]	[0.047,0.127]	[-0.588,0.668]	[0.056,0.150]
<i>Controls</i>				
Base Controls	Yes	Yes	Yes	Yes
Age, Age <sup>2</sup>	No	No	Yes	Yes
SOB	No	No	No	Yes
<i>Instruments</i>				
QOB	Yes	Yes	Yes	Yes
QOB*YOB	No	Yes	Yes	Yes
QOB*SOB	No	No	No	Yes
# instruments	3	30	28	178
Observations	329,509	329,509	329,509	329,509

Table 2: Results for Angrist & Krueger (1991) data. Specifications as in Staiger & Stock (1997):  $Y$  =log weekly wages,  $X$ =years of schooling, instruments  $Z$  and exogenous controls as indicated. QCLR is the quasi-CLR (or GMM-M) confidence set of Kleibergen (2005). Unbiased estimators calculated by averaging over 100,000 draws of  $\zeta$ .

rounded to three decimal places in all specifications. Given that these estimators are more robust to violations of the sign restriction than that with  $c = 0$ , we think it makes more sense to focus on these estimates.

## 6 Conclusion

In this paper, we show that a sign restriction on the first stage suffices to allow finite-sample unbiased estimation in linear IV models with normal errors and known reduced-form error covariance. Our results suggest several avenues for further research. First, while the focus of this paper is on estimation, recent work by Mills et al. (2014) finds good power for particular identification-robust conditional t-tests, suggesting that it may be interesting to consider tests based on our unbiased estimators. By inverting such tests one could obtain confidence sets which are guaranteed to contain the corresponding unbiased point estimates. More broadly, it may be interesting to study other ways to use the knowledge of the first stage sign, both for testing and estimation purposes.

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