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Practical Considerations for Questionable IVs

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Abstract. This paper examines a number of techniques which allow for the construction of bounds estimates based on instrumental variables (IVs), even when the instruments are not valid. The `plausexog` and `imperfectiv` commands are introduced, which implement methods described by Conley et al. (2012) and Nevo and Rosen (2012b) in Stata. The performance of these bounds under a range of circumstances is examined, leading to a number of practical results related to the informativeness of the bounds in different circumstances.

Keywords: IV, instrumental variables, exclusion restrictions, invalidity, plausibly exogenous, imperfect IVs

1 Introduction

Instrumental variables are a work horse estimator in economics as well as in other fields concerned with the causal estimation of relationships of interest. Nonetheless, credible instrumental variables (IVs) are hard to come by. While finding variables which are correlated with an endogenous variable of interest (“relevant” in IV terms) is generally not a challenge, motivating and defending a zero-correlation with unobserved error terms (“validity”) is much less straight-forward.¹

As is well known, validity assumptions in an IV setting are untestable. While partial tests exist (Sargan 1958; Hansen 1982; Kitagawa 2015), these tests are necessary, rather than sufficient, to demonstrate instrumental validity. This often leads to the uncomfortable position where the best estimates for a parameter are based on a strong assumption, for which no definitive proof can be offered.

Given the importance of assumptions regarding the behaviour of unobservables in producing causal estimates in an IV setting, a range of methods have been proposed to preserve causal estimation with weaker assumptions. Work by Manski and Pepper (2000, 2009) loosened the validity assumption replacing strict equalities with (weak) inequalities. Extensions of this work by, among others, Conley et al. (2012); Nevo and Rosen (2012b); Small (2007) propose linear models in an IV framework however with the absence of the traditional IV validity assumption. In each case, with weaker-than-standard assumptions, it is shown that bounds (rather than point estimates) can be produced on an endogenous parameter of interest.

This paper examines a number of recent methodologies for inference with instru-

1. We lay out the classical IV model in section 2 as well as the traditional assumptions leading to consistent estimates of parameters of interest.

ments which (potentially) fail the typical IV validity assumption. In particular, we focus on two recent methodologies which provide bounds on an endogenous variable of interest with as few as one “instrumental variable” which does not necessarily have zero correlation with the unobserved error term.² These methodologies—one from Conley et al. (2012) and one from Nevo and Rosen (2012b)—loosen IV assumptions in different ways, and are relevant to different types of settings in which IVs are suspected not to hold precisely. As we lay out in further detail below, Conley et al. (2012) replace the (exact) exclusion restriction in an IV model with an assumption related to its support or distribution, while Nevo and Rosen (2012b) replace the zero correlation assumption between the instrument and the unobserved error term with an assumption related to the *sign* of the correlation.

IV bounds under weaker-than-standard assumptions are potentially of use in a wide range of applications. Much effort is often spent in empirical work to convincingly argue for the validity of instruments. Nonetheless, the validity of IVs are often questioned. Consider the survey paper of Rosenzweig and Wolpin (2000) which describes a number of “natural” instrumental variables that are not under the control of humans, and hence have been proposed to be valid IVs.³ Among those listed, most have been questioned on various grounds. The use of season of birth (Angrist and Krueger 1991) was suggested to be potentially correlated with a number of relevant correlates (Bound et al. 1995) and then documented to be directly related to maternal characteristics in the US (Buckles and Hungerman 2013). The use of twins (Rosenzweig and Wolpin 1980a,b) was later questioned based on birth spacing and parental responses (Rosenzweig and Zhang 2009) and parental behaviour *in utero* Bhalotra and Clarke (2016), and the use of the gender mix of children (Angrist and Evans 1998) was shown to have other relevant effects on family behaviour (Dahl and Moretti 2008).

However, often critiques of IVs imply minor, rather than major, correlations between instruments and unobserved behaviour. In this paper we introduce two Stata commands which permit for the construction of valid bounds in precisely circumstances like this. These are the `plausexog` module, based on Conley et al. (2012)’s Plausibly Exogenous inference, and `imperfectiv`, based on Nevo and Rosen (2012b)’s Imperfect Instrumental Variables inference. These methods allow for the construction of IV bounds under weaker-than-traditional assumptions. We lay out the basics of each methodology, the usage of each of these commands, and discuss a number of factors to be considered when confronted with questionable IVs. As we lay out further below, each method relaxes IV assumptions in a different way: `plausexog` via the exclusion restriction, and `imperfectiv` via correlations between the IV and the error term. As we show, the relative informativeness of `plausexog` and `imperfectiv` bounds depends on the particular context, with each being particularly suitable in different (invalid) IV circumstances. In

2. There are also an alternative set of methodologies proposing inference in an IV framework without strict validity assumptions, however using more than one (invalid) IV. For example, Small (2007) proposes a case with as few as two instruments, and Kolesár et al. (2015); Kang et al. (2016) describe estimation procedures with many invalid or invalid and valid instruments.

3. In particular they listed 5 outcomes arising from natural (biological or climate) processes that were potentially random and had been used as instruments. These were (i) twin births, (ii) human cloning (monozygotic twinning), (iii) birth date, (iv) gender, and (v) weather events.

what follows of this paper, we document the scope of each procedure, and suggest that these commands should be considered as complements, rather than substitutes, in the applied researcher’s toolbox.

2 Methodology

The habitual Instrumental Variables model is laid out as follows:

$$Y = X\beta + \varepsilon \quad (1)$$

$$X = Z\Pi + V \quad (2)$$

where Y is an outcome variable of interest, X a matrix of (potentially endogenous) treatment variables, and Z a matrix of instruments which are uncorrelated by assumption with the error term ε . Presuming that X contains an endogenous variable (or variables), the parameter vector β is not consistently estimable via OLS. The existence of valid instruments Z which can be excluded from equation 1 thus drives the estimation of the structural parameters of interest β .

Validity is typically presented in one of two formats. The first in terms of the exclusion restriction, or that the instruments Z have no direct effect on Y once purged of their effect on X . The second is in terms of correlations with unobservables: if Z is uncorrelated with ε , instrumental validity is fulfilled. While either condition is appropriate to motivate consistent estimation of parameters in IV models⁴, we consider both here as they provide alternative approaches to conceptualise failures of the underlying assumption in IV.⁵ If it can be credibly argued that the validity assumption holds, two-stage least squared (2sls) estimates of β from equation 1 are consistent.

However, as discussed in the introduction, this assumption is untestable given that it is related to the behaviour of the unobservable ε . Even if instruments are shown to be unrelated to many observable factors, or to pass over-identification tests, this does not provide definitive proof of their validity. This has given rise to a modern literature focused on relaxing these assumptions. Rather than driving estimation and

4. And indeed, their implications are equivalent in the simultaneous equations framework laid out here (additional discussion related to their difference in the potential outcomes interpretation of the Rubin (1974) casual model can be found in Angrist and Pischke (2008, pp. 85-91)). If we consider two structural equations of the form

$$y = \beta_0^a + \beta_1^a X + \varepsilon^a,$$

and

$$y = \beta_0^b + \beta_1^b X + \beta_2^b Z + \varepsilon^b$$

failure of the exclusion restriction means that β_2^b is not equal to zero. However, a non-zero value of β_2^b also implies that $\rho_{Z,\varepsilon^a} \neq 0$ (where ρ is the covariance), and, by definition $\rho_{y,\varepsilon^a} > 0$. Thus, assuming that the exclusion restriction holds in this setup is equivalent to assuming that Z is uncorrelated with the structural error term. And *vice versa*, once the conditional correlation between the instrument and the error term are assumed to be zero, the exclusion restriction assumption is superfluous.

5. In this paper we do not consider with much length the relevance assumption. This assumption is testable, and considerable literature exists on this topic.

inference from dogmatic priors which require strict equalities in the exclusion restriction or correlations, it has been shown that bounds on parameters can be estimated under considerably loosened conditions. Central among these are the bounding process of Conley et al. (2012) and Nevo and Rosen (2012b). While both loosening traditional assumptions to form IV bounds, the precise manner in which this is undertaken in each case is different. The suggestion of Conley et al. (2012) is to relax the exclusion restriction, where rather than assuming that it is exactly equal to zero, some range is allowed for the coefficient on the instrument in the structural equation. They allow the exclusion restriction to fail, but proceed with estimation by restricting the failure to some range. Nevo and Rosen (2012b), on the other hand, document that assuming a *direction* for the covariance between the instrument and the stochastic error ε can result in two-sided bounds for the parameter of interest β . We consider each method as well as the resulting bounds below, before turning to the practicalities of estimation later in this article.

Relaxing the Exclusion Restriction Assumption The classical IV system of equations defined in 1 and 2 is a restricted version of the below:

$$Y = X\beta + Z\gamma + \varepsilon \quad (3)$$

$$X = Z\Pi + V. \quad (4)$$

We arrive at 1 and 2 by imposing the (strong) prior that $\gamma = 0$, resulting in point estimates of the parameter vector of interest β . One way to loosen the IV assumptions is to remove the assumption that γ is precisely equal to zero. A range of literature seeks to restrict the *range* of this unidentified parameter (or parameter vector) γ without assuming that it is exactly equal to zero. Manski and Pepper (2000) document inference in IV settings where the strict equality in $\gamma = 0$ is replaced by a weak inequality, giving “Monotone Instrumental Variables”.⁶ Earlier work by Hotz et al. (1997) propose bounding in an IV setting where the exclusion restriction is assumed to hold for some part of the population, and not hold for others, requiring an estimate or assumption regarding the degree of contamination of the IV. More recent extensions including Small (2007) and Conley et al. (2012) seek to further restrict the range of values for γ while still allowing the exclusion restriction to fail, either by searching for plausible parameters in overidentified systems (Small 2007), or by allowing researchers to specify priors for γ in a range of flexible ways (Conley et al. 2012).

In what remains, when considering relaxations of the exclusion restriction, we will follow the procedure implemented by Conley et al. (2012). This procedure allows for valid inference using an instrumental variable (or variables) even when the exclusion restriction does not hold precisely. They document a number of procedures which can be followed, depending on a researcher’s prior belief regarding the degree of failure of the exclusion restriction, and the amount of structure which the researcher is willing to place on this violation. In particular, assumptions can be made regarding the range of

6. Strictly speaking, Manski and Pepper’s approach does not require 3 and 4, as it is based in a non-parametric setting, where instruments are assumed to monotonically impact conditional expectations, and so involves conditional means rather than covariances.

values that γ can take in 3, regarding the entire distribution for γ , or a fully Bayesian approach can be undertaken, in which as well as a prior for the γ term, priors for each model parameter as well as the distribution of error terms must be provided.

The first of these approaches consists of simply replacing the original exclusion restriction assumption of $\gamma = 0$ with an assumption regarding the minimum and maximum values which γ may take. This allows for circumstances in which γ can be assumed to be entirely positive or negative, or alternatively, overlapping zero. Estimation thus consists of producing confidence intervals on β for a range of models of the following form, where γ_0 refers to values from an (appropriately binned) range $[\gamma_{\min}, \gamma_{\max}]$.

$$(Y - Z\gamma_0) = X\beta + \varepsilon$$

In each case, the above model can be estimated by 2SLS using the transformed dependent variable $Y - Z\gamma_0$. Conley et al. (2012) name this approach the “Union of Confidence Intervals” (UCI) approach, as in practice bounds consist of the union of all confidence intervals in the assumed range of $\gamma_0 \in [\gamma_{\min}, \gamma_{\max}]$.

Additional structure can be placed on assumptions regarding γ to relax the exclusion restriction. If, rather than assuming simple maximum and minimum values for γ , a distributional assumption is made, bounds on the parameter β can be calculated using the entire assumed distribution for γ . This allows, among other things, for more or less weight to be placed on values of γ which are perceived to be more or less likely, for example by placing more weight on values of γ close to zero, and less weight on values of γ further away.⁷ As Conley et al. (2012) document, replacing the assumption that $\gamma = 0$ with an assumption that $\gamma \sim F$ (where F is some arbitrary distribution) implies the following approximate distribution for $\hat{\beta}$:

$$\hat{\beta} \stackrel{a}{\sim} \mathcal{N}(\beta, V_{2SLS}) + A\gamma. \quad (5)$$

Here, the original 2SLS asymptotic distribution is inflated by a second term, where $A = (X'X(Z'Z)^{-1}Z'X)^{-1}(X'Z)$, and γ is assumed to follow some arbitrary distribution F , assumed independent of $\mathcal{N}(\beta, V_{2SLS})$. This approach is called the “Local to Zero” (LTZ) approximation, and treats uncertainty regarding γ and sampling uncertainty as of a similar magnitude.

Practically, estimating bounds on β using the result in 5 can proceed in a number of ways. A simulation-based approach can be used which allows for *any* type of distribution for γ , or, if γ is assumed to have a Gaussian distribution, this leads to a convenient analytical bounds formula for β . In the case that γ is assumed to follow a Gaussian distribution: $\mathcal{N}(\mu_\gamma, \Omega_\gamma)$, bounds on β from 5 simplify to:

$$\hat{\beta} \stackrel{a}{\sim} \mathcal{N}(\beta + A\mu_\gamma, V_{2SLS} + A\Omega_\gamma A').$$

7. Conley et al. (2012) also discuss how this can be housed in the union of confidence interval approach discussed above by giving more or less weight to certain values in the $[\gamma_{\min}, \gamma_{\max}]$ range, however the present approach allows for the flexibility to easily include any distributions for γ , and so we focus on this here.

If a non-Gaussian prior for γ is assumed, Conley et al. (2012) outline a simulation algorithm for calculating bounds on β . This procedure consists of generating a large number of draws of the following quantity, which calculates deviations of $\hat{\beta}$ from β where draws from the assumed γ distribution are included in the second part of the formula:

$$\eta \sim \mathcal{N}(0, V_{2SLS}) + A\gamma.$$

In practice, with a large number of draws of η in hand, confidence intervals on β can be found by subtracting desired quantiles of the η distribution from $\hat{\beta}$ in equation 5. Both the exact and simulation-based method can be implemented using the `plausexog` ado described in further detail later in this article.⁸

Finally, even further structure can be placed on the exclusion restriction if rather than simply assuming a range of values for γ (UCI), or a distribution for γ (LTZ), a full Bayesian procedure is followed. This requires assuming not only a distribution for γ , but also a prior for error terms and other model parameters. We do not go into additional detail regarding this Bayesian procedure here, however direct interested readers to Conley et al. (2012), and computational implementations (in R) as *bayesm* (Rossi 2015).

Relaxing IV Correlation Assumptions The classical IV approach described in 1 and 2 produces consistent estimates of β based on the (unobservable) validity assumption $E[Z\varepsilon] = 0$. Bounds inference in an IV setting can proceed with weaker than classical assumptions by replacing the validity (zero covariance) assumption with an assumption on the sign of the covariance. Nevo and Rosen (2012b) proceed with a linear IV model in which the zero covariance assumption is loosened in this way. Their results extend an earlier line of research from Leamer (1981); Klepper and Leamer (1984); Bekker et al. (1987) and Manski and Pepper (2000). Nevo and Rosen (2012b) document that if replacing the demanding zero covariance assumption with an assumption regarding the sign of the covariance between an IV and the stochastic error, this leads to convenient and easily estimable bounds in the linear IV model.

To define these bounds, we follow Nevo and Rosen (2012b) in using $\rho_{x\varepsilon}$ to signify correlation and $\sigma_{x\varepsilon}$ to signify covariance, and σ_x to signify standard deviation, where subscripts make clear the random variables considered. The traditional IV validity assumption is thus denoted $\rho_{z\varepsilon} = 0$. Nevo and Rosen (2012b) replace this validity assumption with the an assumption regarding only the *direction* of correlation between and instrument Z and the stochastic error term ε in 1:

$$\rho_{x\varepsilon}\rho_{z\varepsilon} \geq 0. \tag{6}$$

8. While it is preferable to use the exact result if a Gaussian prior is assumed for the distribution of γ , a Gaussian prior can also be included using the simulation-based algorithm described in Conley et al. (2012), and, assuming that a large enough number of draws of η are taken, these two approaches return identical bounds. By default, `plausexog` draws 5,000 realizations of η , and this generally leads to very similar bounds in the simulated and closed form approaches with a Gaussian prior. The number of draws of η can be changed by users. Where possible, more draws should always be preferred.

This assumption (Nevo and Rosen (2012b)’s “assumption 3”⁹) thus states that the instrument has (weakly) the same direction of correlation with the omitted error term as the endogenous variable X .

This assumption, combined with a fourth assumption, gives the definition of an “Imperfect Instrumental Variable” as an IV which has the same direction of correlation with the unobserved error term as the endogenous variable of interest x , however is less endogenous than x :

$$|\rho_{x\varepsilon}| \geq |\rho_{z\varepsilon}|. \quad (7)$$

Based on 7, we can define a quantity denoting the relative degree of correlation between the instrument and the error term compared with the same correlation between the original endogenous variable and the stochastic error term. This quantity, which captures how much less flawed the instrument is than the endogenous variable $\lambda^* = \frac{\rho_{z\varepsilon}}{\rho_{x\varepsilon}}$, is not known without further assumptions, however, it is clearly bounded between 0, in the case that the traditional IV assumption holds, and 1 in the case where 7 holds with equality.

Ignoring for now that λ^* is unknown, if it were known a new compound valid instrument could be constructed by forming: $\sigma_X Z - \lambda^* \sigma_Z X$. The logic behind this instrument is that the endogenous components of the original endogenous variable X and the (less) endogenous Z can be cancelled out, and hence

$$E[(\sigma_X Z - \lambda^* \sigma_Z X)\varepsilon] = \sigma_X \sigma_{Z\varepsilon} - \lambda^* \sigma_Z \sigma_{X\varepsilon} = 0,$$

is a valid instrument. Nevo and Rosen (2012b)’s proposal is to replace this above valid instrument, denoted $V(\lambda^*) = (\sigma_X Z - \lambda^* \sigma_Z X)$ with $V(1) = (\sigma_X Z - 1\sigma_Z X)$, the instrument in the limit case implied by 7. While this will not give point estimates on the parameter of interest β , it will allow for the construction of bounds in certain circumstances discussed below.

Consider now the probability limits of three different estimators, β^{ols} , the original estimand of β using endogenous X in a standard linear regression, β_z^{iv} , the 2SLS estimator using the Imperfect IV, and $\beta_{v(1)}^{iv}$, the 2SLS estimator of the transformed variable described above. Based on the above two assumptions in 6 and 7, these parameters are not guaranteed to bound the true parameter β . However, if the instrument is negatively correlated with the endogenous variable, $\sigma_{xz} < 0$, this allows for the construction of upper and lower bounds on the true parameter β . These bounds are described in panel A of table 1. The right-hand panel describes the case in which Nevo and Rosen (2012b)’s Assumption 4 is not maintained, and hence $\beta_{v(1)}^{iv}$ is not used. In this case, the original β_z^{iv} parameter and the OLS estimate β^{ols} bound β , with the upper and lower bounds depending on the assumed correlation between X and ε (and hence Z and ε).¹⁰ However, if the correlation between X and Z is positive, only one sided bounds can be

9. Nevo and Rosen (2012b) make a series of standard assumptions regarding the sampling process and any exogenous covariates included in the model, as assumptions 1 and 2.

10. To see why the IV and OLS parameters bound the true parameter β note that in the simple linear model described in 1-2, we can write $\beta^{ols} = \beta + \frac{\sigma_{x\varepsilon}}{\sigma_x^2}$, and $\beta_z^{iv} = \beta + \frac{\sigma_{z\varepsilon}}{\sigma_{xz}}$. Given that σ_{xz} is assumed negative (a testable assumption), and σ_x^2 is positive, these two parameters bound β .

formed. In the case that assumption 4 (equation 7) is maintained, this leads to a further tightening of the bounds, given that the inconsistent β^{ols} parameter can be replaced by the less-inconsistent $\beta_{v(1)}^{iv}$ parameter.¹¹ Once again however, if the correlation between the endogenous variable and the instrument is not negative, informative bounds cannot be formed, leading to only one sided bounds for β . Both bounds with and without assumption 4 can be produced by the `imperfectiv` ado described later in this paper.

In the discussion up to this point, we have justified the relaxation of the instrumental validity assumption when one imperfect IV is present. However, Nevo and Rosen (2012b) demonstrate that if *more* than one imperfect IV is available, this result can be used to potentially generate tighter bounds¹², and, under an auxiliary assumption, produce two-sided bounds where previously only one-sided bounds were observed. In the simplest case, without further restrictions on the nature of each Imperfect IV (beyond the fact that they each meet assumptions 3 and 4), the bounding procedure consists of a search among all Imperfect IVs and the OLS estimate to generate the tightest set of bounds possible given the assumptions maintained in 6 and 7. This can be seen as a generalisation of Panel A of table 1, where each β^{iv} parameter is replaced with its min (for upper bounds) or max (for lower bounds).

Finally, Nevo and Rosen (2012b) show that if more than one instrument is available, and if one instrument is assumed to be better than another in both relevance and validity, then two sided bounds can be produced, *even* if the original IIVs are positively correlated with the endogenous variable X . Consider two IIVs, Z_1 and Z_2 , where 6 is assumed to hold, $\sigma_{xz_1} > \sigma_{xz_2}$ (Z_1 is more relevant than Z_2), and it is assumed that $\sigma_{\varepsilon z_1} < \sigma_{\varepsilon z_2}$ (Z_1 is less endogenous than Z_2). Then, the production of a new instrument:

$$\omega(\gamma) = \gamma Z_2 - (1 - \gamma) Z_1$$

will lead to two sided bounds so long as $\sigma_{\omega(\gamma)u} \geq 0$ and $\sigma_{\omega(\gamma)x} < 0$. These bounds are described in Panel B of table 1, and are summarised as Nevo and Rosen’s Proposition 5. In practice, Nevo and Rosen (2012b) suggest using a value of $\gamma = 0.5$ to form the re-weighted IIV. In the `imperfectiv` ado, $\gamma = 0.5$ is used by default, and a “better” and “worse” IIV must be indicated by the user to produce bounds in this case.

11. To see why β_z^{iv} and $\beta_{v(1)}^{iv}$ bound the true parameter, we can start from β_z^{iv} and β^{ols} which we know provide bounds. Given that $\beta_{v(1)}^{iv}$ is a weighted average of β_z^{iv} and β^{ols} assuming $\lambda = 1$ (see Nevo and Rosen (2012b) for full details), this estimate will remove part of the bias from the β^{ols} parameter, such that $\beta_{v(1)}^{iv}$ and β^{ols} (weakly) share the same direction of inconsistency, but $\beta_{v(1)}^{iv}$ is less inconsistent.

12. Recent work from Wiseman and Sørensen (2017) suggests under an alternative (implicit) assumption that Nevo and Rosen’s bounds can, in some cases, be further tightened, especially when instruments are weak.

Table 1: Summary of the propositions of Nevo and Rosen (2012)

		Assumption 4		No Assumption 4	
		$\sigma_{xz} < 0$	$\sigma_{xz} > 0$	$\sigma_{xz} < 0$	$\sigma_{xz} > 0$
Panel A: One Instrument					
$\rho_{x\varepsilon} > 0$	$\beta_z^{iv} \leq \beta \leq \beta_{v(1)}^{iv}$	$\beta \leq \min\{\beta_{v(1)}^{iv}, \beta_z^{iv}\}$	$\beta_z^{iv} \leq \beta \leq \beta^{ols}$	$\beta \leq \min\{\beta^{ols}, \beta_z^{iv}\}$	
$\rho_{x\varepsilon} < 0$	$\beta_{v(1)}^{iv} \leq \beta \leq \beta_z^{iv}$	$\beta \geq \max\{\beta_{v(1)}^{iv}, \beta_z^{iv}\}$	$\beta^{ols} \leq \beta \leq \beta_z^{iv}$	$\beta \geq \max\{\beta^{ols}, \beta_z^{iv}\}$	
Panel B: Multiple Instruments with “Proposition 5”					
$\rho_{x\varepsilon} > 0$	-	$\beta_\omega^{iv} \leq \beta \leq \min\{\beta_{z_1}^{iv}, \beta_{z_2}^{iv}, \beta^{ols}\}$	-	$\beta_\omega^{iv} \leq \beta \leq \min\{\beta_{z_1}^{iv}, \beta_{z_2}^{iv}, \beta_{v_1(1)}^{iv}, \beta_{v_2(1)}^{iv}, \beta_{v^*(1)}^{iv}\}$	
$\rho_{x\varepsilon} < 0$	-	$\beta_\omega^{iv} \geq \beta \geq \min\{\beta_{z_1}^{iv}, \beta_{z_2}^{iv}, \beta^{ols}\}$	-	$\beta_\omega^{iv} \geq \beta \geq \min\{\beta_{z_1}^{iv}, \beta_{z_2}^{iv}, \beta_{v_1(1)}^{iv}, \beta_{v_2(1)}^{iv}, \beta_{v^*(1)}^{iv}\}$	
<p>NOTES: Full details of the bounding procedure and assumptions are available in Nevo and Rosen (2012b). Notation is defined in section 2 of this paper. The final case in Panel B for $\rho_{x\varepsilon} < 0$ is not shown in Nevo and Rosen (2012b), however the $\rho_{x\varepsilon} > 0$ case is shown to hold without loss of generality in Nevo and Rosen (2008), giving the reverse case also shown here. In each case, the statement $\rho_{x\varepsilon} > 0$ implies $\rho_{z\varepsilon} > 0$, and $\rho_{x\varepsilon} < 0$ implies $\rho_{z\varepsilon} < 0$, ie 6 is always assumed to hold.</p>					

3 Stata Commands

We describe the basic syntax of two commands below. These two commands are `plausexog`, which implements Conley et al. (2012)’s bounds relaxing the exclusion restriction, and `imperfectiv` which implements the Nevo and Rosen (2012b) bounding procedure by relaxing the traditional validity assumption. We examine the commands in turn in sections 3.1 and 3.2. We also provide extended examples of their syntax and use by replicating empirical examples from Nevo and Rosen (2012b) and Conley et al. (2012) in Appendix 1.

3.1 The `plausexog` command

Syntax

The `plausexog` command is closely related to Stata’s instrumental-variables regression command, with arguments describing the prior expectation of the degree of the violation of the exclusion restriction. The generic syntax of the command is as follows.

```
plausexog method depvar [varlist1] (varlist2 = varlist_iv) [if] [in] [weight]
    [, level(#) vce(vcetype) gmin(numlist) gmax(numlist) grid(#)
    mu(numlist) omega(numlist) distribution(name, params) seed(#)
    iterations(#) graph(varname) graphmu(numlist) graphomega(numlist)
    graphdelta(numlist) *]
```

where *method* must be specified as either `uci` (union of confidence intervals) or `ltz` (local to zero), depending on the desired estimator. The remainder of the syntax follows Stata’s `ivregress` syntax, where first any exogenous variables are specified as *varlist1*, then the endogenous variable(s) as *varlist2*, and final “plausibly exogenous” instruments in *varlist_iv*.

Options

`level(#)` Set confidence level; default is `level(0.95)`.

`vce(vcetype)` determines the type of standard error reported in the estimated regression model, and allows standard errors that are robust to certain types of misspecification. *vcetype* may be `robust`, `cluster clustvar`, `bootstrap`, or `jackknife`.

`gmin(numlist)` Specifies minimum values for γ on plausibly exogenous variables (`uci` only). One `gmin` value must be specified for each plausibly exogenous variable.

`gmax(numlist)` Specifies maximum values for γ on plausibly exogenous variables (`uci` only). One `gmax` value must be specified for each plausibly exogenous variable.

`grid(#)` Specifies number of points (in [`gmin`, `gmax`]) at which to calculate bounds; default is `grid(2)` (`uci` only).

- mu**(*numlist*) Specifies the mean value for the prior distribution of γ , assuming a Gaussian prior and the LTZ approach. One **mu** value must be specified for each plausibly exogenous variable.
- omega**(*numlist*) Specifies the variance value for the prior distribution of γ , assuming a Gaussian prior and the LTZ approach. One **omega** value must be specified for each plausibly exogenous variable.
- distribution**(*name, params*) allows for non-Gaussian priors for the distribution of gamma. When using the **distribution** option, the **mu** and **omega** option do not need to be specified. Bounds based on non-normal distributions for gamma are calculated using the simulation-based algorithm described in Conley et al. (2012, p. 265) and section 2. Accepted distributions names are: **normal**, **uniform**, **chi2**, **poisson**, **t**, **gamma**, and **special**. When specifying any of the first six options, parameters must be specified along with each of these distributions. For **normal**, parameters are the assumed mean and standard deviation; for **uniform**, the parameters are the minimum and maximum; for **chi2** (Chi squared) it is the degrees of freedom; for **poisson** it is the distribution mean, for **t** it is the degrees of freedom; and for **gamma** it is the shape and scale of the assumed distribution. For any assumed distribution of gamma which is not contained in the previous list, **special** can be specified, and a variable can be passed which contains analytical draws from this distribution. If more than one plausibly exogenous variable is used, the relevant parameters must be specified for each plausibly exogenous variable. Note that although a Gaussian prior is allowed in this format, if a Gaussian prior is assumed it is preferable to use the **mu**(#) and **omega**(#) options, as these give an exact, rather than approximate (simulated) set of bounds.
- seed**(#) Sets the seed for simulation-based calculations when using a non-Gaussian prior for the LTZ option. Only required when specifying the **distribution** option.
- iterations**(#) Determines the number of iterations for simulation-based calculations when using a non-Gaussian prior for the LTZ option; default is **iterations**(5000). In Stata IC and Small Stata the number of iterations cannot exceed the maximum matrix size permitted by Stata. As such, these are set to 800 and 100 respectively. The **distribution** option should be used with care in these versions of Stata.
- graph**(*varname*) Indicates that a graph should be produced of bounds over a range of assumptions related to the failure of the exclusion restriction. The *varname* indicates the name of the (plausibly exogenous) Z variable that the user wishes to graph. In the UCI method, confidence intervals will be graphed, while in the LTZ approach both confidence intervals and a point estimate will be graphed over a range of gamma values.
- graphmu**(*numlist*) This option must be used with the LTZ model when a graph is desired. This provides the values for a series of **mu** values for each point desired on the graph. Each point refers to the mean value of γ assuming a Gaussian prior.
- graphomega**(*numlist*) This option must be used with the LTZ model when a graph is desired. Each value for **omega** corresponds to the value in the **graphmu** list, and

specifies the variance of the Gaussian prior at each point.

`graphdelta(numlist)` Allows for the plotting of values on the graph produced above. If not specified, the values in `graphmu` will be plotted on the horizontal axis.

* Any other options documented in [G] **twoway options** are allowed. This overrides default graph options such as title and axis labels.

Returned Objects

`plausexog` is an eclass program, and returns a number of elements in the `e()` list. It returns scalar values for the lower and upper bounds of each endogenous variable as `e(lb_endogname)` and `e(ub_endogname)` respectively, where `endogname` will be the name of the variable in a given application. In the case where Conley et al. (2012)'s LTZ approach is used with an assumption of normality, two matrices are also returned: `e(b)` and `e(V)`. These are the coefficient vector and variance-covariance matrix of the estimated parameters based on the plausibly exogenous model.

3.2 The imperfectiv command

Syntax

The `imperfectiv` command is also closely related to Stata's instrumental-variables regression command, with arguments describing correlation between the endogenous variable and the unobservable to replace the validity of the instruments. The generic syntax of the command is as follows.

```
imperfectiv depvar [varlist1] (varlist2 = varlist_iv) [if] [in] [weight] [,
    level(#) vce(vcetype) ncorr prop5 noassumption4 short ]
```

The syntax follows Stata's `ivregress` syntax, where first any exogenous variables are specified as `varlist1`, then the endogenous variable as `varlist2`, and finally "imperfect" instruments in `varlist_iv`.

Options

`level(#)` Set confidence level; default is `level(0.95)`.

`vce(vcetype)` Determines the type of standard error reported in the estimated regression model, and allows standard errors that are robust to certain types of misspecification. `vcetype` may be `robust`, `cluster clustvar`, `bootstrap`, or `jackknife`.

`ncorr` Specifies that the correlation between the endogenous variable and the unobservable error is assumed negative; by default this correlation is assumed to be positive.

`prop5` Specifies that proposition 5 of Nevo and Rosen (2012b) should be used in the estimation of bounds. If the correlation between the endogenous variable and each

imperfect instrument is positive, the result of the estimation is an interval with only one bound. If there is more than one imperfect instrument, then proposition 5 of Nevo and Rosen can be used to generate two-sided bounds. If `prop5` is specified, the first two instruments specified in `varlist_iv` are used, and it is assumed that the “better” instrument is listed first. Additional discussion is provided in section 2.

noassumption4 Specifies that assumption 4 of Nevo and Rosen (2012) does not hold; by default this is assumed to hold. Assumption 4 states that the correlation between the imperfect instrument and the unobservable is less than the correlation between the endogenous variable and the unobservable.

short Requests that only bounds on the endogenous parameter of interest are displayed in the output table. When **short** is specified, bounds on all (endogenous and exogenous) parameters are still returned in the matrix discussed below.

Returned Objects

`imperfectiv` is an eclass program, and returns a number of elements in the `e()` list. Identically to `plausexog`, it returns scalar values for the lower and upper bounds of each endogenous variable as `e(lb_endogname)` and `e(ub_endogname)` respectively, where `endogname` will be the name of the variable in a given application. A matrix is also returned as `e(LRbounds)` giving the upper and lower bounds on each endogenous and exogenous variable included in the model.

4 Performance Under Simulation

We demonstrate the usage of the `imperfectiv` and `plausexog` programs under a series of simulations. These simulations allow us examine the behaviour of bounds on the (known) endogenous parameter of interest under a series of different assumptions. In particular, we can compare the behaviour of bounds using the Union of Confidence Interval (UCI) and Local to Zero (LTZ) approach of Conley et al. (2012), and with and without the use of Nevo and Rosen (2012b)’s Assumption 4.

We aim to examine performance of bounds under a wide range of situations. To do so, we consider a linear model in which we allow the correlation between an endogenous variable of interest x and the unobserved error term ε to vary (ie varying the degree of endogeneity of the parameter of interest), and in which the correlation between the “instrument” z and the unobserved compound error term varies (varying the quality of the instrument). In particular, we allow for this in the following two-stage set-up:

$$\begin{pmatrix} z \\ \varepsilon \\ \nu \end{pmatrix} \sim \mathcal{N} \left[\begin{pmatrix} 0 \\ 0 \\ 0 \end{pmatrix}, \begin{pmatrix} 1 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 1 \end{pmatrix} \right]$$

$$\begin{aligned} x &= \pi z + \mu \varepsilon + \nu \\ y &= \beta x + \gamma z + \varepsilon. \end{aligned} \tag{8}$$

Here y is a dependent variable, x an endogenous variable of interest, and z is an imperfect, or plausibly exogenous, instrumental variable. Provided that $\mu \neq 0$, β cannot be estimated consistently via an Ordinary Least Squares regression, and provided that $\gamma \neq 0$, instrumental variables estimates of β will not be consistent under standard assumptions. The instrument z and error terms ε and ν are simulated from independent normal distributions. In traditional 2SLS, γ is assumed to be zero, and hence γz is omitted from the final equation. This leads to a compound error term ($\gamma z + \varepsilon$), which we refer to as η below.

Using this structure, we examine the use of and performance of `imperfectiv` and `plausexog` by varying γ (the degree of instrumental invalidity), and μ , (the degree of endogeneity). We fix π at -0.6 in all simulations, ensuring that the instrument is not weak. The performance of `plausexog` following this data generating process (DGP) is documented below:

```
. set obs 1000
obs was 0, now 1000
. foreach var in u z v w {
2.     gen `var' = rnormal()
3. }
. gen x = -0.6*z + 0.33*u + v
. gen y1 = 3.0*x + 0.10*z + u
.
. plausexog uci y1 (x=z), gmin(0) gmax(0.2)
Estimating Conely et al.'s uci method
Exogenous variables:
Endogenous variables: x
Instruments: z
```

Conley et al. (2012)'s UCI results			Number of obs =		1000
Variable	Lower Bound	Upper Bound			
x	2.9297441	3.0572207			
_cons	-.05187203	.06593162			

```
. plausexog ltz y1 (x=z), mu(0.1) omega(0.01)
Estimating Conely et al.'s ltz method
Exogenous variables:
Endogenous variables: x
Instruments: z
```

Conley et al. (2012)'s LTZ results						Number of obs =		1000
	Coef.	Std. Err.	z	P> z	[95% Conf. Interval]			
x	2.987058	.1846939	16.17	0.000	2.625065	3.349052		
_cons	.0106754	.0342024	0.31	0.755	-.0563601	.0777109		

Above we document the use of `plausexog` with the UCI and LTZ option. In each case we “correctly” specify the prior over the violation of the exclusion restriction. In the UCI case the exclusion restriction is allowed to have support $\in [0, 0.2]$, with the true value simulated being 0.1. In the LTZ option, the exclusion restriction is specified to

fall within a normal distribution mean of 0.1 and variance of 0.01. In each case, bounds on the endogenous variable x contain the true parameter $\beta = 3$.

Below we document the use of `imperfectiv` using the same DGP. We first specify that bounds be calculated without assuming that the instrument is “less endogenous” than the endogenous variable, and then in the second case add this assumption:

```
. imperfectiv y1 (x=z), noassumption4
Nevo and Rosen (2012)'s Imperfect IV bounds      Number of obs =   1000
-----
Variable          Lower Bound    Upper Bound
-----
x                  2.699749     3.262712
-----

. imperfectiv y1 (x=z)
Nevo and Rosen (2012)'s Imperfect IV bounds      Number of obs =   1000
-----
Variable          Lower Bound    Upper Bound
-----
x                  2.699749     3.139827
-----

.
```

These examples document performance of `plausexog` and `imperfectiv` under *one particular* DGP. Below, in table 2, we consider a range of DGPs where we vary γ (within each panel), and μ (across each panel). Here, γ refers to the failure of the exclusion restriction with which Conley et al. (2012) are concerned, and the resulting correlations between x and η (the compound error term) and z and η with which Nevo and Rosen (2012b) are concerned are displayed in subsequent columns. Bounds are then documented under 2 cases in Conley et al. (2012) (the UCI and LTZ approach, each with correctly specified priors), and 2 cases in Nevo and Rosen (2012b) (with and without assumption 4). In the case of Nevo and Rosen (2012b), the assumptions for “No A4” will be met providing that the sign on $\rho_{x,\eta}$ and $\rho_{z,\eta}$ are the same, and will be met for “Assumption 4” only if $\rho_{x,\eta} \geq \rho_{z,\eta}$.

The bounds produced in each case on the endogenous variable of interest are presented in table 2. In nearly all simulations, the bounds include the true value of $\beta = 3$. The only cases in which this is not seen is with those in the right-most columns at the bottom of panel A. This is to be expected, given that in this case, the assumptions underlying the bounds (Assumption 4 of Nevo and Rosen (2012b)) are not met, and hence the `imperfectiv` command should correctly have been run with the `noassumption4` option. In each circumstance, the Conley et al. (2012) bounds contain the true parameter, but this is dependent on correctly specifying the prior over γ , as we ensure in table 2. Given that in practice, knowing the true prior for γ is an empirical challenge (see for example Bhalotra and Clarke (2016)), conservative assumptions on γ may be preferred.

In general, while the procedures of both Nevo and Rosen (2012b) and Conley et al. (2012) allow the strong assumptions relating to unobservables in an IV setting to be loosened, bounds estimates still rely on a willingness to specify *something* about the

Table 2: Performance of Various Bounds under Monte Carlo Simulation

γ	$\rho_{x,\eta}$	$\rho_{z,\eta}$	Plausibly Exogenous				Imperfect IV			
			UCI		LTZ $\mathcal{N}(\mu, \sigma^2)$		No A4		Assumption 4	
Panel A: Minor Correlation between x and ε										
0.1	0.22	0.10	[2.952	3.059]	[2.684	3.318]	[2.744	3.230]	[2.744	3.124]
0.2	0.16	0.19	[2.807	3.213]	[2.563	3.439]	[2.584	3.189]	[2.584	3.046]
0.3	0.11	0.28	[2.664	3.365]	[2.468	3.534]	[2.422	3.149]	[2.422	2.968]
0.4	0.06	0.36	[2.521	3.513]	[2.388	3.614]	[2.259	3.109]	[2.259	2.891]
Panel B: Moderate Correlation between x and ε										
0.1	0.45	0.10	[2.957	3.064]	[2.682	3.319]	[2.739	3.376]	[2.739	3.228]
0.2	0.40	0.19	[2.817	3.223]	[2.560	3.441]	[2.574	3.344]	[2.574	3.159]
0.3	0.34	0.28	[2.678	3.378]	[2.465	3.537]	[2.408	3.312]	[2.408	3.090]
0.4	0.29	0.36	[2.540	3.529]	[2.384	3.618]	[2.241	3.280]	[2.241	3.021]
Panel C: Major Correlation between x and ε										
0.1	0.61	0.10	[2.962	3.069]	[2.681	3.321]	[2.735	3.430]	[2.735	3.277]
0.2	0.56	0.19	[2.826	3.232]	[2.558	3.444]	[2.565	3.405]	[2.565	3.217]
0.3	0.51	0.28	[2.692	3.389]	[2.461	3.540]	[2.394	3.381]	[2.394	3.157]
0.4	0.46	0.36	[2.558	3.541]	[2.379	3.623]	[2.223	3.357]	[2.223	3.097]

95% confidence intervals associated with the parameter β are displayed in square parentheses. The true value of β is 3 in the DGP described in (8). The value of γ in each case is displayed in the left-hand column (between 0.1 and 0.4), and the correlation between x and η and z and η inferred in each case is listed in subsequent columns. Here η refers to the compound error term which causes endogeneity and instrumental invalidity. 1000 simulated observations are used. Different panels allow the correlation between the endogenous variable x and the ε term to vary, making x ‘more’ or ‘less’ endogenous. Confidence intervals for the Plausibly Exogenous UCI case are based on a support assumption implying that the true value of γ is at the mean, and hence is $[0, 2\gamma]$. In the LTZ case, the distribution for γ is assumed to be normal, with mean equal to the value of gamma, and variance equal to $\gamma/10$. Confidence intervals for Imperfect IV estimates are based on assumptions that $\rho_{x,\eta} > 0$ and $\rho_{z,\eta} > 0$ in the “No A4” case, and that $\rho_{x,\eta} \geq \rho_{z,\eta} > 0$ in the “Assumption 4” case. The veracity of each assumption can be determined from displayed correlations in columns 2 and 3.

relationship between instruments and unobservables. Ideally, these assumptions should be well founded in a theory related to the nature of failure of IV validity. In the case of Nevo and Rosen (2012b), a willingness to assume that an instrument is positively or negatively related to unobservables may reflect some underlying model of selection into an instrument or of behavioural response to a particular draw of the instrumental variable. Consider briefly two well known instruments in models of human fertility: the gender mix of children, and the occurrence of twin births. In the case of gender mix of births, Dahl and Moretti (2008) document a “demand for sons”, suggesting that investments following sons may depend *positively* on this particular realisation of the instrumental variable. In the case of twins, Bhalotra and Clarke (2016) document a cross-cutting positive selection of twin births, where many (positive) maternal health behaviours *in utero* increase the likelihood of giving live births to twins (even if twin conception is random). Here, assumptions relating to a positive correlation between the instrument and unobservables seems reasonable based on positive correlations between the instrument and many hard-to-measure and frequently unobserved variables.¹³ As discussed above, the willingness to assume a particular range or distribution for the failure of the exclusion restriction is also an empirical challenge. While in the case of Conley et al. (2012) bounds are constructed based on stronger assumptions than just the sign of the correlation, a benefit of this approach is that it allows for the sign to be indeterminate, if for example, one is concerned that instruments may only be “close” to exogenous but not certain of the direction in which failures of validity occur. We return to these considerations below.

Abstracting now from *why* identifying assumptions may be met, table 2 offers a number of lessons regarding the relative performance of Conley et al. and Nevo and Rosen bounds. Firstly, the bounds on the endogenous parameter using Conley et al. (2012)’s `plausexog` procedure are approximately constant across panels (given a particular value for γ), as the degree of endogeneity of x does not impact the estimated bounds. In the case of Nevo and Rosen (2012b), all else constant, bounds are more (less) wide when the independent variable of interest is more (less) exogenous. This owes to the fact that Nevo and Rosen (2012b) use information from the original endogenous variable to form one side of their bounds (when two-sided bounds are formed). In the limit case when assumption 4 is not assumed, the bound on the OLS estimate of β itself is used.

Secondly, it is noted that bounds from Nevo and Rosen (2012b) are always tighter when Assumption 4 is used (in the case shown in table 2, the upper bound on β always falls). Of course, this is not free, but rather a direct result of the assumption that z is less endogenous than x . In the case that this is true, bounds are both tighter and contain the true parameter, but when assumption 4 is not met, bounds are tighter, but

13. More generally, often intuitively, the likely direction of correlation between an observed variable and an unobserved error term is assumed in empirical applications. For example in simple linear models, the well-known omitted variable bias in OLS can be signed if the correlation between an included variable and the unobservable error is assumed. In Nevo and Rosen’s Imperfect IV application, we are concerned with similar correlations between instrumental variables and unobserved errors. Whether or not a reasonable assumption regarding the potential correlation between an instrument and the error term exists, depends entirely on the phenomenon under study.

do *not* contain the true parameter.

Finally, we note that in this case, adding additional structure in the Conley et al. (2012) bounding procedure via the Local to Zero approach actually results in wider bounds. This is a direct result of the parameters assumed in each case. In the UCI case, we allow for a support of $[0, 2 \times \gamma]$ for each implementation, while the LTZ case assumes that $\gamma \sim \mathcal{N}(\gamma, \gamma/10)$, which consistently results in a probability distribution for γ which has a considerable probability mass outside of the values allowed in the UCI approach. This should not be seen as necessarily representative of the use of the UCI and LTZ approaches. Frequently, the LTZ approach leads to tighter bounds, given the additional structure placed on the prior for γ . Indeed, in the above simulations, if we were to use a Gaussian prior in the LTZ approach with an identical variance of a uniform spanning the UCI γ_{min} and γ_{max} values, bounds in the LTZ approach would be tighter than those in the UCI approach. This is a direct result of placing greater weight on values closer to the true value of γ when using the normal prior. Unlike the Nevo and Rosen (2012b) method, the Conley et al. (2012) method allows for a prior that the instrument may be positively related, negatively related, or unrelated with the unobserved error term. However, the additional flexibility of the Conley et al. (2012) method also comes with the caveat that rather than knowing the sign of the correlation between the instrument and the error term, we must assume something about the *magnitude* of the failure of the exclusion restriction.

While Nevo and Rosen (2012b) are based on two assumptions and no further priors are required, (as documented in the two columns of table 2), Conley et al. (2012) bounds are based on parametric priors which can take an unlimited range of values. Thus, if using Conley et al. (2012) bounds, it may be particularly useful or illustrative to visualize bounds based on a range of values for a particular parametric prior¹⁴. This can be achieved using the graphing capabilities of `plausexog`. We document an example of this code below, which produces Figure 1a below.¹⁵

```
. gen x = 0.33*u + 0.6*z + v
. gen y3 = Beta*x + 0.3*z + u
.
. quietly pausexog ltz y3 (x=z), omega(0.01) mu(0.3) graph(x) graphomega(0 0.0
> 225 0.09 0.2025 0.36 0.5625) graphmu(0 0.15 0.3 0.45 0.6 0.75) graphdelta(0 0
> .15 0.3 0.45 0.6 0.75) scheme(sj) ytitle(Estimated {&beta}) xtitle({&delta})
> xlabel(0 "0" 0.2 "0.20" 0.4 "0.40" 0.6 "0.60" 0.8 "0.80") legend(order(1 "Poi
> nt Estimate (LTZ)" 2 "CI (LTZ)")) ylabel(0(1)5)
```

Figure 1a assumes a Gaussian (Normal) prior for γ in the LTZ approach of Conley et al. (2012), however varying the mean and variance. Bounds at each point on the graph are based on the assumption that $\gamma \sim \mathcal{N}(\delta, \delta^2)$. Figure 1b compares the bounds from the Gaussian prior to bounds based on a uniform prior which assumes that $\gamma \sim U(0, 2 \times \delta)$.

14. A comprehensive example of this procedure is provided in the original Conley et al. (2012, p. 267) paper. We show how to replicate a portion of their results using the `plausexog` ado in Appendix 1.2.

15. All code in the paper is made available on one of the authors' websites, currently at www.damianclarke.net/replication/.

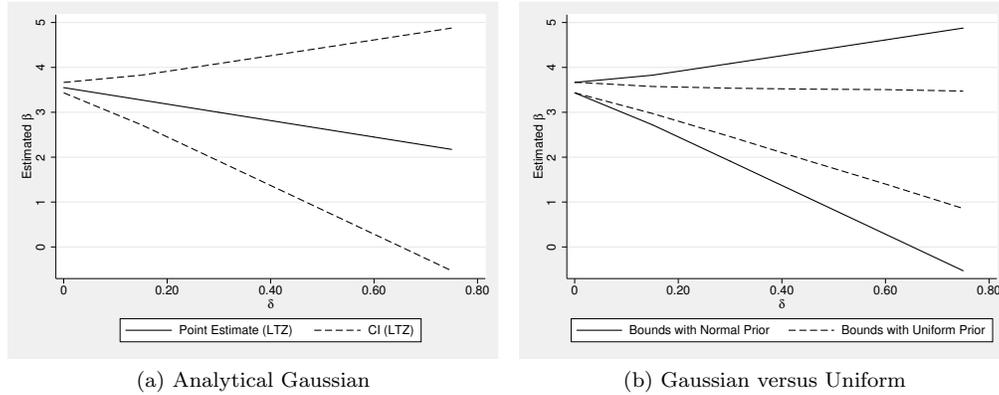


Figure 1: Plausibly Exogenous Bounds Varying Prior Assumptions

The true value for γ is 0.3, and the true value for β is 3. This allows for the comparison of the bounds estimator over a range of priors for γ . We observe (in figure 1a) that the true parameter is contained in the bounds only when the mean of the exclusion restriction is sufficiently high to approach the true value, and that, as in table 2, the bounds grow as the prior allows for additional probability mass on more extreme values of the violation of the exclusion restriction. In each case, classical IV imposing the exact assumption that $\gamma = 0$ would result in confidence intervals considerably above the true population parameter.

5 Conclusion

In this paper we discuss a number of issues involved in the estimation of bounds when examining a causal relationship in the presence of endogenous variables. These types of bounding procedures are likely to be particularly useful given the difficulties inherent in IV estimation, and challenges in convincingly arguing for IV validity, or the exclusion restriction in an IV model.

We introduce two procedures for estimating bounds in Stata: `imperfectiv` for Nevo and Rosen (2012b)’s “Imperfect Instrumental Variable” procedure, and `plausexog` for Conley et al. (2012)’s “Plausible Exogeneity”. In documenting these procedures, we lay out a number of considerations when implementing each bounding process.

Nevo and Rosen (2012b)’s bounds are particularly appropriate when one is convinced of the direction of correlation of an IV with an unobserved error term, but not necessarily its magnitude. The Conley et al. (2012) procedure, on the other hand, is well-suited for situations in which the direction of correlation need not be known (but can be known), but in which the practitioner has some belief over the magnitude of the IV’s importance in the system of interest. All else constant, Nevo and Rosen (2012b) bounds perform

relatively better when the endogenous variable is less correlated with unobservables, while Conley et al. (2012) bounds perform equally well regardless of the correlation between the endogenous variable of interest and unobservables. Finally, while Conley et al. (2012) bounds are often based on more parametric or otherwise stronger assumptions related to the unobservable behaviour of IVs, it is simple to undertake sensitivity testing of estimated bounds' stability to changes in these assumptions, and such sensitivity tests are encouraged when dealing with questionable IVs.

Given that these methodologies loosen IV assumptions in different ways, and are well-suited to different types of (classically invalid) IVs, we suggest that these methodologies should be seen as a complement, rather than a substitute in the empirical researcher's toolbox. The ease of use of each methodology, and their ability to recover parameters under a broad range of failures of IV assumptions, suggests that these procedures should act as a go-to consistency test in the increasingly large number of cases where concerns exist regarding the veracity of instrumental variables.

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Appendices

1 Empirical Examples Using Original Data

We illustrate the performance of each of the `imperfectiv` and `plausexog` programs in Stata by replicating empirical examples from Nevo and Rosen (2012b) and Conley et al. (2012). These use data from the original papers¹⁶ and the syntax of each command as laid out in section 3.

1.1 Nevo and Rosen (2012b)'s Demand for cereal Example

Below we replicate the bounds calculated by Nevo and Rosen (2012b) in their empirical application examining the demand for cereal. We use the `imperfectiv` command described above to calculate bounds. This syntax replicates the results in table 2 of Nevo and Rosen (2012b, p. 667), and in particular columns 3 and 4 where the Imperfect IV methodology is used.

We first show the case where “Assumption 4” is not imposed, and output bounds on both the endogenous and each exogenous variable, and then replicate the results assuming that “Assumption 4” holds. In the second case, we only display the bounds on the endogenous variable of interest using the `short` option to simplify output. We note that in each case the results displayed here are marginally different (at the third decimal place) given that the `imperfectiv` command estimates 2SLS results using Stata’s current `ivregress` command, in place of the now outdated `ivreg` command. If using `ivreg`, bounds replicate those in Nevo and Rosen (2012b) exactly.

(Continued on next page)

16. Both of these datasets are available for public download from the Harvard Dataverse. Refer to Rossi et al. (2012) and Nevo and Rosen (2012a) for full details.

```

. use NevoRosen2012.dta
(Nevo and Rosen's (2012) REStat cereal demand example)

.
. replace addv=addv/10
(986 real changes made)

. local w addv bd1 bd2 bd3 bd4 bd5 bd6 bd7 bd8 bd9 bd10 bd11 bd12 bd13 bd14 bd1
> 5 bd16 bd17 bd18 bd19 bd20 bd21 bd22 bd23 bd24 bd25 dd2 dd3 dd4 dd5 dd6 dd7
> dd8 dd9 dd10 dd11 dd12 dd13 dd14 dd15 dd16 dd17 dd18 dd19 dd20 sf dum
. gen z1=p_bs
. replace z1=p_sf if city==7
(495 real changes made)

.
. imperfectiv y `w' (price=z1 qavgp), prop5 noassumption

```

```
Nevo and Rosen (2012)'s Imperfect IV bounds          Number of obs =    990
```

Variable	Lower Bound	Upper Bound
price	-11.4	-2.363483
addv	.146914	.418325
bd1	.033972	.599707
bd2	.455624	.298904
bd3	.382829	.450718
bd4	.187206	.401833
bd5	.605444	-.492812
bd6	-.125019	-.27417
bd7	.202934	-.19722
bd8	-.530749	-.864742
bd9	-.386157	-.807136
bd10	.894826	.716074
bd11	.4386	.373726
bd12	-.709916	-.558852
bd13	.750836	.23102
bd14	-.010961	-.316191
bd15	-.000558	-.619963
bd16	.405134	-1.042088
bd17	-.521734	-.887364
bd18	-.05868	-.505367
bd19	-.224799	-.370595
bd20	.498622	-.156238
bd21	-.763825	-.882123
bd22	.191673	-.503372
bd23	-.605655	-.555918
bd24	0	0
bd25	.082051	-.559479
dd2	-.11259	.072341
dd3	-.061687	.089817
dd4	-.170849	-.052635
dd5	.129522	.223347
dd6	.105121	.194984
dd7	.195902	.240531
dd8	.122156	.146316
dd9	.101151	.112072
dd10	.081881	.097511
dd11	.035294	.059166
dd12	-.073256	-.054442
dd13	.044194	.029706
dd14	.109479	.090788
dd15	.095593	.097226

```

dd16          -.031882      -.025253
dd17          .092846       .112414
dd18          .089866       .140122
dd19          .105029      .131951
dd20          -.009224      .003433
sf dum       -.097125      -.268704

```

```
. imperfectiv y `w` (price=z1 qavgp), prop5 short
```

```
Nevo and Rosen (2012)'s Imperfect IV bounds          Number of obs =    990
```

Variable	Lower Bound	Upper Bound
price	-11.4	-4.082247

1.2 Conley et al. (2012)'s 401(K) Example

Below we replicate the plausibly exogenous bounds calculated by Conley et al. (2012) in their empirical application examining the effect of participation in 401(k) on asset accumulation. We use the `plausexog` command described in section 3 to calculate local to zero (ltz) bounds.

```

. use Conleyetal2012
(Conely et al's (2012) REStat for 401(k) participation)
. local xvar i2 i3 i4 i5 i6 i7 age age2 fsize hs smcol col marr twoearn db pira
> hown
. plausexog ltz net_tfa `xvar` (p401 = e401), omega(25000) mu(0) level(.95) vce
> (robust) graph(p401) graphmu(1000 2000 3000 4000 5000) graphomega(333333.33 1
> 333333.3 3000000 5333333.3 8333333.3) graphdelta(2000 4000 6000 8000 10000)
Estimating Conely et al.'s ltz method
Exogenous variables: i2 i3 i4 i5 i6 i7 age age2 fsize hs smcol col marr twoearn
> db pira hown
Endogenous variables: p401
Instruments: e401

```

```
Conley et al. (2012)'s LTZ results          Number of obs =    9915
```

	Coef.	Std. Err.	z	P> z	[95% Conf. Interval]	
p401	13222.14	1926.609	6.86	0.000	9446.061	16998.23
i2	962.1541	700.6402	1.37	0.170	-411.0755	2335.384
i3	2190.277	992.1113	2.21	0.027	245.7741	4134.779
i4	5313.626	1420.208	3.74	0.000	2530.069	8097.183
i5	10400.47	2017.663	5.15	0.000	6445.918	14355.01
i6	21859.43	2239.623	9.76	0.000	17469.85	26249.01
i7	62464.83	5871.894	10.64	0.000	50956.12	73973.53
age	-1811.558	536.1392	-3.38	0.001	-2862.372	-760.7445
age2	28.68893	6.712006	4.27	0.000	15.53364	41.84422
fsize	-724.4649	378.4213	-1.91	0.056	-1466.157	17.2273
hs	2761.253	1244.257	2.22	0.026	322.553	5199.952
smcol	2750.739	1643.95	1.67	0.094	-471.3435	5972.821
col	5161.979	1926.959	2.68	0.007	1385.208	8938.749
marr	4453.186	1853.123	2.40	0.016	821.1317	8085.24
twoearn	-15051.59	2125.758	-7.08	0.000	-19218	-10885.18
db	-2750.19	1207.883	-2.28	0.023	-5117.597	-382.783

pira	31667.72	1730.29	18.30	0.000	28276.42	35059.03
hown	4200.889	767.7217	5.47	0.000	2696.182	5705.596
_cons	18929.86	9755.124	1.94	0.052	-189.8365	38049.55

The output displayed above documents bounds on each model parameter. Bounds on the endogenous variable of interest ($p401$) are displayed at the top of the output table, and agree with those displayed in figure 2 of Conley et al. (2012). Below we display the output from replicating the full figure 2 of Conley et al. (2012) with bounds across a range of priors using the LTZ approach, and the graphing capabilities of `plausexog`.

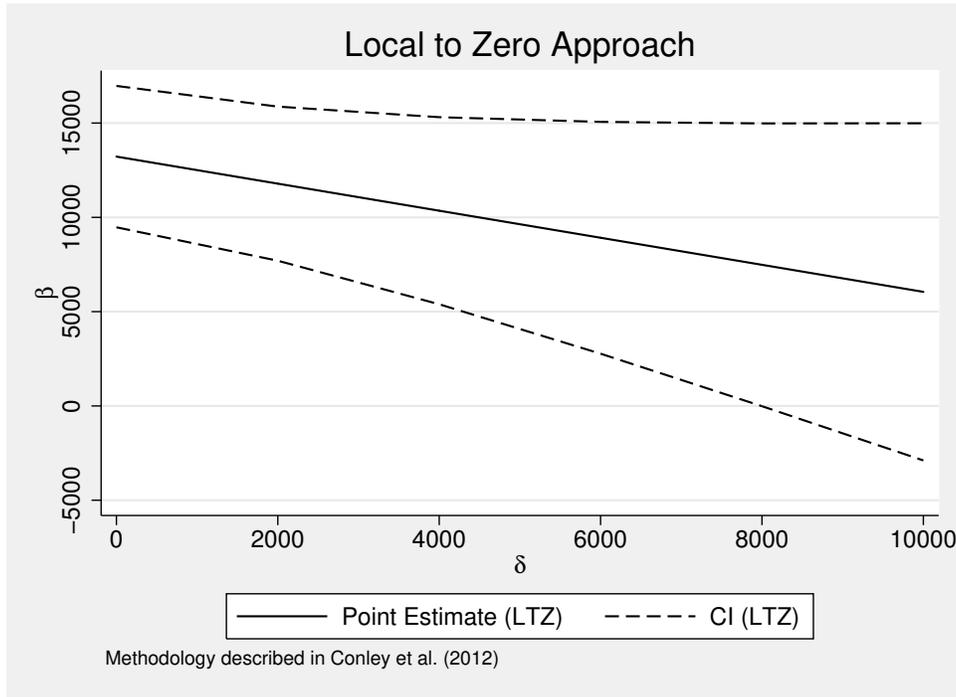


Figure 2: Replicating Figure 2 for 95% Confidence Intervals with Positive Prior