

Montes-Rojas, G. & Galvao, A. F. (2013). Bayesian Endogeneity Bias Modeling (Report No. 13/09).  
London, UK: Department of Economics, City University London.



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**Original citation:** Montes-Rojas, G. & Galvao, A. F. (2013). Bayesian Endogeneity Bias Modeling (Report No. 13/09). London, UK: Department of Economics, City University London.

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**Bayesian Endogeneity Bias Modeling**

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**Discussion Paper Series**  
**No. 13/09**

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# Bayesian Endogeneity Bias Modeling\*

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November 7, 2013

## Abstract

We propose to model endogeneity bias using prior distributions of moment conditions. The estimator can be obtained both as a method-of-moments estimator and in a Ridge penalized regression framework. We show the estimator's relation to a Bayesian estimator.

*Keywords:* Endogeneity; Shrinkage; Ridge regression; Method of moments.

*JEL Classification:* C11, C52

## 1 Introduction

A solution to the problem of endogeneity is to rely on exogenous information, i.e. instrumental variables (IV) (see e.g. Hausman , 1983; Angrist and Krueger , 2001) and proxy variables (see e.g. Olley and Pakes , 1996; Levinsohn and Petrin , 2003; Wooldridge , 2002, 2009) derived from additional exclusion restrictions or equations. In practice, the type of restriction and the point identification strategy chosen determine the model to be used. Alternative approaches are given in Rigobon (2003), Klein and Vella (2010), Chalak and White (2011), and Lewbel (2012). However, in many empirical applications there is disagreement and concern about the exclusion restrictions imposed. A related approach is to seek parameter set

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\*The authors would like to express their appreciation to badi Baltagi, one anonymous referee, Roger Koenker, Stephane Bonhomme, Susumu Imai, and Ehsan Soofi for helpful discussions and comments regarding this paper. All the remaining errors are our solely responsibility.

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identification. The advantage of such methods is that they require weaker assumptions than those needed for consistent point estimation of the parameter of interest.

This paper suggests an alternative strategy to deal with the endogeneity problem. We propose the use of a prior distribution on the endogeneity bias when there is no availability of additional information such as instrumental or proxy variables. By modeling bias, thus, we mean to impose a prior distribution on the amount of endogeneity of the endogenous variables' coefficient estimators and then compute the distribution of the associated parameters of interest. This distribution reflects the researcher's beliefs about endogeneity. The value of our contribution does not only lie in the derived estimators but in the way it proposes to think about endogeneity. In particular, it formalizes the use of prior knowledge about unobservables and their relationship with observables to quantitatively assess the degree of bias. The proposed estimators are constructed using the method-of-moments by imposing the prior distribution of a misspecified moment condition, or equivalently, in a Ridge penalized regression framework, by only penalizing the endogenous variables' coefficients. We show the connection of our estimator with a Bayesian estimator.

The proposed methodology is related to recent developments in the literature. First, it can be interpreted as in Altonji, Elder and Taber (2005a,b, 2008) as a strategy to extract information from observables about the endogeneity bias. They construct an index of observables, which in combination of prior knowledge about the sign of the bias and a condition on the relationship between included (observable) and excluded (non-observable) variables can be used to identify the endogenous variable parameter. In a related work Kiviet (2011) imposes the correlation between the endogenous variable and the innovations. In this case, identification of the true parameters cannot be achieved, but the direction and magnitude of the bias is analyzed instead. Second, the idea of using penalized regression to model endogeneity was proposed by Galvao and Montes-Rojas (2010) in the context of dynamic panel data models where the fixed-effects' shrinkage reduces the dynamic panel bias.

## 2 Regression shrinking of endogenous covariates

Consider the following linear regression model,

$$y_i = x_{1i}\beta_1 + x_{2i}\beta_2 + \epsilon_i, \quad i = 1, 2, \dots, n, \quad (1)$$

where  $x_{1i}$  is a  $1 \times p_1$  vector,  $x_{2i}$  is a  $1 \times p_2$  vector,  $\beta_1$  is a  $p_1 \times 1$  vector, and  $\beta_2$  is a  $p_2 \times 1$  vector.  $x_1$  contains  $p_1$  exogenous regressors and  $x_2$  contains  $p_2$  endogenous regressors. Let  $x = (x_1, x_2)$  and  $\beta = (\beta_1, \beta_2)'$ . In matrix notation  $\mathbf{y} = \mathbf{x}_1\beta_1 + \mathbf{x}_2\beta_2 + \boldsymbol{\epsilon} = \mathbf{x}\beta + \boldsymbol{\epsilon}$  where  $\mathbf{y}, \mathbf{x}, \boldsymbol{\epsilon}$  are the corresponding  $n$ -dimensional vectors or matrices.

We consider the following assumptions:

**Assumption 1.**  $(y_i, x_i)$  is an i.i.d. random sample for  $i = 1, \dots, n$  with  $y_i = x_{1i}\beta_1 + x_{2i}\beta_2 + \epsilon_i$ .

**Assumption 2.**  $\frac{1}{n}\mathbf{x}'_1\mathbf{x}_1 \xrightarrow{p} C_1$  (finite and non-singular),  $\frac{1}{n}\mathbf{x}'_1\mathbf{x}_2 \xrightarrow{p} C_{12}$  (finite),  $\frac{1}{n}\mathbf{x}'_2M_1\mathbf{x}_2 \xrightarrow{p} V_{2.1}$  (finite and non-singular), and  $\frac{1}{n}\mathbf{x}'_1M_2\mathbf{x}_1 \xrightarrow{p} V_{1.2}$  (finite and non-singular), where  $M_j =$

$I_{p_j} - \mathbf{x}_j(\mathbf{x}'_j\mathbf{x}_j)^{-1}\mathbf{x}'_j$ ,  $j = 1, 2$  are the orthogonal projections in least squares.

**Assumption 3.** For any  $i$ ,  $E[x'_{1i}\epsilon_i] = \mathbf{0}$  and  $E[x'_{2i}\epsilon_i] = B_2$ , such that  $\frac{1}{n}\mathbf{x}'_1\epsilon \xrightarrow{p} \mathbf{0}$  and  $\frac{1}{n}\mathbf{x}'_2\epsilon \xrightarrow{p} B_2$ .

Let  $\hat{\beta}^o = (\hat{\beta}_1^o, \hat{\beta}_2^o)'$  be the ordinary least squares (OLS) estimator. Simple OLS algebra and asymptotic calculations show that

$$\begin{aligned}\hat{\beta}_1^o &\xrightarrow{p} \beta_1 - C_1^{-1}C_{12}\delta_2, \\ \hat{\beta}_2^o &\xrightarrow{p} \beta_2 + \delta_2,\end{aligned}$$

where  $\delta_2 = V_{2.1}^{-1}B_2$  is the OLS bias in estimating  $\beta_2$ . The OLS estimator is derived from the set of estimating equations

$$\begin{aligned}E[x'_1\epsilon] &= \mathbf{0}, \\ E[x'_2\epsilon] &= \mathbf{0},\end{aligned}$$

and thus the OLS bias is the result of misspecifying the second moment condition, i.e. wrongly assuming  $x_2$  is exogenous, where in fact,  $E[x'_2\epsilon] = B_2$ , for a general unknown  $B_2$ . In particular, consider the following modified moment condition  $V_{2.1}^{-1}E[x'_2\epsilon] = V_{2.1}^{-1}B_2 = \delta_2$  and let  $\delta_2 = \Lambda\beta_2$ , where  $\Lambda$  is a  $p_2 \times p_2$  diagonal matrix  $\text{diag}\{\lambda_1, \lambda_2, \dots, \lambda_{p_2}\}$ . In many empirical applications the signs of  $\beta_2$ ,  $\delta_2$  and  $C_{12}$  are known, and then the sign of the elements in  $\Lambda$  are also known. Note that  $\Lambda\beta_2$  has the role of  $\delta_2$  above and the greater a parameter  $\lambda_j$  is, the greater the endogeneity problem in  $x_j$  is, for  $j = 1, 2, \dots, p_2$ .  $\Lambda$  can be seen as a *proportional endogeneity tolerance parameter*.

This paper focuses on estimators that impose prior information about  $\delta_2$  or  $\Lambda$ , derived through the moment conditions

$$\begin{aligned}E[x'_1\epsilon] &= \mathbf{0}, \\ V_{2.1}^{-1}E[x'_2\epsilon] &= \Lambda\beta_2.\end{aligned}$$

For convenience we maintain the terminology of *exogenous* and *endogenous* variables. Nevertheless, this could be redefined for the researcher convenience.

The moment conditions can be replaced by the following estimating equations:

$$\mathbf{x}'_1(\mathbf{y} - \mathbf{x}_1\beta_1 - \mathbf{x}_2\beta_2) = \mathbf{0}, \tag{2a}$$

$$(\mathbf{x}'_2M_1\mathbf{x}_2)^{-1}(\mathbf{x}'_2(\mathbf{y} - \mathbf{x}_1\beta_1 - \mathbf{x}_2\beta_2)) = \Lambda\beta_2. \tag{2b}$$

The solution to this problem, if  $(I_{p_2} + \Lambda)$  is non singular, is

$$\begin{aligned}\hat{\beta}_1^\Lambda &= \hat{\beta}_1^o + (\mathbf{x}'_1\mathbf{x}_1)^{-1}(\mathbf{x}'_1\mathbf{x}_2)\Lambda(I_{p_2} + \Lambda)^{-1}\hat{\beta}_2^o, \\ \hat{\beta}_2^\Lambda &= (I_{p_2} + \Lambda)^{-1}\hat{\beta}_2^o.\end{aligned}$$

Let  $\hat{\beta}^\Lambda = (\hat{\beta}_1^\Lambda, \hat{\beta}_2^\Lambda)'$ , one can notice that for this case,

$$\begin{aligned}\hat{\beta}_1^\Lambda &\xrightarrow{p} \beta_1 - C_1^{-1}C_{12}\delta_2 + C_1^{-1}C_{12}(I_{p_2} + \Lambda)^{-1}(\Lambda\beta_2 - \delta_2), \\ \hat{\beta}_2^\Lambda &\xrightarrow{p} (I_{p_2} + \Lambda)^{-1}(\beta_2 + \delta_2).\end{aligned}$$

The proposed estimator  $\hat{\beta}^\Lambda$  also arises as a solution to a penalized Ridge regression version of model (1), where  $\{\lambda_1, \lambda_2, \dots, \lambda_{p_2}\}$  are penalties that apply to the  $p_2$  endogenous regressors only. In this framework  $\hat{\beta}^\Lambda$  is obtained as

$$\hat{\beta}^\lambda = \underset{\beta}{\operatorname{argmin}} \sum_{i=1}^n (y_i - x_i\beta)^2 + (\mathbf{x}'_2 M_1 \mathbf{x}_2) \sum_{j=1}^{p_2} \lambda_j \beta_{2j}^2. \quad (3)$$

Interpreting  $\Lambda$  as a penalty parameter allows for applications to other contexts in which our ignorance and lack of identification about some parameters determine that their estimation involves a burden for the desirable properties of the full model. Note that for the Ridge regression method, the  $\lambda_j$ s are assumed to be non-negative, and thus they are interpreted as shrinkage parameters, that is, larger values of  $\lambda$  shrinks the corresponding parameter estimates towards zero. However, for our purposes we could consider  $\lambda >< 0$  in order to account for different types of bias.

One important case in (3) is when  $\Lambda = \operatorname{diag}\{\lambda\}$  and  $\lambda > 0$  is the same positive scalar for all endogenous variables. In this case, all endogenous variables regression parameters are subject to the same penalty, that is, they shrink according to  $\lambda$ . This is the case of Galvao and Montes-Rojas (2010) where all fixed-effects parameters are reduced in a similar fashion. There are two important special cases of the above method. First, the model with no shrinkage, and secondly, the model that totally shrinks  $\beta_2$ . In the former case, if  $\lambda = 0$ , then  $\hat{\beta}^\circ$  is the standard OLS estimator where both  $x_1$  and  $x_2$  are used. In this case, the endogeneity in  $x_2$  produces an *endogeneity bias* in the estimators of  $\beta_1$  and  $\beta_2$ . In the second case, if  $\lambda \rightarrow \infty$ , then  $\hat{\beta}^\infty$  is the estimator where only  $x_1$  is used and  $\beta_2$  is set to 0. This will generate an *omitted variable bias* for the estimation of  $\beta_1$ .

### 3 Imposing a prior on endogeneity

In practice the estimator above is infeasible because  $\Lambda$  (or  $\delta_2$ ) is unknown and the parameters cannot be identified. If  $\Lambda$  is known the parameters are point identified, while if  $\Lambda$  is known to belong to a defined set the parameters are set identified by the range of the tuning parameter. We propose a Bayesian estimator where we impose a prior on the tuning parameter and compute the estimate of the parameters of interest based on the penalized regression. The prior of the tuning parameter thus implicitly defines a weighted set for identification and produces a posterior distribution for an estimator of  $\beta$ . As in a Bayesian context, the properties of the derived estimator depends on how accurate the prior is.

Our uncertainty on  $\beta_2$  (and  $\delta_2$ ) is modeled by  $\lambda$ , the proportional endogeneity tuning parameter. By modeling bias, thus, we mean to impose a prior distribution on  $\delta_2$ , the amount of endogeneity, through the parameter  $\lambda$ , and then compute the distribution of the associated parameters of interest. As an example, suppose that  $\lambda \sim \chi_1^2$ , so that it has mean equal to 1. Then  $\delta_2$  is of the same magnitude as  $\beta_2$ , that is,  $E[\hat{\beta}_2^\circ] = \beta_2 + \delta_2 = 2\beta_2$ .

The distribution of  $\hat{\beta}^\lambda = (\hat{\beta}_1^\lambda, \hat{\beta}_2^\lambda)$  can then be obtained from the distribution of  $\hat{\beta}^\circ$  and  $\lambda$ . Define  $h(\beta|y, x)$  as the density function of this estimator. A point estimate can be based

on the mean (i.e.  $E_\lambda[\hat{\beta}^\lambda]$ ), median (i.e.  $Q_{\hat{\beta}^\lambda}(0.5)$ , where  $Q(\cdot)$  is the quantile function) or any other statistic of interest such as mode. As a practical matter,  $h(\cdot)$  can be simulated by using the asymptotic normality of  $\hat{\beta}^\circ$  and the distribution of  $\lambda$ , which is assumed to be independent of  $(y, x)$ .

## 4 Connection with a Bayesian estimator

In order to see the connection of our estimator with a Bayesian estimator, consider the loss function in a standard Ridge regression problem

$$L(\beta, \lambda) = \sum_{i=1}^n (y_i - x_i \beta)^2 + \lambda |\beta|^2.$$

Consider now a Bayesian estimator based on assuming a Gaussian likelihood for  $\epsilon \equiv y - x\beta$ ,  $p(y, x|\beta)$  where for simplicity the variance of  $\epsilon$  is assumed to be known and equal to 1:

$$p(y, x|\beta, \lambda) \propto \exp\left(-1/2 \sum_{i=1}^n (y_i - x_i \beta)^2\right).$$

Now consider a prior on  $(\beta, \lambda)$  given by

$$p(\beta, \lambda) = \phi(\beta; 0, 1/\lambda)g(\lambda) \propto \lambda g(\lambda) \exp\left(-\frac{1}{2}\lambda |\beta|^2\right),$$

where  $\phi(\cdot; m, v)$  is the Gaussian density function with mean  $m$  and variance  $v$ . Here  $\beta$  is assumed to have a Gaussian prior and  $\lambda$  a density function  $g(\lambda)$ . By Bayes theorem, the posterior can be obtained by

$$p(\beta, \lambda|y, x) = p(y, x|\beta, \lambda)p(\beta, \lambda) \propto \lambda g(\lambda) \exp\left(-\frac{1}{2}L(\beta, \lambda)\right).$$

Define  $y_\lambda = [y:0]$  and  $x_\lambda = [x:\sqrt{\lambda}]$ , and note that  $\hat{\beta}^\lambda$  is the OLS estimator of a regression of  $y_\lambda$  on  $x_\lambda$ . Then using (Zellner, 1971, p.66) results

$$p(\beta, \lambda|y, x) \propto \lambda g(\lambda) \exp\left(-\frac{1}{2}\left((y_\lambda - x_\lambda \hat{\beta}^\lambda)'(y_\lambda - x_\lambda \hat{\beta}^\lambda) + (\beta - \hat{\beta}^\lambda)'x'_\lambda x_\lambda(\beta - \hat{\beta}^\lambda)\right)\right).$$

Note that for fixed  $\lambda$ , the maximum value of  $p(\beta, \lambda|y, x)$  corresponds to  $\beta = \hat{\beta}^\lambda$ . The posterior distribution of  $\beta$ ,  $p(\beta|y, x)$ , requires integrating out  $\lambda$  which depends on the assumed distribution  $g(\cdot)$ . Note that  $p(\beta|y, x)$  is different from  $h(\beta|y, x)$ , the density function of our estimator. In Figure 1, we illustrate this idea. Assume for simplicity that  $\beta$  is unidimensional and consider a given joint posterior distribution of  $\beta$  and  $\lambda$ . For each  $\beta$ ,  $p(\beta|y, x)$  is

obtained by computing the mean over all possible values of  $\lambda$ . For instance, for  $\beta_1$ ,  $p(\beta_1|y, x)$  corresponds to the integration over  $\lambda$  along the vertical line at  $\beta_1$ , and the same for  $\beta_2$  and  $\beta_3$ . However,  $h(\beta|y, x)$  is obtained by computing the mean over those values of  $\lambda$  for which  $\beta = \hat{\beta}^\lambda$ , i.e. only those for which for a given  $\lambda$  it finds the maximum density. For  $\beta_1$  that only corresponds to  $\lambda_1$ ,  $\beta_2$  to  $\lambda_2$  and  $\beta_3$  to  $\lambda_3$ , as in the dashed line. Thus, the distribution of our proposed estimator uses the most likely  $\beta$  for each  $\lambda$ .

[FIGURE 1 HERE]

## 5 Conclusions

This paper proposed a novel way of dealing with endogeneity bias when there is no additional information such as instrumental or proxy variables. In particular, a prior is imposed on the endogeneity bias and the resulting estimators can be constructed both as a method-of-moments estimator or in a Ridge penalized regression framework.

Several extensions could be proposed for future research. First, the parameters modeling the endogeneity bias could depend on both prior information and observable variables. Second, this model could be applied to the  $l_1$  penalization case using the asymptotic results in Knight and Fu (2000), and could be extended to model selection with several endogenous regressors. Third, the proposed model could be applied to forecasting where there is uncertainty about some parameter values and prior information would be imposed.

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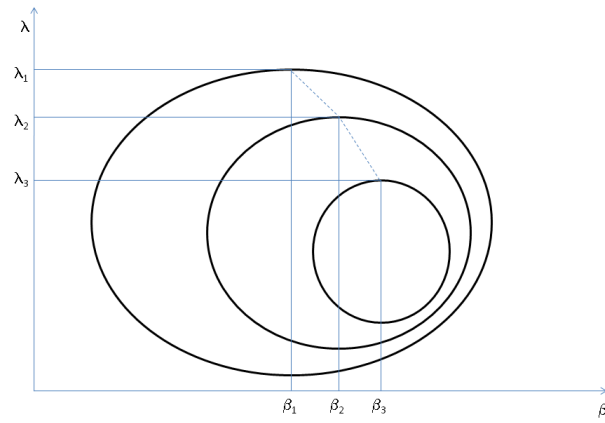


Figure 1: Bivariate posterior distribution of  $(\beta, \lambda)$