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## When Outcome Heterogeneously Matters for Selection – A Generalized Selection Correction Estimator

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Arndt Reichert  
Harald Tauchmann

## **When Outcome Heterogeneously Matters for Selection**

A Generalized Selection Correction Estimator

# Imprint

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Arndt Reichert and Harald Tauchmann<sup>1</sup>

# When Outcome Heterogeneously Matters for Selection – A Generalized Selection Correction Estimator

## Abstract

*The classical Heckman (1976, 1979) selection correction estimator (heckit) is misspecified and inconsistent if an interaction of the outcome variable and an explanatory variable matters for selection. To address this specification problem, a full information maximum likelihood estimator and a simple two-step estimator are developed. Monte-Carlo simulations illustrate that the bias of the ordinary heckit estimator is removed by these generalized estimation procedures. Along with OLS and the ordinary heckit procedure, we apply these estimators to data from a randomized trial that evaluates the effectiveness of financial incentives for weight loss among the obese. Estimation results indicate that the choice of the estimation procedure clearly matters.*

*JEL Classification:* C24, C93

*Keywords:* Selection bias; interaction; heterogeneity; generalized estimator

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# 1 Introduction

The Heckman (1976, 1979) selection correction (heckit) estimator is a workhorse of applied econometrics, commonly used for removing possible bias due to selection on unobservables.<sup>1</sup> In many applications, selection into the subsample of observations with observed outcome is directly affected by the value of the outcome variable itself. Think, for instance, of estimating a wage equation. Here, wages are only observed for individuals who have accepted a wage offer. Yet, the likelihood of accepting the offer increases with the offered wage. The regular specification of the heckit estimator implicitly accounts for a possible impact of the outcome on selection, given that all exogenous variables enter the selection part of the model.<sup>2</sup>

However, unlike the above case, the regular heckit estimator is misspecified and biased if the offered pay is of differential relevance depending on individual characteristics such as gender. That is, heterogeneous effects of the outcome variable on selection are not accommodated by the regular heckit approach. In order to allow for heterogeneity with respect to a certain individual characteristic, any selection correction estimator must take the interaction of the outcome variable with the relevant covariate into account. The present paper develops generalizations of the regular heckit estimator that overcome the inconsistency of the ordinary heckit model in the presence of heterogeneous effects of the outcome on selection. In particular, we suggest a full information maximum likelihood (FIML) estimator and a computationally very simple two-step approach.

We test the performance of the suggested estimators using Monte Carlo simulations. We also apply the estimators to data gathered from a randomized field experiment, which was conducted to examine the effectiveness of financial incentives to induce weight loss in obese individuals. This experiment represents an exemplary application of the proposed generalized selection-correction estimators because the design of the monetary rewards makes favorable outcomes more likely to be reported than unfavorable ones. Thus, a link between the outcome variable and the probability of observing the outcome is first of all expected for participants who were offered a reward for weight loss.

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<sup>1</sup>Though it has been criticized for being very vulnerable to various kinds of misspecification (e.g. Puhani, 2000; Grisdal, 2001), and less restrictive semi-parametric alternatives have been proposed (e.g. Ichimura and Lee, 1991; Ahn and Powell, 1993); see Vella (1998) for a survey.

<sup>2</sup>The commonly used tobit (type 1) (Tobin, 1958) model represents an extreme case with selection *exclusively* depending on the outcome.

The remainder of the paper is organized as follows. Section 2 develops the generalized heckit estimators. Section 3 compares the performance of the different estimators using a Monte Carlo experiment. Section 4 provides a real data application and Section 5 concludes.

## 2 A Generalized Heckit Model

Consider a familiar linear regression model, where the focus of the econometric analysis is on estimating the coefficient vector  $\beta$ :

$$Y_i = \beta' X_i + \varepsilon_i. \quad (1)$$

Here  $i$  indexes observations, and  $Y_i$ ,  $\varepsilon_i$ , and  $X_i$  denote the outcome variable, a random error, and the vector of exogenous explanatory variables, respectively. The latter includes the variable  $D_i$ , which is of special relevance to the analysis.

However,  $Y_i$  is observed only for a subsample of observation. Selection into this subsample, indicated by  $S_i = 1$ , is modeled as suggested by Heckman (1979). Yet, besides a  $K$ -dimensional vector  $Z_i$  that includes  $X_i$  and some further exogenous variables,  $Y_i$  as well as the interaction term  $Y_i D_i$  are allowed to enter the selection equation:

$$S_i = \begin{cases} 1 & \text{if } \theta' Z_i + \tau Y_i + \gamma Y_i D_i + v_i > 0 \\ 0 & \text{else.} \end{cases} \quad (2)$$

As in the ordinary heckit model, joint normality  $N(0, 0, \sigma_\varepsilon^2, \sigma_v^2, \sigma_{\varepsilon v})$  is assumed for the error terms  $\varepsilon_i$  and  $v_i$ .  $\theta_1, \dots, \theta_K$ ,  $\tau$ , and  $\gamma$  denote unknown coefficients. Substituting  $Y_i$  by (1) and rearranging terms leads to

$$S_i = \begin{cases} 1 & \text{if } \tilde{v}_i > -\alpha' Z_i - \gamma \beta' X_i D_i \\ 0 & \text{else} \end{cases} \quad (3)$$

$$\tilde{v}_i = v_i + (\tau + \gamma D_i) \varepsilon_i, \quad (4)$$

where  $\alpha_k = \theta_k + \tau \beta_k$  holds for any regressor  $k$  included in  $X_i$  and  $\alpha_k = \theta_k$  holds for those variables that enter  $Z_i$  but do not enter  $X_i$ . Evidently, the coefficient  $\tau$  has no



impact on the general structure of the model.<sup>3</sup> Hence, for the special case  $\gamma = 0$ , (1), (3), and (4) represent the standard Heckman (1979) selection model.

For  $\gamma \neq 0$ , however, the model deviates from the standard case for two reasons: (i) a full set of interaction terms  $X_i D_i$  enters the selection equation and (ii), more important,  $D_i$  enters the error  $\tilde{v}_i$ , rendering the error variance-covariance structure heterogeneous with respect to  $D_i$ :

$$\text{var}(\tilde{v}_i|D_i) = \sigma_v^2 + 2(\tau + \gamma D_i)\sigma_{\epsilon v} + (\tau + \gamma D_i)^2 \sigma_\epsilon^2 \quad (5)$$

$$\text{cov}(\epsilon_i, \tilde{v}_i|D_i) = \sigma_{\epsilon v} + (\tau + \gamma D_i)\sigma_\epsilon^2. \quad (6)$$

Ignored heteroscedasticity in the probit and, hence, in the selection part of the heckit model, is well known to render probit estimation inconsistent (Wooldridge, 2002; Harvey, 1976). Thus, a generalized estimator is required.

## 2.1 FIML Estimation

In order to develop a FIML estimator that accounts for the model structure, with no loss of generality, we introduce the normalization

$$\sigma_v^2 + 2\tau\sigma_{\epsilon v} + \tau^2\sigma_\epsilon^2 = 1. \quad (7)$$

That is, we assume standard normality for  $\tilde{v}_i$  conditional on  $D_i = 0$ . This is equivalent to the familiar normalization required for identifying the coefficients of any probit model. We re-parameterize as follows:

$$\rho \equiv \text{cor}(\epsilon_i, \tilde{v}_i|D_i = 0) = \frac{\sigma_{\epsilon v}}{\sigma_\epsilon} + \tau\sigma_\epsilon. \quad (8)$$

Then the individual log-likelihood  $l_i$  reads as

$$l_i = \begin{cases} \log \Phi \left( \frac{-\alpha' Z_i - \gamma \beta' X_i D_i}{\sqrt{1 + 2\rho\sigma_\epsilon \gamma D_i + \sigma_\epsilon^2 \gamma^2 D_i^2}} \right) & \text{if } S_i = 0 \\ \log \Phi \left( \frac{\alpha' Z_i + \gamma \beta' X_i D_i + (Y_i - \beta' X_i) \left( \frac{\rho}{\sigma_\epsilon} + \gamma D_i \right)}{\sqrt{1 - \rho^2}} \right) & \text{if } S_i = 1. \\ -\frac{1}{2} \left( \frac{Y_i - \beta' X_i}{\sigma_\epsilon} \right)^2 - \log \left( \sigma_\epsilon \sqrt{2\pi} \right) & \end{cases} \quad (9)$$

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<sup>3</sup>Effectively,  $\tau$  only changes the unknown error variance-covariance structure, which is subject to estimation. Hence,  $\tau$  is not identified.

See Appendix A.1 for how (9) is derived from the log-likelihood function of the ordinary heckit model. Besides the coefficient vectors  $\alpha$  and  $\beta$ , the scalar parameters  $\gamma$ ,  $\sigma_\varepsilon$ , and  $\rho$  are subject to estimation.<sup>4</sup> Note that  $D_i$  may either be continuous, a count, or binary.

The model is straightforwardly transferred to the case where the effect of  $Y_i$  on selection differs across  $M + 1$  mutually exclusive groups, indexed by  $m = 0, \dots, M$ . For group membership being indicated by a set of binary indicators  $D_{0i}, \dots, D_{Mi}$ , the log-likelihood conditional on  $D_{mi} = 1$  is identical to (9), besides  $D_i$  is substituted by the value one and  $\gamma$  is replaced by  $\gamma_m$ .<sup>5</sup> Here,  $\gamma_0$  has to be restricted to zero in order to render the model identified.

## 2.2 Two-Step Estimation

The model (9) is, however, difficult to fit and may cause problems in the optimization procedure. Yet, for a binary variable  $D_i$  and, more general, group-wise heterogeneity, a computationally very simple two-step estimator is available. Here, the heterogeneity in the selection mechanism is accounted for by estimating group-wise probit models at the first stage. For each group  $m$ , a specific coefficient vector  $\alpha_m$  is estimated, where the coefficients attached to  $D_{1i}, \dots, D_{Mi}$  need to be restricted to the value of zero. At the second stage a vector of group-specific inverse Mills-ratios  $\lambda(\cdot)$  enter as additional regressors

$$Y_i = \beta' X_i + \sum_{m=0}^M \delta_m \lambda(\hat{\alpha}'_m Z_i) D_{mi} + \tilde{\varepsilon}_i \quad \text{if } S_i = 1. \quad (10)$$

The attached coefficients  $\delta_m$ , subject to estimation, capture  $\sigma_\varepsilon \text{cor}(\varepsilon_i, \tilde{v}_i | D_{mi} = 1)$ . Two-step estimation comes, however, to the cost of efficiency loss. In the present case, it is not only genuinely less efficient than FIML, like, e.g., the two-step estimator for the ordinary heckit model. It also ignores many parameter restrictions that stem from the structural model, inflating the number of parameters subject to estimation by  $M(K - 1) - M^2$ . Moreover, (10) may suffer from near-collinearity of correction terms and group indicators. On the other hand, two-step estimating in-

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<sup>4</sup>Technically,  $\text{atanh}(\rho)$  and  $\log(\sigma_\varepsilon)$  are estimated in the optimization procedure in order to avoid a bounded valid parameter space.

<sup>5</sup>Typically, all dummies  $D_{mi}$ , except for  $D_{0i}$  indicating the reference category, enter  $X_i$  and  $Z_i$ .

volves less assumptions about the selection mechanism than FIML and, hence, also accommodates types of heterogeneity in selection that render (9) misspecified.

### 3 Monte Carlo Analysis

In order to illustrate the performance of the FIML and the two-step estimators and to compare them with those of ordinary heckit and simple OLS estimation, we run a Monte-Carlo (MC) experiment, where the endogenous variables  $Y_i$  and  $S_i$  are generated according to (1) and (2). The exogenous variables, i.e. the vector  $Z_i$ , are drawn once and then kept fixed. We draw the binary indicator  $D_i$  from the  $B(1, 0.5)$  distribution and two continuous control variables from the uniform  $U(-1, 1)$  distribution. One of the latter is excluded from the vector  $X_i$ , while  $D_i$  enters (2) not only through  $Z_i$  but also interacted with  $Y_i$ . For all coefficients  $\beta_k$  and  $\theta_k$ , we choose the value of one, except for the constant terms, which both are set to zero. With respect to the variance-covariance matrix of the normal errors, we choose  $\sigma_\epsilon^2 = 2$ ,  $\sigma_v^2 = 1$ , and  $\sigma_{\epsilon v} = 0.75$ . We run six different simulations, varying the experimental setup with respect to: (i)  $\gamma$ , for which we try the values  $-1$ ,  $0$ , and  $1$ ; and (ii)  $\tau$ , for which we try the two values consistent with (7), i.e.,  $-0.75$  and  $0$ . The sample size is 10 000 and the size of the simulations is 2 000 repetitions. Our focus is on the estimators' performance in estimating the coefficients  $\beta$ . Hence, for each estimator, we report estimates for  $\text{bias}(\hat{\beta})$  and  $\text{MSE}(\hat{\beta})$ .

As predicted by theory, MC-results display no significant (warranted by simulation based tests on joint unbiasedness of  $\hat{\beta}$ ) bias for the FIML and the Two-Step estimator, while OLS is biased in any simulation; see Table 1. Furthermore, the ordinary heckit estimator does not exhibit a significant bias for  $\gamma = 0$ , while it is severely biased for  $\gamma \neq 0$ . Focussing on the coefficient attached to  $D_i$ , depending on the sign of  $\gamma$ , an upward or an downward bias may occur. Interestingly, for  $\gamma \neq 0$ , the ordinary heckit does not perform much better than OLS in terms of the estimated bias. In simulation (vi), it even perform worse. This means, correcting parametrically for selection bias but misspecifying the selection mechanism may not be an improvement compared to simply ignoring selectivity. As expected, in terms of the estimated MSE, Two-Step estimation performs worse than FIML.<sup>6</sup>

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<sup>6</sup>For simulations based on a small sample ( $N = 400$ ), this shortcoming of two-step estimation becomes even more prominent. There, in terms of the MSE, two-step estimation may even be outperformed by the biased ordinary heckit estimator.

Table 1: Monte-Carlo Simulation Results

	FIML		Two-Step		Ordinary Heckit		OLS	
	Bias	MSE	Bias	MSE	Bias	MSE	Bias	MSE
<b>simulation (i): <math>\gamma = -1; \tau = -0.75</math></b>								
<i>D</i>	0.001	0.002	0.005	0.024	-0.496	0.248	-0.756	0.573
control	0.000	0.001	0.001	0.001	-0.035	0.002	0.927	0.861
constant	-0.001	0.003	-0.001	0.004	0.272	0.077	-0.506	0.257
<b>simulation (ii): <math>\gamma = -1; \tau = 0</math></b>								
<i>D</i>	0.001	0.004	0.001	0.009	-1.223	1.496	-1.225	1.501
control	0.000	0.001	0.000	0.001	-0.197	0.040	0.841	0.709
constant	0.000	0.002	0.001	0.004	0.612	0.379	0.494	0.244
<b>simulation (iii): <math>\gamma = 0; \tau = -0.75</math></b>								
<i>D</i>	-0.002	0.002	0.000	0.008	-0.003	0.001	0.082	0.008
control	0.001	0.001	0.001	0.001	0.001	0.001	1.082	1.172
constant	0.002	0.002	0.000	0.004	0.002	0.002	-0.512	0.263
<b>simulation (iv): <math>\gamma = 0; \tau = 0</math></b>								
<i>D</i>	0.000	0.002	0.002	0.004	0.001	0.002	-0.268	0.073
control	0.002	0.001	0.002	0.002	0.002	0.001	0.737	0.545
constant	0.000	0.002	-0.001	0.004	0.000	0.002	0.515	0.266
<b>simulation (v): <math>\gamma = 1; \tau = -0.75</math></b>								
<i>D</i>	-0.001	0.003	-0.002	0.006	0.724	0.526	0.811	0.659
control	0.000	0.001	0.000	0.001	-0.209	0.045	0.842	0.709
constant	0.001	0.002	0.002	0.004	-0.336	0.117	-0.501	0.251
<b>simulation (vi): <math>\gamma = 1; \tau = 0</math></b>								
<i>D</i>	-0.001	0.002	-0.002	0.004	0.238	0.058	-0.099	0.011
control	-0.002	0.001	-0.001	0.002	-0.058	0.005	0.600	0.361
constant	0.002	0.002	0.001	0.004	-0.102	0.013	0.544	0.297

Notes: results based on 2 000 replications; sample size  $N = 10\,000$ ; exogenous variables drawn once and then kept fixed; true coefficient values:  $\beta_D = 1$ ,  $\beta_{control} = 1$ , and  $\beta_{const} = 0$ .

Even for  $\gamma = 0$  (simulations iii and iv), FIML exhibits an MSE that just marginally exceeds the MSE of the ordinary heckit model.

## 4 Real Data Application

We further apply the estimators to data from a randomized trial; see Augurzky et al. (2012) for a detailed description and a comprehensive empirical analysis. This experiment aims at analyzing the effectiveness of financial incentives for assisting obese individuals in losing body weight. By the end of a rehab hospital stay, 698 overweight individuals were set an individual weight-loss target (6 to 8 percent of current body weight), which they were prompted to realize within four months. Participants were then randomly assigned to two incentive groups and one control group. While, contingent on success, a reward of up to €150 (group 150) and €300 (group 300), respectively, was offered to members of the treatment groups, the con-

trol group received no financial incentive. Rewards were offered as a function of the degree of target achievement, i.e., participants who lost some weight but failed to realize the weight-loss target received less than the maximum reward. After four months, participants were requested to visit an assigned pharmacy for verifying actual weight loss. Yet, a substantial number of participants failed to show up at the weigh-in. More precisely, 178 individuals selected themselves out of the trial, while 520 complied and attended the weigh-in. The compliance rate varied substantially between groups. While for the control group it was 66.5 percent, it was 72.9 and 84.3 percent for group 150 and group 300, respectively. This nicely meets our prior expectation of the probability of reporting weight being affected by the interaction of actual weight loss and group membership, as only those who were both successful and members of one of the treatment groups had an financial incentive to attend the weigh-in.

In the present empirical analysis, the degree of target achievement, i.e., actual weight-loss divided by targeted weight loss, serves as dependent variable. Indicators for group membership are the key explanatory variables, with the control group serving as reference. Besides these, age and indicators for being female and being born in Germany enter the regression equation as controls. A further dummy indicating that a participant had to visit a nearby pharmacy, i.e., one within the same zip-code area as the place of residence, exclusively enters the selection equation. This exclusion restriction is justified by travel time representing a likely determinant for the decision whether or not to show up at the weigh-in. Yet, there is no obvious link to success.

Table 2 displays regression results for FIML, two-step, ordinary heckit, and OLS. Test results do not clearly argue for selection on unobservables since the estimate for  $\rho$  does not significantly deviate from zero, neither for ordinary heckit nor for FIML estimation. This equivalently holds for the two-step approach, where the group-specific Mill's ratios are jointly insignificant. Yet, conditional on selection correction, both FIML and two-step are clearly favored over ordinary heckit ( $p$ -values 0.03 and 0.01).

Focussing on the estimated treatment effects, the choice of estimation method clearly matters. OLS and ordinary heckit, both, yield that receiving a financial incentive increases the success rate by roughly 40 to 50 percentage points. Yet, the amount of the financial reward seems to be immaterial. FIML and two-step estimation of the generalized model, however, yield a different picture. For the latter, no

Table 2: Results for Weight-Loss Experiment

	FIML		Two-Step		Ordinary Heckit		OLS	
	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
<b>Main equation:</b>								
€150	0.215	0.143	0.238	0.373	0.429**	0.091	0.408**	0.088
€300	0.470**	0.154	0.272	0.392	0.506**	0.104	0.452**	0.086
age	-0.001	0.004	-0.001	0.006	-0.003	0.004	-0.005	0.004
female	-0.160**	0.079	-0.147*	0.085	-0.172**	0.077	-0.178**	0.076
native	0.020	0.088	0.020	0.091	0.013	0.087	0.008	0.087
$\delta_{\text{control}}$	-	-	0.059	0.687	-	-	-	-
$\delta_{\text{€150}}$	-	-	0.470	0.425	-	-	-	-
$\delta_{\text{€300}}$	-	-	0.762	0.903	-	-	-	-
constant	0.340	0.255	0.409	0.577	0.445	0.302	0.655**	0.201
<b>Selection equation:</b>								
€150	-0.077	0.206	-	-	0.200	0.124	-	-
€300	0.521*	0.277	-	-	0.595**	0.133	-	-
age	0.020**	0.005	-	-	0.019**	0.005	-	-
female	0.115	0.128	-	-	0.076	0.116	-	-
native	0.039	0.143	-	-	0.057	0.129	-	-
nearby pharmacy	0.332**	0.122	-	-	0.309**	0.108	-	-
$\gamma_{\text{€150}}$	1.206**	0.484	-	-	-	-	-	-
$\gamma_{\text{€300}}$	0.142	0.495	-	-	-	-	-	-
constant	-0.817**	0.326	-	-	-0.741**	0.278	-	-
<b>Selection equation control group:</b>								
age	-	-	0.016**	0.008	-	-	-	-
female	-	-	-0.042	0.198	-	-	-	-
native	-	-	0.075	0.216	-	-	-	-
nearby pharmacy	-	-	0.300*	0.179	-	-	-	-
constant	-	-	-0.569	0.489	-	-	-	-
<b>Selection equation €150 group:</b>								
age	-	-	0.024**	0.008	-	-	-	-
female	-	-	0.023	0.195	-	-	-	-
native	-	-	-0.018	0.208	-	-	-	-
nearby pharmacy	-	-	0.537**	0.183	-	-	-	-
constant	-	-	-0.836**	0.424	-	-	-	-
<b>Selection equation €300 group:</b>								
age	-	-	0.019**	0.009	-	-	-	-
female	-	-	0.308	0.231	-	-	-	-
native	-	-	0.175	0.274	-	-	-	-
nearby pharmacy	-	-	-0.056	0.215	-	-	-	-
constant	-	-	-0.077	0.522	-	-	-	-
$\sigma_\epsilon$	0.836**	0.034	-	-	0.802**	0.036	0.795	-
$\rho$	0.199	0.252	-	-	0.260	0.271	-	-

Notes: \*\* significant at 5%; \* significant at 10%; total number of obs. is 698; for 178 obs. weight-loss information is missing.

significant incentive effect is found. Here, the inefficiency of two-step estimation is underpinned by rather large standard errors. For FMIL, the estimated treatment effect for group 300 is similar to its counterpart from OLS and ordinary heckit estimation. Yet, the estimated effect for group 150 is substantially smaller and even becomes statistically insignificant. Hence, on basis of FIML, one concludes that the amount of the reward matters for weight loss. The estimates for  $\gamma_{\text{€150}}$  and  $\gamma_{\text{€300}}$  bear the expected positive sign. However – contrary to expectations – the latter is

much smaller. Moreover, a large standard error renders the estimate for  $\gamma_{\in 300}$  statistically insignificant. This may be explained by the small number of dropouts in group 300, which makes the identification of  $\gamma_{\in 300}$  difficult.

## 5 Conclusions

In this article we demonstrate that the classical Heckman (1976, 1979) selection correction estimator is misspecified and inconsistent when an interaction of the outcome and an explanatory variables matters for selection. Randomized trials assessing the effects of an incentive scheme, may serve as a typical example for this kind of sample selection problems. A FIML and a simple two-step estimator that both address this specification problem are developed. Monte-Carlo simulations illustrate that the bias of the ordinary Heckman (1976, 1979) estimator is removed by these generalized estimation procedures. Finally, the suggested estimators are applied to data from a randomized trial that evaluates the effectiveness of financial incentives for assisting obese in their attempt to lose weight. Estimation results indicate that the choice of the estimation procedure clearly matters.

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## A Appendix

### A.1 Generalizing the Log-Likelihood Function

In order to generalize the log-likelihood function of the ordinary heckit model (see e.g. Amemiya, 1985, p. 386), we augment the index function  $\alpha'Z_i$  by  $\gamma\beta'X_iD_i$  and replace the scalar parameters  $\sigma_v^2$  and  $\sigma_{\varepsilon v}$  by the functions (5) and (6), respectively:

$$l_i = \begin{cases} \log \Phi \left( \frac{-\alpha'Z_i - \gamma\beta'X_iD_i}{\sqrt{\text{var}(\tilde{v}_i|D_i)}} \right) & \text{if } S_i = 0 \\ \log \Phi \left( \frac{\alpha'Z_i + \gamma\beta'X_iD_i + (Y_i - \beta'X_i) \left( \frac{\text{cov}(\varepsilon_i, \tilde{v}_i|D_i)}{\sigma_\varepsilon^2} \right)}{\sqrt{\text{var}(\tilde{v}_i|D_i) \left( 1 - \frac{\text{cov}(\varepsilon_i, \tilde{v}_i|D_i)^2}{\sigma_\varepsilon^2 \text{var}(\tilde{v}_i|D_i)} \right)}} \right) & \text{if } S_i = 1. \\ -\frac{1}{2} \left( \frac{Y_i - \beta'X_i}{\sigma_\varepsilon} \right)^2 - \log(\sigma_\varepsilon \sqrt{2\pi}) & \end{cases} \quad (11)$$

Then we apply the normalization (7) to (5), and eliminate  $\tau$  and  $\sigma_{\varepsilon v}$  by entering (8) into the equation, yielding

$$\begin{aligned} \text{var}(\tilde{v}_i|D_i) &= 1 + 2\gamma(\sigma_{\varepsilon v} + \tau\sigma_\varepsilon^2)D_i + \sigma_\varepsilon^2\gamma^2D_i^2 \\ &= 1 + 2\rho\sigma_\varepsilon\gamma D_i + \sigma_\varepsilon^2\gamma^2D_i^2, \end{aligned} \quad (12)$$

which is nonnegative, by  $\rho$  being bounded to the  $[-1, 1]$  interval. Further, using (6) and, once more, eliminating  $\tau$  and  $\sigma_{\varepsilon v}$  by entering (8) into the equation yields

$$\frac{\text{cov}(\varepsilon_i, \tilde{v}_i|D_i)}{\sigma_\varepsilon^2} = \frac{\sigma_{\varepsilon v} + (\tau + \gamma D_i)\sigma_\varepsilon^2}{\sigma_\varepsilon^2} = \frac{\rho}{\sigma_\varepsilon} + \gamma D_i. \quad (13)$$

Finally, using (13) and (12) we simplify

$$\begin{aligned} \text{var}(\tilde{v}_i|D_i) \left( 1 - \frac{\text{cov}(\varepsilon_i, \tilde{v}_i|D_i)^2}{\sigma_\varepsilon^2 \text{var}(\tilde{v}_i|D_i)} \right) &= \text{var}(\tilde{v}_i|D_i) - \sigma_\varepsilon^2 \left( \frac{\rho}{\sigma_\varepsilon} + \gamma D_i \right)^2 \\ &= \text{var}(\tilde{v}_i|D_i) - \rho^2 - 2\rho\sigma_\varepsilon\gamma D_i - \sigma_\varepsilon^2\gamma^2D_i^2 \\ &= 1 - \rho^2, \end{aligned} \quad (14)$$

and substitute (12), (13), and (14) into (11), yielding (9).