

# Comparing Logit and Probit Coefficients between Models and Across Groups

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# Introduction

- We are used to estimating models where an observed, continuous independent variable,  $Y$ , is regressed on one or more independent variables, i.e.

$$Y = \alpha + \sum X \beta + \varepsilon, \quad \varepsilon \sim N(0, \sigma^2)$$

- Since the residuals are uncorrelated with the  $X$ s, it follows that

$$\begin{aligned} V(Y) &= V(\alpha + \sum X \beta) + V(\varepsilon) \\ &= \text{Explained Variance} + \text{Residual Variance} \end{aligned}$$

- As you add explanatory variables to a model, the variance of the observed variable  $Y$  stays the same in OLS regression. As the explained variance goes up, the residual variance goes down by a corresponding amount.

- But suppose the observed  $Y$  is not continuous – instead, it is a collapsed version of an underlying unobserved variable,  $Y^*$
- Examples:
  - Do you approve or disapprove of the President's health care plan? 1 = Approve, 2 = Disapprove
  - Income, coded in categories like \$0 = 1, \$1- \$10,000 = 2, \$10,001-\$30,000 = 3, \$30,001-\$60,000 = 4, \$60,001 or higher = 5

- For such variables, also known as limited dependent variables, we know the interval that the underlying  $Y^*$  falls in, but not its exact value
- Binary & Ordinal regression techniques allow us to estimate the effects of the  $X$ s on the underlying  $Y^*$ . They can also be used to see how the  $X$ s affect the probability of being in one category of the observed  $Y$  as opposed to another.

- The latent variable model in binary logistic regression can be written as

$$y^* = \alpha + \sum X \beta + \varepsilon, \quad \varepsilon \sim \text{Standard Logistic}$$

If  $y^* \geq 0$ ,  $y = 1$

If  $y^* < 0$ ,  $y = 0$

In logistic regression, the errors are assumed to have a standard logistic distribution. A *standard logistic distribution* has a mean of 0 and a variance of  $\pi^2/3$ , or about 3.29.

- Since the residuals are uncorrelated with the Xs, it follows that

$$V(y^*) = V(\alpha + x\beta) + V(\varepsilon_{y^*}) = V(\alpha + x\beta) + \pi^2 / 3 = V(\alpha + x\beta) + 3.29$$

- Notice an important difference between OLS and Logistic Regression.
  - In OLS regression with an observed variable Y,  $V(Y)$  is fixed and the explained and unexplained variances change as variables are added to the model.
  - But in logistic regression with an unobserved variable  $y^*$ ,  $V(\varepsilon_{y^*})$  is fixed so the explained variance and total variance change as you add variables to the model.
  - This difference has important implications. Comparisons of coefficients between nested models and across groups do not work the same way in logistic regression as they do in OLS.

# Comparing Logit and Probit Coefficients across Models

```
. use http://www.nd.edu/~rwilliam/xsoc73994/statafiles/standardized.dta  
. logit ybinary x1, nolog
```

```
Logit estimates                Number of obs   =           500  
                               LR chi2(1)          =           161.77  
                               Prob > chi2         =           0.0000  
Log likelihood = -265.54468     Pseudo R2      =           0.2335
```

```
-----+-----  
      ybinary |      Coef.   Std. Err.      z    P>|z|     [95% Conf. Interval]  
-----+-----  
          x1 |   .7388678   .072961    10.13   0.000   .5958668   .8818687  
       _cons |  -.0529777   .105911    -0.50   0.617  -.2605593   .154604  
-----+-----
```

```
. logit ybinary x2, nolog
```

```
Logit estimates                Number of obs   =           500  
                               LR chi2(1)          =           160.35  
                               Prob > chi2         =           0.0000  
Log likelihood = -266.25298     Pseudo R2      =           0.2314
```

```
-----+-----  
      ybinary |      Coef.   Std. Err.      z    P>|z|     [95% Conf. Interval]  
-----+-----  
          x2 |   .4886751   .0482208   10.13   0.000   .3941641   .5831861  
       _cons |  -.0723833   .1058261    -0.68   0.494  -.2797986   .135032  
-----+-----
```

```
. logit ybinary x1 x2, nolog
```

```
. logit ybinary x1 x2, nolog
```

```
Logit estimates                               Number of obs   =           500
                                                LR chi2(2)      =           443.39
                                                Prob > chi2     =           0.0000
Log likelihood = -124.73508                    Pseudo R2       =           0.6399
```

ybinary	Coef.	Std. Err.	z	P> z	[95% Conf. Interval]
x1	1.78923	.1823005	9.81	0.000	1.431927 2.146532
x2	1.173144	.1207712	9.71	0.000	.9364369 1.409851
_cons	-.2144856	.1626906	-1.32	0.187	-.5333532 .1043821

Usually, when we add variables to a model (at least in OLS regression), the effects of variables added earlier goes down. However, in this case, we see that the coefficients for x1 and x2 increase (seemingly) dramatically when both variables are in the model, i.e. in the separate bivariate regressions the effects of x1 and x2 are .7388678 and .4886751, but in the multivariate regressions the effects are 1.78923 and 1.173144, more than twice as large as before. This leads to two questions:

1. If we saw something similar in an OLS regression, what would we suspect was going on? In other words, in an OLS regression, what can cause coefficients to get bigger rather than smaller as more variables are added?
2. In a logistic regression, why might such an interpretation be totally wrong?

```
. corr, means
```

```
(obs=500)
```

Variable	Mean	Std. Dev.	Min	Max
y	5.51e-07	3.000001	-8.508021	7.981196
ybinary	.488	.5003566	0	1
x1	-2.19e-08	2	-6.32646	6.401608
x2	3.57e-08	3	-10.56658	9.646875

	y	ybinary	x1	x2
y	1.0000			
ybinary	0.7923	1.0000		
x1	0.6667	0.5248	1.0000	
x2	0.6667	0.5225	<b>0.0000</b>	1.0000

- **x1 and x2 are uncorrelated! So suppressor effects cannot account for the changes in coefficients.**
- **Long & Freese's listcoef command can add some insights.**

```
. quietly logit ybinary x1
. listcoef, std
```

logit (N=500): Unstandardized and Standardized Estimates

Observed SD: .50035659  
**Latent SD: 2.3395663**

Odds of: 1 vs 0

ybinary	b	z	P> z	bStdX	bStdY	bStdXY	SDofX
x1	<b>0.73887</b>	10.127	0.000	1.4777	<b>0.3158</b>	0.6316	2.0000

```
. quietly logit ybinary x2
. listcoef, std
```

logit (N=500): Unstandardized and Standardized Estimates

Observed SD: .50035659  
**Latent SD: 2.3321875**

Odds of: 1 vs 0

ybinary	b	z	P> z	bStdX	bStdY	bStdXY	SDofX
x2	<b>0.48868</b>	10.134	0.000	1.4660	<b>0.2095</b>	0.6286	3.0000

```
. quietly logit ybinary x1 x2
. listcoef, std
```

logit (N=500): Unstandardized and Standardized Estimates

Observed SD: .50035659

**Latent SD: 5.3368197**

Odds of: 1 vs 0

ybinary	b	z	P> z	bStdX	bStdY	bStdXY	SDofX
x1	<b>1.78923</b>	9.815	0.000	3.5785	<b>0.3353</b>	0.6705	2.0000
x2	<b>1.17314</b>	9.714	0.000	3.5194	<b>0.2198</b>	0.6595	3.0000

- Note how the standard deviation of  $y^*$  fluctuates from one logistic regression to the next; it is about 2.34 in each of the bivariate logistic regressions and 5.34 in the multivariate logistic regression.
- It is because the variance of  $y^*$  changes that the coefficients change so much when you go from one model to the next. In effect, the scaling of  $Y^*$  is different in each model. By way of analogy, if in one OLS regression income was measured in dollars, and in another it was measured in thousands of dollars, the coefficients would be very different.

- Why does the variance of  $y^*$  go up? Because it has to. The residual variance is fixed at 3.29, so improvements in model fit result in increases in explained variance which in turn result in increases in total variance.
- Hence, comparisons of coefficients across nested models can be misleading because the dependent variable is scaled differently in each model.

- **How serious is the problem in practice?**
  - Hard to say. We easily found dozens of recent papers that present sequences of nested models. Their numbers are at least a little off, but without re-analyzing the data you can't tell whether their conclusions are seriously distorted as a result.
  - Several attempts of our own using real world data have failed to raise major concerns with the comparisons
  - We asked several authors for copies of their data, but most were unwilling or unable to do so.

- One author, Ervin (Maliq) Matthew, did graciously provide us with the data used for his paper “Effort Optimism in the Classroom: Attitudes of Black and White Students on Education, Social Structure, and Causes of Life Opportunities” (Sociology of Education 2011 84:225-245)
- The paper contains potentially problematic statements such as “The effect of race on the dependent variable is even stronger once GPA, SES, and sex are controlled for (Model 2), indicating that when blacks and whites have equal GPAs and family SES, blacks are more likely to agree with this statement.”
- In practice, however, we found that any potential errors were modest, with estimates being only slightly affected by solutions we discuss later. For example, his Table 7 modestly understates how much the effect of race declines as controls are added.

- **Nonetheless, researchers should realize that**
  - **Increases in the magnitudes of coefficients across models need not reflect suppressor effects**
  - **Declines in coefficients across models will actually be understated, i.e. you will be understating how much other variables account for the estimated direct effects of the variables in the early models.**
  - **Distortions are potentially more severe when added variables greatly increase the pseudo  $R^2$  statistics, as the variance of  $Y^*$  will increase more when that is the case.**

- **What are possible solutions?**
  - Just don't present the coefficients for each model in the first place. Researchers often present chi-square contrasts to show how they picked their final model and then only present the coefficients for it.
  - Use y-standardization. With y-standardization, instead of fixing the residual variance, you fix the variance of  $y^*$  at 1. This does not work perfectly, but it does greatly reduce rescaling of coefficients between models.
    - Listcoef gives the y-standardized coefficients in the column labeled bStdy, and they hardly changed at all between the bivariate and multivariate models (.3158 and .2095 in the bivariate models, .3353 and .2198 in the multivariate model).

- The Karlson/Holm/Breen (KHB) method (Papers are available or forthcoming in both *Sociological Methodology* and *Stata Journal*) shows promise
  - According to KHB, their method separates changes in coefficients due to rescaling from true changes in coefficients that result from adding more variables to the model (and does a better job of doing so than y-standardization and other alternatives)
  - They further claim that with their method the total effect of a variable can be decomposed into its direct effect and its indirect effect.

- We would add that, when authors estimate sequences of models, it is often because they want to see how the effects of variables like race decline (or increase) after other variables are controlled for. The KHB method provides a parsimonious and more accurate way of depicting such changes.
- We'll first present a simple example showing the relationship between diabetes, race & weight.

# khb example 1

```
. webuse nhanes2f, clear  
. khb logit diabetes black || weight
```

Decomposition using the KHB-Method

```
Model-Type:  logit                Number of obs   =   10335  
Variables of Interest: black      Pseudo R2       =    0.02  
Z-variable(s): weight
```

diabetes	Coef.	Std. Err.	z	P> z	[95% Conf. Interval]	
-----						
black						
Reduced	.6038012	.1236714	4.88	0.000	.3614098	.8461926
Full	.5387425	.1241889	4.34	0.000	.2953368	.7821483
Diff	.0650587	.0132239	4.92	0.000	.0391403	.0909771
-----						

- Possible interpretation of results
  - In the line labeled Reduced, only black is in the model. .6038 is the total effect of black.
  - However, blacks may have higher rates of diabetes both because of a direct effect of race on diabetes, and because of an indirect effect: blacks tend to be heavier than whites, and heavier people have higher rates of diabetes.
  - Hence, the line labeled Full gives the direct effect of race (.5387) while the line labeled Diff gives the indirect effect (.065)

## Khb Example 2

- Matthew (2011; see Table 7, p. 240) examines the determinants of how likely a student is to feel they will have a job he or she enjoys (0 = 50 percent or lower; 1 = better than 50 percent).
- In the first model, race (0 = white, 1 = black) is the only independent variable. The estimated effect of race is  $-.510$ .
- In the final model controls are added for GPA, SES, and others. The effect of race declines to  $-.471$ , an apparent  $-.039$  drop.
- The khb method shows that the decline is actually about twice as large. Again this is at least partly because the variance of  $y^*$  becomes greater as more variables are added, causing coefficients to increase.

```
. khb logit jobenjoy race || gpa ses sex educjob educimportant luckimportant sbprevent
```

Decomposition using the KHB-Method

```
Model-Type:  logit                      Number of obs   =   6731
Variables of Interest: race              Pseudo R2       =   0.08
Z-variable(s): gpa ses sex educjob educimportant luckimportant sbprevent
```

jobenjoy	Coef.	Std. Err.	z	P> z	[95% Conf. Interval]	
race						
Reduced	-.5727334	.10607	-5.40	0.000	-.7806269	-.3648399
Full	-.4833004	.1095584	-4.41	0.000	-.6980309	-.26857
Diff	-.089433	.0349898	-2.56	0.011	-.1580117	-.0208542

# Comparing Logit and Probit Coefficients across groups

- We often want to compare the effects of variables across groups, e.g. we want to see if the effect of education is the same for men as it is for women
- Both OLS and logistic regression assume that error variances are the same for both groups
- When that assumption is violated in OLS, the consequences are often minor: standard errors and significance tests are a bit off but coefficients remain unbiased.
- But when a binary or ordinal regression model incorrectly assumes that error variances are the same for all cases, the standard errors are wrong and (unlike OLS regression) the parameter estimates are wrong too.

- As Hoetker (2004, p. 17) notes, “in the presence of even fairly small differences in residual variation, naive comparisons of coefficients [across groups] can indicate differences where none exist, hide differences that do exist, and even show differences in the opposite direction of what actually exists.”
- **Explanation.** Suppose that  $y^*$  were observed, but our estimation procedure continued to standardize the variable by fixing its residual variance at 3.29. How would differences in residual variability across groups affect the estimated coefficients?
  - In the examples, the coefficients for the residuals reflect the differences in residual variability across groups.
  - Any residual that does not have a coefficient attached to it is assumed to already have a variance of 3.29

Case 1: True coefficients are equal, residual variances differ

	Group 0	Group 1
True coefficients	$y_i^* = x_{i1} + x_{i2} + x_{i3} + \varepsilon_i$	$y_i^* = x_{i1} + x_{i2} + x_{i3} + 2\varepsilon_i$
Standardized Coefficients	$y_i^* = x_{i1} + x_{i2} + x_{i3} + \varepsilon_i$	$y_i^* = .5x_{i1} + .5x_{i2} + .5x_{i3} + \varepsilon_i$

In Case 1, the true coefficients all equal 1 in both groups. But, because the residual variance is twice as large for group 1 as it is for group 0, the standardized  $\beta$ s are only half as large for group 1 as for group 0. *Naive comparisons of coefficients can indicate differences where none exist.*

Case 2: True coefficients differ, residual variances differ

	Group 0	Group 1
True coefficients	$y_i^* = x_{i1} + x_{i2} + x_{i3} + \varepsilon_i$	$y_i^* = 2x_{i1} + 2x_{i2} + 2x_{i3} + 2\varepsilon_i$
Standardized Coefficients	$y_i^* = x_{i1} + x_{i2} + x_{i3} + \varepsilon_i$	$y_i^* = x_{i1} + x_{i2} + x_{i3} + \varepsilon_i$

In Case 2, the true coefficients are twice as large in group 1 as in group 0. But, because the residual variances also differ, the standardized  $\beta$ s for the two groups are the same. Differences in residual variances obscure the differences in the underlying effects. *Naive comparisons of coefficients can hide differences that do exist.*

Case 3: True coefficients differ, residual variances differ even more

	Group 0	Group 1
True coefficients	$y_i^* = x_{i1} + x_{i2} + x_{i3} + \varepsilon_i$	$y_i^* = 2x_{i1} + 2x_{i2} + 2x_{i3} + 3\varepsilon_i$
Standardized Coefficients	$y_i^* = x_{i1} + x_{i2} + x_{i3} + \varepsilon_i$	$y_i^* = \frac{2}{3}x_{i1} + \frac{2}{3}x_{i2} + \frac{2}{3}x_{i3} + \varepsilon_i$

In Case 3, the true coefficients are again twice as large in group 1 as in group 0. But, because of the large differences in residual variances, the standardized  $\beta$ s are smaller for group 0 than group 1. Differences in residual variances make it look like the Xs have smaller effects on group 1 when really the effects are larger. *Naive comparisons of coefficients can even show differences in the opposite direction of what actually exists.*

## Example: Allison's (1999) model for group comparisons

- Allison (Sociological Methods and Research, 1999) analyzes a data set of 301 male and 177 female biochemists.
- Allison uses logistic regressions to predict the probability of promotion to associate professor.

*Table 1: Results of Logit Regressions Predicting Promotion to Associate Professor for Male and Female Biochemists (Adapted from Allison 1999, p. 188)*

<i>Variable</i>	<i>Men</i>		<i>Women</i>		<i>Ratio of Coefficients</i>	<i>Chi-Square for Difference</i>
	<i>Coefficient</i>	<i>SE</i>	<i>Coefficient</i>	<i>SE</i>		
Intercept	-7.6802***	.6814	-5.8420***	.8659	.76	2.78
Duration	1.9089***	.2141	1.4078***	.2573	.74	2.24
Duration squared	-0.1432***	.0186	-0.0956***	.0219	.67	2.74
Undergraduate selectivity	0.2158***	.0614	0.0551	.0717	.25	2.90
<i>Number of articles</i>	<i>0.0737***</i>	<i>.0116</i>	<i>0.0340**</i>	<i>.0126</i>	<i>.46</i>	<i>5.37*</i>
Job prestige	-0.4312***	.1088	-0.3708*	.1560	.86	0.10
Log likelihood	-526.54		-306.19			
Error variance	3.29		3.29			

*\*p < .05, \*\*p < .01, \*\*\* p < .001*

- As his Table 1 shows, the effect of number of articles on promotion is about twice as great for males (.0737) as it is for females (.0340).
- If accurate, this difference suggests that men get a greater payoff from their published work than do females, “a conclusion that many would find troubling” (Allison 1999:186).
- BUT, Allison warns, women may have more heterogeneous career patterns, and unmeasured variables affecting chances for promotion may be more important for women than for men.
  - Put another way, the error variance for women may be greater than the error variance for men
  - This corresponds to the Case I we presented earlier.
  - Unless the residual variability is identical across populations, the standardization of coefficients for each group will also differ.

# Allison's solution for the problem

- Ergo, in his Table 2, Allison adds a parameter to the model he calls delta. Delta adjusts for differences in residual variation across groups.

*Table 2: Logit Regressions Predicting Promotion to Associate Professor for Male and Female Biochemists, Disturbance Variances Unconstrained* (Adapted from Allison 1999, p. 195)

<i>Variable</i>	<i>All Coefficients Equal</i>		<i>Articles</i>	
	<i>Coefficient</i>	<i>SE</i>	<i>Coefficient Unconstrained</i>	<i>SE</i>
Intercept	-7.4913***	.6845	-7.3655***	.6818
Female	-0.93918**	.3624	-0.37819	.4833
Duration	1.9097***	.2147	1.8384***	.2143
Duration squared	-0.13970***	.0173	-0.13429***	.01749
Undergraduate selectivity	0.18195**	.0615	0.16997***	.04959
Number of articles	0.06354***	.0117	0.07199***	.01079
Job prestige	-0.4460***	.1098	-0.42046***	.09007
$\delta$	-0.26084*	.1116	-0.16262	.1505
<i>Articles x Female</i>			-0.03064	.0173
Log likelihood	-836.28		-835.13	

\* $p < .05$ , \*\* $p < .01$ , \*\*\* $p < .001$

- The delta-hat coefficient value  $-.26$  in Allison's Table 2 (first model) tells us that the standard deviation of the disturbance variance for men is 26 percent lower than the standard deviation for women.
  - This implies women have more variable career patterns than do men, which causes their coefficients to be lowered relative to men when differences in variability are not taken into account, as in the original logistic regressions.

- Allison's final model shows that the interaction term for Articles x Female is NOT statistically significant
- Allison concludes "The apparent difference in the coefficients for article counts in Table 1 does not necessarily reflect a real difference in causal effects. It can be readily explained by differences in the degree of residual variation between men and women."

# Problems with Allison's Approach

- Williams (2009) noted various problems with Allison's approach
- Allison says you should first test whether residual variances differ across groups
- To do this, you contrast two models:
  - In both cases, the coefficients are constrained to be the same for both groups
  - but in one model the residual variances are also constrained to be the same, whereas in the other model the residual variances can differ.
- Allison says that if the test statistic is significant, you then allow the residual variances to differ

- The problem is that, if the residual variances are actually the same across groups but the effects of the Xs differ, the test can also be statistically significant!
- Put another way, Allison's test has difficulty distinguishing between cross-group differences in residual variability & differences in coefficients.
  - Hence, his suggested procedure can make it much likely that you will conclude residual variances differ when they really don't
  - Erroneously allowing residual variances to differ can result in mis-estimates of other group differences, e.g. differences in coefficients can be underestimated

- Also, Allison's approach only allows for a single categorical variable in the variance equation. The sources of heteroskedasticity can be more complex than that; more variables may be involved, & some of these may be continuous
- Keele & Park (2006) show that a mis-specified variance equation, e.g. one in which relevant variables are omitted, can actually be worse than having no variance equation at all.

- Finally, Allison's method only works with a dichotomous dependent variable
  - Models with binary dvs that allow for heteroskedasticity can be difficult to estimate
  - Ordinal dependent variables contain more information about  $Y^*$
- Williams (2009, 2010) therefore proposed a more powerful alternative

# A Broader Solution: Heterogeneous Choice Models

- **Heterogeneous choice/ location-scale models explicitly specify the determinants of heteroskedasticity in an attempt to correct for it.**
- **These models are also useful when the variability of underlying attitudes is itself of substantive interest.**

# The Heterogeneous Choice (aka Location-Scale) Model

- Can be used for binary or ordinal models
- Two equations, choice & variance
- Binary case :

$$\Pr(y_i = 1) = g\left(\frac{x_i\beta}{\exp(z_i\gamma)}\right) = g\left(\frac{x_i\beta}{\exp(\ln(\sigma_i))}\right) = g\left(\frac{x_i\beta}{\sigma_i}\right)$$

- Allison's model with delta is actually a special case of a heterogeneous choice model, where the dependent variable is a dichotomy and the variance equation includes a single dichotomous variable that also appears in the choice equation.
- Allison's results can easily be replicated with the user-written routine `oglm` (Williams, 2009, 2010)

```

. * oglm replication of Allison's Table 2, Model 2 with interaction added:
. use "http://www.indiana.edu/~jslsoc/stata/spex_data/tenure01.dta", clear
(Gender differences in receipt of tenure (Scott Long 06Jul2006))
. keep if pdasample
(148 observations deleted)
. oglm tenure female year yearsq select articles prestige f_articles, het(female)

```

```

Heteroskedastic Ordered Logistic Regression      Number of obs   =      2797
                                                LR chi2(8)      =      415.39
                                                Prob > chi2     =      0.0000
Log likelihood = -835.13347                    Pseudo R2       =      0.1992

```

	Coef.	Std. Err.	z	P> z	[95% Conf. Interval]	
-----						
tenure						
female	-.3780597	.4500207	-0.84	0.401	-1.260084	.5039646
year	1.838257	.2029491	9.06	0.000	1.440484	2.23603
yearsq	-.1342828	.017024	-7.89	0.000	-.1676492	-.1009165
select	.1699659	.0516643	3.29	0.001	.0687057	.2712261
articles	.0719821	.0114106	6.31	0.000	.0496178	.0943464
prestige	-.4204742	.0961206	-4.37	0.000	-.6088671	-.2320813
f_articles	-.0304836	.0187427	-1.63	0.104	-.0672185	.0062514
-----						
lnsigma						
female	.1774193	.1627087	1.09	0.276	-.141484	.4963226
-----						
/cut1	7.365285	.6547121	11.25	0.000	6.082073	8.648497
-----						

```

. display "Allison's delta = " (1 - exp(.1774193)) / exp(.1774193)
-.16257142

```

- As Williams (2009) notes, there are important advantages to turning to the broader class of heterogeneous choice models that can be estimated by `oglm`
  - Dependent variables can be ordinal rather than binary. This is important, because ordinal vars have more information and hence lead to better estimation
  - The variance equation need not be limited to a single binary grouping variable, which (hopefully) reduces the likelihood that the variance equation will be misspecified

- Williams (2010) also notes that, even if the researcher does not want to present a heterogeneous choice model, estimating one can be useful from a diagnostic standpoint
  - Often, the appearance of heteroskedasticity is actually caused by other problems in model specification, e.g. variables are omitted, variables should be transformed (e.g. logged), squared terms should be added
  - Williams (2010) shows that the heteroskedasticity issues in Allison's models go away if  $\text{articles}^2$  is added to the model

```

. use "http://www.indiana.edu/~jslsoc/stata/spex_data/tenure01.dta", clear
(Gender differences in receipt of tenure (Scott Long 06Jul2006))
. keep if pdasample
(148 observations deleted)
. * hetero effect becomes insignificant when articles^2 is added to model
. oglm tenure i.female year c.year#c.year select articles prestige c.articles#c.articles, het(i.female)

```

```

Heteroskedastic Ordered Logistic Regression      Number of obs   =      2797
                                                LR chi2(8)      =      440.07
                                                Prob > chi2     =      0.0000
Log likelihood = -822.79102                    Pseudo R2       =      0.2110

```

tenure		Coef.	Std. Err.	z	P> z	[95% Conf. Interval]	
-----							
tenure							
1.female		-.5612041	.2926285	-1.92	0.055	-1.134746	.0123373
year		1.739173	.1933258	9.00	0.000	1.360261	2.118084
c.year#c.year		-.1265911	.0162677	-7.78	0.000	-.1584751	-.094707
select		.1710519	.0504271	3.39	0.001	.0722167	.2698872
articles		.1533931	.0244003	6.29	0.000	.1055694	.2012168
prestige		-.454951	.0936162	-4.86	0.000	-.6384354	-.2714666
c.articles#c.articles		-.0026412	.0007213	-3.66	0.000	-.004055	-.0012274
-----							
lnsigma							
1.female		.141633	.1377843	1.03	0.304	-.1284193	.4116853
/cut1		7.40805	.648316	11.43	0.000	6.137374	8.678726
-----							

. \* You don't need the female\*articles interaction terms either

```
. oglm tenure i.female year c.year#c.year select articles prestige c.articles#c.articles  
i.female#(c.articles c.articles#c.articles)
```

Ordered Logistic Regression

Number of obs = 2797  
LR chi2(9) = 439.05  
Prob > chi2 = 0.0000  
Pseudo R2 = 0.2105

Log likelihood = -823.3041

tenure	Coef.	Std. Err.	z	P> z	[95% Conf. Interval]	
1.female	-.2588495	.3094906	-0.84	0.403	-.8654399	.3477409
year	1.649057	.1651954	9.98	0.000	1.32528	1.972834
c.year#c.year	-.11984	.0142338	-8.42	0.000	-.1477378	-.0919423
select	.1583254	.0466906	3.39	0.001	.0668136	.2498372
articles	.1492724	.0295934	5.04	0.000	.0912703	.2072745
prestige	-.4386555	.0898018	-4.88	0.000	-.6146637	-.2626472
c.articles#c.articles	-.0025455	.0009236	-2.76	0.006	-.0043558	-.0007352
female#c.articles						
1	-.007599	.0450029	-0.17	0.866	-.0958031	.0806051
female#c.articles#c.articles						
1	.0001025	.0013542	0.08	0.940	-.0025517	.0027568
/cut1	7.091958	.5479358	12.94	0.000	6.018024	8.165893

# Problems with heterogeneous choice models

- Models can be difficult to estimate, although this is generally less problematic with ordinal variables
- While you have more flexibility when specifying the variance equation, a mis-specified equation can still be worse than no equation at all
- But the most critical problem of all may be...

# Problem: Radically different interpretations are possible

- An issue to be aware of with heterogeneous choice models is that radically different interpretations of the results are possible
  - Hauser and Andrew (2006), for example, proposed a seemingly different model for assessing differences in the effects of variables across groups (where in their case, the groups were different educational transitions)
  - They called it the *logistic response model with proportionality constraints* (LRPC):

$$\log_e \left( \frac{p_{ij}}{1 - p_{ij}} \right) = \beta_{j0} + \lambda_j \sum_k \beta_k X_{ijk}$$

- Instead of having to estimate a different set of coefficients for each group/transition, you estimate a single set of coefficients, along with one  $\lambda_j$  proportionality factor for each group/ transition ( $\lambda_1$  is constrained to equal 1)
  - The proportionality constraints would hold if, say, the coefficients for the 2<sup>nd</sup> group were all 2/3 as large as the corresponding coefficients for the first group, the coefficients for the 3<sup>rd</sup> group were all half as large as for the first group, etc.

# Models compared

$$\Pr(y_i = 1) = g\left(\frac{x_i\beta}{\exp(z_i\gamma)}\right) = g\left(\frac{x_i\beta}{\exp(\ln(\sigma_i))}\right) = g\left(\frac{x_i\beta}{\sigma_i}\right)$$

$$\log_e\left(\frac{p_{ij}}{1-p_{ij}}\right) = \beta_{j0} + \lambda_j \sum_k \beta_k X_{ijk}$$

- Hauser & Andrew note, however, that “one cannot distinguish empirically between the hypothesis of uniform proportionality of effects across transitions and the hypothesis that group differences between parameters of binary regressions are artifacts of heterogeneity between groups in residual variation.” (p. 8)
- Williams (2010) showed that, even though the rationales behind the models are totally different, the heterogeneous choice models estimated by oglm produce identical fits to the LRPC models estimated by Hauser and Andrew; simple algebra converts one model’s parameters into the other’s
- Williams further showed that Hauser & Andrew’s software produced the exact same coefficients that Allison’s software did when used with Allison’s data

```
. * Hauser & Andrew's original LRPC program
. * Code has been made more efficient and readable,
. * but results are the same. Note that it
. * actually estimates and reports
. * lambda - 1 rather than lamba.
. program define lrpc02
1.     tempvar theta
2.     version 8
3.     args lnf  intercepts lambdaminus1 betas
4.     gen double `theta' = `intercepts' + `betas' + (`lambdaminus1' * `betas')
5.     quietly replace `lnf' = ln(exp(`theta')/(1+exp(`theta'))) if $ML_y1==1
6.     quietly replace `lnf' = ln(1/(1+exp(`theta'))) if $ML_y1==0
7. end
```

```

. * Hauser & Andrews original LRPC parameterization used with Allison's data
. * Results are identical to Allison's Table 2, Model 1
. ml model lf lrpc02 ///
>     (intercepts: tenure = male female, nocons) ///
>     (lambdaminus1: female, nocons) ///
>     (betas: year yearsq select articles prestige, nocons), max nolog
. ml display

```

```

Log likelihood = -836.28235
Number of obs   =      2797
Wald chi2(2)    =      180.60
Prob > chi2     =      0.0000

```

tenure	Coef.	Std. Err.	z	P> z	[95% Conf. Interval]	
-----+-----						
intercepts						
male	-7.490506	.6596634	-11.36	0.000	-8.783422	-6.197589
female	-6.230958	.6205863	-10.04	0.000	-7.447285	-5.014631
-----+-----						
lambdaminus1						
female	-.2608325	.1080502	-2.41	0.016	-.4726069	-.0490581
-----+-----						
betas						
year	1.909544	.1996937	9.56	0.000	1.518151	2.300936
yearsq	-.1396868	.0169425	-8.24	0.000	-.1728935	-.1064801
select	.1819201	.0526572	3.45	0.001	.0787139	.2851264
articles	.0635345	.010219	6.22	0.000	.0435055	.0835635
prestige	-.4462074	.096904	-4.60	0.000	-.6361357	-.2562791
-----+-----						

- But, the theoretical concerns that motivated their models and programs lead to radically different interpretations of the results.
  - According to Allison's theory (and the theory behind the heterogeneous choice model) apparent differences in effects between men and women are an artifact of differences in residual variability.
  - Once these differences are taken into account, there is no significant difference in the effect of articles across groups, implying there is no gender inequality in the tenure process.

- Someone looking at these exact same numbers from the viewpoint of the LRPC, however, would conclude that the effect of articles (and every other variable for that matter) is 26 percent smaller for women than it is men.
- Those who believed that the LRPC was the theoretically correct model would likely conclude that there is substantial gender inequality in the tenure promotion process.
- For any given problem, strong substantive arguments might be made for one perspective or the other.
- Researchers using any of these models should realize, however, that there is often if not always a radically different interpretation that, empirically, fits the data just as well.

# Long's solution

- Long (2009) looks at these same sorts of problems, but proposes a different analytical approach. He says
  - “An alternative approach [to Allison]... uses predicted probabilities. Since predicted probabilities are unaffected by residual variation, tests of the equality of predicted probabilities across groups can be used for group comparisons without assuming the equality of the regression coefficients of some variables... Testing the equality of predicted probabilities requires multiple tests since group differences in predictions vary with the levels of the variables in the model.”

# A simple example of Long's technique

```
. use "http://www.indiana.edu/~jslsoc/stata/spex_data/tenure01.dta", clear
(Gender differences in receipt of tenure (Scott Long 06Jul2006))
. keep if year <= 10
(148 observations deleted)
. * Basic model - articles only
. logit tenure articles i.male i.male#c.articles, nolog
```

Logistic regression

```
Number of obs = 2797
LR chi2(3) = 121.58
Prob > chi2 = 0.0000
Pseudo R2 = 0.0583
```

Log likelihood = -982.04029

tenure	Coef.	Std. Err.	z	P> z	[95% Conf. Interval]	
articles	.0471351	.0104974	4.49	0.000	.0265605	.0677097
1.male	-.2198428	.1853876	-1.19	0.236	-.5831959	.1435102
male#c.articles						
1	.0552514	.0148436	3.72	0.000	.0261585	.0843444
_cons	-2.501162	.140056	-17.86	0.000	-2.775667	-2.226657

```
. margins, dydx(male) at(articles=(0(1)50)) vsquish
```

```
Conditional marginal effects      Number of obs   =      2797  
Model VCE      : OIM
```

```
Expression      : Pr(tenure), predict()  
dy/dx w.r.t.    : 1.male  
1._at          : articles      =      0  
2._at          : articles      =      1  
[output deleted]  
51._at         : articles      =     50
```

```
-----
```

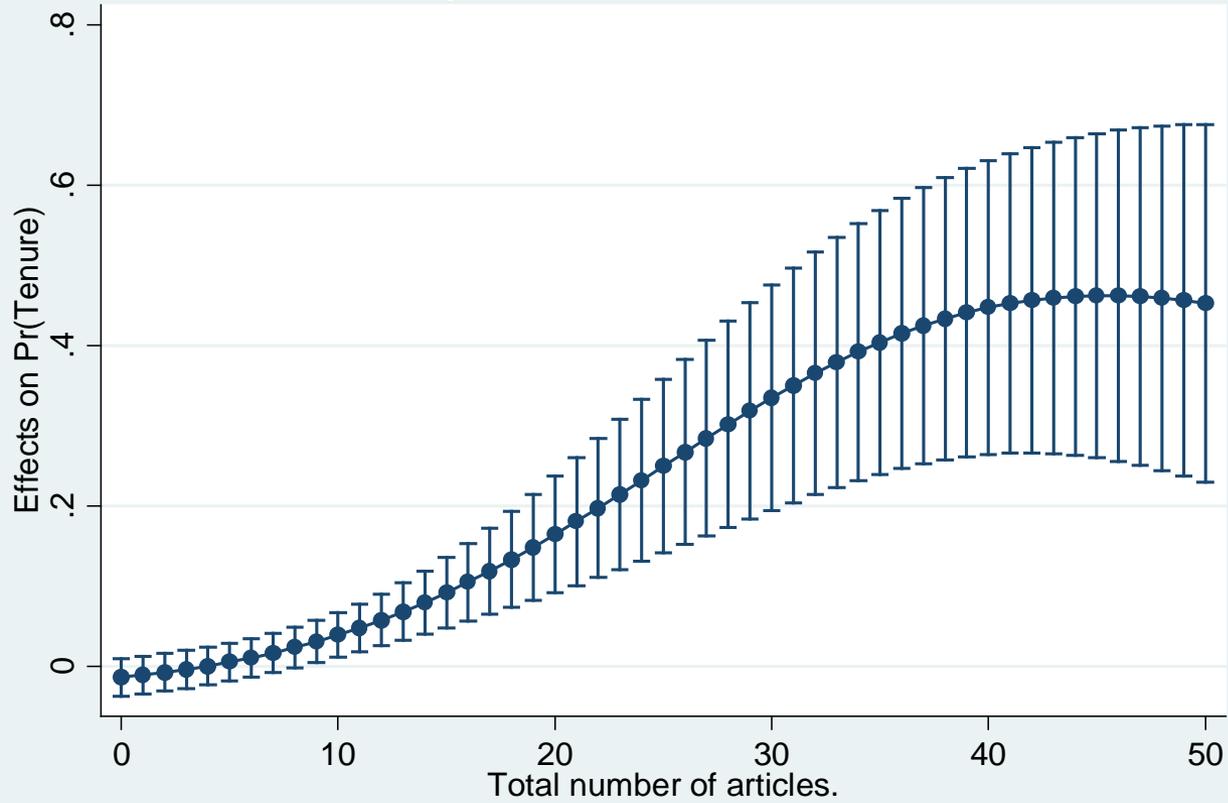
		Delta-method			[95% Conf. Interval]	
		dy/dx	Std. Err.	z	P> z	
1.male						
	_at					
	1	-.0140315	.0120717	-1.16	0.245	-.0376916 .0096285
	2	-.0111948	.0120559	-0.93	0.353	-.0348239 .0124343
	[output deleted]					
	50	.4562794	.1118036	4.08	0.000	.2371485 .6754104
	51	.4527383	.1139979	3.97	0.000	.2293066 .67617

```
-----
```

Note: dy/dx for factor levels is the discrete change from the base level.

```
. marginsplot
```

Conditional Marginal Effects of 1.male with 95% CIs



- This simple example shows that the predicted probabilities of tenure for men and women differ little for small numbers of articles; indeed the differences are not even statistically significant for 8 articles or less.
- The differences become greater as the number of articles increases. For example, a women with 40 articles is predicted to be 45 percent less likely to get tenure than a man with 40 articles.

- The analyses can be further extended by adding more variables to the model, and/or by doing various subgroup analyses, e.g. comparing women at high-prestige universities with men at high prestige Universities
- As Long says, this can lead to “more complex conclusions on how groups differ in the effect of a variable.”
- If you are lucky, the differences in predicted probabilities may disappear altogether, e.g. variables added to the model may be able to account for the initially observed group differences.
- But if they don't...

# Critique of Long

- The predictive margins produced by Long's approach might be seen as a sort of high-tech descriptives. They illustrate the predicted differences between groups after controlling for other variables.
- Description can be very useful. In this case we see that the predicted probabilities of tenure differ dramatically by gender and the number of articles published.
- Once such differences in predicted probabilities are discovered, policy makers may decide that some sort of corrective action should be considered.

- At the same time, Long's approach may be frustrating because it doesn't try to explain why the differences exist.
  - Are the differences due to the fact that men are rewarded more for the articles they publish?
  - Or, are they due to the fact that residual variability differs by gender? Perhaps women's careers are disrupted more by family or other matters.

- From a policy standpoint, we would like to know what is causing these observed differences in predicted probabilities
  - If it is because women are rewarded less for each article they write, we may want to examine if women's work is not being evaluated fairly
  - If it is because of differences in residual variability, we may want to further examine why that is. For example, if family obligations create more career hurdles for women than they do men, how can we make the workplace more family-friendly?
  - But if we do not know what is causing the differences, we aren't even sure where to start if we want to eliminate them.

- Long defends his approach by arguing:
  - For many things, like his report on women in science for the NAS, predictions were of much more interest than was the slope of articles or unobserved heterogeneity.
  - using other information, e.g. on the history of women in science, may resolve issues far more effectively than the types of assumptions that are needed to be able to disentangle differences in coefficients and unobserved heterogeneity
  - there are times when predictive margins provide more insights than simple answers to yes no hypotheses. For example, there can be cases where, overall the lines for men and women are the same (can't reject they are equal), yet they differ significantly when testing equality at a particular case. Both are valid, but overreliance on one, omnibus test is not a good thing in general.

- Further, as we have seen, when we try to explain group differences, the coefficients can be interpreted in radically different ways.
  - Two researchers could look at the exact same set of results, and one could conclude that coefficients differ across groups while another could say that it is residual variability that differs.
- Given such ambiguity, some might argue that you should settle for description and not strive for explanation (at least not with the current data).

- Others might argue that you should go with the model that you think makes most theoretical sense, while acknowledging that alternative interpretations of the results are possible.
- At this point, it is probably fair to say that the descriptions of the problem may be better, or at least more clear-cut, than the various proposed solutions.

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For more information, see:

<https://www3.nd.edu/~rwilliam>