RESEARCH NOTES

AN APPLICATION OF THE ESTIMATED DEPENDENT VARIABLE APPROACH: TRADE UNION MEMBERS' SUPPORT FOR ACTIVE LABOR MARKET POLICIES AND INSIDER-OUTSIDER POLITICS

Moira Nelson

Methodological responses to the issues associated with multilevel data quickened in recent years. Multilevel models, as a generalization of regression techniques, perform at least as good or better than regression models (Gelman, 2006; Gelman & Hill, 2007); at the same time, beyond explicitly modeling the multilevel nature of a given data structure, multilevel models exhibit differences that practitioners should observe. Recent accounts recommend the "estimated dependent variable" approach where the number of first level variable is high (often individuals) and the number of second-level variables is low (often countries, interviewers, election periods, *etc.*) (Franzese, 2005; Huber, Kernell, & Leoni, 2005; Jusko & Shively, 2005; Lewis & Linzer, 2005).

Given the prevalence of such a data structure in international public opinion research, this article sets out to make the estimated dependent variable more accessible by examining the effect of contextual effects on union members' preferences over active labor market policies. The note proceeds as follows: the first section outlines the estimated dependent variable approach; the next section clarifies the theoretical puzzle of trade union member support for active labor market policies as well as the data; the third section includes the analysis; the final section concludes.

ADVANTAGES OF THE ESTIMATED DEPENDENT VARIABLE APPROACH

An explicit way to incorporate contextual effects into statistical models includes hierarchical linear models (Raudenbush & Bryk, 2001; Snijders & Bosker, 1999; Steenbergen & Jones, 2002). When building a multilevel model, there are three main decisions to consider: how many levels do you want to model, how many predictors do you want to include, and do you want to model level one slopes, intercepts,

or both as a function of contextual variables (Luke, 2004)? When assessing the effect of the level two variables on the level one variables constitutes the main focus, scholars often use random effects models such as the "means as outcomes" model and the "slopes and intercepts as outcomes" model as well as Bayesian estimation models. The estimated dependent variable (EDV) approach provides an alternative that is easier to employ than Bayesian models and addresses a small sample problem that frequently frustrates estimation of random effects models.

Much research interested in testing contextual effects of level two variables on level one variables runs into the situation of too few level two observations. The small sample problems of MLE models are well-documented (e.g. Hox & Maas, 2001; Hu, Bentler, & Kano, 1992; Lee & Song, 2004). Of particular importance to the discussion here is that random effects models demand at least 50 level-two units when assessing the effect of contextual variables (Maas & Hox, 2004). The EDV approach addresses this problem by estimating the level-1 effects in the first stage without any level-2 predictors. In this way, the EDV approach capitalizes on the typically high number of level-1 observations, because this first stage leads to reliable estimates of the level one coefficients. Although the number of level-1 observations would ideally be equal across contextual (level 2) units, MLE works well even with unbalanced datasets (Raudenbush, 1988). To cope with difference variances between contextual groups, however, one can estimate feasible GLS (Jusko & Shively, 2005; Lewis & Linzer, 2005).

The coefficients from the level one analysis become the dependent variable in the second stage of the EDV approach and can be regressed on contextual variables using OLS, which does not rely on the same asymptotic assumptions as MLE. There are no hard and fast rules for identifying a minimum sample size for the second stage OLS analysis. Montenegro (2001) recommends a conservative 10 observations per parameter, although other analyses of the EDV approach often used about 20–30 observations with about five parameters (Huber, Kernell, & Leoni, 2005; Jusko & Shively, 2005).

The central advantage of the EDV approach is that these coefficients are easy to understand. For example, Bowers and Drake (2005) show that coefficients from the level one analysis can be valuable in analyzing relationships graphically when asymptotic assumptions of MLE do not hold due to few level two units. Moreover, evidence that this approach performs just as well as a pooled estimation technique (Duch & Stevenson, 2005) suggests that there is not necessarily a high price for the user-friendly advantage of the EDV approach. Of course this may change in the near future, given the growing number of statistical packages making the estimation of single-step methods more accessible (Beck, 2005).

TRADE UNION MEMBERSHIP AND SUPPORT FOR ACTIVE LABOR MARKET POLICIES

The present study applies the EDV approach to the analysis of trade union members' preferences over active labor market policies. Trade union members historically support generous social policy, and, as a subset of this group, active labor market policies play a central role in the present economic environment. Active labor market

policies aim at facilitating labor market participation and generally include job counseling, public job creation policies, retraining programs, and wage subsidies to private firms.

Contextual effects portend to increase trade union members' support for active labor market policies in two ways. First, in contexts in which job insecurity is high, trade union members are hypothesized to support active labor market policies to a high degree either out of concern for their own employment prospects or for the prospects of other workers, for whom they feel responsible. Indicators of job insecurity include high unemployment, low employment protection, and rapid deindustrialization.

Beyond conditions that increase job insecurity, the organizational resources of trade unions increase the capacity to influence members' preferences. Organizational resources, in the form of high levels of union density and high rates of union participation in policy making, are hypothesized to lead to stronger trade union support, because trade unions have more channels through which to inform members about the advantages of active labor market policies for the workers, whether trade union members or not. Measures for organizational resources include union density and a variable for union participation in policy-making.

Contrary to these conventional hypotheses, however, the insider-outsider theory creates opposite hypotheses about the effects of job insecurity and organizational resources on trade union members' preferences. For theoretical clarification, scholars within the insider-outsider approach refute the assumption within classical economics of competitive labor markets and contend that stably employed individuals are able to reduce job competition (and thereby increase their personal job security) by manipulating their superior knowledge of their work conditions to the detriment of those seeking employment (Lindbeck & Snower, 1988, 1990, 2002). Given the evidence that trade union members are generally securely employed (*e.g.* Polacheck & McCutcheon, 1983), trade union members are considered insiders, who should exhibit a preference for supporting policies that secure their insider position and opposing those policies that erode it.

For this reason, trade union members' role as labor market insiders provides them with an incentive to oppose active labor market policies. High spending on active labor market policies promotes the competitiveness of labor market outsiders, reducing insiders' advantageous position. Therefore, union member opposition will increase when job insecurity rises, which occurs when labor markets are more flexible, such as when employment protection is low, or tighter, such as when unemployment or deindustrialization are high.

Also, given that trade union members may oppose active measures, the conventional hypothesis that trade unions bolster member support may in fact be reversed. Trade unions may indeed caution their members that active labor market policies lead to a deterioration of working conditions by undermining the traditional full-time, permanent contract. High union density and increased participation in policy-making may therefore have the effect of dampening trade union member support.

To summarize, the theoretical expectations for trade union members' preferences over active labor market policies are unclear. On the one hand, trade union members constitute the traditional support base for generous social policies. Within contexts of

heightened job insecurity and strong organizational resources, trade union members should increase their support for active measures in order to improve their own employability and that of workers more generally. On the other hand, trade union members represent labor market insiders who are able to secure their own job security by reducing competition. Since supporting active labor market policies would increase competition for their jobs, trade union members should oppose these policies. The following analysis provides an empirical test of this theoretical quandary by explicitly modeling the role of contextual factors in explaining trade union members' preferences over active labor market policies.

ANALYSIS

The public opinion data employed in this study comes from three waves (1985, 1990, and 1996) of the International Social Survey Programme's Role of Government surveys.[1] During these three rounds of the ISSP's Role of Government Survey, individuals were asked about their preferences towards job creation policies. Specifically, individuals were asked: "Here are some things that the government might do for the economy. Circle one number for each action to show whether you are in favor of it or against it. Here, Government financing of projects to create new jobs." The response categories range from 1 ("strongly disagree") to 5 ("strongly agree").[2] The most individuals responded "Agree" (46 percent), whereas a slightly smaller group answered "Strongly Agree" (31 percent); the remaining 23 percent was distributed among the remaining three answer categories to create a single-peaked distribution. Although a total of 39 countries participated in the three rounds of the Role of Government survey, high standards of comparability of the initial analysts as well as their interests in particular questions restricted the total number of countries included in the final dataset to 22.

The chosen question is preferred for two reasons. First, by invoking the need of government funds to finance job creation programs, respondents are confronted with the opportunity cost of program financing, namely lower taxes or funding for alternative policies. In doing so, this question arguably taps into respondents' actual spending preferences rather than normative beliefs about the role of government. Second, this question addresses the goal of creating jobs and does so without specifying a particular policy tool, which avoids conflating preferences for job creation with preferences over other issues. To clarify, additional questions regarding the ways that the government could improve the economy included, for example, controlling prices and wages and cutting government spending. The implications of these measures on job growth, however, are not straightforward, and therefore preferences over job creation cannot be ascertained. On the other hand, one question indeed asks respondents if the government should support declining industry to protect jobs. Although the question directly invokes job creation, governmental protection of

[2] The five answer categories include: 1 "Strongly Agree"; 2 "Agree"; 3 "Neutral"; 4 "Disagree"; 5 "Strongly Disgree"; 8 "Can't Choose"; and 9 "NA." Values 8 and 9 were switched to missing values and the scale of the remaining response variables was reversed to facilitate interpretation.

declining industry is largely recognized as a hazard to economic growth (*e.g.* Hufbauer & Elliott, 1994). As such, this question specification portends to negatively bias the responses. In short, although similar survey questions exist which could be used to construct a composite index of measures aimed at improving the economy (*e.g.* Bean & Papadakis, 1998), the phrasing of the chosen question best captures this article's specific interest in support over active labor market policies.

Turning to the independent variables, on the individual level, gender, income, education, age, and unemployment status are included as controls. The main independent variable of interest is a dummy variable for union membership. Since I centre all non-dummy variables, the baseline category is an employed, 40 years old male, non-union member with average income. The dependent variable is the five-pronged ordered categorical variable constructed from the ISSP question described above and labeled *ALMP* for active labor market policies in the equation below. As a result of missing values for the independent variables in the second, or contextual, level, Spain is dropped from the analysis. In addition, data on Germany includes West Germany for 1985 and 1990 and both West and East Germany for 1996. The sample is constrained to those in the labor force (specifically the employed and the unemployed).

The contextual variables used include unemployment levels, employment protection, deindustrialization, union density, and union participation in policy making. The variables for unemployment and union density are lagged by one year. Deindustrialization is measured following Iversen and Cusack (2000) as 100 minus the percent of the workforce employed in the industrial and agricultural sectors. In order to capture dynamics, the variable is measured as the average change in deindustrialization over the previous five year period. All three variables are derived from Huber and Stephens Comparative Welfare States Dataset.[3] Employment protection is a four-point measure provided by OECD Labour Force Statistics Database. Finally, the variable measuring union participation in policy-making comes from the index provided by Traxler, Blaschke, and Kittel (2001).[4]

The methodological technique is the two-stage approach to model contextual factors. In the first step, ordered probit regressions for each country-year are performed.[5] In the second step, the coefficients for the union membership variable in each probit regression, β_{1jt}, become the dependent variable. In this step, the coefficients for union membership are regressed on the contextual variables of interest.

---

[3] http://www.lisproject.org/publications/welfaredata/welfareaccess.htm.
[4] The scores for this variable vary between 0 and 1 and represent an average of union participation in 12 different policies areas of which 2 are categorized as general activities and 10 are categorized as specialized activities. The generalized activities include influence over national government or parliamentary bodies with regard to labour market activities and representation of union interests on national corporatist institutions. The specialized activities include whether unions participate or implement (or participate in the implementation of) in five different programs including industrial policy programs, regional development programs, public occupational programs and active labour market policies, research and development programs, and quality control programs and/or standardization of product. The generalized activities are weighted by 9 since they represent the central dimensions of union participation, whereas specialized activities are weighted by 8. For more information refer to (Traxler, Blaschke, & Kittel, 2001).
[5] The individual level analysis was also conducted using OLS regression. The results were substantively the same.

The first step is ordered probit analysis where the coefficients are estimated for each country separately. The model is

$$ALMP_{ijt} = \alpha_{ijt} + M_{ijt}\beta_{1jt} + I_{ijt}\beta_{2jt} + E_{ijt}\beta_{3jt} + A_{ijt}\beta_{4jt} + G_{ijt}\beta_{5jt} + U_{ijt}\beta_{6jt} + \varepsilon_{ijt}$$
$$i = 1, \ldots, I, \, i = 1, \ldots, J, \, t = 1, \ldots, T \tag{1}$$

where j presents the separate country estimations and i and t stand for individual and year, respectively. The key variable in the analysis is M, which stands for union membership. The remaining variables are controls and include the following: I for income, E for education, A for age, G for being female and U for unemployment status.

I will briefly discuss the hypotheses for the control variables. Low paid, less educated individuals should face higher labor market risk and therefore be more supportive of active labor market policies. The effect of age is ambiguous, because older workers face a relatively shorter window to find new employment if they lose their job, although they are also more likely than younger workers to be protected by "last in, first out" laws. Women should demonstrate higher support for active measures that facilitate labor market integration since they confront the risk of taking career breaks to have children. The unemployed should support active measures to improve the success of their job search efforts. The directional hypotheses are all confirmed by the individual level analysis. The results for age demonstrate lower support among older workers.

To set up the second level, or contextual, analysis, a bit more discussion of the union membership variable is warranted. In the individual country-year estimations shown above, union membership is only negatively related to support for active measures in one case (Italy 1996), and it is not significant. In every other case, union membership is positively related to preferences for active measures. This positive relationship in 21 of 22 cases is statistically significant in 11 of 22 country-years.

Predicted probabilities are a way to clarify the effect of union membership on the probability of falling into each answer category for each country-year in the sample. For example, in Australia in 1985, holding all other variable at their mean (or minimum for dichotomous variables), union membership has the effect of increasing the probability of an individual switching into the "strongly agree" category by 7.15 percent. Turning to the "strongly agree" category, union members always have a higher predicted probability of choosing this answer than non-members. The converse is true for the neutral options and the disagree categories—union members are less likely to fall into these categories than non-members. The agree category seems to be an in between category, where union members are sometimes more likely to fall into this category and sometimes less likely to fall into this category than non-members. This is most likely the case in countries where union members have a high probability of answering "strongly agree."

The average change in predicted probability is the absolute value of each response category divided by the total number of response categories. The average change provides a measurement of the total weight or influence of trade union membership on individuals' preferences towards active policies.

In light of the empirical analysis, the average change statistic represents not only the magnitude of the effect of trade union membership, but also the direction of

this effect. Having established a positive relationship between union membership and support for active measures, the average change in predicted probability can be seen to represent the strength of the union membership variable in increasing support for active labor market policies. Given the similarity in substantive meaning between the average change in predicted probability and the coefficient for union membership, it is not surprising that the correlation between the two is .94.

This review of the coefficients of union membership and their substantive effect on individuals' probability of supporting active labor market policies sets the stage for the second part of the analysis. In this second part, the coefficients for union membership are regressed on the aforementioned of contextual variables. This second step, then, measures the effect of the contextual variables on the probability that union membership increases support for active labor market policies.

The models in the second step are corrected using OLS techniques with White's heteroskedastic consistent standard errors. This method has been shown to be superior under many circumstances (Lewis & Linzer, 2005).[6]

Following the theoretical discussion, I run both a "job insecurity" model and an "organizational resources" model. The model for the job insecurity preferences,

$$\beta_{jt} = \Phi_0 + DC_{jt}\Phi_2 + UL_{jt}\Phi_1 + EPL_{jt}\Phi_3 + \varepsilon_{jt}, \tag{2}$$

measures the salience of union support for active labor market policies, which are captured by the coefficients of union membership determined from the results in model (1). In this model, β_{jt} is the coefficient for union membership from model (1), DC is changes in deindustrialization, UL is level of unemployment, and EPL is employment protection. A second model captures the effect of organizational structures on trade union members' preferences. The model

$$\beta_{jt} = \Phi_0 + UP_{jt}\Phi_1 + DEN_{jt}\Phi_2 + \varepsilon_{jt}, \tag{3}$$

includes UP, a measure of union participation in policy-making, and DEN, a measure of union density.

The results are listed below in Table 1. In regards to the job insecurity dimension, trade union member support appears to increase in contexts of rapid deindustrialization and low employment protection. High unemployment, however, leads to lower support. Turning to the organizational resources model, the results provide support for the claim that union density increases support for active labor market policies. The actual participation of union organizations in policy-making, however, does not significantly increase union member support for active measures.
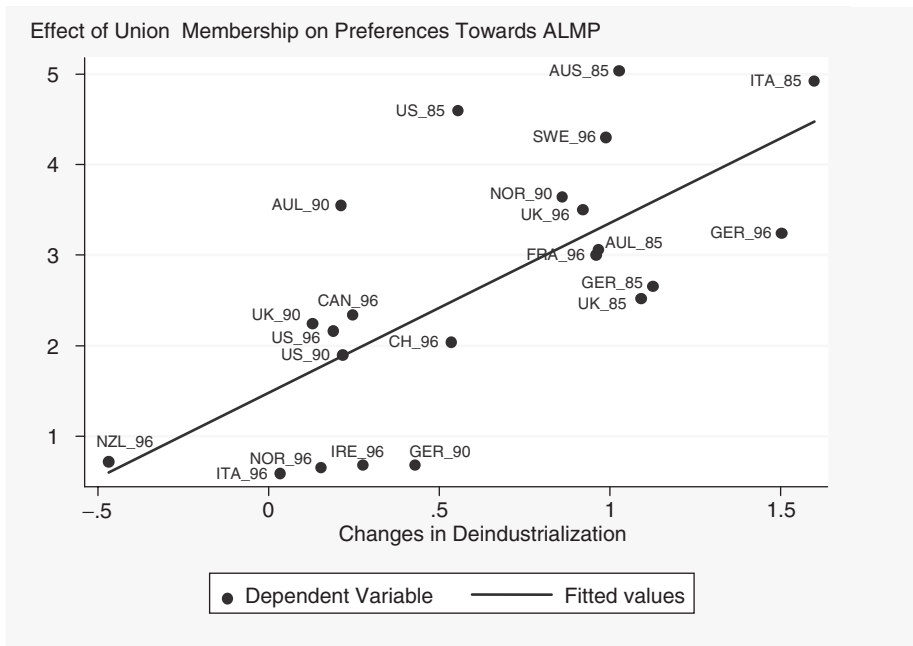
In order to interpret the substantive effects of these results, it is useful to look at the country-years with the lowest and highest values of the variable in question and the corresponding predicted probabilities on the dependent variable. With regards to deindustrialization, the lowest rate of deindustrialization is found in New Zealand in 1996 and the highest rate of deindustrialization is found in Italy in 1985. The overall

---

[6] On the other hand, White's standard errors may be negatively biased in small samples. To test for possible heteroskedasticity, I use the Breusch–Pagan test. The null hypothesis of constant variance is affirmed in both models included in Table 1.

TABLE 1 **Determinants of trade union preferences over active labor market policies**

	Job insecurity	Organizational resources
Change in Deindustrialization	0.179 (0.030)**	
Unemployment	−0.019 (0.008)*	
Employment Protection	−0.029 (0.015)[†]	
Union Participation		−0.014 (0.057)
Union Density		0.002 (0.001)*
Constant	0.269 (0.062)**	0.098 (0.067)
Observations	22	22
Adjusted R^2	0.5253	0.0344

Note: [†]$p < .10$, *$p < .05$, **$p < .01$. Robust standard errors are given in parentheses.

FIGURE 1 **Deindustrialization and union member preferences towards ALMP**

Effect of Union Membership on Preferences Towards ALMP

difference in the rate of deindustrialization between these two country years is 2.06 percent with a corresponding difference in the effect of union membership on support for active measures of 4.2 percent. A graph of the coefficients for union membership and the rates of deindustrialization are plotted in Figure 1 for greater clarification.

CONCLUSION AND DISCUSSION

The results do not provide conclusive support for either traditional theories, which predict that trade union members will be advocates of redistributive social policy, or the insider-outsider theory, which expects trade union members to oppose integrating social policies as a result of their advantaged labor market position. Traditional theories hypothesizing strong trade union support for social policies do garner the most support. Nevertheless, the results for unemployment suggest caution in assuming unconditional trade union member support.

A possible explanation that would resolve this inconsistency lies in the effect of employment protection and deindustrialization on the long-term employment chances of trade union members. Although deindustrialization and low employment protection increase job insecurity in the short term, these conditions may not reduce trade union member support until the aggregate supply of jobs in the labor market decreases and therefore makes finding new employment more difficult as well.

To conclude, the specific nature of the data employed in this research note mirrors much of the data in international public opinion research where the number of level one units is high, but the number of level two units, such as countries, regions, electoral periods, or interviewers, is small. Estimating the contextual parameters with MLE in such cases will underestimate the standard errors. Bayesian approaches do remain a viable alternative, although estimating the necessary priors is an arduous task. In addition to addressing the issue of low number of level two units head on, the EDV approach also offers a modeling technique that provides a relatively straight-forward strategy for assessing the relationship between level one and level two predictors that could also potentially be used to fit a subsequent one-step multilevel model.

ACKNOWLEDGEMENT

REFERENCES

Bean, C., & Papadakis, E. (1998). A comparison of mass attitudes towards the welfare state in different institutional regimes, 1985–1990. *International Journal of Public Opinion Research, 10,* 211–236.

Beck, N. (2005). Multilevel analyses of comparative data: A comment. *Political Analysis, 13,* 457–458.

Bowers, J., & Drake, K. W. (2005). EDA for HLM: Visualization when probabilistic inference fails. *Political Analysis, 13,* 301–326.

Duch, R. M., & Stevenson, R. (2005). Context and the economic vote: A multilevel analysis. *Political Analysis, 13*, 387–409.

Franzese, R. J. (2005). Empirical strategies for various manifestations of multilevel data. *Political Analysis, 13*, 430–446.

Gelman, A. (2006). Prior distributions for variance parameters in hierarchical models. *Bayesian Analysis, 3*, 515–533.

Gelman, A., & Hill, J. (2007). *Data Analysis Using Regression and Multilevel/hierarchical Models*. Cambridge: Cambridge University Press.

Hox, J. J., & Maas, C. J. M. (2001). The accuracy of multilevel structural equation modeling with pseudobalanced groups and small samples. *Structural Equation Modeling, 8*, 157–174.

Hu, L. T., Bentler, P. M., & Kano, Y. (1992). Can test statistics in covariance structure analysis be trusted? *Psychol Bull, 112*, 351–362.

Huber, J. D., Kernell, G., & Leoni, E. L. (2005). Institutional context, cognitive resources and party attachments across democracies. *Political Analysis, 13*, 365–386.

Hufbauer, G. C., & Elliott, K. A. (1994). *Measuring the Costs of Protection in the United States*. Peterson Institute.

Iversen, T., & Cusack, T. (2000). The causes of welfare state expansion: deindustrialization or globalization? *World Politics, 52*, 313–349.

Jusko, K. L., & Shively, W. P. (2005). Applying a two-step strategy to the analysis of cross-national public opinion data. *Political Analysis, 13*, 327–344.

Lee, S. Y., & Song, X. Y. (2004). Evaluation of the Bayesian and maximum likelihood approaches in analyzing structural equation models with small sample sizes. *Multivariate Behavioral Research, 39*, 653–686.

Lewis, J. B., & Linzer, D. A. (2005). Estimating regression models in which the dependent variable is based on estimates. *Political Analysis, 13*, 345–364.

Lindbeck, A., & Snower, D. J. (1988). *The Insider-Outsider Theory of Employment and Unemployment*. Cambridge, Mass: MIT Press.

Lindbeck, A., & Snower, D. J. (1990). *Interactions Between the Efficiency Wage and Insider-Outsider Theories*. Sweden Institute for International Economic Studies: Stockholm.

Lindbeck, A., & Snower, D. J. (2002). *The Insider-Outsider Theory: A Survey*. Unpublished manuscript, Bonn.

Luke, D. A. (2004). *Multilevel Modeling*. Thousand Oaks: Sage University Press.

Maas, C., & Hox, J. (2004). Robustness issues in multilevel regression analysis. *Statistica Neerlandica, 58*, 127–137.

Montenegro, A. (2001). On sample size and precision in ordinary least squares. *Journal of Applied Statistics, 28*, 603–605.

Polacheck, S., & McCutcheon, E. (1983). Union effects on employment stability: A comparison of panel versus cross-sectional data. *Journal of Labor Research, 4*.

Raudenbush, S. W. (1988). Educational applications of hierarchical linear models: A review. *Journal of Educational Statistics, 13*, 85–116.

Raudenbush, S. W., & Bryk, A. S. (2001). *Hierarchical Linear Models: Applications and Data Analysis Methods*. London: Sage.

Snijders, T. A. B., & Bosker, R. J. (1999). *Multilevel Analysis. An Introduction to Basic and Advanced Multilevel Modeling*. London: Sage.

Steenbergen, M. R., & Jones, B. S. (2002). Modeling multilevel data structures. *American Journal of Political Science, 46*, 218–237.

Traxler, F., Blaschke, S., & Kittel, B. (2001). *National Labour Relations in Internationalized Markets: A Comparative Study of Institutions, Change and Performance*. Oxford: Oxford University Press.

BIOGRAPHICAL NOTE

Moira Nelson (PhD, North Carolina—Chapel Hill) is a post-doctoral fellow at the Bremen International Graduate School of Social Sciences. Her research interests include the origins and development of social policy as well as the political economy of skill formation.

Address correspondence to Moira Nelson, Lenbachstrasse 7a, D-10245 Berlin, Germany, E-mail: moira@email.unc.edu