**Appendix : Expanded References**

(The complete reference list appears at the end of these appendices.)

**A Subject-Organized Reference List for Applied Spatial-Modeling in Political Science:**

On Policy-Innovation Diffusion Among US States: Crain 1966; Walker 1969, 1973; Gray 1973; Knoke 1982; Caldiera 1985; Lutz 1987; Berry & Berry 1990; Case et al. 1993; Berry 1994; Rogers 1995; Mintrom 1997ab; Brueckner 1998; Mintrom & Vergari 1998; Mossberger 1999; Berry & Berry 1999; Godwin & Schroedel 2000; Balla 2001; Mooney 2001; Wejnert 2002; Coughlin et al. 2003; Bailey & Rom 2004; Boehmke & Witmer 2004; Daley & Garand 2004; Grossback et al. 2004; Mencken 2004; Berry & Baybeck 2005; Garrett et al. 2005; Costa-Font & Ons-Novell 2006; Karch 2006; Rincke 2006; Shipan & Volden 2006; Volden 2006; Werck et al. 2006; Woods 2006; Volden et al. 2007.

On Inter/Cross-National Policy-Innovation Diffusion: Schneider & Ingram 1988; Rose 1993; Bennett 1997; Dolowitz & Marsh 2000; True & Mintrom 2001; Tews et al. 2003; Jensen 2004; Meseguer 2004, 2005; Brooks 2005, 2007; Gilardi 2005; Gilardi et al. 2005; Murillo & Schrank 2005; Weyland 2005; Braun & Gilardi 2006; Linos 2006; Parys 2006; Ermini & Santolini 2007; Moscone et al. 2007.

On Institutional/Regime Diffusion: Dahl’s 1971 classic *Polyarchy*, e.g., implicitly references international interdependence among the eight causes of democracy he lists; Starr’s 1991 “Democratic Dominoes” and Huntington’s 1991 *Third Wave* accord it a central role; Beissinger 2007 and Bunce & Wolchik 2006, 2007, *inter alia*, emphasize it in the context of post-communist democratic transitions in Eastern Europe, and Hagopian & Mainwaring 2005 among others in the Latin American context; finally, O’Loughlin et al. 1998, Brinks & Coppedge 2006, and Gleditsch & Ward 2006, 2007 estimated empirically the extent, paths, and/or patterns of international diffusion of democracy. Kelejian et al. 2007 give institutional diffusion general theoretical and empirical treatment.

Empirical Attention to the Inherent Interdependence of International Relations… is most extensive in the work of Ward, Gleditsch, and colleagues—Shin & Ward 1999; Gleditsch & Ward 2000; Gleditsch 2002; Ward & Gleditsch 2002; Hoff & Ward 2004; Gartzke & Gleditsch 2006; Salehyan & Gleditsch 2006; Gleditsch 2007—and, in a different way, in Signorino and colleagues— Signorino 1999, 2002, 2003; Signorino & Yilmaz 2003; Signorino & Tarar 2006.

On Diffusion in Comparative & International Political Economy, and Globalization: Simmons & Elkins 2004 and Simmons et al. 2006, e.g., stress cross-national diffusion as the main force behind recent economic liberalizations, as do Eising 2002; Brune et al. 2004; Brooks 2005, 2007; Jordana & Levi-Faur 2005; Way 2005; Lazer 2006; Prakash & Potoski 2006; Brune & Guisinger 2007; and many others. Empirical work on globalization-induced interdependencies are far too numerous even to cite. Just a list of recent works emphasizing those that recognize explicitly that interdependence implies effects of some units outcomes on others—and still a small subset at that—would include Genschel 2002; Guler et al. 2002; Hays 2003; Franzese & Hays 2003, 2004ab, 2005ab, 2006abc, 2007abcd, 2008abcd, 2009abc; Badinger et al. 2004; Basinger & Hallerberg 2004; Hays et al. 2005; Heichel et al. 2005; Henisz et al. 2005; Holzinger & Knill 2005; Knill 2005; Polillo & Guillén 2005; Elkins et al. 2006; Jahn 2006; Lee & Strang 2006; Manger 2006; Swank 2006; Baturo & Grey 2007; Cao 2007; Cao et al. 2007; Coughlin et al. 2007; Garretsen & Peeters 2007; Mosley & Uno 2007; Mukherjee & Singer 2007.

On Interdependence of Legislators’ Votes (Modeled Spatially): See, for example, Lacombe & Shaughnessy 2005.

On Interdependence of Citizens’ Votes (Modeled Spatially): See, for example, Huckfeldt & Sprague 1991; O’Laughlin et al. 1994; Pattie & Johnston 2000; Beck et al. 2003; Calvo & Escolar 2003; Kim et al. 2003; Schofield et al. 2003; Lacombe & Shaughnessy 2007.

On Interdependence of Election Outcomes (Modeled Spatially): See, for example, Shin & Agnew 2002, 2007; Hiskey & Canache 2005; Wing & Walker 2006; Kayser 2007.

On Interdependence of Candidate Qualities, Contributions, or Strategies: See, for example, Goldenberg et al. 1986; Mizruchi 1989; Krasno et al. 1994; Cho 2003; Gimpel et al. 2006.

For Spatial Models of the Interdependence of the probabilities and outcomes of coups: e.g., Li & Thompson 1975; Of Riots: e.g., Govea & West 1981; Of Civil Wars: e.g., Murdoch & Sandler 2004, Buhaug & Rød 2006; Of Revolutions: e.g., Brinks & Coppedge 2006.

On Interdependence in Treaty Signing: see, e.g., Murdoch et al. 2003.

On Interdependence in Terrorist Origins and Targets: see, e.g., Brathwaite & Li 2008.

Contextual/Neighborhood Effects in Micro-Behavioral Studies: Huckfeldt & Sprague (1993) review the large literature on contextual/neighborhood effects in political behavior; as do Sampson et al. (2002) and Dietz (2002) for sociology. Recent analyses that stress interdependence include Straits 1990; O’Loughlin et al. 1994; Knack & Kropf 1998; Liu et al. 1998; Braybeck & Huckfeldt 2002ab; Beck et al. 2002; McClurg 2003; Huckfeldt et al. 2005; Cho & Gimpel 2007; Cho & Rudolph 2007.

On Interdependent Social-Movements: see, e.g., McAdam & Rucht 1993; Conell & Cohn 1995; Giugni 1998; Strang & Soule 1998; Biggs 2003; Browning et al. 2004; Andrews & Biggs 2006; Holmes 2006; Swaroop & Morenoff 2006.

On Interdependence in Violence and Crime: see, e.g., Grattet et al. 1998; Myers 2000; Baller et al. 2001; Morenoff et al. 2001; Villareal 2002; Baker & Faulkner 2003; Oberwittler 2004; Bhati 2005ab; Mears & Bhati 2006.

On Interdependence in (Microeconomic) Utilities: see, e.g., Akerloff 1997; Postlewaite 1998; Glaeser & Scheinkman 2000; Manski 2000; Brock & Durlauf 2001; Durlauf 2001; Glaeser et al. 2003; Yang & Allenby 2003; Sobel 2005; Ioannides 2006; Soetevent 2006.

On Interdependence in Macroeconomic Performance: see, e.g., Fingleton 2003; Novo 2003; Kosfeld & Lauridsen 2004; Maza & Villaverde 2004; Kelejian et al. 2006; Mencken et al. 2006.

On Interdependence in Technology, Marketing, and other Firm Strategies: see, e.g., ; Abramson & Rosenkopf 1993; Geroski 2000; Strang & Macy 2001; Holloway 2002; Bradlow 2005; Autant-Berard 2006; Mizruchi et al. 2006.

On Interdependence in Fertility, Birthweight, Child development, or Child Poverty: see, e.g., Tolnay 1995, Montgomery & Casterline 1996; Morenoff 2003; Sampson et al. 1999; Voss et al. 2006.

On Interdependence in ordainment of women: Chaves 1996; in right-wing extremism: Rydgren 2005, in marriage: Yabiku 2006, in (sub)national identity: Lin et al. 2006; in obesity: Christakis & Fowler 2007; and in research faculty: Weinstein 2007.

**Appendix : The Econometric Problem and Estimation Strategies; The Spatial-Error, The Conditional-Spatial, and the Time-Lagged Spatial-Lag Probit-Model**

Methods for properly estimating and analyzing models of interdependent qualitative or limited dependent variables (henceforth: *QualDep* models) have received significant attention in the spatial-econometric literature recently. Most of this research considers the spatial-probit model with inter­dependence in the latent-variable, i.e., in the unobserved argument to the probit-modeled probability of a binary outcome.[[1]](#footnote-1) Models of spatial sample-selection (spatial Tobit or Heckit: McMillen 1995, Smith & LeSage 2004, Flores-Lagunes & Schnier 2006), spatial multinomial-probit (McMillen 1995, Bolduc et al. 1997), and spatial discrete-duration (Phaneuf & Palmquist 2003), all of which closely resemble the spatial probit, have also been suggested, as have models of interdependent survival (Hays & Kachi 2009) or of survival with spatial “frailty” (i.e., error components: Banerjee et al. 2004, Darmofal 2007) and of spatial counts (e.g., Bhati 2005, Franzese & Hays 2009a), including a zero-inflated-count model (e.g., Rathbun & Fei 2006). Spatial probit is far the most-common *S-QualDep* model in applied research, however.[[2]](#footnote-2)

Several estimation strategies have been suggested for the spatial-probit model. McMillen (1992) suggested an EM algorithm, which first rendered the spatial-probit’s non-additively-separable log-likelihood (see below) estimable, but the strategy also did not provide standard-errors for the crucial spatial-dependence parameter and required arbitrary parameterization of the heteroscedasticity that dependence induces (see below). McMillen (1995) and Bolduc et al. (1997) applied simulated-likelihood strategies to estimate their spatial-multinomial-probit models, and Beron et al. (2003) and Beron & Vijverberg (2004) advanced a recursive-importance-sampling (RIS) estimator in that line. LeSage (1999, 2000) introduced a Bayesian strategy of Markov-Chain-Monte-Carlo (MCMC) by Metropolis-Hastings-within-Gibbs sampling. (LeSage & Pace 2009 corrects a crucial error in the earlier formulations of the estimator.) Fleming (2004) reviews these two families and simpler approximation strategies allowing spatial interdependence in linear or nonlinear probability models estimable by nonlinear least-squares,[[3]](#footnote-3) generalized linear-models, or generalized linear-mixed-models. Pinkse & Slade’s (1998) two-step GMM estimator for spatial-error probit has seen some use in the literature, as has McMillen’s (2005) GMM for linearized spatial-lag logit or probit and Pinkse et al.’s (2006) one-step (continuously updating) GMM for spatial-probit, but the first is inconsistent for the spatial-lag model and all three, being instrumental-variable estimations of linear approximations around zero interdependence, work well only in large samples with weak interdependence. The RIS and Bayesian strategies do not have these limitations[[4]](#footnote-4) and (so) have dominated recent applications.

Section II formally describes the spatial-, temporal-, and spatiotemporal-probit models, and Section III the MSL-RIS strategy for estimating it. This appendix gives the spatial-error probit-model, explains the Bayesian-MCMC estimation strategy for spatial-*QualDep* models, and adds some discussions of technicalities that arose in the corresponding sections of the text.

The spatial-error version of the probit model is slightly simpler, taking the form:

  ,

with , and having the marginal probabilities:

  ,

where **x***i* is the *i*th row of **X**. As with spatial-lag probit, these *ui* are heteroskedastic and interdependent, so the probability derives from the *i*th marginal distribution of a multivariate cumulative-normal with means **0** and variance-covariance , so spatial-error probit models entail most of the same estimation and interpretation complications as spatial-lag models. (Mixed spatial-lag/spatial-error models are also possible, but they have received little attention.) In the spatial-error model, because the interdependence operates only through  and not all of , the position of the *i*th observation on the sigmoidal probit-function depends on the entire vector  but only on that observation’s independent-variable values, .

Special circumstances might allow standard-probit estimation of spatial-lag models, but we view these as highly unlikely. E.g., Anselin (2006) notes that, in a conditional version of

  ,

 could be estimated by , the spatially weighted average of actual outcomes in units *j*, without introducing endogeneity problems only under stringent conditions that ensure other units’ observations *j* are not jointly determined with those of *i*, and that “coding methods ensure that the sample does not contain these neighbors” (Anselin 2006). This means that any units *j* from which diffusion to any *i* in the sample is non-negligible (at any order spatial-lag) must be excluded from the sample but used in constructing the  spatial lag for the retained observations *i*. Alternatively, all *i*’s neighboring *j* according to  must be exogenous to *i* for all *i* in the sample; i.e., feedback must be directional and orderable from *j*’s to *i*’s only, severing feedback from *i* back to itself. Relatedly, while some substantive-theoretical contexts might suggest that interdependence propagates through the actual outcome rather than the latent variable, a simultaneous such model is not generally possible because, indirectly via feedback,  would generate  but also, directly,  is generated by .[[5]](#footnote-5) Conditions like those described above allow direct inclusion of  because they sever such indirect generation of  by . These limitations are usually prohibitive practically, though contexts where such directional ordering and such omissions of certain *j* may be defensible are imaginable. Swank (2006, 2007), e.g., argues that U.S. tax policies exclusively lead others’ tax policies—the U.S. is *the unmoved mover*, so to speak—and he excludes all U.S. data from the left-hand side of his empirical models, reserving those U.S. data solely for the role of spatial lag. If valid, arguments and sample-exclusions such as these would allow standard-probit estimation.

The text focuses on the unconditional, simultaneous spatial-lag models. It ignores the spatial-error and conditional spatial-lag models because they are typically less plausibly (spatial-error) or implausibly (conditional-spatial) applicable and because they raise lesser (spatial-error) or no (conditional-spatial) estimation complications. It also does not discuss the time-lagged spatial-lag model because the conditions described there for the practical adequacy of the strategy seem restrictive for many social-science applications and because, even if otherwise adequate, the strategy evades little of the estimation complications, which arise even for merely time-lagged binary-dependent-variables (as just discussed). The text also ignores tests of the adequacy of time-lagged spatial-lag models or specification tests of spatial-*lag* vs. spatial-*error* vs. non-spatial models here, though these tests are important to consider, especially given the complexity and computational intensity of valid estimation strategies for full, simultaneous spatially, temporally, or spatiotemporally autoregressive probit.[[6]](#footnote-6) For starts on these discussions, we refer the reader to Pinkse & Slade (1998), Pinkse (1999), Kelejian & Prucha (2001), and, for a relatively recent review, Anselin (2006). The text focuses on unconditional, simultaneous spatial-probit estimation by MSL-RIS and its comparison to standard-probit estimation with the endogenous spatial-lag,, included as a regressor, which latter is current standard-practice in empirical work where interdependence of binary outcomes is addressed.

**Appendix : The Bayesian-MCMC Estimator for Simultaneous Spatial-Lag Probit**

LeSage (1999, 2000) suggests using Bayesian Markov-Chain-Monte-Carlo (MCMC) methods to surmount the estimation complications introduced by the *n*-dimensional cumulative-normal in the spatial-probit likelihood (posterior). The basic idea of Monte Carlo (simulation) methods is simple:[[7]](#footnote-7) if one can characterize the joint distribution (likelihood or posterior) of the quantities of interest (parameters), then one can simply sample (take random draws) from that distribution and calculate the desired statistics in those samples. With sufficient draws, the sample statistics can approximate the population parameters they aim to estimate arbitrarily closely.[[8]](#footnote-8) In basic Monte-Carlo simulation, the draws are independent and the target distribution is specified directly. In MCMC, each draw is dependent on the previous one in a manner that generates samples with properties mirroring those of the joint population using just the conditional distribution of each parameter. This is useful where the joint distribution is not expressible directly or, as with spatial probit, where its complexity makes direct sampling from the joint distribution prohibitively difficult and/or time-consuming.

We can describe Gibbs sampling, the simplest and most-common of the MCMC family, thusly: Given distributions for each parameter conditional on the other parameters, one can cycle through draws from those conditional distributions, eventually reaching a *convergent* state past which point all subsequent draws will be from the targeted posterior joint-distribution. To elaborate: first express the distribution for each parameter conditional on all the others, then choose (arbitrary) starting values for those parameters and draw a new value for the first parameter conditional on the others’ starting values. Then, conditional on this new draw of the first parameter and starting values for the rest, draw a new value for the second parameter from its conditional distribution. Continue thusly until all parameters have their first set of drawn values, then return to the first parameter and draw its second simulated value conditional on the others’ first draws. Cycle thusly for some large number of iterations, and, under rather general conditions, the limiting (asymptotic) distribution of this set of parameter draws is the desired joint posterior-distribution. Thus, after having gathered some very large set of parameter-vector values by this process, discard some large initial set of draws (the *burn-in*) and base inferences on sample statistics from the remaining set of parameter vectors. A typical burn-in might be 1000 draws, and inferences might be based on the next 5000 or 10,000. Also, since each draw is conditional on the previous drawn values, autocorrelation typically remains, so “thinning” the post-burn-in sample by using every, say, third or fifth draw may boost efficiency.

The drawbacks of MCMC may be obvious from what we have said and declined to say. First, no universal tests exist to verify that *convergence* has occurred, so a burn-in may appear sufficient in that the next 5000 drawn parameter-vectors seem to follow some circumscribed bounds and behavior of some unknown target distribution (i.e., the sampler may seem to have *settled down*) only to have the 5001st leap into a new range and proceed toward convergence elsewhere. Second, despite their Markov-Chain origins, adjacent draws are asymptotically serially uncorrelated, but this *asymptopia* may not arrive within practical limits, and thinning may be insufficient help or too computationally costly. Third, the starting values are likewise asymptotically irrelevant, assuming the supplied set of conditional distributions properly could come from a valid joint distribution, but, as the previous two caveats imply, starting values may matter short of convergence, arrival at which is not verifiable.[[9]](#footnote-9) These issues concern careful researchers, and many diagnostics and tests for non-convergence, serial correlation, or starting-value sensitivity, and numerous strategies for ameliorating them, exist (all imperfect, but useful still). However, the concerns do not outweigh the remarkably flexible utility of the Gibbs sampler, either in general or specifically in its application to spatial-probit estimation.

All but one of the conditional distributions for the spatial-probit-model parameters (given below) are standard, so the Gibbs sampler is useful for them. The crucial spatial-lag-coefficient, , has the lone non-standard conditional-distribution; for it, Metropolis-Hastings sampling is used. Metropolis-Hastings differs from Gibbs sampling in the former’s *seeding* or *jump* distribution from which values are drawn and then accepted or rejected as the next sampled parameters, depending on how they compare to a suitably transformed expression of the target distribution.[[10]](#footnote-10) The Bayesian spatial-probit estimator (LeSage1999, 2000) uses Metropolis-Hastings for  within the Gibbs sampler procedure for the other parameters. Of course, this step adds some to the estimator’s computational intensity.

With this brief introduction to Bayesian MCMC estimation by Gibbs and Metropolis-Hastings sampling, we now introduce their application to the spatial-probit model. We follow LeSage (2000) to state the likelihood in terms of the latent outcome, —an additional conditional distribution will later apply the measurement equation to convert unobserved  to observed [[11]](#footnote-11)—for the spatial-lag model as:

  ,

where . (The likelihood for spatial-error probit model is the same but with , where *ρ* here is ’s *λ*.) Diffuse priors yield joint posterior-density:

  .

One can now derive the conditional posterior densities for  for the sampler. First:

  .

Notice that conditioning on  allows  to be subsumed into the constant of proportionality and that implies , a standard distribution facilitating the Gibbs sampler. Next,

  ,

where, in spatial lag,  and , and, in spatial error,  and . The conditional multivariate-normality of  allows the Gibbs sampler for it also, but *ρ* has non-standard conditional distribution, requiring Metropolis-Hastings sampling:

  ,

with  defined as given above for the spatial-error and the spatial-lag models.[[12]](#footnote-12)

Finally, LeSage (1999, 2000) erroneously added the conditional distribution, namely a truncated normal, that translates  to , as a *univariate* truncated normal:

  ,

where  is the predicted value of  (the *i*th element of  for spatial-lag or of  for spatial-error models) and the variance of  is  with  the *i*th element of . In addition to producing inconsistent estimates, this mistake, which earlier versions of this paper followed, gave the false impression that the Bayesian MCMC estimation-strategy was much simpler and faster than the classical simulated-likelihood (RIS) strategy. Lesage & Pace (2009) corrects the mistake, replacing *univariate* with the properly multivariate truncated normal distribution:

 .[[13]](#footnote-13)

That is, the Bayesian MCMC estimator must also confront the multidimensional-normal integration that is the major complication raised by (inter)dependence in probit models.

Since the choice rule is , the cutpoint that gives *p*(*yi*=1), *μi*, depends on all the *yj*\*. The stochastic-component draws must therefore come from the nonspherical truncated multivariate normal (TMVN) with variance-covariance **Σ** and bounds (‑∞,*μi*) for *yi*=1 and (*μi*,+∞) for *yi*=0. Following Geweke (1991) on drawing from a TMVN, the correct Bayesian MCMC estimator for the spatial-probit model adds another *m* step Gibbs sampler within the overall sampler, drawing each cutpoint, *zi*, conditional on all the *z~i*, from the conditional distributions for this *n*-variate TMVN. This parallels closely the computation intensity of the classical RIS strategy, which must also simulate the integration of this same cumulative, nonspherical TMVN (and uses the related Geweke-Hajivassiliou-Keane (GHK) simulator to do so).[[14]](#footnote-14)

With all the conditional distributions, we can implement MCMC to estimate the model thus:[[15]](#footnote-15)

1. Use expression to draw  using starting values .
2. Use ,, and expression to draw .
3. Use , , and expression to draw  by Metropolis-Hastings sampling.
4. Subloop: Use an *m*-step Gibbs sampler to sample the outcomes, **z**, using the conditional distributions from the multivariate censoring distribution and , , and .
5. Return to step 1 incrementing the subscript counters by one.

After a sufficient burn-in—our simulation and application experiences so far suggest at least 1000 is advisable—the distributions of *σ*, **β**, and *ρ* will have reached convergence and subsequent draws on the parameters may be used to give their estimates (as means or medians of some large number of draws) and estimates of their certainty (as standard deviations or percentile ranges).[[16]](#footnote-16)

Notice that exactly the same equations and procedures apply for the temporal-autorgressive probit model, substituting *ϕ***A** for *ρ***W** (as separately derived by Beck et al. 2001).

**Appendix : Effect Calculations–The Brute-Force Simulation Method (Hays 2009)**

A simpler expedient evades integration of the *n*-dimensional multivariate-normal by drawing reduced-form disturbances  and coefficients from the multivariate posterior or sampling distribution for .[[17]](#footnote-17) Calculate  and using  and for some fixed **X**1 and **X**0, then simply apply the standard measurement equation to convert those to vectors of ones and zeros,  and . For a large number of draws from the distribution of reduced-form disturbances and a given set of coefficients, the averages of  and will be  and , and  will be the desired vector of estimated effects, and the variance-covariance of those differences, produced by repeated draws of  will be the variance-covariance of those estimated effects.[[18]](#footnote-18) In other words, we can use the model to generate counterfactual values of the dependent variable for a given set of **X** and **W**, and estimate the conditional probabilities of interest. This technique can be used to generate probabilities and effects conditional on observed outcomes in other units, for example  and . In many social-science applications, these kinds of counterfactuals are particularly interesting substantively. The probabilities/relative frequencies are ratios of quadrant counts from a 2-dimensional graph where the axes represent the negative of the reduced form cutpoints for units *i* and *j*, the *i*th and *j*th elements of the vector . Once we have the parameter estimates and a specific counterfactual, the computation costs of proceeding thusly are relatively low (see Figure 1). A single draw from the reduced-form disturbances for units *i* and *j*, the *i*th and *j*th elements of the vector , identifies an x-y coordinate (a point) located in one of the quadrants. If this point is in quadrant I, for example, the reduced-form disturbances for both *i* and *j* are above the negative of their respective reduced-form cutpoints and . Conditional relative frequencies, which are ratios of quadrant counts, provide estimates for the conditional probabilities of interest. Specifically, the probabilities are estimated by:



.

Examples of this approach to counterfactual analysis are in the empirical application of section VI.

**Figure 1 – Estimating Spatial Effects via Simulation**



In interpretation, as in estimation, the issues raised by temporal auto-dependence in binary-choice models are analogous to those of spatial interdependence. The procedure we propose for interpreting spatiotemporal effects is likewise analogous. To calculate the spatiotemporal responses across *N* units, say 50, over *T* periods, say 20, to a hypothetical shock—say that the first unit gets a draw in the first period below its cutpoint (and so has outcome *y11*=1), as opposed to above (*y11*=0)—start with a vector of *N*×*T*=50×20=1000 standard-normal draws. Transform those white-noise residuals into the spatiotemporally dependent shocks by premultiplying that 1000x1 vector by the spatiotemporal multiplier, (**I**-*ρ***W**-*φ***L**)-1, and likewise transform the *NT*×1 vector **Xβ** into the spatiotemporally dependent cutpoints for every unit in every period premultiplying by (**I**-*ρ***W**-*φ***L**)-1. This gives one simulated case of the *N* paths over *T* periods. Repeat the procedure say 1000 times to generate 1000 sets of *N* paths over *T* periods, and separate those 1000 sets into those in which the hypothetical in question holds and those in which it does not—in this case, into the set of responses in which unit 1 gets draws above its cutpoint in period 1, that’s set 1, and set 2 of draws for which unit 1 got a draw below its cutpoint in period 1. The responses in all countries (including unit 1) to this hypothetical are then the differences between the path of that unit’s share of above-its-cutpoint draws when unit 1 got an above-cutpoint first-period draw and the path of that unit’s share of above-its-cutpoint draws when unit 1 got a below-cutpoint first-period draw.[[19]](#footnote-19) To estimate the uncertainties of these estimated response paths by parametric simulation, the whole procedure is repeated many times drawing a vector of parameters from their estimated joint distribution, which is asymptotic-normal with asymptotic mean of the estimated parameter vector and asymptotic variance-covariance of the estimated parameter-vector estimates’ variance-covariance by the properties of MSL.

**Brute-Force Method: Simulations**

Ultimately, we are not interested in the parameter estimates *per se*, but rather in the effects that they imply. We start with the first difference − for the immediate (first/same-period) spatial effect (i.e., the effect of a time *t* change in the outcome for unit *j* on the time *t* probability that we observe a particular outcome for unit *i*). We report the quadrant counts from our Monte Carlos in Table 2. The first row of results provides the actual distribution of simulation-outcomes across the four quadrants. Given the true parameters and weights matrices from Experiments #1 and #2, the reduced-form disturbances of units *i* & *j* lie in Quadrant I, i.e., above the negative of their respective reduced-form cutpoints, implying *y*=1 in both *i* & *j*, 15.9% of the time. The analogous true percentages for Quadrants II, III, and IV are 18.4%, 42.4%, and 23.4%. The next row reports the mean estimated-count produced by our estimator across 10,000 samples. Comparing rows one and two speaks to the bias or prediction-accuracy of our estimation strategy. The third row provides the actual standard deviation in the estimated count, which relates to the efficiency of our strategy. Overall, our strategy for estimating conditional outcomes seems to perform well.

Table 2: Simulation Results (100 Trials)

|  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- |
|  |  | **Quadrant I** | **Quadrant II** | **Quadrant III** | **Quadrant IV** |
| **Experiment #1:** n=50, t=5, φ=0.5, ρ=0.25,  |  |  |  |  |  |  |  |  |
|  | ***True Count (per 10,000)*** | ***1589*** | ***1839*** | ***4243*** | ***2335*** |
|  | Mean Estimated Count | 1575 | 1811 | 4182 | 2431 |
|  | Std Dev of Estimates | 381 | 113 | 575 | 215 |
| **Experiment #2:** n=50, t=20, φ=0.5, ρ=0.25, |  |  |  |  |  |  |  |  |
|  | Mean Estimated Count | 1514 | 1787 | 4363 | 2335 |
|  | Std Dev of Estimates  | 198 | 74 | 309 | 105 |
| **Experiment #3:** n=150, t=5, φ=0.5, ρ=0.25, |  |  |  |  |
|  | ***True Count (per 10,000)*** | ***1474*** | ***6599*** | ***1660*** | ***267*** |
|  | Mean Estimated Count | 1318 | 6994 | 1481 | 206 |
|  | Std Dev of Estimates | 207 | 319 | 226 | 51 |
| **Experiment #4:** n=150, t=20, φ=0.5, ρ=0.25, |  |  |  |  |
|  | Mean Estimated Count | 1219 | 7123 | 1487 | 169 |
|  | Std Dev of Estimates | 107 | 195 | 131 | 29 |

**Brute-Force Method: Application**

Historically, the border between Guinea-Bissau and Senegal has been a breeding ground for instability. Guinea-Bissau served as training ground for *Mouvement des Forces Democratiques de Casamance* (MFDC) fighters and a conduit to funnel arms into the decade’s (1990s) long Casamance conflict. Most view the 1998-99 civil war in Guinea-Bissau as an outgrowth of these same tensions, with Senegalese forces ultimately fighting on both sides of the conflict (Humphreys & Mohamed 2005).[[20]](#footnote-20) We explore the extent to which the conflict in Senegal affected the onset of civil war in Guinea-Bissau, or, stated differently to highlight the conditional probability of the counterfactual: in the absence of the Casamance conflict, how likely was conflict in Guinea-Bissau?

In terms of the model, the question becomes: given that Senegal’s reduced-form disturbance is above/below (negative) its reduced-form cutpoint, what is the probability that Guinea-Bissau’s reduced-form disturbance will be above/below (negative) its reduced-form cutpoint? To answer this question, we sample the reduced-form disturbances; specifically, we draw 10,000  errors for each state-year in the sample, yielding a  matrix of *i.i.d.* standard-normal disturbances. Then we multiply this disturbance matrix by the spatial multiplier, giving . Since the counterfactual question involves the participation of Senegal and Guinea-Bissau specifically, we take just that bivariate slice of the resulting 1434-dimensional multivariate distribution, although the procedure being described here produces the entire vector of all states’ responses to the hypothetical.

The vector of reduced-form cutpoints is calculated as . A country experiences civil war if its reduced-form disturbance is greater than negative its reduced-form cutpoint. Figure 2 plots the bivariate pair of these simulated reduced-form disturbances corresponding to Senegal and Guinea-Bissau. Given their covariates, the reduced-form cutpoints are ‑.3157 for Senegal and ‑1.207 for Guinea-Bissau, so these countries experience conflict when their reduced-form disturbances exceed .3157 and 1.207 (as the lines indicate). In these simulations, Guinea-Bissau has a civil war 11.07% of the time when Senegal is peaceful—i.e., 11.07% of points left of Senegal’s cutpoint lie above Guinea-Bissau’s cutpoint—and 13.93% of the time when Senegal is at war—13.93% of points right of Senegal’s line lie above Guinea-Bissau’s. Thus, the model estimates suggest that Senegal’s conflict increased the risk of war in Guinea-Bissau by 2.86%. To calculate our uncertainty about these effects estimates, we sample parameter-estimates from their estimated sampling distribution; doing so reveals the effect-sizes at the 5th and 95th percentiles are 0.80% and 4.92%.

**Figure 2 – Scatterplot of Reduced-Form Disturbances & Cutpoints (Guinea-Bissau & Senegal)**



**Appendix : Effect Calculations–further discussion of delta method v. parametric simulation**

Since the latent-variable model is a spatial linear-regression, estimated effects in terms of  and their certainties would derive exactly as in that case, which we have discussed elsewhere (Franzese & Hays 2004, 2007, 2008ab, Hays et al. 2010). We review here, beginning with the cross-sectional effects, which are identical to the first/same-period effects in a panel, we have:

  .[[21]](#footnote-21)

Thus, denoting the *i*th column of  as  and their estimates as  and , the estimated effect of explanatory variable *k* in unit *i*, , on the outcomes in all units, *i* and all *j*, is  which is simply, . The standard-error calculation, using the delta method approximation, is

  ,

The vector  is the *i*th column of . Since  is an inverse matrix, the derivative in equation is .

The marginal dynamic response paths (i.e., the period-by-period increments) and their delta-method approximate standard-errors can be derived by analogously differencing:

  ,

and the long-run-steady-state effects of a permanent shock derive from differencing:

  .

The cumulative response-paths, finally, are found by recursive substitution using

  ,

where the superscript *cf* refers to the *counterfactual* series of shocks to which the *N* units’ responses are tracked. The delta-method calculations can be unwieldy, so standard errors can also come from *parametric simulation*. Draw many vectors of parameters from their estimated means and variance-covariance (MSL parameter-estimates are asymptotically normal), and calculate the estimated effects or response paths to the desired counterfactual for each draw, using the average and standard deviation across the draws as the estimate and its standard error.[[22]](#footnote-22)

Several issues regarding the application of delta-method asymptotic linear-approximation merit cautionary mention here, the intrinsic appeal of analytic solutions notwithstanding. First, deriving from a linearization, the certainty estimates only approximate validly in some proximity of the estimated nonlinear expression, and we do not know in general how small a range. Being asymptotic, they only approximate validly for large samples, how large also being unknown, and they are in any event an approximation. Finally, using the approximately estimated standard errors to generate confidence intervals and hypothesis tests in the usual manners assumes multivariate normality of the parameter estimates. Although all maximum-likelihood estimates are at least asymptotically normal, sample-size concerns may arise, perhaps especially regarding estimates involving $\hat{ρ}$, which is where the spatial complications tend to arise also. Given all this, the asymptotic linear-approximations we have previously recommended may have been larger than need be even in the linear-regression context.

Even in these spatial linear-regression contexts, though, simulation strategies for calculating effects and responses and associated certainty estimates—i.e., sampling coefficient estimates from s with the estimated, and calculating the quantities of interest and their certainty estimates from those draws—is often easier and as or more effective.[[23]](#footnote-23) Given that the nonlinearities in the estimates of interest in spatial probit are more severe and that asymptotic normality may be more distant,[[24]](#footnote-24) we especially stress simulation methods here.

**Appendix VI: Results for Figure 1**

The numeric results for the experiments represented in Figure 1 in the text are reported below:

**Experiment # 1 –**

|  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
| *φ*=0.3 *ρ*=0.1 | **t** | **t+1** | **t+2** | **t+3** | **t+4** | **t+5** | **t+6** | **t+7** | **t+8** | **t+9** |
| *True Effect* | *0.0552* | *0.0258* | *0.0108* | *0.0042* | *0.0016* | *0.0006* | *0.0002* | *0.0001* | *0.0000* | *0.0000* |
| Mean Effect Estimate | *0.0489* | *0.0220* | *0.0089* | *0.0034* | 0.0012 | 0.0005 | 0.0002 | 0.0001 | 0.0000 | 0.0000 |
| STD | 0.0230 | 0.0113 | 0.0053 | 0.0025 | 0.0012 | 0.0006 | 0.0003 | 0.0001 | 0.0001 | 0.0001 |
| SE | 0.0244 | 0.0117 | 0.0052 | 0.0023 | 0.0010 | 0.0005 | 0.0002 | 0.0001 | 0.0001 | 0.0000 |

**Experiment # 2 –**

|  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
| *φ*=0.3 *ρ*=0.25 | **t** | **t+1** | **t+2** | **t+3** | **t+4** | **t+5** | **t+6** | **t+7** | **t+8** | **t+9** |
| *True Effect* | *0.1349* | *0.0704* | *0.0322* | *0.0137* | *0.0056* | *0.0022* | *0.0009* | *0.0003* | *0.0001* | *0.0001* |
| Mean Effect Estimate | 0.1172 | 0.0569 | 0.0252 | 0.0103 | 0.0046 | 0.0019 | 0.0012 | 0.0005 | 0.0007 | 0.0003 |
| STD | 0.0222 | 0.0163 | 0.0102 | 0.0082 | 0.0064 | 0.0058 | 0.0052 | 0.0049 | 0.0045 | 0.0043 |
| SE | 0.0261 | 0.0159 | 0.0096 | 0.0061 | 0.0042 | 0.0032 | 0.0027 | 0.0025 | 0.0023 | 0.0022 |

**Experiment # 3 –**

|  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
| *φ*=0.5 *ρ*=0.1 | **t** | **t+1** | **t+2** | **t+3** | **t+4** | **t+5** | **t+6** | **t+7** | **t+8** | **t+9** |
| *True Effect* | *0.0626* | *0.0439* | *0.0292* | *0.0186* | *0.0115* | *0.0069* | *0.0041* | *0.0024* | *0.0014* | *0.0008* |
| Mean Effect Estimate | 0.0562 | 0.0382 | 0.0243 | 0.0148 | 0.0088 | 0.0051 | 0.0029 | 0.0017 | 0.0010 | 0.0006 |
| STD | 0.0190 | 0.0134 | 0.0092 | 0.0063 | 0.0043 | 0.0029 | 0.0020 | 0.0014 | 0.0010 | 0.0007 |
| SE | 0.0245 | 0.0172 | 0.0116 | 0.0076 | 0.0049 | 0.0031 | 0.0020 | 0.0013 | 0.0008 | 0.0006 |

**Experiment # 4 –**

|  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
| *φ*=0.5 *ρ*=0.25 | **t** | **t+1** | **t+2** | **t+3** | **t+4** | **t+5** | **t+6** | **t+7** | **t+8** | **t+9** |
| *True Effect* | *0.1507* | *0.1162* | *0.0851* | *0.0597* | *0.0407* | *0.0272* | *0.0179* | *0.0116* | *0.0075* | *0.0049* |
| Mean Effect Estimate | 0.1358 | 0.0996 | 0.0688 | 0.0456 | 0.0294 | 0.0186 | 0.0117 | 0.0073 | 0.0046 | 0.0028 |
| STD | 0.0182 | 0.0156 | 0.0126 | 0.0097 | 0.0073 | 0.0053 | 0.0038 | 0.0026 | 0.0018 | 0.0013 |
| SE | 0.0187 | 0.0153 | 0.0120 | 0.0091 | 0.0067 | 0.0049 | 0.0035 | 0.0025 | 0.0017 | 0.0012 |

**Appendix VII: Characteristics of Weights Matrices**

**Number of Links in Percentage Terms**

|  |  |  |
| --- | --- | --- |
| **Number of Links** | **US States (Monte Carlos)** | **Sub-Saharan Africa (Illustration)** |
| **0** | **4** | **0** |
| **1** | **2** | **5** |
| **2** | **6** | **15** |
| **3** | **20** | **23** |
| **4** | **24** | **15** |
| **5** | **20** | **19** |
| **6** | **18** | **15** |
| **7** | **2** | **4** |
| **8** | **4** | **3** |

**Appendix VIII: Bibliography**

Note: This is a complete list of sources we’ve found which may be useful for researchers including, and extending upon, those works we explicitly reference in the main text.

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1. See, e.g., McMillen 1992, 1995, 2005; Bolduc et. al. 1997; Pinkse & Slade 1998; LeSage 1999, 2000, LeSage&Pace 2004, 2009; Beron et al. 2003; Beron & Vijverberg 2004. Spatial logit has also been suggested (e.g., Dubin 1997; Lin 2003; Autant-Bernard 2006), but spatial probit dominates the methodological and applied literatures, perhaps because the *n*-dimensional normal is relatively easier to manage than the *n*-dimensional extreme-value distribution. [↑](#footnote-ref-1)
2. E.g., Holloway et al 2002, Beron et al 2003, Coughlin et al 2003, Murdoch et al 2003, Novo2003, Schofield et al 2003, Garrett et al. 2005, Lacombe & Shaughnessy 2005, Autant-Bernard 2006, Rathbun&Fei 2006, Mukherjee&Singer 2007. [↑](#footnote-ref-2)
3. Even the linear-probability model becomes nonlinear in parameters given the spatial multiplier, . [↑](#footnote-ref-3)
4. The instrumented-approximation approaches, on the other hand, are massively more efficient computationally, with estimation times orders of magnitude faster, which can be a dominant consideration in samples of thousands, plus. [↑](#footnote-ref-4)
5. The requirement applies to any simultaneous feedback among endogenous qualitative variables, as Heckman (1978) noted for a system of two endogenous equations, at least one being qualitative and modeled by a latent variable crossing a threshold. He states: “A necessary and sufficient condition for [sensibility of such a system of endogenous latent-variable equations is] that the probability of the event *di*=1 is not a determinant of the event… …[This] principal assumption essentially requires that the latent variable *y*\* and not the measured variable *y* appears [on the right-hand side of the] structural equation” (pp. 936-7). The same limitation does not quite obtain for temporal dependence, however. Since time is unidirectional, one may be able to rely on pre-determinedness of *y*t‑1, i.e., the indirect feedback from *y*t to *y*t‑1 does not occur (given sufficiently full and accurate specification of the temporal dynamics). Still, conditions for proper identification of just a temporally dynamic model with lagged binary-dependent-variables remain less than straightforward (see, e.g., Chamberlain 1993, Honore & Kyriazidou 2000). [↑](#footnote-ref-5)
6. Monte Carlo simulation exploring the sensitivity of the time-lagging spatial-dependence strategy to validity of the lagged-interdependence-only assumption, to the periodicity-matching assumption, and to the empirical adequacy of the spatiotemporal dynamic model and tests thereof are also important analyses that remain for the future. [↑](#footnote-ref-6)
7. Our simple introduction draws heavily from Gill’s (2002) wonderful text on Bayesian methods. [↑](#footnote-ref-7)
8. The *population parameters* thusly *arbitrarily closely approximated* are usually some *estimates* in an application, like spatial-probit parameter-estimates, not the *true parameters* (a foreign concept in Bayesian terminology anyway). [↑](#footnote-ref-8)
9. The conditional distributions must also be expressible & sufficiently tractable to make so many draws a practicality. [↑](#footnote-ref-9)
10. To elaborate: to sample from some non-standard density *f(∙)*, let *x*0 be the current draw from *f(∙)*, beginning with an arbitrary starting value. Consider a candidate next value, *x*1, for *x* given by *x*1=*x*0+*cZ* with *Z* being drawn from a standard-normal distribution and *c* a given constant. Then, we assign a probability of accepting this candidate as the next value of *x* in our MCMC as *p*=min{1, *f(x1)/ f(x*0*)*}. I.e., we draw from a Uniform(0,1) distribution, and, if *U<p*, the candidate *x*1 becomes the next *x*; if *U>p* *x* remains *x*0. Metropolis-Hastings is thus one type of *rejection sampling*. [↑](#footnote-ref-10)
11. This enables LeSage to express the spatial-Tobit model by this same likelihood, adding a conditional distribution later to generate latent variables *z* for censored observations instead of one to generate *y*=(0,1) for the probit. [↑](#footnote-ref-11)
12. Anselin (1988) shows that the minimum and maximum eigenvalues of a standardized spatial-weight matrix, **W**, bound *ρ* to 1/λmin<*ρ<*1/λmax. Adding this constraint to the rejection sampling should be beneficial. [↑](#footnote-ref-12)
13. If we have correctly generated the multivariate analogue to the erroneously univariate expression in Lesage (2000) and Smith & LeSage (2004), spatial Tobit would replace with:. [↑](#footnote-ref-13)
14. While one doesn’t need near as many *m* on this Gibbs-within-Gibbs/MH sampler as the thousands recommended for outer Gibbs/MH sampler, but even *m*=10 for, say, a sample of the 3000 US counties yields 30,000 draws within each of the outer thousands of draws. For instance, LeSage & Pace (2009) report that, for just *m*=1 and merely 1000 outer draws for the 3000 US counties, their “relatively slow laptop” required 45 minutes for one spatial-probit estimation. [↑](#footnote-ref-14)
15. In assigning diffuse priors to the parameters, LeSage (2000) also relaxes the assumption of homoskedasticity in **ε**, allowing V(*ε*) to vary arbitrarily by observation *i*. This allows exploration of variation in model fit and identification of and robustness to potential outliers, but creates as many parameters to estimate as observations. LeSage circumvents that issue by specifying an informative prior for those relative-variance parameters, specifically one suggested by Geweke (1993) that, *inter alia*, has the useful property of yielding a distribution of **ε** consistent with a probit choice-model as the Gewekian-distribution parameter, *q*, goes to infinity, and that at *q*≈7.5 yields a choice-model approximating logit. The posterior-estimates of *q*, may therefore be used to test logit versus probit (versus un-named possibilities *q*≠7.5 and *q*≠∞).

Allowing arbitrary relative-variance requires the additional (informative) Gewekian prior and a (diffuse) hyper-prior on its parameter, *q*; produces more complicated expressions for the conditional distributions of *σ*, *ρ*, **β**; and adds a conditional distribution (fortunately standard: χ2q+1) for the relative variances, *υ*i. The steps below would now also include conditioning on starting values for, and then the previous draws of, **υ**, and a step inserted between 2 and 3 would draw the next **υ** from χ2q+1 conditional on the current *σ*, *ρ*, **β**. Notice that setting the hyper-prior for *q* determinately to a large number (or 7.5) yields spatial probit (or logit) without heteroscedasticity/outlier-robustness. [↑](#footnote-ref-15)
16. Thinning may also be advisable, although we have not yet explored that or found relevant discussion in the literature. [↑](#footnote-ref-16)
17. If one wishes to include estimated inherent-uncertainty as well as estimation-uncertainty in these counterfactuals, then one should also draw **ε** from its independent-normal distributions, adding it to the  in the next term. [↑](#footnote-ref-17)
18. Beron & Vijverberg (2004) calculate the *marginal* effect of **x***i* on the probability *yi*=1, avoiding the multivariate integral. We would also be interested in *d*[p(*yi*=1)]/*d***x***j*, but this would require conditioning as noted above. Beron & Vijverberg (2004) argue that it is inappropriate to condition thusly because the *yj* respond endogenously to the changes in **X**, but this claim seems unnecessarily restrictive since, as we just explained, once we estimate the model, we can easily sample from the distribution of disturbances using the reduced-form, generate *y*’s according to the measurement equation, and calculate exactly these conditional frequencies. [↑](#footnote-ref-18)
19. Assuming stationarity, these differences should fade for all units, including unit 1, going forward in time (because we do not make the hypothetical shock permanent). [↑](#footnote-ref-19)
20. Specifically, the Senegalese government sent troops to support the Vieira regime, while MFDC sent forces to support revolutionary Ansoumane Mané. [↑](#footnote-ref-20)
21. The **Ly\*** term is not involved in and so drops out of Δ**y**t\*/Δ**x**t, and so can be omitted here. [↑](#footnote-ref-21)
22. Appendix IV contains further discussion of delta method versus parametric simulation in this context. [↑](#footnote-ref-22)
23. Note: the average of simulated quantities of interest and their standard deviation will not generally coincide exactly with the quantity of interest calculated at the ML parameter estimates and their (Delta Method approximated) standard errors. The former are  and , with  whereas the latter are  and . By definition of *maximum likelihood* and its invariance property, the latter should correspond to a modal estimate of the quantity of interest and the estimated asymptotic variance of the linear-approximation to that modal estimate, whereas the former is the average and variance of the quantity of interest calculated at draws from the multivariate normal sampling/posterior distribution. [↑](#footnote-ref-23)
24. In fact, the (kernel of the) posterior/likelihood joint-density of the parameters is not normal (due to the term  term), and, of the posterior/likelihood conditional-distributions, only that of **β** is exactly normal. [↑](#footnote-ref-24)