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WEB APPENDIX

Expanded References

(These lists are supplemental to the citations in the text and follow the order of the notes in which the abbreviated lists in the text appear. Reference list appears at the end of this Appendix.)

On policy-innovation diffusion/interdependence in the U.S., see also:
Walker 1973; Knoke 1982; Lutz 1987; Berry & Berry 1990,1999; Case et al. 1993; Rogers 1995; Mintrom 1997ab; Mintrom & Vergari 1998; Mossberger 1999; Godwin&Schroedel 2000; Balla 2001; Mooney 2001; Bailey & Rom 2004; Daley & Garand 2004; Grossback et al. 2004; Volden 2006.

On comparative/international policy diffusion/interdependence, see also:
Rose 1993, Meseguer 2005; Gilardi 2005.

On citizen-vote diffusion/interdependence, see also:
O'Laughlin et al. 1994; Schofield et al. 2003; Pattie & Johnston 2000; Beck et al. 2002ab; Calvo & Escolar 2003; Kim et al. 2003.

On election-outcomes diffusion/interdependence, see also:
Shin & Agnew 2007; Hiskey & Canache 2005; Wing & Walker 2006.

On diffusion/interdependence of candidate qualities, strategies, or contributions, see also:
Mizruchi 1989; Krasno et al. 1994.; Cho 2003; Gimpel et al. 2006.

On civil-war diffusion/interdependence, see also:
Murdoch & Sandler 2002; Buhaug & Rød 2006; Salehyan & Gleditsch 2006.

On contextual/neighborhood effects from interdependence perspective, see also (and in this case, even this is a still highly abbreviated list):
Knack & Kropf '98; Liu et al. 1998; Braybeck & Huckfeldt 2002ab; Beck et al. 2002ab; McClurg 2003.

On interdependence and/in social movements, see also:
Conell & Cohn 1995; Giugni 1998; Strang & Soule 1998; Biggs 2003; Browning et al. 2004; Andrews & Biggs 2006; Holmes 2006.

On interdependence/globalization and the international diffusion of economic liberalization, see also (and in this case, even this is still a highly abbreviated list):
Brune et al. 2004; Elkins et al. 2006; Brooks 2005, 2007; Jordana & Levi-Faur 2005; Way 2005; Lazer 2006; Prakash & Potoski 2006; Simmons et al. 2006.

On globalization-induced interdependencies in economic and other policies more generally, even a still radically abbreviated list is lengthy:

Basinger & Hallerberg 2004; Henisz et al. 2005; Elkins et al. 2006; Jahn 2006; Swank 2006; Baturo & Grey 2007; Cao 2007; Mosley & Uno 2007; Guler et al. 2002; Badinger et al. 2004; Heichel et al. 2005; Holzinger & Knill 2005; Knill 2005; Polillo & Guillén 2005; Lee & Strang 2006; Manger 2006; Cao et al. 2007; Coughlin et al. 2007; Garretsen & Peeters 2007.

On the intellectual-historical genesis of *Galton's Problem*:

Galton originally raised the issue thus: "[F]ull information should be given as to the degree in which the customs of the tribes and races which are compared together are independent. It might be that some of the tribes had derived them from a common source, so that they were duplicate copies of the same original. ...It would give a useful idea of the distribution of the several customs and of their relative prevalence in the world, if a map were so marked by shadings and colour as to present a picture of their geographical ranges" (Galton 1889, as quoted in Darmofal 2007). We find further historical context in http://en.wikipedia.org/wiki/Galton's_problem: "In [1888], Galton was present when Sir Edward Tylor presented a paper at the Royal Anthropological Institute. Tylor had compiled information on institutions of marriage and descent for 350 cultures and examined the correlations between these institutions and measures of societal complexity. Tylor interpreted his results as indications of a general evolutionary sequence, in which institutions change focus from maternal to paternal lines as societies grow more complex. Galton disagreed, noting that similarity between cultures could be due to borrowing, could be due to common descent, or could be due to evolutionary development; he maintained that without controlling for borrowing and common descent one cannot make valid inferences regarding evolutionary development. Galton's critique has become the eponymous *Galton's Problem* (Stocking 1968:175), as named by Raoul Naroll (1961, 1965), who proposed [some of] the first statistical solutions."

On distinguishing spatial-econometric & spatial-statistical approaches to spatial analysis:

Methodologically, two approaches to spatial analysis can be discerned: spatial statistics and spatial econometrics (Anselin 2006). The distinction rests, on the one hand, on the relative emphasis in spatial-econometric approaches to theoretical models of interdependence processes wherein space may often have broad meaning, well beyond geography and geometry to encompass all manner of social, economic, or political connection that induces effects from outcomes in some units on outcomes in others (Brueckner 2003; Beck et al. 2006). The spatial-lag regression model plays a starring role in that tradition (Hordijk 1974; Paelinck and Klaassen 1979; Anselin 1980, 1988, 1992; Haining 1990; LeSage 1999). According to Anselin, theory driven models deal with substantive spatial correlation (2002). This approach has its own method to model specification and estimation. The importance of spatial interdependence is tested using Wald tests and the unrestricted spatial lag model. On the other hand, spatial-error models, analysis of spatial-correlation patterns, spatial kriging, and spatial smoothing, e.g., characterize the more-exclusively data-driven approaches and the typically narrower conception of space in solely geographic/geometric terms in the longer spatial-statistics tradition (initially inspired by Sir Galton's famous comments at the 1888 meetings of the Royal Anthropological Society, and reaching crucial methodological milestones in Whittle 1954; Cliff & Ord 1973, 1981; Besag 1974; Ord 1975; Ripley 1981; Cressie 1993). According to Anselin, this kind of modeling is driven by data problems such as measurement error. The spatial correlation is viewed as a *nuisance*. This approach has its own method to model specification and estimation. The presence of spatial effects is tested using Lagrange multiplier tests and the restricted non-spatial lag model.

Formal-Theoretical Model of Capital-Tax Competition (Persson & Tabellini 2000:ch. 12):

We gave a general theoretical formulation of interdependence above, following Brueckner (2003); here we follow Persson & Tabellini (2000:ch.12) to elaborate formally how tax competition specifically implies spatial interdependence. The model's essentials are these. In two jurisdictions (e.g., countries), denote the domestic and foreign capital-tax rates τ_k and τ_k^*. Individuals can invest in either country, but foreign investment incurs *mobility costs*. Governments use revenues from taxes levied (by the source, not the residence, principle) on capital and on labor to fund a fixed spending-level.[1] Individuals differ in their relative labor-to-capital endowment, e^i, and make labor-leisure, l and x, and savings-investment, $s=k+f$ (k=domestic; f=foreign), decisions to maximize quasi-linear utility, $\omega=U(c_1)+c_2+V(x)$, over leisure and consumption in the model's two periods, c_1 and c_2, subject to a time constraint, $1+e^i=l+x$, and to period-1 and -2 budget constraints, $1\text{-}e^i=c_1+k+f+\equiv c_1+s$ and $c_2=(1\text{-}\tau_k)k+(1\text{-}\tau^*_k)f\text{-}M(f)+(1\text{-}\tau_l)l$.

The equilibrium economic choices of citizens i in this model are as follows:

$$s = S(\tau_k) = 1 - U_c^{-1}(1-\tau_k) \qquad (1),$$

$$f = F(\tau_k, \tau_k^*) = M_f^{-1}(\tau_k - \tau_k^*) \qquad (2),$$

$$k = K(\tau_k, \tau_k^*) = S(\tau_k) - F(\tau_k, \tau_k^*) \qquad (3).$$

With labor, $L(\tau_l)$, leisure, x, and consumption, c_1, c_2, implicitly given by these conditions, this leaves individuals with indirect utility, W, defined over the policy variables, capital and labor tax-rates, of:

$$W(\tau_l,\tau_k) = U\left(1-S(\tau_k)\right) + \left(1-\tau_k\right)S(\tau_k) + \left(\tau_k - \tau_k^*\right)F(\tau_k,\tau_k^*)\; M\left(F(\tau_k,\tau_k^*)\right) + \left(1-\tau_l\right)L(\tau_l) + V\left(1-L(\tau_l)\right) \quad (4).$$

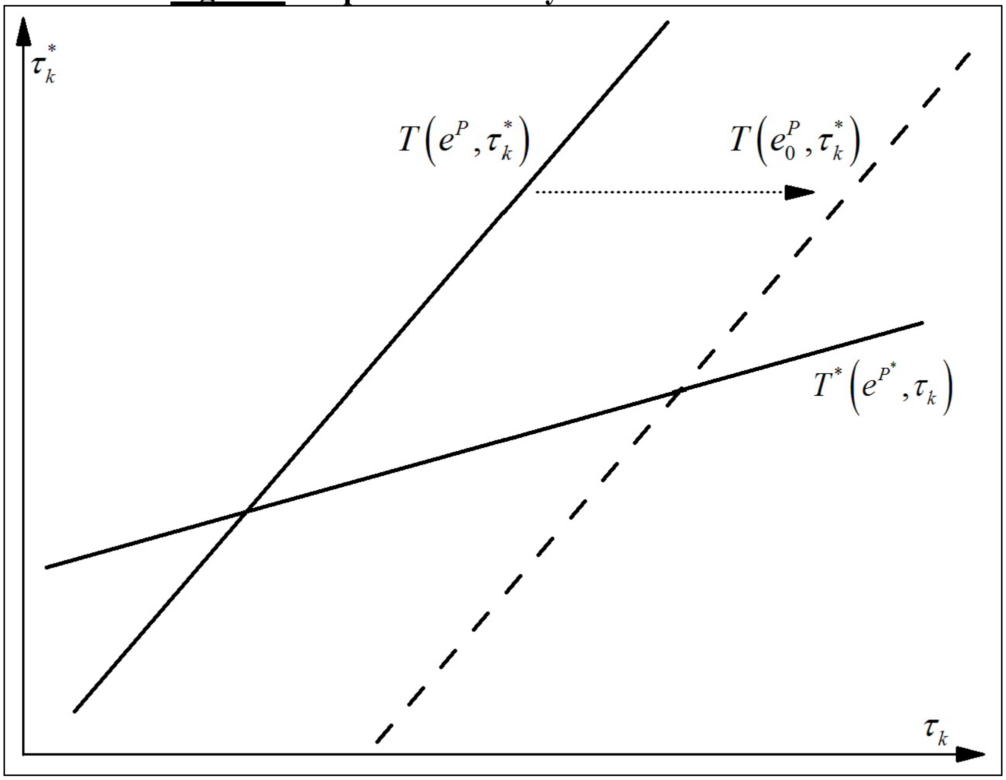
Facing an electorate with these preferences, using a Besley-Coate (1997) citizen-candidate model wherein running for office is costly and citizens choose whether to enter the race by an expected-utility calculation, some citizen-candidate will win and set tax rates to maximize its own welfare. The model's stages are: 1) both countries hold elections, 2) elected citizen-candidates set their countries' tax rates, and 3) all private economic decisions are made. In this case, the candidate who enters and wins will be the one with labor-capital endowment, e^P, such that s/he will implement this *Modified Ramsey Rule*:

$$\frac{S(\tau_k^p)-e^p}{S(\tau_k^p)}\left[1+\varepsilon_l(\tau_l^p)\right] = \frac{L(\tau_l^p)+e^p}{L(\tau_l^p)}\left[1+\frac{S_\tau(\tau_k^p)+2F_\tau^*(\tau_k^{p*},\tau_k^p)\tau_k}{S_\tau(\tau_k^p)}\right] \qquad (5).$$

The *Ramsey Rule* from public economics indicates the optimality of equating the marginal net-benefit elasticities of alternative instruments, here: labor and capital taxes. Essentially, one equates elasticities of labor-supply with respect to labor tax and of savings-investment with respect to capital tax: roughly, the bracketed terms. The modifications involve the weighting toward capital-tax as citizen-candidates' labor-endowments, e^p, increase, and the complications in the right brackets reflecting the foreign- and domestic-investment tax-elasticities to domestic capital-taxes, which push the other way.

[1] N.b., government consumption is also entirely wasted; i.e., it enters no one's utility function.

<u>Figure 1</u>: Capital-Tax Policy Reaction-Functions



Most crucially, (5) shows *optimal capital-tax-rates for domestic policymakers to be a function of capital tax-rates abroad, τ^*_k*. Optimal capital-tax-rates for foreign policymakers depend on domestic capital-tax-rates, τ^*, analogously. I.e., (5) gives best-response functions like those specified generically above; specifically: $\tau_k = T(e^P, \tau_k^*)$ and $\tau_k^* = T^*(e^{P^*}, \tau_k)$ for foreign and domestic policymakers. In words: domestic capital-tax rates depend on domestic policymakers' endowments and foreign capital-tax rates (and so on foreign policymakers' endowments); i.e., capital-tax policy is strategically interdependent. These reaction-function slopes, $\partial T/\partial \tau_k^*$ and $\partial T^*/\partial \tau_k$, can be positive or negative because, while higher foreign tax-rates induce capital flow into the domestic economy, this increased tax-base may induce domestic policymakers to lower rates or, instead, to raise rates given the reduced elasticity of this base. Figure 1 graphs the reaction functions assuming they slope positively. The illustrated comparative-static shows an increase in the domestic policymaker's labor-endowment (a leftward shift?), which shifts the domestic function outward and thereby raises equilibrium capital-tax rates *in both countries*.

On the Omitted-Variable Biases of non-spatial OLS & Simultaneity Biases of Spatial OLS:

If researchers omit the spatial lag necessary to reflect the actual interdependence of their data, their OLS coefficient-estimates for domestic, exogenous-external, or context-conditional factors will suffer standard omitted-variable biases (OVB). The formula for OVB is **Fβ** where **F** is the matrix of coefficients from regressing the omitted on the included variables and **β** is the vector of (true) coefficients on the omitted variables:

$$\mathrm{OVB}\begin{bmatrix} \hat{\beta}_{\mathrm{OLS}} \\ \hat{\varphi}_{\mathrm{OLS}} \end{bmatrix} = \rho \times \left(\mathbf{Q}_1'\mathbf{Q}_1\right)^{-1}\mathbf{Q}_1'\mathbf{W}\mathbf{y}, \text{ where } \mathbf{Q}_1 \equiv \begin{bmatrix} \mathbf{X} \\ \mathbf{M}\mathbf{y} \end{bmatrix} \qquad (6).$$

($\hat{\rho}_{\mathrm{OLS}} \equiv 0$ by construction, which is biased by $-\rho$.) Thus, insofar as spatial lags covary with the non-spatial regressors, which is highly likely if domestic conditions correlate spatially and is certain for common exogenous-external shocks, OLS overestimates domestic, exogenous-external, or context-conditional effects while ignoring spatial interdependence. This applies equally to qualitative analyses that ignore interdependence sources of their observed phenomena.

However, OLS estimation of models including spatial lags (or qualitatively observing the spatial correlation of outcomes across units or tracing putative diffusion processes) entails an endogeneity— spatial lag, **Wy**, covaries with residual, **ε**—and so simultaneity bias/inconsistency. Simply: spatial lags are weighted averages of outcomes in other units; they thus place some observations' left-hand sides on others' right-hand sides. Simpler still, by example: Germany causes France, but France also causes Germany: textbook simultaneity. The asymptotic simultaneity biases (SB) are:

$$\mathrm{SB}\begin{bmatrix}\hat{\rho} & \hat{\varphi} & \hat{\boldsymbol{\beta}}\end{bmatrix}' = \left(\mathbf{Q}'\mathbf{Q}\right)^{-1}\mathbf{Q}'\boldsymbol{\varepsilon}, \text{ where } \mathbf{Q} \equiv \begin{bmatrix}\mathbf{Wy} & \mathbf{My} & \mathbf{X}\end{bmatrix} \qquad (7).$$

In the case where **X** contains just one exogenous explanator, **x**, we can rewrite these biases thus:

$$\mathrm{SB}\begin{bmatrix}\hat{\rho}\\ \hat{\varphi}\\ \hat{\beta}\end{bmatrix} = \frac{1}{|\boldsymbol{\Psi}|}\begin{bmatrix}\operatorname{cov}(\mathbf{Wy},\boldsymbol{\varepsilon})\times\operatorname{var}(\mathbf{My})\times\operatorname{var}(\mathbf{x})\\ -\operatorname{cov}(\mathbf{Wy},\boldsymbol{\varepsilon})\times\operatorname{cov}(\mathbf{Wy},\mathbf{My})\times\operatorname{var}(\mathbf{x})\\ -\operatorname{cov}(\mathbf{Wy},\boldsymbol{\varepsilon})\times\operatorname{cov}(\mathbf{Wy},\mathbf{x})\times\operatorname{var}(\mathbf{My})\end{bmatrix}, \text{ where } \boldsymbol{\Psi} = \operatorname{plim}\left(\frac{\mathbf{Q}'\mathbf{Q}}{n}\right) \qquad (8).$$

In the likely common case of positive interdependence and positive covariance of spatial-lag and exogenous regressors, for instance, one would overestimate the interdependence-strength, $\hat{\rho}$, and correspondingly underestimate temporal dynamics, $\hat{\varphi}$, and domestic/external/contextual effects, $\hat{\beta}$.

On statistical estimation of spatial-lag models:

The instrumental-variables (IV), two-stage-least-squares (2SLS), generalized-method-of-moments (GMM) family of estimators relies on the spatial structure of the data to instrument for the endogenous spatial lag. On the assumption that what we call *cross-spatial endogeneity*, **y**'s in some units cause **x**'s in others, does not exist, instruments comprised of **WX** are ideal by construction. Cross-spatial endogeneity may seem highly unlikely in many contexts, perhaps, until one realizes that combinations of vertical connections from $\mathbf{y}_i$ to $\mathbf{y}_j$ and horizontal ones from $\mathbf{y}_j$ to $\mathbf{x}_j$ (the usual sort of endogeneity) combine to give the offending diagonal ones from $\mathbf{y}_i$ to $\mathbf{x}_j$. As usual, there are no magic instruments in empirical analysis.

Conditional likelihood functions for spatio-temporal-lag models are straightforward extensions of standard spatial-lag likelihood functions, which, in turn, add only one mathematically and conceptually small complication (albeit a computationally intense one) to the likelihood functions for standard linear-normal models. To see this, first rewrite the spatial-lag model with the stochastic component on the left:

$$\mathbf{y} = \rho\mathbf{Wy} + \mathbf{X}\boldsymbol{\beta} + \boldsymbol{\varepsilon} \Rightarrow \boldsymbol{\varepsilon} = \left(\mathbf{I} - \rho\mathbf{W}\right)\mathbf{y} - \mathbf{X}\boldsymbol{\beta} \equiv \mathbf{Ay} - \mathbf{X}\boldsymbol{\beta} \qquad (9).$$

Assuming *i.i.d.* normality, the likelihood function for **ε** is then the typical linear-normal one:

$$L(\boldsymbol{\varepsilon}) = \left(\frac{1}{\sigma^2 2\pi}\right)^{\frac{NT}{2}} \exp\left(-\frac{\boldsymbol{\varepsilon}'\boldsymbol{\varepsilon}}{2\sigma^2}\right) \qquad (10),$$

which, in this case, will produce a likelihood in terms of **y** as follows:

$$L(\mathbf{y}) = |\mathbf{A}|\left(\frac{1}{\sigma^2 2\pi}\right)^{\frac{NT}{2}} \exp\left(-\frac{1}{2\sigma^2}\left(\mathbf{Ay} - \mathbf{X}\boldsymbol{\beta}\right)'\left(\mathbf{Ay} - \mathbf{X}\boldsymbol{\beta}\right)\right) \qquad (11).$$

This still resembles the typical linear-normal likelihood, except that the transformation from **ε** to

y is not by the usual factor, 1, but by |**A**|=|**I**- ρ**W**|.[2] Written in $(N\times 1)$ vector notation, the spatio-temporal-model conditional-likelihood is mostly conveniently separable into parts, as seen here:

$$\mathrm{Log} f_{\mathbf{y}_t,\mathbf{y}_{t-1},\ldots,\mathbf{y}_2|\mathbf{y}_1} = -\frac{1}{2}N(T-1)\log\left(2\pi\sigma^2\right)+(T-1)\log\left|\mathbf{I}-\rho\mathbf{W}\right|-\frac{1}{2\sigma^2}\sum_{t=2}^{T}\boldsymbol{\varepsilon}_t'\boldsymbol{\varepsilon}_t \qquad (12),$$

where $\boldsymbol{\varepsilon}_t = \mathbf{y}_t - \rho\mathbf{W}_N\mathbf{y}_t - \phi\mathbf{I}_N\mathbf{y}_{t-1} - \mathbf{X}_t\boldsymbol{\beta}$.

We note that the unconditional (exact) likelihood function, retaining the first time-period observations as non-predetermined, is more complicated (Elhorst 2001, 2003, 2005).[3]

$$\begin{aligned}\mathrm{Log} f_{\mathbf{y}_t,\ldots,\mathbf{y}_1} = &-\frac{1}{2}NT\log\left(2\pi\sigma^2\right)+\frac{1}{2}\sum_{i=1}^{N}\log\left(\left(1-\rho\omega_i\right)^2-\phi^2\right)+(T-1)\sum_{i=1}^{N}\log\left(1-\rho\omega_i\right)\\ &-\frac{1}{2\sigma^2}\sum_{t=2}^{T}\varepsilon_t'\varepsilon_t-\frac{1}{2\sigma^2}\varepsilon_1'\left(\left(\mathbf{B}-\mathbf{A}\right)'\right)^{-1}\left(\mathbf{B}'\mathbf{B}-\mathbf{B}'\mathbf{A}\mathbf{B}^{-1}\left(\mathbf{B}'\mathbf{A}\mathbf{B}^{-1}\right)'\right)^{-1}\left(\mathbf{B}-\mathbf{A}\right)^{-1}\varepsilon_1\end{aligned} \qquad (13)$$

where $\boldsymbol{\varepsilon}_1 = \mathbf{y}_1 - \rho\mathbf{W}_N\mathbf{y}_1 - \phi\mathbf{I}_N\mathbf{y}_1 - \mathbf{X}_1\boldsymbol{\beta}$. When T is small, the first observation contributes greatly to the overall likelihood, and the unconditional likelihood should be used to estimate the model. In other cases, the more compact conditional likelihood is acceptable for estimation purposes.

We explained above that model specifications that omit spatial lags assume zero interdependence by construction and show (analytically and in simulations) that this induces omitted-variable bias that inflates estimates of the effects of non-spatial model-components. Note that, in the present substantive context, this means that most extant studies of globalization, having neglected spatial lags, will have overestimated the effects of domestic and exogenous-external conditions while effectively preventing any globalization-induced interdependence from manifesting empirically. Conversely, we also showed that standard regression estimates of models with spatial lags suffer simultaneity biases, which have become more common recently among researchers interested in interdependence, and which our previously analyses as summarized above show to have vastly improved upon previous non-spatial estimation strategies, will nonetheless tend to inflate interdependence-strength estimates at the expense of domestic and exogenous-external explanators. We also shown there that the spatial-ML approach just described effectively redresses the simultaneity issues.

Finally, the issue of stationarity arises in more-complicated fashion in spatio-temporal dynamic models than in purely temporally dynamic ones. Nonetheless, the conditions and issues arising in the former are reminiscent but not identical to those arising in the latter. Defining $\mathbf{A}=\phi\mathbf{I}$, $\mathbf{B}=\mathbf{I}-\rho\mathbf{W}$, and ω as a characteristic root of $\mathbf{W}$, the spatio-temporal process is covariance stationary if

$$\left|\mathbf{A}\mathbf{B}^{-1}\right|<1 \qquad (14),$$

or, equivalently, if

[2] This difference does complicate estimation somewhat. The complication is that the determinant, |**A**|, involves ρ and so must be calculated at each iteration of the likelihood-maximization routine. Two strategies that simplify the problem are using an eigenvalue approximation for the determinant (Ord 1975) and maximizing a concentrated likelihood function (Anselin 1988). We discuss both of these procedures, and estimation more generally, elsewhere (Franzese & Hays 2004, 2006a, 2007cd, 2008ab).

[3] Note that the same condition that complicates ML estimation of the spatio-temporal lag model, namely the first set of observations is stochastic, also invalidates the use of OLS to estimate a model with a temporally lagged spatial lag. (Franzese & Hays 2006a, 2007c present this likelihood). Hence, asymptotically, this consideration offers no econometric reason to prefer S-OLS over S-ML estimation of spatio-temporal-lag models or the converse.

$$\begin{cases} |\phi| < 1 - \rho\omega_{\max}, & \text{if } \rho \geq 0 \\ |\phi| < 1 - \rho\omega_{\min}, & \text{if } \rho < 0 \end{cases} \tag{15}$$

For example, in the case of positive temporal dependence and positive, uniform spatial dependence ($\rho > 0$ and $w_{ij} = 1/(N-1)\ \forall\ i \neq j$), stationarity requires simply that $\phi + \rho < 1$. In fact, the maximum characteristic root is +1 for any row-standardized **W**.

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