Supplemental Information File

The World Economy, Political Control, and Presidential Success

Ryan E. Carlin & Timothy Hellwig

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A. Revising the Selection Model of Rational Retrospective Voting for Latin America

In the text we argue that the model of rational retrospective voting must be reconceived to fit political and economic context of Latin America. Since length restrictions prevent us from developing this argument in full, we do so in this online appendix. The following discussion is informed by Duch and Stevenson (2008, chapter 5) which is, in turn, informed by Alesina and Rosenthal's (1995) signal extraction model of rational retrospective voting. Adapting the model for Latin America of the late 20th and early 21st centuries requires two innovations: i) incorporating policy into the competency signal and ii) incorporating policy into the number of decisions made by electorally dependent decision makers. We extract and expand upon those components of the model that pertain to our innovation and refer the reader to the above sources for the full derivations.

i. Incorporating Policy in the Competency Signal

The starting point of the model is the growth equation, which follows an expectations-augmented Phillips curve:

$$y_{it} = \bar{y} + \pi_{it} - \pi_{it}^{e} + \eta_{it} \qquad (1)$$

where y_{it} is the rate of economic growth at time *t* under incumbent party *i*; $\bar{y}$ is the natural rate of growth; π_{it} is the inflation rate, π_{it}^{e} is the natural rate of inflation that voters expect; and η_{it} is a random shock to the economy. This random shock then consists of two parts, such that

$$\eta_{it} = \varepsilon_{it} + \xi_{t} \qquad (2)$$

where ε_{it} is the "competency shock" attributable to the incumbent government's managerial competence; and the other part of the shock contributing to growth, ξ_t, is "exogenous" and is not dependent on the incumbent administration. These two parts of the growth shock prove to be key to determining the extent to which the voter assigns observed economic outcomes to the incumbent and, hence, sanctions the sitting government based on it.

In our conception of this model, the exogenous component, ξ_t, is treated the same as in Duch and Stevenson. However, we conceive of the competency shock in a different way. In the model, as advanced by Alesina and Rosenthal and Duch and Stevenson, the competency shock, ε_{it}, is comprised of shocks that inform voters about the incumbent's capacity to execute policy. Policy itself lies outside of ε_{it} and is captured in full by π_{it}, the incumbent's choice of inflation rate. Further, Duch and Stevenson make the simplifying assumption that the inflation rate does not vary across incumbents. They assume voters know that incumbent politicians will pick the level of inflation that maximizes their expected utility. This means voters "are never surprised by the government's inflation policy" and all politicians will choose the same economic policy of a zero inflation rate" (Duch and Stevenson 2008, 134). Equation (1) thus reduces simply to

$$y_{it} = \bar{y} + \eta_{it}. \qquad (3)$$

In our adaptation of the competency model for Latin America, we maintain that these simplifying assumptions are not credible. While the assumption that economic policy can be fully captured by the inflation rate in advanced capitalist democracies may be credible, this is not the case for developing and transitioning economies of Latin America. In the context of Western democracies, a 2% annual inflation rate communicates a great deal about incumbent performance. In an idealized version of the United States, for example, it would suggest the executive has met the inflation target by passing a budget and signing into law various bills (and vetoing others) to align fiscal policy with the interest rates the Federal Reserve Bank sets to govern monetary policy. Any adjustments in other policy areas are minor and gradual. Citizens, in turn, can determine with some accuracy how responsible the incumbent is for this outcome.

In Latin America, a 2% rate of annual inflation alone communicates far less about incumbent performance. On one hand, exposure to global economic forces, as CZ posit, may be determinative

and low-information voters may simply – and wrongly – fully credit the incumbent for the outcome. On the other hand, we argue, achieving 2% inflation may reflect not only complementary fiscal and monetary policies but also a mixture of mutually reinforcing choices over policy reform (versus status quo) in other policy areas such as trade, labor, privatization/industrial, financial markets, and taxation. And while external actors and lenders influence leaders' choice sets, ultimately only they take policy decisions. Application of the selection model of retrospective economic voting to Latin America ought to allow the competency shock to capture these decisions.[1]

Thus, as we describe in the main text, we assume that the market orientation of the incumbent's policies range between statist and neoliberal poles. These policy mixes imply different levels of inflation tolerance, such that $\pi_{it} \geq 0$. Policy regimes also vary widely across areas such as fiscal policy, industrial policy, labor market policy, and so on, i.e., those policies we incorporate into our summary *Policy Regime* index. Furthermore, these policy mixes imply *different degrees of incumbent control over the economy*, a key point on which we elaborate below.

Thus, unlike the competency model developed by Duch and Stevenson for a set of advanced capitalist democracies, our adapted model allows politicians to vary in terms of policy.[2] This means that in our revised model, the competency shock, ε_{it}, depends on the government's economic policy.

[1] Whether or not this is useful beyond Latin America is worthy of consideration but is beyond the scope of this rejoinder.
[2] Indeed, as mentioned in the text, for the countries in our sample variation in policy certainly is greater than variation in competence.

ii. Incorporating policy in the number of decisions made by EDDs

Duch and Stevenson extend the Alesina-Rosenthal model to provide a clearer substantive interpretation of the competency signal. This extension is valuable in that it provides observable implications of the model which can be assessed with data. They distinguish between two types of decision-makers: "electorally dependent decision makers" (EDDs) and nonelectorally dependent decision makers (NEDDs). The decisions made by these actors contribute to economic growth through the random shock, η_{it}. Augmenting equation (3), Duch and Stevenson rewrite the growth equation (their eq. 5.13) as

$$y_{it} = \bar{y} + \sum_{l=1}^{\alpha} \omega_{ilt} + \sum_{l=1}^{\beta} \psi_{lt} \qquad (4)$$

where ω_{ilt} is the growth shock associated with the *l*th decision of the EDD, indexed by *i*; and ψ_{lt} is the growth shock associated with the *l*th decision of the NEDD. In this model, α is the number of decisions made by EDDs and β is the number made by NEDDs.

From (4), Duch and Stevenson (2008, 141-145) formally derive a voter utility function whereby the effect of observed economy is conditioned by $\frac{\alpha\sigma_{\mu}^2}{\alpha\sigma_{\mu}^2+\beta\sigma_{\psi}^2}$. This "competency signal" controls how much information about the actions of incumbents voters can extract from the observed economy. As α (the number of decisions made by EDDs) increases while β (the number made by NEDDs) holds constant, the strength of the competency signal increases. This is an important insight. Electoral contexts, and the strength of rational retrospective voting, vary based on the number of decisions under consideration, and by whether they are made by EDDs or NEDDs.

In our application to Latin America, we are interested chiefly in α.[3] The number of decisions made by EDDs ranges across contexts for several reasons. Duch and Stevenson provide the example of presidential versus parliamentary regimes. If presidential regimes have larger numbers of elected decisions that affect the economy (on grounds that decisions are more dispersed in separation of powers systems), it follows that $\alpha_{presidential} > \alpha_{parliamentary}$. Relaxing the competency model to include policy regimes provides another source of variation in EDDs. A characteristic of Washington Consensus neoliberal orthodoxy is that it removes economically important decisions from the ambit of control by electorally dependent decision makers. In place of EDDs, decisions are made by technocratic elites and/or to align with pre-determined macro-economic targets, as well as uncoordinated market actors. In contrast, statist policy regimes imply government activism in numerous policy areas and, hence, a greater number of EDDs involved in many decisions. Thus it follows that $\alpha_{statism} > \alpha_{neoliberal}$, yielding a higher competency signal in statist regimes (when values of *Policy Regime* are low) than in neoliberal regimes (*Policy Regime* high).

[3] In contrast, by emphasizing the importance of LCSEs, Campello and Zucco (2016) implicit focus on β, the number made by NEDDs, in their analysis of economic conditions and presidential success in Latin America.

B. Replication and Extension: CZ Table 1

Table A1 provides parameter estimates for model diagnostics reported in the first two rows of Table 1 in the text. Missing values for *Policy Regime* are recovered using linear interpolation where possible and, for those countries without data (Panama and the United States), by imputation. Table A2 provides parameter estimates for model diagnostics reported in the bottom two rows of Table 1.

Table A1. Models with *Neoliberalism* imputed – samples same as CZ

	GDP Growth		Log Inflation		Unemployment		Okun Index		Hanke Index	
	M1	M2	M3	M4	M5	M6	M7	M8	M9	M10
Lagged DV	0.25**	0.24**	0.72**	0.67**	0.81**	0.81**	0.80**	0.80**	0.01	0.03
	(0.04)	(0.04)	(0.03)	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
GET index	0.17	0.22	-0.57	0.36	0.84	-2.01	-0.07	-0.18*	-1.35	-5.07**
	(0.20)	(0.19)	(0.68)	(0.54)	(1.41)	(1.28)	(0.10)	(0.09)	(1.09)	(0.91)
LSCE	0.15		-0.91		18.19**		1.09**		8.58+	
	(0.89)		(3.02)		(6.06)		(0.41)		(4.83)	
GET × LSCE	0.54+		-1.92*		-5.61**		-0.36**		-2.46	
	(0.28)		(0.95)		(1.94)		(0.13)		(1.55)	
Neoliberalism		0.17+		-1.56**		0.05		-0.07		3.63**
		(0.09)		(0.34)		(0.68)		(0.05)		(0.53)
GET × Neoliberalism		-0.08		0.07		-0.35		0.00		-0.96**
		(0.06)		(0.21)		(0.47)		(0.03)		(0.36)
Intercept	1.97**	2.58**	27.00**	27.34**	18.32**	35.68**	2.03**	2.93**	12.20**	21.36**
	(0.62)	(0.66)	(3.42)	(3.43)	(4.58)	(6.06)	(0.37)	(0.47)	(2.84)	(3.63)
N	544	544	544	544	505	505	505	505	448	448
R^2	0.17	0.17	0.65	0.66	0.79	0.79	0.87	0.87	0.34	0.42
BIC	3022.58	3023.76	4354.56	4335.61	4689.71	4703.93	1986.50	1997.45	3863.05	3813.72
RMSE	3.50	3.49	11.90	11.64	22.44	22.65	1.54	1.55	15.92	14.99

Note: Cells report OLS coefficients with standard errors in parentheses. All models include country fixed effects. Column headings report model dependent variable. ** p < 0.01. * p < 0.05

Table A2. Models with *Neoliberalism* not imputed – samples differ from CZ

	GDP Growth		Log Inflation		Unemployment		Okun Index		Hanke Index	
	M1	M2	M3	M4	M5	M6	M7	M8	M9	M10
Lagged DV	0.29**	0.29**	0.70**	0.64**	0.79**	0.79**	0.78**	0.78**	0.01	0.02
	(0.05)	(0.05)	(0.03)	(0.04)	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)	(0.02)
GET index	-0.25	-0.08	-1.27	0.12	3.42	-1.90	0.05	-0.16	-0.66	-6.96**
	(0.39)	(0.39)	(1.41)	(1.37)	(2.76)	(2.62)	(0.19)	(0.18)	(2.24)	(1.91)
LSCE	-0.12		-2.93		17.98*		1.37**		-4.40	
	(1.10)		(4.03)		(7.50)		(0.52)		(6.26)	
GET × LSCE	0.83+		-3.23+		-8.25*		-0.57*		-1.28	
	(0.50)		(1.86)		(3.51)		(0.24)		(2.84)	
Neoliberalism		0.19		-1.95**		0.23		-0.08		4.12**
		(0.13)		(0.51)		(0.90)		(0.06)		(0.70)
GET × Neoliberalism		-0.16		0.56		-0.28		-0.01		-1.23**
		(0.11)		(0.39)		(0.82)		(0.06)		(0.60)
Intercept	1.92*	2.24**	29.40**	28.48**	24.74**	43.06**	2.21**	3.41**	26.00**	24.18**
	(0.83)	(0.76)	(4.49)	(4.08)	(6.40)	(7.42)	(0.47)	(0.57)	(4.20)	(4.37)
N	410	410	410	410	391	391	391	391	348	348
R^2	0.14	0.15	0.64	0.66	0.77	0.77	0.87	0.86	0.33	0.42
BIC	2331.97	2332.51	3394.53	3391.74	3700.55	3712.01	1603.64	1530.49	3077.73	3034.75
RMSE	3.68	3.66	13.45	13.07	24.21	24.42	1.66	1.67	17.54	16.38

Note: Cells report OLS coefficients with standard errors in parentheses. All models include country fixed effects. Column headings report model dependent variable. ** p < 0.01. * p < 0.05

C. Measuring the Policy Regime

We measure policy regimes through a pair of indices tapping the degree to which the economy accords with market fundamentals and state's ability to offset negative market externalities. To start, we rely on Lora's (2012) indices of structural reform in five areas: trade, financial markets, tax reform, private sector ownership, and labor markets. The trade index gauges the removal of tariffs and other trade barriers. The financial liberalization index combines metrics of bank reserve ratios, interest rates, taxes on financial transactions, and bank supervision. The tax index looks at national tax policy legislation. The privatization index taps levels of private investment in infrastructure projects in the transport, telecommunications, energy, and water sectors, and reflects the accumulated value of the privatizations, net of nationalizations, as a percentage of GDP. Lastly, the labor market index gauges flexibility in hiring and firing, social security contributions, and minimum wage levels.

Nearly all Latin American countries made neoliberal reforms in the 1990s, especially in trade and finance, and to lesser degrees in taxes and privatization. In contrast, labor market reforms were fewer, more limited, and often a hodgepodge of pro-market and statist policies—a reality Lora's labor market index obscures. We address this by developing an unambiguously statist measure of reforms to worker welfare based on protections from joblessness, sickness, and old age. Specifically, the worker welfare index combines Lora's index of social security contributions and other taxes and payroll contributions with his measure of the minimum wage. We exclude three sub-indices of Lora's Labor Reform Index – expected costs of firing a worker, hiring flexibility, and flexibility in working hours – from our *Worker Welfare* index for several reasons. For one, they a qualitatively distinct from labor market reforms as social protections. Second, hiring flexibility is measured discretely, complicating index construction. Lastly, Lora's Labor Reform index loads

on neither the neoliberal nor the statist policy regime dimension. To incorporate the state's willingness to offset negative market externalities through a broad range of fiscal interventions, following Kurtz and Brooks (2008) we include data on final government consumption as a share of GDP from the World Bank's World Development Indicators. The measure includes all government expenditures (including payrolls) save military expenditures classified as capital formation.

These six indicators tap distinct orientations of economic policy regimes. We expect trade, finance, taxation, and privatization to mark *orthodox* orientations, and worker welfare and government consumption to reflect *statist* orientations. With annual measures available from the mid-1980s to 2009 for most of Latin America, we use dynamic factor modeling (DFM) to generate summary indexes that tap orientations to *Orthodoxy* and *Statism* across policy regimes.

D. Dynamic Factor Models

Found mainly in macroeconomics and psychology, applications of dynamic factor models in political science are rare. Dynamic techniques are better suited than standard factor-analytic methods to the panel nature of our data (i.e., multiple observed indicators for multiple countries over a series of years). "Dynamic" implies that at every time point, the estimation of each index incorporates information from the entire sample of available data, thus rendering smoothed indices. We estimate the factor loadings with Tripodis and Zirogiannis' (2015) algorithm for two reasons. First, it accounts for how the indicators vary within and between countries over time. Second, it is appropriate for relatively short time dimensions ($T < 50$). Since country-specific estimates would preclude cross-national comparison, we assume the estimated factor loadings do not differ across countries.

E. Applying the Dyads Ratio Algorithm

As noted in the text, we construct measures of presidential approval in countries by applying Stimson's (1991) dyads ratio algorithm to survey marginals (4492 survey marginals from 139 series total). The method assumes that to the extent a given data time series is a valid indictor of presidential approval, the ratio of any two values within the series is a *relative* indicator of approval. The algorithm uses all such dyadic ratios within a given series to estimate approval values at regular time intervals. To combine *N* time series for a country *i* into a single measure, each raw series undergoes this transformation, resulting in *N* dyads-ratio series. If these *N* dyads-ratio series are relative indicators of presidential approval, they should co-vary where they have temporal overlap and this common variance should tap a single latent construct – presidential approval. From this covariance, we compute validity estimates for each of *N* input series to estimate the best single series of latent approval. Exponential smoothing on the resulting series removes random fluctuation due to sampling error and sharpens the estimates.

Evidence suggests this approach is valid and reliable. Following Erikson *et al.* (2002) we calculate approval as $\frac{\%\ positive\ rating}{(\%\ positive\ rating + \%\ negative\ rating)}$. Measuring approval as the total percentage of positive ratings or net approval (positive rating minus negative rating) changes the results reported below very little. By excluding neutral and ambivalent responses this approach eases comparisons across series. On average, a single dimension theorized to be presidential approval, accounts for 87.56% of the variance in our measurement models, with a low of 70.83% (Mexico) and a high of 95.25% (Brazil). Most input series load highly (≥ 0.90) on the latent factor. Meeting conservative criteria for confirmatory factor analysis bolsters our confidence in the validity and reliability of our measures of presidential approval. For the analyses reported in Table A3 in this

SI file, we aggregate series on a monthly level; for the analyses reported in Table 3 in the main text we use quarterly data.

F. Replication and Extension: CZ Table 2

Table A3. Replication of Table 2

Reelection	Model 1 Cl. SE	Model 2 FE	Model 3 RE	Model 4 Cl. SE	Model 5 Cl. SE
GET index	0.152	0.025	0.134	0.103	0.158
	(0.179)	(0.385)	(0.347)	(0.158)	(0.180)
GET index × LSCE	1.000*	1.384**	1.064*	0.893*	1.009*
	(0.392)	(0.597)	(0.523)	(0.375)	(0.405)
Incumbent ran				2.453**	
				(0.778)	
Ideology = right					0.104
					(0.500)
Intercept	-0.806*		-0.855*	-0.981*	-0.904
	(0.402)		(0.417)	(0.405)	(0.624)
LSCE	0.159		0.186	-0.027	0.175
	(0.514)		(0.537)	(0.522)	(0.523)
Countries	18	16	18	18	18
N	106	96	106	106	106
Wald Chisqr	11.73**	17.94	10.29*	22.98**	11.44*
BIC	143.67	181.42	147.77	136.22	148.31

Note: Cells report logit coefficients with standard errors in parentheses. In Model 2, Guatemala and Panama drop out of the sample because they perfectly predict success. ** $p < 0.01$. * $p < 0.05$

Table A4. Replication of Table 2, substituting interpolated version of Neoliberalism for LSCE

Reelection	Model 1 Cl. SE	Model 2 FE	Model 3 RE	Model 4 Cl. SE	Model 5 Cl. SE
GET index	0.657*	0.694*	0.656*	0.422	0.629*
	(0.325)	(0.289)	(0.297)	(0.313)	(0.314)
GET index × Neoliberalism	-0.110	-0.141	-0.114	-0.125	-0.015
	(0.124)	(0.114)	(0.099)	(0.137)	(0.097)
Incumbent ran				2.612**	
				(0.778)	
Ideology = right					-0.324
					(0.475)
Intercept	-0.567+		-0.588*	-0.818*	-0.280
	(0.337)		(0.291)	(0.371)	(0.545)
Neoliberalism	-0.011		-0.004	0.037	-0.015
	(0.092)		(0.100)	(0.124)	(0.097)
Countries	18	16	18	18	18
N	106	96	106	106	106
Wald Chisqr	12.36**	18.61	8.07*	22.23**	14.63**
BIC	147.50	185.54	151.75	137.47	151.84

Note: Cells report logit coefficients with standard errors in parentheses. In Model 2, Guatemala and Panama drop out of the sample because they perfectly predict success. ** $p < 0.01$. * $p < 0.05$

Table A5. Replication of Table 2, substituting non-interpolated version of Neoliberalism for LSCE

Reelection	Model 1 Cl. SE	Model 2 FE	Model 3 RE	Model 4 Cl. SE	Model 5 Cl. SE
GET index	1.134*	1.013*	1.134*	0.895+	1.089*
	(0.525)	(0.448)	(0.442)	(0.487)	(0.509)
GET index × Neoliberalism	-0.199	-0.194	-0.199	-0.269	-0.214
	(0.184)	(0.169)	(0.144)	(0.194)	(0.178)
Incumbent ran				2.295**	
				(0.618)	
Ideology = right					-0.328
					(0.515)
Intercept	-0.520		-0.520+	-0.711+	-0.225
	(0.374)		(0.275)	(0.376)	(0.572)
Neoliberalism	-0.011		-0.088	-0.022	-0.092
	(0.092)		(0.107)	(0.151)	(0.133)
Countries	17	16	17	17	17
N	85	77	85	85	85
Wald Chisqr	6.53*	13.66	7.21*	15.91**	7.09
BIC	122.73	160.33	127.18	118.09	126.94

Note: Cells report logit coefficients with standard errors in parentheses. Panama is missing due to missing data on *Neoliberalism*. In Model 2, Guatemala and Panama drop out of the sample because they perfectly predict success. ** $p < 0.01$. * $p < 0.05$

G. Monthly Time-Series Analyses of GET and Presidential Popularity in 16 Countries

Table A6

i. LCSEs

	ARG	BRA	CHL	COL	ECU	PER	URY	VEN
GET Index	0.11	***1.51***	***-0.52***	***2.26***	***1.84***	***1.73***	0.37	-0.12
	(0.28)	***(0.57)***	***(0.25)***	***(0.60)***	***(0.52)***	***(0.50)***	(0.23)	(0.10)
Lag popularity	0.96	0.91	0.92	0.89	0.86	0.81	0.92	0.70
	(0.02)	(0.02)	(0.02)	(0.02)	(0.03)	(0.04)	(0.02)	(0.04)
Intercept	2.32	4.07	4.76	-2.77	7.15	3.68	4.37	14.89
	(0.91)	(1.14)	(1.35)	(0.93)	(1.59)	(1.12)	(1.25)	(1.99)
R2	0.92	0.91	0.92	0.99	0.87	0.84	0.89	0.51
N (months)	322	322	282	110	305	149	322	322
ADF	-2.7	-2.59	-2.72	-4.72	-2.86	-2.33	-2.84	-7.34
p-val	<.01	0.01	<.01	<.01	<.01	0.01	<.01	<.01
Br-Godfrey	7.42	4.38	0.83	4.45	5.87	5.63	3.79	4.91
p-val	0.12	0.36	0.95	0.34	0.21	0.23	0.44	0.30

ii. Non LSCEs

	CRI	DOM	SLV	GTM	HND	MEX	PAN	PRY
GET Index	-0.24	***1.19***	0.04	0.23	-0.46	-0.14	-0.09	***2.40***
	(0.30)	***(0.57)***	(0.31)	(0.32)	(0.35)	(0.15)	(0.98)	***(0.84)***
Lag popularity	0.96	-0.08	0.92	0.93	0.92	0.77	0.90	0.80
	(0.02)	(0.09)	(0.02)	(0.02)	(0.02)	(0.06)	(0.04)	(0.04)
Intercept	2.38	-2.04	5.42	3.96	4.78	6.62	6.19	9.41
	(1.00)	(1.03)	(1.43)	(1.21)	(1.37)	(1.88)	(2.62)	(2.20)
R2	0.92	0.05	0.86	0.86	0.87	0.80	0.82	0.76
N (months)	322	108	322	313	322	296	124	204
ADF	-2.62	-2.45	-4.02	-3.04	-3.4	-3.29	-2.96	-3.34
p-val	<.01	<.01	<.01	<.01	<.01	<0.01	<.01	<.01
Br-Godfrey	4.51	0.27	6.99	8.69	1.8	19.41	4.43	1.39
p-val	0.34	0.99	0.14	0.07	0.77	<0.01	0.35	0.85

Note: The dependent variable in each column is the presidential approval for the country noted. Cells report parameter estimates with standard errors in parentheses. Estimates reported in bold-italicized font indicate $p \leq .05$, two-tailed test. All series assessed for stationarity prior to estimation. In cases of non-stationarity (present for Chile, Colombia, Dominican Republic, Mexico, Nicaragua, Peru, Uruguay, and Venezuela), series were pre-filtered using ARFIMA modeling techniques. The model for Mexico includes a lagged endogenous variable at t-2 (not shown).

H. Replication of Table 3. Time Series Analysis for the Effect of GET on Popularity

Table A7. Brazil

	M1	M2	M3	M4	M5
Lag popularity	0.863	0.879	0.856	0.855	0.875
	(0.030)	(0.028)	(0.030)	(0.030)	(0.029)
GET Index	2.163	1.796	2.157	2.189	1.850
	(0.584)	(0.546)	(0.586)	(0.586)	(0.556)
Intercept	3.817	3.398	3.986	3.999	3.519
	(2.011)	(1.918)	(2.017)	(2.043)	(1.893)
Pollsters Indicators	Yes	Yes	Yes	Yes	Yes
N (months)	310	310	310	310	310
R2	0.908	0.913	0.904	0.902	0.916
Augmented Dickey-Fuller	-12.711	-12.892	-13.369	-13.231	-12.647
p-value	<.01	<.01	<.01	<.01	<.01
Box-Pierce	1.214	0.003	2.411	0.015	5.628
p-value	0.271	0.956	0.121	0.904	0.018
Breusch-Godfrey	1.693	0.004	3.478	0.021	7.797
p-value	0.193	0.949	0.062	0.885	0.005

Note: Table reports estimates from linear regression models with standard errors in parentheses. Models are estimated on five different imputed datasets.

Table A8. Mexico

	M1	M2	M3	M4	M5
GET Index	-2.644	-2.559	-2.199	-4.028	-1.767
	(3.022)	(3.238)	(3.079)	(2.888)	(3.113)
MA1	-0.507	-0.517	-0.538	-0.535	-0.525
	(0.050)	(0.049)	(0.050)	(0.050)	(0.051)
AR1	0.958	0.957	0.960	0.958	0.959
	(0.022)	(0.023)	(0.022)	(0.022)	(0.022)
Intercept	60.947	61.301	60.235	60.972	60.189
	(4.774)	(4.950)	(4.769)	(4.655)	(5.019)
Pollsters Indicators	Yes	Yes	Yes	Yes	Yes
No. Obs.	287	287	287	287	287
Log Likelihood	-900.097	-912.927	-904.593	-908.616	-908.243
Augmented Dickey-Fuller	-11.560	-12.058	-11.959	-12.203	-11.715
p-value	<.01	<.01	<.01	<.01	<.01
Box-Pierce	0.003	0.454	0.017	0.178	0.003
p-value	0.955	0.500	0.898	0.673	0.960
Breusch-Godfrey	2.850	3.135	3.192	2.5365	2.508

Note: Table reports estimates from linear regression models with standard errors in parentheses. Models are estimated on five different imputed datasets.

I. Cases in Analyses of Presidential Approval

Table A9. Cases Included in the Analyses for Table 3

Country	Years	No. Quarters
Argentina	1990-2009	79
Bolivia	1999-2009	42
Brazil	1990-2009	77
Chile	1990-2009	74
Colombia	1994-2009	60
Costa Rica	1990-2009	79
Dominican Republic	2004-2009	23
Ecuador	1990-2009	79
El Salvador	1990-2009	79
Guatemala	1990-2009	79
Honduras	1994-2009	61
Mexico	1990-2009	79
Nicaragua	1991-2009	73
Paraguay	1999-2009	42
Peru	1997-2009	48
Uruguay	1990-2009	77
Venezuela	1999-2009	42
Total N		1093

Note: differences in coverage across countries due to missing data for *Approval*, *Neoliberalism*, or both.

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