The World Economy, Political Control, and Presidential Success

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Abstract

In a recent *Journal of Politics* article, Daniela Campello and Cesar Zucco argue that Latin American voters credit and blame presidents for economic outcomes unambiguously exogenous to their policy choices, a claim which undermines broadly-held understandings of how voters hold politicians accountable. While we concur on the importance of the global economy to politics in the region, we challenge their study on two grounds. First, we question their designation of economies as fitting a low-savings commodity-exporting (LSCE) profile as the key factor shaping the influence of world markets on presidential success. Second, we contend that the impact of the world economy on domestic politics depends instead on the degree of control exercised by national policymakers. In particular, the region's unprecedented experiment with economic policy regimes informs voters' assignment of responsibility to policymakers. A series of analyses provides support for our argument.

Key words: economic reforms, responsibility, accountability, selection model, Latin America

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Supplementary material for this article is available in the appendix in the online edition. Replication files are available in the JOP Data Archive on Dataverse (https://dataverse.harvard.edu/dataverse/jop).

What drives presidential success – in public standing and at the polls – is a venerable question. In a 2016 article in *The Journal of Politics*, Daniela Campello and Cesar Zucco (CZ) test a model of rational voting in Latin America. Unlike in Europe's highly integrated economies, CZ reason that in many Latin American contexts voters lack sufficient information about their country's engagement in world markets to benchmark domestic economic outcomes and, thus, are unlikely to "distinguish competence from chance in their assessment of governments' performance" (591). Marshalling an array of evidence in support of this hypothesis, the authors infer the success of many Latin American presidents derives not from their competence as policymakers but from global economic shocks "unambiguously exogenous" to it (589). CZ's provocative conclusion circumscribes the generalizability of retrospective voting as a mechanism for popular control of elected agents and undercuts mounting evidence of economic accountability in Latin America.

This rejoinder offers a critical but constructive response. We agree with CZ that Latin American voters operate in very different contexts and that classic theories of political behavior best travel to the region when we take these differences seriously (Carlin, Singer and Zechmeister 2015). Yet we disagree with CZ about which regional differences are essential for testing the model of rational retrospective voting. Rather than jettisoning the model for Latin America, we propose an alternative formulation of it.

Like CZ, we assume voters punish and reward governments "exclusively for economic outcomes they *control*" (2016, 591, emphasis added). This requires voters to extract the anticipated from the unanticipated components of economic performance when evaluating policymaker performance. CZ maintain that the relative impacts of both elected and non-elected economically-relevant actors influence this signal extraction problem (cf. Duch and Stevenson 2008). Whereas this version of the model of rational retrospective voting gains much of its empirical purchase from

variation in the impact of *non-elected* policymakers, our revised model emphasizes variation in the policies espoused by national *elected* policymakers. Non-elected actors (international lending agencies, transnationally-mobile investors, etc.) have certainly influenced Latin American economies and, in turn, public perceptions. Yet we argue elected policymakers also shape accountability in the region: the greater their control of policy levers, the more easily voters can assign responsibility for economic outcomes.

Following work on the politics of economic policy (Frieden 1991), we posit the most obvious and relevant signal Latin American voters can extract concerns not their leaders' managerial competence but their control.[1] When economic policies tilt in a statist direction, incumbent control over the economy increases vis-à-vis global forces, simplifying the task of attributing responsibility; when policy skews neoliberal, incumbent control over the economy weakens vis-à-vis global forces, complicating responsibility attribution. If true, Latin Americans need not evaluate presidents on luck-of-the-draw world economic conditions because unprecedented economic policy experimentation in the region since 1980 informs voters about presidents' policymaking control and enables them to judge economic responsibility on that basis. If, upon assessment, Latin American voters fail to use control signals to discount global economic conditions, it would validate CZ's pessimistic conclusions. If not, then a reformulated version of the model – grounded in control rather than competence – may very well extend to the region.

To this theoretical innovation we add a measurement reappraisal. CZ classify some Latin American economies as "low-savings commodity-exporting" (LSCE). They find that the world economy influences incumbent evaluations in countries which match that static LSCE designation.

[1] The importance of policy control to presidential success has long been recognized (Tufte 1978).

Our re-analysis, however, reveals over-time heterogeneity that highlights the dynamism of economic policy across *and within* LSCE and non-LSCE countries. Thus informed, we construct a dynamic indicator of incumbent economic control based on policy regimes. When we employ it to re-analyze presidential success, we find voters do not simply reward presidents for favorable global conditions in 10 Latin American LSCE countries. Instead, voters throughout the region use information about world economic conditions in conjunction with information about their country's economic policy regime to assign responsibility for economic outcomes to the president. More statist policy regimes signal that national policymakers are in control and, thus, lead voters to assign responsibility to them. Neoliberal policies, on the other hand, signal a weaker degree of political control over the economy, be it global or domestic.

CZ are correct that the standard competence-based model of retrospective voting is not applicable to Latin America but jettison it in exchange for a blindly retrospective view of voter decision-making (Achen and Bartels 2016). We, however, propose a reformulated model, retooled with a mechanism linking economic policy to signals of incumbent control rather than signals of incumbent competence. Though applied to a specific setting, our findings engage larger debates about the roles of voter knowledge and policymaker capacity for democratic accountability.

Retrospective Voting, Exogenous Forces, and the Competency Signal

CZ conceptualize rational retrospective voting according to the selection model introduced by Alesina and Rosenthal (1995) and elaborated by Duch and Stevenson (2008). The model assumes economic performance influences incumbent support only to the extent that voters can extract an incumbent "competency signal" from the din of "exogenous shocks." So voters first identify and

then discount those components of their country's economic performance exogenous to the incumbent's policymaking competence based on the discount rate $\frac{\sigma_\mu^2}{\sigma_\mu^2+\sigma_\xi^2}$, where σ_μ^2 indicates the variance in the incumbent's competency shock (from a normal distribution of policymaker decisions) and σ_ξ^2 is the variance in the exogenous shock (from a normal distribution of external factors beyond the incumbent's control). The discount rate "controls how much information about incumbent competence voters can extract from observed movements in the economy" (Duch and Stevenson 2008, 137).[2]

Compared to Europe's highly-integrated context (where the model was substantiated), CZ argue extracting competency signals is far more difficult in Latin America given its inward-looking models of economic development, limited degrees of political and economic integration, and low access to information. Therefore, they reason Latin American voters cannot "distinguish competence from chance when assessing government's management of the economy" (591). So rather than voters assessing the (in)actions of national policymakers, CZ assert that presidential success is driven by the world economy. They operationalize global "good economic times" (GET) with an index combining international commodity prices and international interest rates. They hypothesize this index predicts presidential success in countries whose global economic orientation can be characterized as "low-savings-commodity-exporting" ("LSCE"). "Low-savings" countries' reliance on international capital highly exposes them to fluctuations in international interest rates. "Commodity-exporting" economies depend on prices set in international markets.

CZ test their hypothesis in three steps. Their first analysis suggests GET affects domestic economic performance in Latin America's LSCE countries only. A second analysis links GET to

[2] Duch and Stevenson (2008, 131-138) provide a formal presentation of the competency signal.

presidential re-election success but, again, only in 10 LSCE cases. A third shows GET predicts presidential popularity in Brazil, which is classified LSCE, but not in Mexico, which is not. Rather than discounting information about the world economy, Latin Americans in LSCE contexts appear to rely on it to sanction or reward politicians. Hence, CZ conclude the selection model's predictions fail in Latin America because voters are ill-informed and receive weak signals about national economy's relative performance. The upshot is that many voters "do not separate chance from competence when evaluating their presidents" (600). If they are correct, then Latin American presidents had better be "lucky" (ride favorable exogenous shocks) because it does not pay to be "good" (pursue credible, competent economic policies).

Incumbent Competency and the Importance of the Policy Regime

The world economy certainly influences most Latin American economies (e.g., Baker 2009; Alcañiz and Hellwig 2011), warranting CZ's emphasis on "good economic times." Missing from CZ's study, however, is the role of policy choice in shaping presidential success. We argue that policy choices signal presidents' control which voters, in turn, use to discount economic performance and attribute responsibility. Duch and Stevenson (2008, 133) foreshadow this point, noting the unobserved economic impact of the incumbent's behavior is "not constant over time or administration." Moreover, they expect the competency signal to weaken as a greater share of the decisions shaping economic conditions reflects the actions of non-elected, as opposed to elected, leaders. CZ recognize this but nevertheless reject the application of the competency model to Latin America out of hand. This decision ignores the potential for Latin American presidents to shape the competence shock (represented by σ_{μ}^2), even in the face of a large exogenous shock (σ_{ξ}^2).

In their supplementary analyses, CZ consider the influence of competence by way of "merit" and find it does not affect presidential reelection. Their merit measure, however, suffers from two problems. First, it represents the residuals from a model that regresses economic growth on GET. Their rationale is that any fluctuations in growth unaccounted for by GET can be assigned to policymaker competence. This assumption is untenable; many factors driving growth both lie outside of GET and are separate from policymaker actions (e.g., entrepreneurialism, job creation by local businesses, legal decisions, or policies imposed by agreements with international lenders). Second, "merit" as constructed does not capture *policy* orientations, which we maintain are the most crucial component of variation signaled to voters.

We call for a reconsideration of the competency signal. The competency signal is the product of the signal from the observed economy and the variance of the distribution of competence shocks associated with decisions made by elected policymakers.[3] Expanding the number, or range, of choices at the policymaker's disposal with respect to the economy should therefore increase this product. As Duch and Stevenson (2008, 142) put it, "this variance should be larger in countries in which [elected decision-makers] make more of the economic decisions." Any assessment of responsibility thus requires a consideration of variation in the exogenous shock—i.e., actions taken by decision-makers not dependent on the electorate—*relative* to variation in the distribution of competence shocks attributed to the elected policymaker.

Thus, rather than rule out a competence-driven model for Latin America, we revise its empirical realization with two alterations, summarized here and detailed in the SI. First, whereas previous empirical assessments of the competency model emphasize variation in the distribution of

[3] Duch and Stevenson's (2008) shorthand is "electorally dependent decision makers" (EDDs).

shocks from *non*-electorally dependent decision makers, we focus on the actions of *elected* decision makers, namely executives. Second, we expand the set of decisions made by elected decision makers to include policy. The model, as developed by Alesina and Rosenthal (1995), assigns incumbent action to competence alone. For contemporary Latin America, however, this assumption is too restrictive. In the region, variation in executive action is chiefly signaled to the electorate not by how well policies are executed but through shifting the orientation of policy to address changing internal and external conditions (e.g. Stokes 2001).

Elaborating on this second point, Latin American executives have confronted the exhaustion of import-substituting industrialization policies and, in many places, societal backlash against neoliberal orthodoxy. Reform is the norm rather than the exception in much of the region (Kurtz and Brooks 2008; Lora 2012; Morley, Machado and Pettinato 1999). Presidents must also consider how their peripheral status in the world economy impacts domestic outcomes. Whereas policymakers in advanced capitalist economies target inflation rates consistent with their domestic economic goals, in Latin America they must calibrate an inflation target to align their own domestic goals *conditional* on the capacity of policymakers in the developed world to achieve *their* stated economic goals. And per our first point above, with complex environments come a multiplicity of possibilities: the region's residents may, and indeed have, embraced a wide range of policy regimes. Presidents who successfully navigate these complexities become national heroes; those who fail are driven from office (cf. O'Donnell 1994, Pérez-Liñán 2007).

Incumbents in peripheral economies demonstrate their competence to the extent they can buffer the domestic economy from pernicious fluctuations in world economic conditions and take advantage of the opportunities they present. Hence the mixture of policies and reforms Latin American presidents adopt—their economic *policy regime*—cannot be excluded from retrospective

models of political accountability. If we are correct, incumbent policymaking decisions—captured for contemporary Latin America by choice of policy regime along the statist-orthodox continuum—should condition the world economy's impact both on domestic economic outputs and on presidential success.[4]

Stable Economic Traits or Dynamic Economic Policy Orientations?

The concepts of policymaker control and choice lie at the heart of our conception of accountability in Latin America's political economies. This contrasts with CZ's perspective, which assigns primacy to whether and how national economies are integrated into world markets. Any support CZ find for their theoretical propositions hinges on their designation of 10 of the 18 Latin American economies as "LSCE." Theoretically, CZ justify the LSCE designation by arguing that GET's influence depends on the country's "mode of insertion into the world economy" (592). Operationally, CZ inductively classify countries as LSCE (or not) by plotting their share of commodity exports against their debt service as a share of total exports averaged 1980-2012 (594, Figure 2). By this rationale, voters evaluate their presidents based on global economic information in the 10 Latin American economies deemed sufficiently exposed to the world economy.

According to the authors' static binary operationalization, a country like Brazil is *always* LSCE and a country like Mexico *never* qualifies as LSCE. This ignores massive variation in policy regimes and global economic insertion over this 32-year period in LSCE's and non-LSCEs (Frieden 1991; Kurtz and Brooks 2008; Lora 2012; Morley, Machado and Pettinato 1999). Figure 1 displays box plots of the commodity exports (top) and debt service (bottom) variables from 1980 (or year

[4] We substantiate this argument in the SI.

of first democratic election thereafter) to 2012. For some economies, such as Colombia, mean values for both measures are informative. Others, however, exhibit considerable over-time variation on commodities (e.g., Paraguay), debt service (e.g., Brazil), or both (e.g., Mexico).

[FIGURE 1 HERE]

We argue that long-run patterns of savings and exports convey less about the economy's exposure to global economic forces under a given president, let alone his or her policymaking competence, than the orientation of the economic policy regime does. And this is crucial to resolving voters' signal-extraction problem. By engaging in activist economic policies, policymakers provide voters with information to link the observed economy to the president's performance, just as rational retrospective voting models predict. The deeper incumbents intervene in markets, the more responsibility for outcomes they receive. Alternatively, when incumbents opt for a more market-oriented strategy, they signal their removal from highly discretionary policy decisions. Voters, in turn, update their responsibility attributions and increasingly de-link incumbent evaluations from the economy, be they domestic outcomes (as standard models of economic voting posit), world conditions (as CZ claim), or some combination (e.g., Kayser and Peress 2012).

The upshot is that even in a region where signals about the workings of the domestic economy are relatively weak, voters can still separate policy choice from chance and evaluate presidents accordingly. Motivated to incorporate incumbent policy decisions into the selection model of rational retrospective voting, we now construct a dynamic indicator of policy regime orientation in Latin America countries. Then we use it to reassess the linkages between economic conditions and presidential success.

Since the 1980s economic policy regimes in Latin America have ranged from ISI statism to

orthodox neoliberalism and myriad forms of heterodoxy in between. The exhaustion of statist import-substitution development models spawned economic policy change following the 1980s debt crisis. After short-term stabilization packages largely failed, Latin American presidents implemented neoliberal structural reforms unilaterally or via the mechanism of loan conditionality. By reducing the state's role, neoliberal orthodoxy made interventionist policies such as price and capital controls, credit subsidies, and social insurance programs untenable. However, policy change was uneven across sectors and countries. Liberalization in trade, foreign investment, and privatization was aggressive but timid in taxation and labor (Lora 2012). And many leaders embraced heterodox strategies combining economic openness *and* state production promotion, welfare protections, investments in human capital formation, countercyclical spending, and supply-side interventions (Avelino *et al*. 2005).

To approximate these nuanced policy regimes, we rely on Lora's (2012) indices of structural reform in five areas: trade, financial markets, tax reform, private sector ownership, and labor markets.[5] Following Kurtz and Brooks (2008), we incorporate the state's willingness to offset negative market externalities through fiscal interventions by including a measure of final government consumption as a share of GDP from the World Bank. Together, these indicators capture distinct orientations of economic policy regimes. We expect trade, finance, taxation, and privatization to tap

[5] Labor market reforms often mixed pro-market and statist policies—a reality Lora's labor market index obscures by combining social security contributions and other taxes and payroll contributions with a measure of the minimum wage. We address this by developing an unambiguously statist measure of reforms that increase economic control of labor markets via interventions that alter wage equilibria by offering protections from joblessness, sickness, and old age.

market orthodoxy, and worker welfare and government consumption to reflect statist policies. With annual measures available from the mid-1980s to 2009 for most of Latin America, we use dynamic factor modeling to generate summary indexes tapping degrees of neoliberal and statist policy orientations (see SI for details).

[FIGURE 2 HERE]

Figure 2 displays our estimates of policy orientation over time. Trends in *Orthodoxy* (long dashed lines) show a region-wide pro-market shift with respect to taxes, trade, finance, and regulation. However, *Statism* trends (short dashed lines) in government spending and worker protections—two areas in which Latin American policymakers retained or increased influence (cf. Kurtz and Brooks 2008)—are not uniform. To produce a more general orientation of economic policy, we subtract *Statism* from *Orthodoxy*. The resulting series, which we label *Neoliberalism*, appears as the dark solid lines in the figure. *Neoliberalism's* peaks and valleys are most obvious in emblematic cases such as Bolivia, Ecuador, Nicaragua, Peru, and Venezuela where market-tilting reforms provoked instability. Smoother liberalizing trajectories are found in Chile, Colombia, Costa Rica, Honduras, Mexico, and Uruguay.

With this dynamic indicator in hand, we perform a series of analyses. Following CZ, we first model domestic economic outcomes as a function of GET. We find the conditional effects of executive policy control rival those of the LSCE classification on GET's impact on domestic economic indicators. We proceed to replicate CZ's analysis of incumbent candidate reelection modeled as a function of GET and LSCE. Extending this, we demonstrate how policy choices and the domestic economy provide a more plausible set of results. Lastly, we report findings from time series analyses of presidential approval from countries across the region.

The Effect of the World Economy on Domestic Economies

CZ first model the relationship between *GET* and the domestic economic outcomes conditional on *LSCE* status using annual panel data, country fixed effects, and lagged dependent variables. They regress five domestic economic variables—growth, inflation, unemployment, and two economic discomfort indices—on *GET*, an *LSCE* dummy, and their interaction. For four of five domestic economic indicators *GET* wields a greater influence in the LSCE sample. They springboard from this conditional relationship to their analyses of presidential success.

[TABLE 1 HERE]

In similar motivating spirit, we first replicate CZ's models (reported in the SI). We then re-estimate the same model specification substituting *Neoliberalism* for *GET* on the sample of country-years for which we have policy regime data (see Figure 2). Rather than comparing effect sizes, our interest here pertains to which measure—*LSCE* or *Neoliberalism*—best fits the data.[6] Thus Table 1 compares model results using the Bayesian Information Criterion (BIC) and residual mean square error (RMSE), with full models reported in the SI. Lower BICs suggest *Neoliberalism* outperforms *LSCE* in models for inflation and the Hanke index. For models which specify unemployment and the Okun index (which simply sums unemployment and inflation), *LSCE* better fits the data. Given the nature of the policy tools at governments' disposal, it is not surprising that our policy measure fares better for the growth and prices models, whereas the LSCE designator matters relatively more for jobs.[7] For the *Growth* model, results are ambiguous. We can conclude that the

[6] *LSCE* is a dummy variable; *Neoliberalism* is continuous.

[7] See note 6 on the weakness of labor market reforms in the region.

mix of national policies moderates the effects of the international economy on the domestic economy. While modest, this finding provides a warrant to proceed to our central task of assessing how policy choices affect the translation of economic conditions into presidential success.

The World Economy and Presidential Elections

CZ argue that Latin American voters are less likely than their European counterparts to distinguish competence from chance when assigning the incumbent responsibility for economic outcomes. They hypothesize that "voters do not discount exogenous factors when evaluating presidents by examining whether factors beyond incumbents' control—but that have a great deal of influence on domestic economies—determine presidential success" (593). They test this hypothesis by examining presidential reelection in 121 elections in 18 Latin American countries. Their results suggest exogenous factors affect the probability of reelection – but only in the LSCEs, where world economic factors have "a great deal of influence" (*ibid*).

Our argument instead emphasizes the role of the president's control over policy decisions as captured by the policy regime. By engaging in activist, statist policies, presidents provide voters with information to link outcomes to incumbent actions. Voters thus evaluate governments based on the observed economy. This case accords with predictions of the selection model of rational retrospective voting. But under more market-oriented regimes, policymaker discretion is reduced and citizens rely less on the economy as a gauge of presidential success. We test this claim with a re-analysis of CZ's reelection data and an original analysis of presidential approval throughout Latin America.

We reanalyze CZ's models of candidate reelection (Table 2, 598). After replicating the results from their Table 2 (reported in the SI file), we select CZ's most comprehensive specification

(Model 4) and report our estimates in Table 2, Model 1, below.[8] Informed by the variation displayed in Figure 1, we then re-estimate this model substituting a continuous measure of LSCE for CZ's binary and time-invariant one, *LSCE continuum* (Model 2).[9] In Model 3 we condition *GET* on our preferred measure, *Neoliberalism*, in order to gauge the influence of the policy regime on presidential success. Model 4 includes both CZ's *LSCE* and our *Neoliberalism* measure. Lastly, Model 5 restricts the estimates to CZ's LSCEs.[10] The negatively-signed coefficient on *GET index* × *Neoliberalism* suggests the impact of international commodity prices and US interest rates on reelection weakens under more neoliberal policy regimes.

[TABLE 2 & FIGURE 3 HERE]

Care must be exercised when interpreting interaction results. The nonlinear and unconditional nature of the logit coefficient estimate means that inferences about the effects on the bounded 0-1 dependent variable cannot be inferred from significance levels (Ai and Norton, 2003). To assess our claims more reliably, we graph re-election probabilities as a function of model covariates.

Figure 3A uses estimates from Model 2 to chart the impact of *GET* on the probability of reelection as the continuous LSCE measure ranges across its in-sample values. While there is some indication that the positive influence of the world economy on presidential reelection success is

[8] As Table A3 in the SI shows, CZ's Model 4 provides the best overall fit among their specifications, based on pseudo R^2 and BIC diagnostics.

[9] We construct *LSCE continuum* by creating z-scores for the commodity exports and debt service summarized in Figure 1 and then averaging the two measures.

[10] With only seven countries in the data set (Costa Rica, Dominican Republic, El Salvador, Guatemala, Honduras, Mexico, Paraguay) we are unable to estimate a model on non-LSCEs only.

stronger when countries approximate LSCEs, this difference is not statistically significant. The next two graphs use Model 4 estimates. Figure 3B compares effects using CZ's time-invariant LSCE indicator. Consistent with CZ, the international economy influences presidential reelection success in LSCEs but not otherwise. Figure 3C illustrates how policy regimes shape the political effects of the world economy. Consistent with our argument that voters use information about policy control when evaluating political incumbents, the graph shows economic conditions matter only in cases where the policy orientation is sufficiently activist, i.e. statist. Finally, Figure 3D graphs the effects using Model 5 estimates and reveals the impact of policy regime is even stronger when we consider LSCEs alone.[11] This is testimony to the importance of policy-based variation *within* these ostensibly similarly endowed national economies.

In sum, we find that *Neoliberalism* conditions the world economy's impact on electoral prospects as much as, if not more than, *LSCE*. These re-analyses call into question the singular impact of LSCE status on elections. An economy's reliance on the world economy is not the only or even most critical factor connecting global economic conditions to national politics. Policy orientation—which can and does vary over time and space, signaling incumbent economic control—also matters.

The World Economy and Presidential Popularity

To further test the relationship between the world economy and the president's public standing, CZ estimate time-series models of monthly presidential approval in Brazil, classified as LSCE, and in Mexico, not classified as LSCE. They show that fluctuations in *GET* influence presidential

[11] The smaller sample size in Model 4 produces higher levels of uncertainty about the estimates.

popularity in Brazil but not in Mexico. The result is taken as further evidence that the world economy drives presidents' fortunes but only in economies that rely on global demand for commodities and investment inflows.

The popularity function analyses provide the strongest test of CZ's argument since they permit assessments of public sentiment at regular intervals over long periods rather than isolated at election times. Analyses of popularity are also less affected by confounding institutional factors, such as the separation of powers and the electoral system, which enter voters' calculi. But CZ's analysis is limited in least three respects. First, it is unclear how well results from two countries generalize across the region. Second, their research design is time-invariant with respect to economic policy orientation. Evidence of over-time variation notwithstanding (see Figure 1 above), this assumes that Brazil rates as an LSCE for the duration of the 25-year period under analysis (1987-2012) and that Mexico similarly does not.[12] Third, by ignoring policy regimes, CZ's analysis holds constant the amount of economically relevant actions taken by electorally-dependent decision makers.

We address these shortcomings in two ways. As an initial check we extend CZ's analysis of monthly presidential approval ratings beyond Brazil and Mexico to 16 countries: 8 LSCEs and 8 non-LSCEs.[13] Data come from the Executive Approval Project (Carlin et al. 2016) and are described in the SI. To assess the influence of *GET* on popularity, we estimate a series of partial

[12] CZ's Figure 2 calls this assumption into question for Brazil by implying its exposure to commodity prices (vertical axis) are actually below the regional mean. Of course, during certain years this gauge of exposure to the world economy may be even lower than that reported in the figure.

[13] We lack a sufficient number of survey marginals to construct monthly popularity series for Bolivia and Nicaragua.

adjustment autoregressive distributed lag models, or what CZ refer to as LagDV models. Model estimates show that *GET* carries a positive and statistically significant coefficient in six of the 16 models (see SI Table A6). This includes four LSCEs: Brazil, Colombia, Ecuador, and Peru but also two non-LSCEs, the Dominican Republic and Paraguay. The effect of *GET* is insignificant in all other cases except Chile (an LSCE), where *GET*'s influence on presidential approval is *negative*. Collectively, these single-country regressions further imply that LSCE or non-LSCE designation is not the critical factor linking economic conditions to public evaluations of the president.

Second, we consider our alternative explanation—that the influence of global economic conditions on approval depends on signals that include policy control—by modeling presidential approval as a function of *GET*, *LSCE*, and *Neoliberalism*. In order to place the policy regime data on a common temporal scale and to capture cross-national differences, we employ data at quarterly (rather than monthly) intervals. Such data are available for 17 countries over a 20-year period from the Executive Approval Database (Carlin *et al.* 2016).[14] Models control for several other factors which might shape approval. These include GDP per capita growth (*Growth*); a series coded 1 for quarters following elections of new executives to account for elevated approval at the beginning of their terms (*Honeymoon*); a series of interventions for scandals involving the president (*Scandal*); an index of political *Corruption* calculated from country expert surveys of corruption perceptions in various aspects of the government (Coppedge et al. 2016); and a measure for international and domestic *Terrorist Attacks* per quarter.[15]

[14] As noted, Lora's indices exclude Panama (non-LSCE).

[15] *Growth* is from the World Bank's World Development Indicators, *Scandal* is from Pérez-Liñán (2007) and updated in Carlin *et al.* (2015); *Corruption* is from the Varieties of Democracy project; and *Terrorist attacks* is from the Global Terrorism Database (2017).

We estimate a set of partial adjustment autoregressive distributed lag models. Lagged dependent variables account for potential time dependence within panels and, as such, resemble fixed-effects models in implying that the parameters on exogenous predictors represent within-country effects (Beck and Katz 2011). Models include country fixed effects to account for unobserved unit heterogeneity. And to address differences in the persistence of error processes across countries, we include panel-specific auto-regressive terms. Panel unit-root tests reveal no indication of non-stationarity (Im *et al.* 2003). Panel-corrected standard errors address any heteroskedasticity or contemporaneously correlated disturbances induced by the panel data structure.[16]

[TABLE 3 & FIGURE 4 HERE]

Table 3 reports results. Model 1 shows that approval increases when growth increases and during honeymoons, declines during times of scandal, and responds positively to *GET*. Model 2 assesses whether *GET*'s influence depends on a country's LSCE classification. In LSCEs the estimate on *GET* is 1.37 with standard error 0.74; for the comparison group the coefficient on *GET* is 0.86 with standard error 0.70. Model 3 replaces CZ's binary LSCE measure with a continuous LSCE measure. Neither model uncovers a difference based on country orientation toward the world economy. Model 4 conditions *GET* on *Neoliberalism*. Here policy regime orientation does

[16] Regarding dynamics, the first lag of the dependent variable accounts for the series' autoregressive properties. Since models estimated with an additional lag are substantively identical, and the coefficient on the additional lag is insignificant, estimates in Table 3 are unbiased by residual serial autocorrelation. Regarding cross-unit heterogeneity, debates persist about pairing fixed effects with lagged dependent variables (Nickell 1981; Plümper and Troeger 2007). However, as *T* increases, the bias fixed effects induce is minimal.

indeed condition whether publics evaluate presidents based on world economic conditions. This finding persists when we reintroduce LSCE in Model 5. From these estimates we graph the marginal effect of *GET* across values of *Neoliberalism*. The resulting Figure 4 shows the performance of world markets only influences presidential approval in the more statist cases, where policymakers possess sufficient control over the levers of policy.

[FIGURE 5 HERE]

Figure 5 shows that this effect is independent of LSCE status by leveraging the time-series nature of the data to forecast the longer-term effects of policy reforms on presidential approval. For both graphs, forecasts depict two states of the world: in one, policies are highly neoliberal, in another they approach statism. Using Model 5 estimates, we then simulate the influence of a positive one-standard deviation shock to *GET* on *Approval* for LSCEs (panel A) and non-LSCEs (panel B).[17] The same pattern prevails for both: a booming world economy drives up presidential success but the magnitude of this effect is governed by the policy regime—stronger in more statist regimes and weaker in cases closer to market orthodoxy. Model 6 in Table 3 underlines this conclusion, showing that *Neoliberalism* matters for *GET*'s impact *even within* the subset of LSCEs.

Conclusion

"Presidential Success and the World Economy" (Campello and Zucco 2016) is intriguing. By linking presidential success ultimately to the performance of world markets, the article casts a long

[17] We set *Approval* at time *t* to 50%; *Growth* to its mean; *Honeymoon*, *Scandal*, and *Terrorist attack* at their modes (0). Simulations are performed using the dynsim package in STATA (Williams and Whitten 2011).

shadow over work on the economy and political accountability in Latin America. Like all good research, it raises as many questions as it addresses. It challenges us to consider just how global economic conditions influence economic accountability. Unlike previous research on rational retrospective voting, CZ suggest voters don't discount. Instead they "get it wrong" by attributing responsibility for exogenous economic shocks to national politicians. The argument and evidence presented in this rejoinder, however, imply CZ's pessimism may be overstated. World economy conditions do shape presidential success, but so do the actions of elected policymakers. This is true even in LSCEs, countries CZ deem especially susceptible to global economic fluctuations.

By considering not only the range of exogenous shocks, as CZ do, but also the set of decisions contributing to the range of incumbent control over policy, our study provides an important corrective to CZ's provocative conclusions. Namely, it suggests elected policymakers' actions matter as long as they possess a degree of control over the levers of policy and can, thus, be credited for capitalizing on strong global conditions. When control is high, as in regimes with a statist mix of policies, policymakers can harness economic conditions and voters hold them to accounts. But when control is low, as in regimes toward the market orthodoxy end of the spectrum, policymaker control is weak, and voters thus decouple their political evaluations from the economic conditions they observe.

A wide range of cross-national evidence—from the domestic economy to vote shares to popularity ratings—supports our claims. This rejoinder, therefore, advances our understanding not only of the role of the global economy but also of the role of executive control in selection models of retrospective economic voting. An implication is that the "unambiguously exogenous" shocks that drive economic conditions in LSCEs may not be exogenous at all but are instead shaped by the orientation of the policy regime. Both matter for presidential success, whereas the relevance of

the LSCE classification may well be overstated. By taking seriously the types of policies that voters use to extract executive competence signals, our evidence reveals a stronger role for national policymaking than CZ's theoretical and empirical contributions allow.

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Biographical Statement

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Table 1. Model Diagnostics for Regressing Domestic Economic Indicators on GET

	GDP Growth		**Log Inflation**		**Unemployment**		**Okun Index**		**Hanke Index**	
Neoliberalism imputed	LSCE	Neolib.	LSCE	Neolib.	LSCE	Neolib.	LSCE	Neolib.	LSCE	Neolib.
BIC	**3022.581**	3023.755	4354.564	**4335.606**	**4689.709**	4703.934	**1986.503**	1997.448	3863.049	**3813.723**
RMSE	3.499	**3.486**	11.902	**11.640**	**22.444**	22.646	**1.544**	1.553	15.923	**14.986**
N	544	544	544	544	505	505	505	505	448	448
Neoliberalism not imputed	LSCE	Neolib.	LSCE	Neolib.	LSCE	Neolib.	LSCE	Neolib.	LSCE	Neolib.
BIC	**2331.968**	2332.511	3394.532	**3376.080**	**3700.551**	3712.010	**1603.644**	1613.834	3077.725	**3034.745**
RMSE	3.681	**3.661**	13.450	**13.072**	**24.211**	24.415	**1.658**	1.669	17.540	**16.377**
N	410	410	410	410	391	391	391	391	348	348

Note: cells report BIC and RMSE statistics from regressing the domestic economic indicator noted in the first column on GET conditioned either by *LSCE* or by *Neoliberalism*. Full regression results are reported in the supplementary information file Tables A1 and A2.

Table 2. Predicting Incumbent Candidate Reelection

	Model 1 CZ Table 2 Model 4	Model 2 Continuous LSCE	Model 3 Neoliberalism	Model 4	Model 5 LSCEs only
GET index	0.103	0.631**	0.895^{+}	0.312	1.337*
	(0.158)	(0.205)	(0.487)	(0.519)	(0.627)
GET index × LSCE	0.893*			1.076+	
	(0.375)			(0.555)	
GET index × LSCE continuum		0.171			
		(0.281)			
GET index × Neoliberalism			-0.269	-0.276^{+}	-0.421*
			(0.194)	(0.158)	(0.176)
Incumbent ran	2.453**	2.505**	2.295**	2.272**	3.114**
	(0.778)	(0.759)	(0.618)	(0.551)	(0.734)
LSCE	-0.027			-0.023	
	(0.522)			(0.660)	
LSCE continuum		0.336			
		(0.275)			
Neoliberalism			-0.022	-0.031	0.060
			(0.151)	(0.151)	(0.181)
Intercept	-0.981*	-1.028**	-0.711^{+}	-0.695	-0.690
	(0.405)	(0.297)	(0.376)	(0.558)	(0.463)
Pseudo R^2	.19	.18	.16	.18	.26
Log Likelihood	-56.45	-57.37	-47.94	-46.95	-26.71
BIC	136.22	138.06	118.09	125.01	73.27
Countries	18	18	17	17	10
N	106	106	85	85	53

Note: Cells report parameter estimates from OLS regression with standard errors clustered by country in parentheses. ** $p \leq .01$, * $p \leq .05$, $^{+}$ $p \leq .10$, two-tailed test for unconditional coefficients.

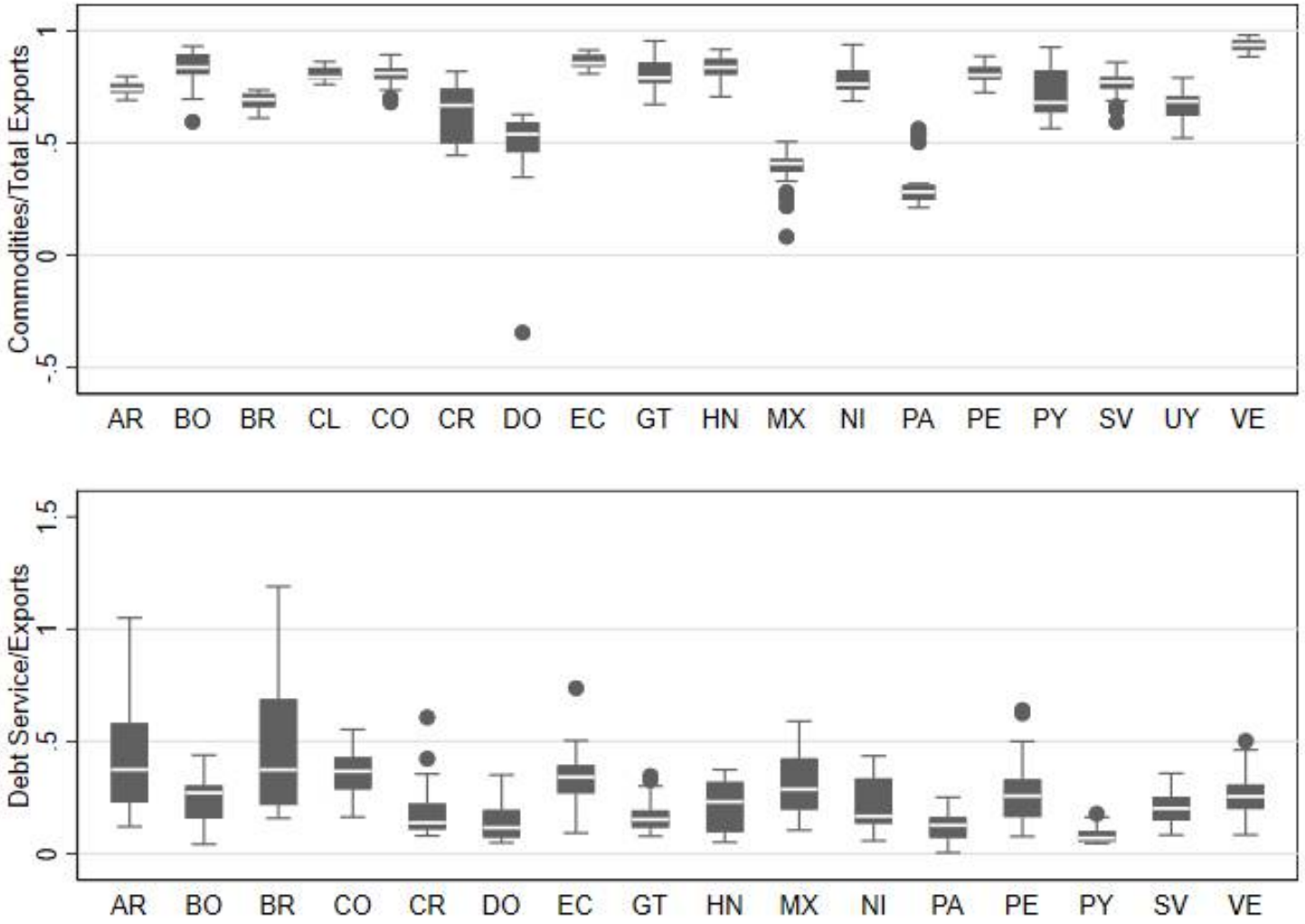
Table 3. Effects of GET Index on Presidential Approval, Conditional on Economic Structure

	Model 1 Baseline	Model 2 LSCE	Model 3 LSCE Continuous	Model 4 Neoliberalism	Model 5 Both	Model 6 LSCEs only
GET index$_t$	1.181**	0.863	1.228**	1.981**	1.686**	2.269**
	(0.371)	(0.696)	(0.394)	(0.638)	(0.770)	(0.824)
LSCE$_t$		-2.881			-3.659	
		(2.243)			(2.237)	
GET index$_t$ × *LSCE*$_t$		0.508			0.512	
		(0.893)			(0.898)	
LSCE continuum$_t$			0.029			
			(0.732)			
GET index × *LSCE cont.*$_t$			0.487			
			(0.769)			
Neoliberalism$_t$				0.086	0.071	0.125
				(0.220)	(.221)	(.252)
GET index$_t$ × *Neoliberalism*$_t$				-0.476*	-0.482*	-0.590**
				(0.212)	(.212)	(.222)
Growth$_t$	0.278**	0.276**	0.275**	0.260**	0.257**	0.297**
	(0.078)	(0.078)	(0.080)	(0.078)	(0.079)	(0.091)
Honeymoon$_t$	7.417**	7.428**	7.424**	7.388**	7.402**	11.501**
	(1.005)	(1.003)	(1.004)	(1.003)	(1.001)	(1.410)
Scandal$_t$	-2.630*	-2.612*	-2.646*	-2.522*	-2.504*	-2.892*
	(1.110)	(1.111)	(1.108)	(1.108)	(1.110)	(1.319)
Corruption$_t$	1.486	0.696	0.880	3.756	3.076	-2.404
	(5.267)	(5.603)	(5.511)	(5.273)	(5.506)	(6.747)
Terrorist attacks$_t$	0.003	0.002	0.004	0.007	0.006	-0.021
	(0.016)	(0.016)	(0.017)	(0.017)	(0.017)	(0.017)
Approval$_{t-1}$	0.784**	0.782**	0.784**	0.773**	0.771**	.762**
	(.022)	(.021)	(0.022)	(.022)	(.024)	(.026)

Intercept	11.486*	12.647*	12.539*	10.660*	11.776*	9.014**
	(4.754)	(5.403)	(5.192)	(4.756)	(5.345)	(4.639)
Countries	17	17	17	17	17	10
N	1093	1093	1093	1093	1093	651
R^2	.87	.87	.87	.87	.87	.86

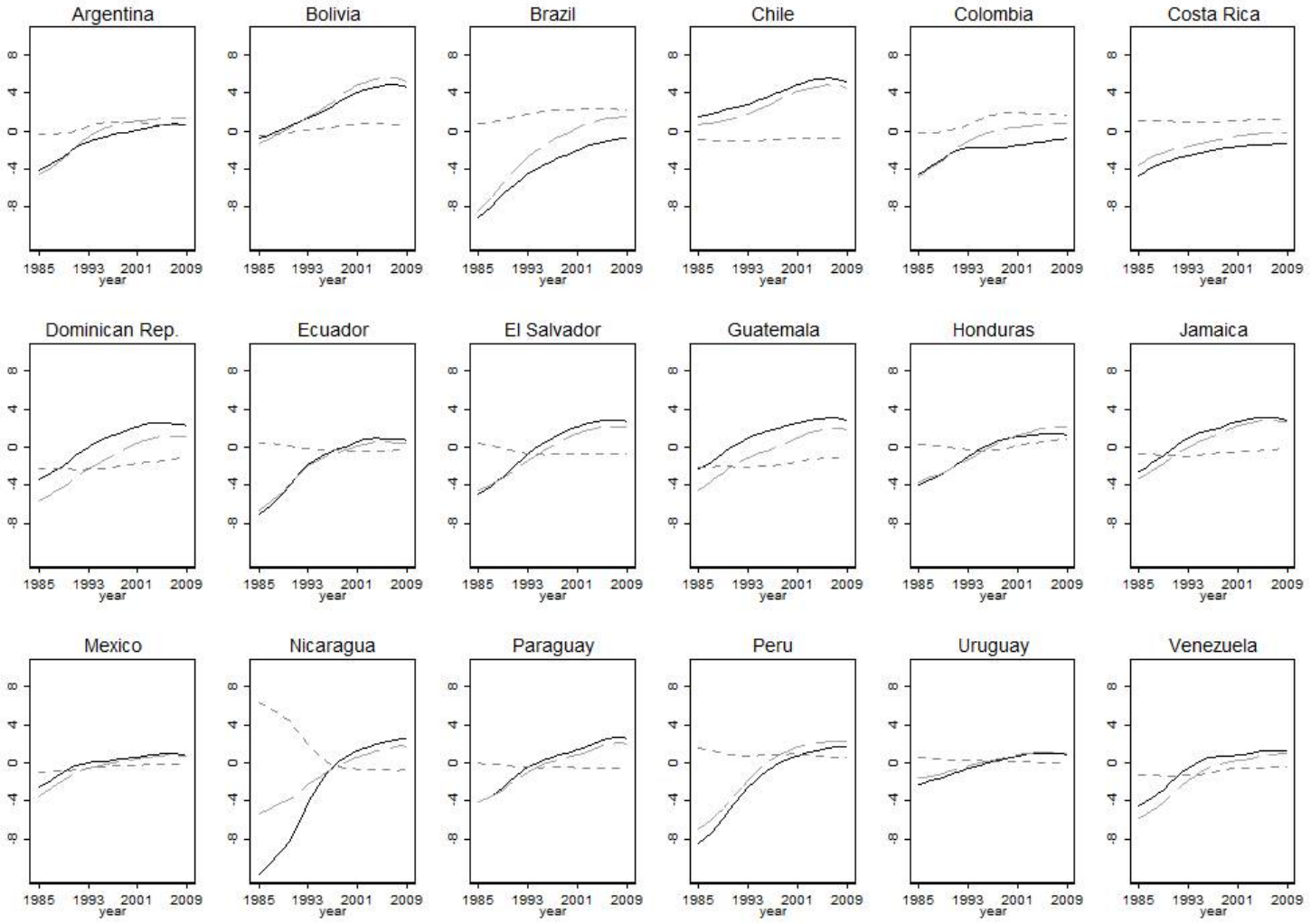
Note: Cells report parameter estimates with panel-corrected standard errors in parentheses. Standard errors are adjusted for panel-specific AR(1) processes. All models include country fixed effects. ** $p \leq .01$, * $p \leq .05$, $^{+}$ $p \leq .10$, two-tailed test.

Figure 1. Exposure to commodity prices and international interest rates



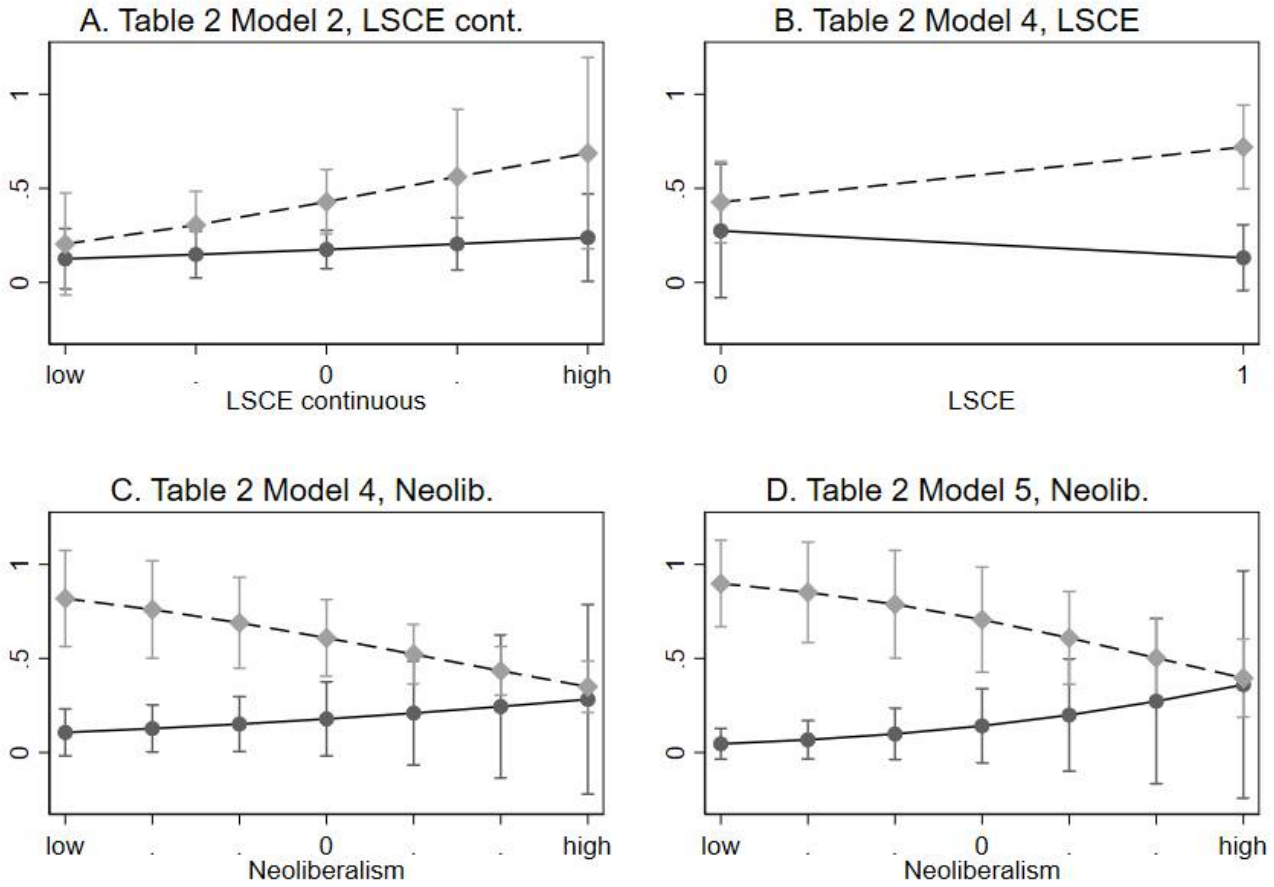
The graph at the top displays commodity exports divided by sum of merchandise and service exports for each country. The bottom graph reports debt service as a share of total exports.

Figure 2. Economic Policy Regimes in 18 Latin American and Caribbean Countries, 1985-2009



Note: Long-dashed lines report *Othodoxy*, dynamic factor scores from Lora's (2012) trade, finance, taxation, and privatization indices. Short-dashed lines report *Statism*, dynamic factor scores from *Worker Welfare* and government consumption. Solid line is *Neoliberalism*, equal to the difference in *Orthodoxy* and *Statism*.

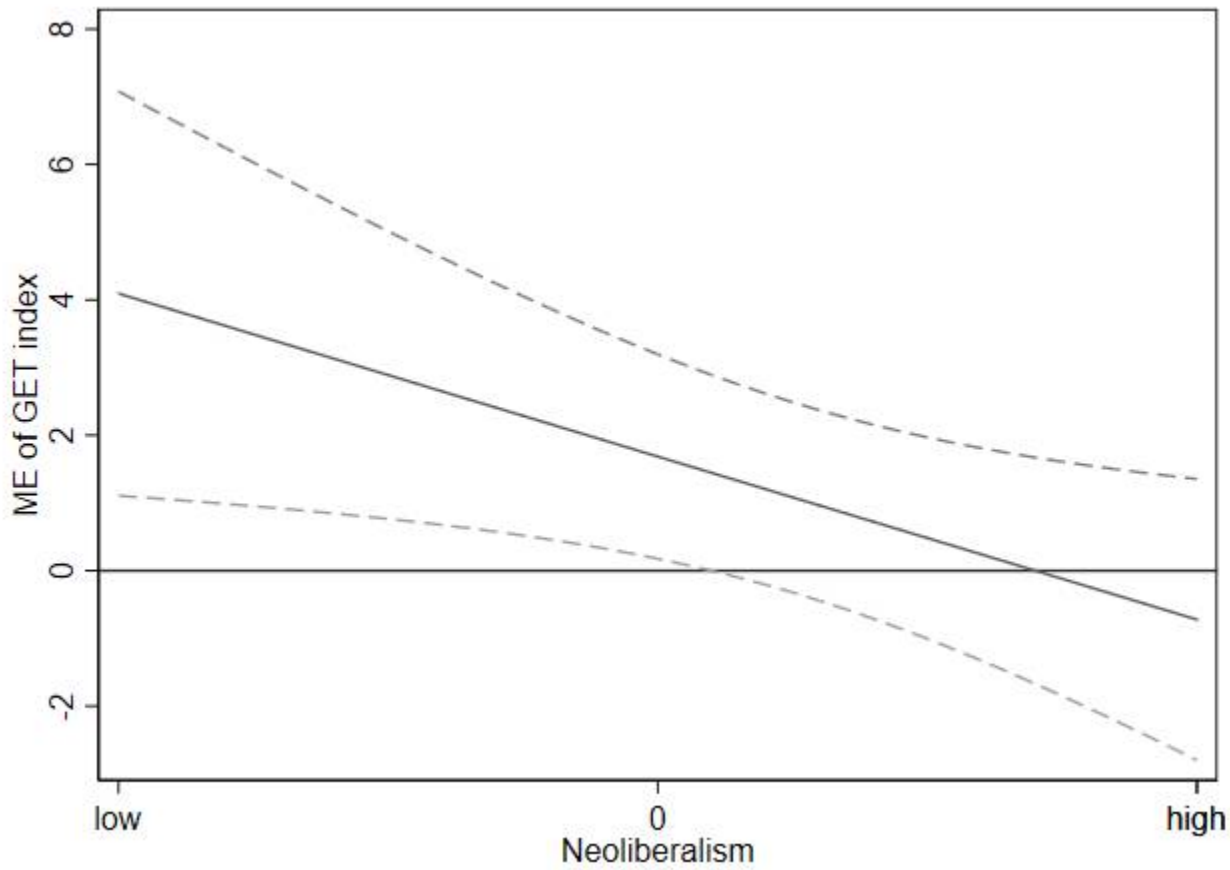
Figure 3. Effect of Policy Regime on the Probability of Presidential Reelection



Dashed grey line = GET good; Solid black line = GET bad

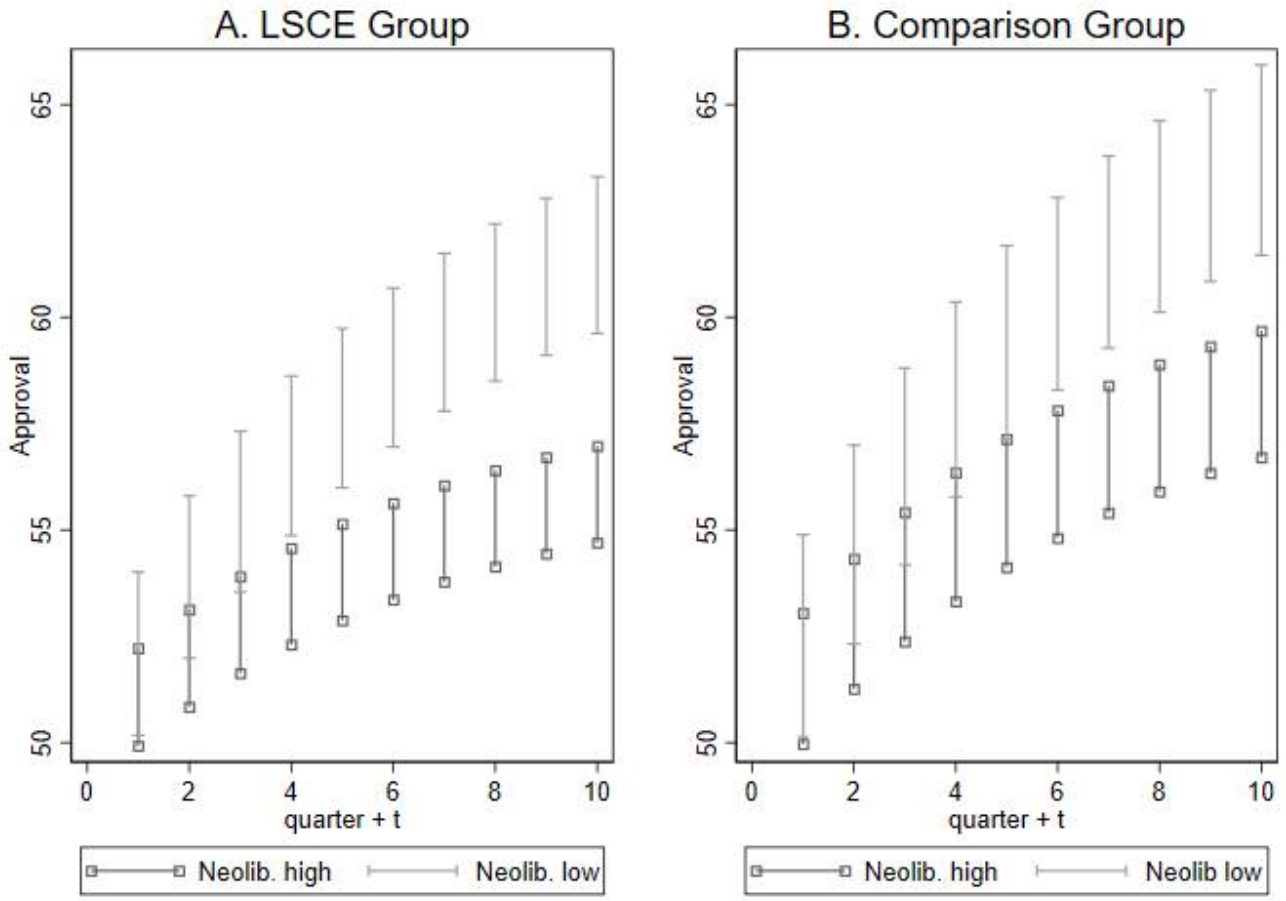
Note: Graphs display expected probability of presidential reelection as LSCE or the policy regime varies. Graphs produced using estimates from Table 2. GET values set to "bad" and "good" as given by one standard deviation below and one above the mean of the GET index, Vertical bars represent 95% confidence intervals.

Figure 4. Conditional Effects of GET on Presidential Approval as the Policy Regime Varies



Note: Graph reports marginal effects of a one-unit increase in *GET* on *Approval* over the sample range of *Neoliberalism*. Estimates are from Table 3 Model 4. Dashed represent 95% confidence intervals.

Figure 5. Forecasting the Effects of GET on Presidential Approval in Two Policy Regimes



Note: Graphs produced using estimates from Table 3 Model 5. *Neoliberalism* is set to its 90th percentile in-sample value to illustrate a "neoliberal policy" scenario and to its 10th percentile value for a "statist policy" orientation. Vertical bars represent 95% confidence intervals.

Supplemental Information File

The World Economy, Political Control, and Presidential Success

Ryan E. Carlin & Timothy Hellwig

This document includes the following:

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 subindices 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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