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Abstract

Presidential approval in Peru depends on economic outcomes. However, voters are unable to distinguish between outcomes resulting from economic policies and those caused by exogenous shocks. Estimation results from seven Fractional Cointegrated VAR (FCVAR) models suggest that presidential approval increases with the monetary policy interest rate, the terms of trade, and manufacturing employment; and decreases with the nominal PEN/USD exchange rate and inflation volatility. Additionally, a Principal Components Analysis (PCA) conducted over a large set of macroeconomic indicators points to a greater influence of external over domestic factors in explaining presidential approval; i.e., economic outcomes that determine the dynamics of presidential approval are not under presidential control in Peru. It can be argued that these findings identify a significant source of political instability and a considerable challenge to democratic governance. To the authors' best knowledge, this is the first application of fractional cointegration analysis to political economy in Latin America.

JEL Codes: C32, D72.

Keywords: Economic Voting, Fractional Cointegration, Political Economy, Macroeconomics, Latin America, Peru.

Resumen

La aprobación presidencial en el Perú depende del estado de la economía. Sin embargo, los ciudadanos no son capaces de distinguir entre los resultados económicos determinados por las políticas económicas y aquellos que son consecuencia de choques externos. Los resultados de la estimación de siete modelos VAR Fraccionalmente Cointegrados (FCVAR) sugieren que la aprobación presidencial aumenta con la tasa de interés de política monetaria, los términos de intercambio y el empleo manufacturero; y disminuye con el tipo de cambio nominal PEN/USD y la volatilidad de la inflación. Adicionalmente, un Análisis de Componentes Principales (PCA) sobre un amplio conjunto de indicadores macroeconómicos sostiene que la influencia de los factores externos sobre la aprobación presidencial es mayor que la de los factores domésticos; ie., los resultados económicos que determinan la dinámica de la aprobación presidencial no están bajo el control del presidente en Perú. Esta es la primera aplicación del análisis de cointegración fraccional para el estudio de la economía política en América Latina, según el conocimiento del autor.

Clasificacion JEL: C32, D72.

Palabras Claves: Voto Económico, Cointegración Fraccional, Economía Política, Macroeconomía, América Latina, Perú.

Presidential Approval in Peru: An Empirical Analysis Using a Fractionally Cointegrated VAR[1]

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1 Introduction

Peru's history of high-level corruption raises concerns about the effectiveness of national democratic governance in choosing the most qualified politicians. Although elected on the pledge to overturn previous authoritarian practices, all Presidents since 2001 are currently under investigation for allegations of accepting bribes from Odebrecht, a Brazilian construction company, in exchange for support to win public sector tenders. According to a survey endorsed by Peru's electoral authority[4], the average voter does not have a party affiliation and becomes concerned with politics only during elections. Voters mainly care about whether a candidate is perceived as corrupt and, secondly, about economic policies geared to reduce unemployment and poverty. The influence of economic performance on political approval creates a democratic mechanism whereby voters can promote sound economic policies, as Presidents rarely implement policies that may damage their public image. Along these lines, the economic voting approach suggests that political preferences respond to changes in economic variables. However, the prevalence of information asymmetries and weak institutions are arguments against the validity of economic voting in Peru and Latin America.

There is no agreement in the literature on whether voters consider economic performance in making their choices. The determinants of political preferences change with time and across countries. Additionally, particular characteristics, such as persistent autocorrelation of political survey data, make hypothesis testing difficult and, therefore, papers about the same country differ in their conclusions. Hence, more empirical research is required to verify the many theories that attempt to explain voters' behavior. Towards this end, Byers, Davidson, and Peel (1997) model polling data as an aggregation of persistent and temporary preferences (by committed and uncommitted voters, respectively) that results in a fractional process. This paper presents evidence that some Peruvian macroeconomic series, such as manufacturing employment, the policy rate of the Central Reserve Bank of Peru (BCRP), and the terms of trade can be modeled as fractional processes. This set of fractional time series allows estimation of Fractionally Cointegrated VAR (FCVAR) models to

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[4] <observaigualdad.jne.gob.pe/pdfs/recursos/PERFIL_ELECTORAL_EN_EL_PERU.pdf >

assess the economic voting hypothesis econometrically. To the authors' best knowledge, this is the first application of fractional cointegration analysis to political economy in Latin America.

Johansen and Nielsen (2012) develop the required estimation and inference framework for FCVAR models. As an extension of Johansen's (1995) Cointegrated VAR (CVAR), FCVAR includes the same inference capabilities: estimation of the cointegration vectors, if any; and hypothesis testing, but under fractional orders of cointegration. In a significant contribution to political science in Canada, Jones, Nielsen, and Popiel (2014) use the FCVAR to show that political parties' approval ratings hold a long-run relationship with economic variables. They further reject the hypothesis of approval ratings being weakly exogenous and dependent on the U.S. economy.

This paper finds that, in 2001-2018, presidential approval in Peru increased with the interest rate, the terms of trade, and employment in the manufacturing sector; and decreased with the nominal PEN/USD exchange rate, inflation volatility, and corruption perception. Furthermore, presidential approval adjusts to changes in macroeconomic variables. In turn, interest rates and external economic variables are weakly exogenous, as expected from discretionary monetary policy and international markets, respectively. In the authors' view, this research contributes to understanding the complex dynamics of political preferences in Peru. Additionally, a Principal Components Analysis (PCA) points to a greater influence of external over domestic factors in explaining presidential approval. According to current political economy thought, globalization blurs voters' ability to differentiate between economic outcomes resulting from sound policies and those caused by external shocks. Specifically, this paper suggests that, due to reasons such as information asymmetries and incipient regional integration in Latin America, presidential approval ratings in Peru respond to exogenous factors such as the behavior of commodity prices. This has implications for political accountability and poses a challenge to Peru's democratic governance. Altogether, economic variables are still significant for presidential approval .

The rest of the paper is organized as follows. Section 2 discusses theoretical developments on the determinants of voting behavior and arguments in favor of the validity of economic voting in Peru. It also highlights the debate about the mechanism through which economic voting implies political accountability. Section 3 presents the FCVAR methodology. Section 4 contains the empirical analysis and discusses important findings. Section 5 describes other estimation results. Section 6 concludes.

2 Literature Review

2.1 Economic Voting and Political Accountability

Economic voting explains changes in political preferences as a response to varying economic conditions. The literature provides evidence of economic voting for a vast number of countries. The first studies in this field, such as Downs (1957), Fair (1978), and Kramer (1971), assume that voters maximize their expected utility taking into account past economic performance. This means that voters are rational and, despite information asymmetries and party affiliations, their decisions are based on empirical data[5]. However, it is difficult to reach an agreement on which variables should be included in voters' utility function due to inconclusive statistical results.

In Latin America, the literature outlines the particular characteristics of the political system

[5] For a comprehensive and detailed survey of the theoretical aspects of the political economy of dynamic elections, see Duggan and Martinelli (2017).

that facilitate economic voting. First, Roberts and Wibbels (1999) point out that the lack of political party institutionalization causes weak ideological identification, which in turn fails to stabilize political preferences in a changing economic context. Similarly, Ratto (2013) suggests that political preferences are driven by two factors: a stable component (ideology) and a rather volatile one (economic performance). For example, Chile's presidential approval during the last 12 years has remained high despite economic deceleration; see Perello (2015). One explanation might be the importance of ideology and partisanship among left-leaning voters in Chile, as both administrations in that period were of that persuasion. In the case of Peru, Osorio (2015) studies Ollanta Humala's government and concludes that voters are driven by ideology only during elections because of low partisanship. Weak political party institutionalization has two effects in Peru. First, voters' preferences are governed by elements such as leaders' personality and communication skills. As a result, Presidents tend to lose support a few months after being elected. Second, a lack of voters' commitment makes their preferences unstable; i.e., basically short-term economic conditions are considered when assessing the President's approval.

Further studies on the topic investigate how voters value economic performance. Voters need a comparison reference to decide whether an economic outcome is good or bad. Traditionally, political economy considers a temporal criterion; i.e., retrospective and prospective voting. Retrospective voters penalize bad decisions; i.e., they refresh their preferences according to past political performance. Both Barro (1973) and Ferejohn (1986) describe voting behavior as penalizing or rewarding politicians by modeling a utility function including only past events. On the other hand, prospective voting is aligned with rational expectations. According to this theory, individuals do not make systematic mistakes and include every new piece of information available; see, for example, Rogoff and Sibert (1988) and Cukierman and Meltzer (1989). For further information on the ongoing discussion about this temporal dimension, see Lewis-Beck and Stegmaier (2000) and Duch and Stevenson (2008). These authors highlight that the number of lags required in the model specification to avoid autocorrelation is also subject to debate: more lags are needed under retrospective voting because political preferences are formed using information from many periods in the past. In the case of Latin America, the structural reforms that established democracy also caused high unemployment and inflation. Although economic outcomes improved rapidly, the reelection of many Presidents during that period raises questions about the validity of economic voting. Nevertheless, Ratto (2013) finds empirical evidence in favor of economic voting and against theories suggesting that uncertainty deprives voters of economic guidance in young democracies; and highlights that economic voting works as a democracy-reinforcing mechanism; i.e., citizens re-elect Presidents provided that they act according to their interests. The fact that economic voting prevails in Latin America means that voters are able to choose the best policymakers and punish bad ones for their mistakes through democratic practice.

Voters also compare economic outcomes in a spatial dimension. Globalization poses a challenge to economic voting, as economic outcomes can be caused by either domestic policies or the global environment. Powell and Whittenn (1993) stress that clarity of responsibility plays a central role in forming political sentiment; but in an economy exposed to external shocks, the relationship between presidential approval and economic outcomes becomes weaker. Later studies interpret external factors as a second comparative dimension for evaluating economic outcomes. First, in case voters are able to discount exogenous factors, economic voting maintains its role as a democratic mechanism. Kayser and Peress (2012) show that voters benchmark the domestic economy against an international baseline. Thus, voters do not penalize the President for bad economic results when other countries are also experiencing poor economic performance. Although benchmarking allows the identification of political responsibility, the authors argue that it occurs due to pre-benchmarked information disseminated by the media, instead of a highly sophisticated voting behavior. Therefore, it is to be expected that, as economic integration increases, comovements between the domestic economy and its benchmark reduce approval responsiveness to economic performance; see Kayser and Peress (2016). However, these studies were carried out in countries with consolidated democracies, highly-educated voters, and regionally integrated economies; thus, their conclusions are not necessarily valid in the Latin American political context. For example, Navia and Osorio (2015) assert that presidential systems in Latin America tend to blame the executive power for current economic context; nevertheless, voters cannot tell whether good economic results are caused by domestic policies or external shocks.

Recent research in the region disregards international benchmarking in voting and outlines the prevalence of information asymmetries. Campello and Zucco (2016) show that in a large subset of Latin American countries, including Peru, presidential popularity and reelection prospects are driven by exogenous factors (i.e., not controlled by politicians) such as commodity price behavior. This finding negatively affects political accountability in the region: Presidents who do not feel responsible for the results of their policies tend to adopt a rent-seeking behavior. In addition, there is evidence that media scandals become more relevant in a context of bad economic performance. In a cross-sectional study of 18 Latin American countries, Carlin, Love, and Martinez-Gallardo (2014) argue that political accountability is conditioned by the state of the economy. Hence, the region's exposure to volatile international markets reinforces political instability. Furthermore, even though presidential approval responds to economic changes, the democratic mechanism promoted by retrospective economic voting does not occur.

This paper provides empirical evidence that economic voting hypothesis is valid for Peru, consistently with current literature; see Stokes (1996), Arce and Carrion (2010), Maldonado and Pimentel (2013), and Osorio (2015). Peruvian voters are prone to punish political incumbents based on short-term economic performance. Weak partisanship softens long-term factors such as ideology. As mentioned, external influence and uncertainty about its impact on voters should also be considered. In particular, the Peruvian economy has been exposed to external shocks since the 1990s because of its export orientation. However, even though economic performance is actually determined to a considerable extent by the international markets (see Dancourt, 2017; and Rodriguez et al., 2018), there is no evidence that voters are aware of it. Further research on these issues can make a significant contribution to political science.

2.2 Fractional Cointegration and Political Economy

Empirical research in political science describes economic voting using error correction models; i.e., political preferences adjust to changes in economic variables; see Brody and Page (1975) and Beck (1991). The first theoretical development using error correction modeling is Ostrom and Smith (1992). According to these authors, voters are mainly concerned about quality of life outcomes, which establishes a long-run relationship between political preferences and the economy. However, ordinary and extraordinary political events deviate voters' assessment about the economy. Extraordinary events such as corruption scandals are included in the cointegration vectors because they have a permanent effect on public sentiment. Ordinary events and other exogenous deviations, on the other hand, are corrected or mainstreamed through media coverage. Regarding the individual series, the study strongly rejects the hypothesis that political approval is stationary and assumes it is a unit root process. Later, Beck (1992) highlights that political approval series seem to behave as stationary processes when assessing long periods, whereas a random walk offers a better fit for short-term analysis. Modeling approval rating as a random walk implies that all past information is already contained in the present, which is plausible for financial series (efficient market hypothesis). However, it is questionable that presidential approval today is determined by, for instance, two-year-old events. Lastly, since political economy argues that movements in economic outcomes are reflected in political support and not vice versa, the researcher must only assess the hypothesis of political support being error-correcting.

Box-Steffensmeier and Tomlinson (2000) explain the benefits of modeling political support data as fractional processes. According to the literature there are two kinds of voters. Committed voters make rigorous decisions based on information from both the past and present. On the other hand, less sophisticated respondents are driven by current impressions about political performance. The aggregation of these series is best modeled as a fractional process. Fractional integration allows for long memory in stationary time series. Additionally, a fractional cointegration framework does not require parent series to be $I(1)$ nor residuals to be stationary. From this more general notion of cointegration, it is possible to affirm that if parent and residual series are integrated of order (d) and (d'), respectively, where $d - d' < 0.5$, then the two series are fractionally cointegrated; see Dueker and Startz (1998).

Treisman (2011) models presidential approval as a fractional process to analyze extensively the prevalence of economic factors over opinion manipulation in Russia. The pronounced difference between Yeltsin's and Putin's popularity is attributed to the latter's effective political strategies. The author estimates various fractionally integrated single equation models, including dummy variables for high-impact events such as the occupation of Chechnya, the attack on Kosovo, and Yeltsin's illness; and a set of macroeconomic variables. The model shows that Putin's higher popularity is mainly correlated with improving economic conditions in Russia. In contrast, manipulation strategies do not fulfill their goals. Thus, the common long-run trend between presidential approval and realized economic performance holds even under information uncertainty.

More recently, Jones, Nielsen, and Popiel (2014) show that Canadian prospective voters respond to current macroeconomic conditions by using a FCVAR model. Since the FCVAR does not assume any integration order for the political variable, it is expected to provide better estimates and more precise inference. Using aggregated polling data, it shows that Canada's Liberal Party benefits from high unemployment and a low interest rate, whereas the Conservative Party gains support in the opposite manner. It finds that political support and unemployment correct deviations from equilibrium, while the interest rate is weakly exogenous. However, as mentioned in Beck (1992), finding that unemployment is error-correcting might not be compatible with political economy theory.

In conclusion, there is theoretical and empirical evidence that economic performance determines political preferences. However, since common estimation techniques are not able to incorporate the mixed behavior of political time series, empirical research on economic voting yields inconclusive results. The fractional cointegration analysis offers better tools to avoid model misspecification. Furthermore, it allows to describe the short-run dynamics between political and economic variables.

This paper aims to show that presidential approval ratings hold a long-run relationship with macroeconomic performance in Peru. It is worth highlighting that the analysis focuses on how, and to what extent, macroeconomic variables drive political support for the current President (the incumbent) and not on how voters choose from a set of candidates.

3 Methodology

This paper uses an FCVAR model to identify and assess the long-run relations between presidential approval ratings and macroeconomic time series. The estimation and inference tools were developed by Johansen (2008) and Johansen and Nielsen (2010, 2012, 2014). The FCVAR can be better understood using the CVAR approach developed by Johansen (1995):

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \sum_{i=1}^{k} \Gamma_i \Delta Y_{t-i} + \epsilon_t, \tag{1}$$

where Y_t is a column vector of n variables, α and β are $n \times r$ matrices, r is the number of cointegrating vectors, and $0 < r < n$. The r cointegrating vectors are included in each column of β, which implies that the long-run relationships are given by the elements of $\beta' Y_t$. The coefficients for the speed of adjustment towards the equilibrium are contained in α. Lastly, Γ represents the system's short-run behavior.

As shown in Johansen and Nielsen (2012) and Jones et al. (2014), the whole extension can be illustrated by replacing Δ and L by their fractional counterparts Δ^b and $L_b = 1 - \Delta^b = 1 - (1 - L)^b$:

$$\Delta^b Y_t = \alpha\beta' L_b Y_{t-1} + \sum_{i=1}^{k} \Gamma_i L_b^i \Delta^b Y_t + \epsilon_t. \tag{2}$$

The FCVAR model (see Johansen, 2008), is derived after (2) is applied to $Y_t = \Delta^{d-b} X_t$:

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_{t-1} + \sum_{i=1}^{k} \Gamma_i L_b^i \Delta^d X_t + \epsilon_t. \tag{3}$$

Specifically, Johansen and Nielsen (2014) consider the simplified case with $d = b$:

$$\Delta^d X_t = \alpha L_d (\beta' X_{t-1} + \rho') + \sum_{i=1}^{k} \Gamma_i L_d^i \Delta^d X_t + \epsilon_t. \tag{4}$$

Johansen and Nielsen (2012) point out that $\beta' X_t + \rho'$ is a zero-mean process of fractional order zero, where $\beta' X_t$ has the same interpretation as equation (1). Note that since $\Delta^d 1 = 0$, $L_d \rho' = \rho'$. The authors demonstrate that when $0 < r < n$, X_t is fractional of order d and cofractional of order $d - b$. This means that $\beta' X_t$ is fractional of order $d - b$. Since fractional differentiation is defined as an infinite time series, calculation under finite samples is not possible. It can be assumed that X_t is zero before $t = 1$ to avoid this issue. However, according to Johansen and Nielsen (2014), this assumption would introduce a bias that can be avoided by including a level parameter μ:

$$\Delta^d (X_t - \mu) = L_d \alpha\beta' (X_t - \mu) + \sum_{i=1}^{k} \Gamma_i L_d^i \Delta^d (X_t - \mu) + \epsilon_t. \tag{5}$$

As mentioned before, FCVAR provides the same inference benefits as CVAR. It includes a set of tests to determine how many long-run relationships exist. Moreover, adjustment responses to deviations from equilibrium and short-run behavior can be modeled and estimated. Lastly, it is also possible to assess the model fit by testing the residuals for serial correlation.

The hypotheses for the long-run coefficients, β, are given by $\beta = H\varphi$, where, given that s is the number of freely varying parameters, H is an $n \times s$ matrix that specifies the restrictions to be tested and φ is an $s \times r$ matrix of freely varying parameters. When $r = 1$, that is, when there is only one cointegrating relationship, the same restriction is imposed and the degrees of freedom of the test are calculated as follows: $df = (n - s)r$. If $r > 1$, different restrictions can be imposed on different columns of β; i.e., $\beta = (H_1\varphi_1, \ldots, H_2\varphi_2)$, where each H_i matrix is of dimensions $n \times s_i$. The degrees of freedom will in turn be $df = \sum_{i=1}^{r}(n - r - s_i + 1)$.

Analogously, the hypotheses on α are expressed as $\alpha = A\psi$, where, given that m is the number of freely varying parameters, the known $n \times m$ matrix A contains the restricted coefficients for the speed of adjusment. Lastly, ψ is an $m \times r$ matrix of freely varying parameters with $m \geq r$ and the degrees of freedom of the test are $df = (n - m)r$.

4 Empirical Analysis

This study uses the Matlab software package *à la* Nielsen and Popiel (2016). It allows the calculation of estimated values and test statistics at each step of the FCVAR estimation. First, the number of lags is chosen as a result of a general-to-specific procedure. Beginning from a high lag order $k_{\max}$ to $k = 0$, the significance of the highest-order lag coefficient is tested. If it is not significant, it is dropped and the model is re-estimated. This procedure is repeated until the null hypothesis is rejected. A Ljung-Box Q-test, $Q_{\hat{\varepsilon}}(h)$, for $h = 12$ lags, a serial correlation test, and information criteria AIC and BIC are calculated at each step. Order k is selected so the error terms are not correlated and the information criteria are minimum. Next, the FCVAR estimation procedure under the simplified case with $d = b$ described in the former section is carried out.

The hypothesis tests are specified as follows: $\mathcal{H}_1^d$ tests $d = 1$; $\mathcal{H}_i^\beta$ tests whether the long-run coefficient of the *ith* variable in the model is not significant ($\beta_i = 0$); and $\mathcal{H}_i^\alpha$ tests for weak exogeneity ($\alpha_i = 0$). Seven models are constructed to test the economic voting hypothesis for Peru.

4.1 Data

Previous studies have been flexible: there is no agreement on which economic variables should be taken as explanatory of political support. Thus, a set of series of good indicators of macroeconomic performance in Peru are combined. First, monetary policy controls inflationary pressures by acting in response to demand and supply shocks. Therefore, the interest rate contains information about the general state of the economy. Second, local demand, unemployment, and manufacturing employment represent domestic economic activity. Third, Peru's economic exposure to external shocks, including nominal exchange rates and terms of trade, is taken into account. Fourth, inflation volatility is considered to have a negative influence on economic growth; see Judson and Orphanides (1999). Therefore, a Generalized Autoregressive Conditional Heteroskedasticity (GARCH) model is estimated to include inflation volatility in the analysis. Fifth, after considering a large set of macroeconomic indicators, a PCA is performed to summarize economic activity in a small number of relevant factors. Finally, the Google Trends corruption index is included.

Presidential approval ratings are obtained from monthly national surveys conducted by IPSOS[6] in August 2001-February 2018. This information is public and available online[7]. Next, monthly macroeconomic time series for the same period are obtained from the Central Reserve Bank of Peru (BCRP) website. Three groups of variables are distinguished: (i) domestic demand: manufacturing employment; (ii) the BCRP monetary policy rate; and (iii) external shocks: nominal exchange rate (PEN/USD) and Peru's terms of trade. All macroeconomic indicators are plotted in Figure 1. The recent economic history of Peru can be divided in a high-growth period, favorable terms of trade, and currency appreciation (2002-2008); the aftermath of the global financial crisis (2009); and, finally, a period of weak output growth (2010-2018). In consequence, the sample analyzed is deemed useful to avoid biased conclusions.

The political variable is first converted to log-odds. Following Jones, Nielsen, and Popiel (2014), the logit transformation of variables measured as percentages allows for unbounded error terms in the models used. On the other hand, Figure 2a (solid red) indicates the presence of political cycles in presidential approval ratings series. The period of study covers the administrations of Alejandro Toledo, Alan García (his second term), Ollanta Humala, and Pedro Pablo Kuczynski (his first and a half year in power). Each vertical line indicates election dates. As documented in the literature (see Chappell and Keech, 1985; and Nielsen, 2014), there is a "honeymoon" characterized by high ratings at the beginning of each mandate, which smoothly decreases thereafter.

Actual ratings can be decomposed in a cyclical component (C_t^{cycle}) and a non-cyclical one (pa_t):

$$pa_t^{actual} = C_t^{cycle} + pa_t. \tag{6}$$

In order to account for both effects, equation (6) is estimated using the following specification:

$$C_t^{cycle} = \theta_1 \tau_t + \theta_2 \tau_t^2, \tag{7}$$

where τ is the time since the last election measured in months. The term τ^2 is included to capture changes in the political cycle's effect over time. Standard error terms are recovered to obtain the filtered presidential approval series: pa_t. Both coefficients estimated from equation (7), $\hat{\theta}_1 = -0.095$ and $\hat{\theta}_2 = 0.001$, are significant at the 5% level. Moreover, $\bar{R}^2 = 0.507$ and the F test indicate that the parameters are jointly significant with 99% confidence. The filtered presidential approval series is shown in Figure 2a (dashed blue). Finally, Figure 2b shows the filtered presidential approval series (dashed blue) and the corruption index in log-odds (solid red).

A univariate analysis is then done to show the appropriateness of the FCVAR framework. First, Figure 3a shows that the autocorrelation function for the political variable decays hyperbolically. This suggests that political preferences have a long memory: next month's ratings are commonly similar to current ones. Moreover, Figure 3b evidences the existence of a long-run stochastic trend, since the spectral density plot is concentrated around the zero frequency.

Next, Table 1 shows the estimated integration orders, $\hat{d}$, and a Ljung-Box Q-test for each series from the model's univariate version, i.e. Fractional AR (FAR); see Jones, Nielsen, and Popiel (2014). It is worth noting that in all variables, with the exception of inflation volatility and corruption perception, the estimated integration orders are fractional and serial correlation can be avoided by using at most two lags. Inflation volatility is fractional stationary; therefore, $\hat{d}$ is close to zero. Additionally, FAR-estimated residuals for corruption perception are uncorrelated with

[6]IPSOS is a well-known local polling agency.

[7]< www.ipsos.com/es-pe/news-and-polls/overview?page=0 >

$k = 3$. Table 1 also shows that the integration orders estimated using Geweke and Porter-Hudak (1983) with bandwidths $m = T^{0.5}$, $m = T^{0.6}$ and $m = T^{0.7}$are fractional and consistent with the ones estimated using FAR estimates. Presidential approval ratings have a mixed behavior, since $\hat{d}$ is fractional and increases with m. In sum, given the features of the analyzed time series, the FCVAR model appears to be more suitable than the standard CVAR for avoiding misspecification errors.

4.2 Model 1: Presidential Approval, Interest and Nominal Exchange Rates

A system including presidential approval, the interest rate, and the nominal PEN/USD exchange rate is assessed initially. Table 2 presents the relevant results for the model. First, the lag selection criterion determines $k = 0$. Next, rank tests start with the hypothesis $r = 0$, which is rejected against $r = 3$. It is not possible to reject $r = 1$, since the corresponding p value $= 0.919$.

The model is estimated with $k = 0$ and $r = 1$. This results in an estimated integration order $\hat{d} = 1.160$ with standard error 0.042. Moreover, the Ljung-Box Q-test, $Q_{\hat{\varepsilon}}(12)$, indicates that there is no sign of serial correlation in the residuals. The long-run relationship is normalized by assigning a value of one to the coefficient of the political variable ($\beta_1 = 1$). The estimated coefficients for the interest and nominal exchange rates are $\hat{\beta}_2 = 0.132$ and $\hat{\beta}_3 = -4.801$, respectively. The signs of the coefficients imply that political approval increases with the interest rate and decreases with the nominal exchange rate.

As mentioned in Section 4, inflation remained low during the period of analysis due to efficient inflation targeting. Thus, high interest rates are a monetary policy response to positive demand pressures and reflect that the economy is doing well. By contrast, a nominal exchange rate depreciation causes higher liability costs in Peru because of bank dollarization; see Dancourt (2017). The fact that both variables hold a long-run relationship with presidential approval imply that they determine a convergence towards political equilibrium. For example, negative exogenous shocks on the model, such as political scandals, have a limited impact during appreciation periods because presidential approval in equilibrium is high.

The appropriateness of the FCVAR approach is confirmed by rejecting $\mathcal{H}_1^d$ (16.367 with p value $= 0.000$). The hypotheses $\mathcal{H}_1^\beta$,$\mathcal{H}_2^\beta$, and $\mathcal{H}_3^\beta$ are rejected, indicating that all three variables enter in the long-run relationship (p-values 0.000, 0.008, and 0.000, respectively). The $\mathcal{H}_1^\alpha$ is also rejected (19.107 with p-value 0.000), suggesting that the political variable is error-correcting (i.e., it adjusts to changes in the long-run equilibrium). Changes in the interest and nominal exchange rates are corrected through changes in presidential approval. The opposite occurs with the interest rate: tests on α fail to reject $\mathcal{H}_2^\alpha$. The exogenous behavior of the interest rate is consistent with the fact that it is fixed in a discretionary manner by the monetary authority, so it does not move when the other variables change. On the other hand, nominal exchange rates are influenced by international currency markets and discretionary BCRP interventions. However, nominal exchange rates do respond to interest rate changes through the interest rate parity; thus, $\mathcal{H}_3^\alpha$ is borderline rejected (p-value of 0.044).

Summing up, the interest rate is weakly exogenous, while both approval ratings and exchange rates are error-correcting. All the tests are specified following (6) and (7), respectively, where $(p, s, r) = (3, 2, 1)$ and $(p, m, r) = (3, 2, 1)$. The model is estimated again using $\alpha_2 = 0$ as a restriction to account for the long-run exogeneity of the policy interest rate. The $\hat{\beta}$ coefficients change slightly but keep the same signs as in the unrestricted model.

Finally, the negative sign of the adjustment coefficient for political approval may be interpreted as follows: a 1% decrease in the interest rate is corrected by an increase in the error term of the cointegration vector, ν_t, to maintain the long-run equilibrium. Next, the increase in ν_t is responded by a decrease in presidential approval (-0.218) that pushes the system back to equilibrium through the error- correction mechanism. In other words, changes in economic variables result in deviations of presidential approval from its long-run equilibrium (hereafter referred to as political equilibrium).

4.3 Model 2: Presidential Approval, Interest Rate and Terms of Trade

The second model includes the terms of trade and excludes the nominal exchange rate. The general-to-specific estimation procedure indicates that this system does not require any lags, so $k = 0$ is chosen. Table 3 shows the tests and the estimation results. First, the cointegration test rejects the null hypothesis $r = 0$. The existence of one cointegration vector is selected, since it is not possible to reject $r = 1$ (p value $= 0.616$). The estimation of the system with $k = 0$ and $r = 1$ implies that $\hat{d} = 1.160$ with standard error 0.041. The Ljung-Box-Q-test fails to reject the null hypothesis of no serial correlation ($Q_{\hat{\varepsilon}}(12) = 125.730$ with p value $= 0.117$). The estimated coefficients of the interest rate and the terms of trade are 0.180 and 1.584, respectively, both statistically significant. The interpretation for the interest rate's positive effect is the same as for Model 1: effective inflation targeting by the BCRP establishes a link between high rates and good macroeconomic outcomes. Next, Peru's dependence on commodity and metal prices since 2002 has been documented in the literature; see Dancourt (2017), Rodriguez et al. (2018). Increases in the terms of trade correspond to positive external shocks; therefore, its positive coefficient is plausible. On the one hand, specifying the system as an FCVAR model produces an inference gain, since $\mathcal{H}_1^d$ rejects $d = 1$ (16.659 with p value $= 0.000$). On the other hand, the specification tests ($\mathcal{H}_1^\beta$) indicate that the political variable enters the cointegration relation (12,335 with a p-value=0.000). The interest rate and the terms of trade also enter the long-run relationship. From $\mathcal{H}_2^\alpha$ and $\mathcal{H}_3^\alpha$, both the interest rate and the terms of trade turn out to be weakly exogenous (p-values of 0.954 and 0.308, respectively); i.e., they are not error-correcting. As in the previous model, since the interest rate is fixed by the monetary authority and the terms of trade depend solely on international markets, this result is reasonable. Changes in either the interest rate or the terms of trade are corrected by changes in presidential approval because $\mathcal{H}_1^\alpha$ is rejected (p value $= 0.001$ and $\hat{\alpha}_1 = -0.194$). The model is re-estimated taking into account the restrictions $\alpha_2 = \alpha_3 = 0$. The parameters estimated for the restricted model do not differ significantly from the unrestricted ones. There is only a minor decrease in the coefficients estimated for the interest rate and the terms of trade. In addition, the speed of adjustment of the political variable increases slightly ($\hat{\alpha}_1 = -0.196$).

Lastly, Models 1 and 2 suggest that external shocks influence political preferences through exchange rates and the terms of trade, respectively. External influence on presidential approval has two major consequences. First, it might cause political ambiguity, as pointed out in Navia and Osorio (2015) and Campello and Zucco (2016). Specifically, a favorable international context increases the long-run presidential approval equilibrium in Peru, regardless of the current president's ability to conduct domestic economic policies. Second, Ratto (2013) claims that the democratic mechanism of economic voting is weakened by political ambiguity, as citizens can be rewarding bad economic policies that go unperceived because of a favorable external shock.

4.4 Model 3: Presidential Approval, Interest Rate, and Manufacturing Employment

The next model combines presidential approval ratings, the policy interest rate, and manufacturing employment. The results are shown in Table 4. According to Akaike's Information Criterion, the best alternative needs $k = 1$. The cointegration rank test rejects $r = 0$, but a p value $= 0.457$ indicates that it is not possible to reject $r = 1$. The estimation of the model with $k = 1$ and $r = 1$ results in $\hat{d} = 1.152$ with standard error 0.056. The absence of serial correlation is inferred from $Q_{\hat{\varepsilon}}(12) = 129.169$ (p value $= 0.081$). The cointegration vector has a coefficient of 0.339 for the interest rate and 3.253 for employment. Manufacturing performance is a good indicator of overall economic activity, as it has the second largest share in GDP after services. This result provides grounds to suggest that approval ratings can also reach a long-run equilibrium with solely domestic factors.

The CVAR framework is disregarded as a result of rejecting $\mathcal{H}_1^d$ (5.342 with p value $= 0.021$). Additionally, specification tests confirm that the three variables enter the cointegration relationship (see strong rejections for $\mathcal{H}_1^\beta, \mathcal{H}_2^\beta$ and $\mathcal{H}_3^\beta$). Weak exogeneity tests fail to reject $\mathcal{H}_2^\alpha$ (1.822 with p-value=0.177). As explained in previous subsections, the interest rate's weak exogeneity is consistent with a discretionary monetary policy. Presidential approval is error-correcting ($\mathcal{H}_1^\alpha$ is rejected with a p value $= 0.008$). Moreover, the adjustment coefficient of manufacturing employment is positive ($\hat{\alpha}_3 = 0.003$) and different from zero ($\mathcal{H}_3^\alpha$ has a p value $= 0.005$). As an example, consider an increase in the interest rate. This must be followed by a decrease in ν_t in order to maintain the long-run equilibrium. Therefore, in the short run, presidential approval increases because of its negative adjustment coefficient and pushes the system back to equilibrium. In contrast, $\hat{\alpha}_3 > 0$ reduces manufacturing employment and moves the system away from equilibrium, which is plausible given the negative impact of the interest rate on employment. A restricted model is estimated fixing $\alpha_2 = 0$. In general, the parameters do not vary. Only the speed of adjustment of the political variable increases considerably (from $\hat{\alpha}_1 = -0.098$ to $\hat{\alpha}_1 = -0.131$), whereas the coefficient of the interest rate decreased from 0.339 to 0.239. On the other hand, the absolute magnitudes of the coefficients for the speed of adjustment suggest that presidential approval responds to exogenous shocks faster than to manufacturing employment.

Finally, it is important to highlight that the FCVAR approach shows an advantage over previous studies for Peru, which failed to find any significant effect of employment on presidential approval. Stokes (1996) concludes that employment is negatively correlated with presidential approval by using a multinomial logit model. First, an estimation bias can be introduced from not considering endogeneity among explanatory variables. Second, the hypothesis testing results are debatable, since autocorrelation in the political variable is rejected and its long memory behavior is not considered. In contrast, FCVAR accounts for error correction dynamics and fractional orders of integration, respectively. However, it cannot be discarded that analyzing a short period under political reform explains the biased results. Additionally, Maldonado and Pimentel (2013) run an OLS regression and come to the same conclusion. The problems arising from OLS estimation are similar to the ones in Stokes (1996). However, the fact that this paper obtains robust opposite results for the same period seems to confirm the estimation mistakes in those works. In conclusion, there are two benefits from estimating FCVAR Model 3 to analyze presidential approval. First, it considers the long memory of the political support variables. Second, it allows studying the dynamics of presidential approval over long periods, thereby avoiding biased results.

4.5 Model 4: Presidential Approval and Inflation Volatility (Uncertainty)

The literature on economic voting suggests that high inflation and political support are negatively related; see Powell and Whitten (1993). Nevertheless, this hypothesis has been commonly proven for hyperinflationary economies; see for example Stokes (1996). In the case of Peru, it is not possible to identify a long-run relationship between inflation and the political variable during the period of study using the FCVAR approach. A further examination of inflation dynamics, in order to test the economic voting hypothesis, implies estimating its volatility.

A GARCH model is estimated to obtain a measure of inflation to gauge uncertainty. We use monthly CPI inflation from August 2000 to February 2018. First, the number of lags in the mean equation is selected using both inflation autocorrelation functions and Akaike's Information Criterion. Only the lags that are significant at the 5% level are included. LM autocorrelation tests are then run over the squared residuals. The null hypothesis of non-serial autocorrelation is rejected at the 5% level of significance only for the first lag (p value = 0.041). A GARCH (1,1) model is estimated, since it provides a better fit; see Bollerslev (1986):

$$\pi_t = \gamma_0 + \gamma_1\pi_{t-1} + \gamma_2\pi_{t-5} + \gamma_3\pi_{t-8} + \gamma_4\pi_{t-12} + \epsilon_t, \tag{8}$$

$$h_t = \mu + \alpha_1\epsilon_{t-1}^2 + \beta_1 h_{t-1}. \tag{9}$$

Finally, the conditional variance estimates ($\hat{h}_t$) are recovered and an additional FCVAR model is estimated to assess the joint dynamics between inflation volatility (uncertainty) and presidential approval ratings.

The results from estimating Model 4, including presidential approval ratings and estimated inflation volatility, are shown in Table 5. First, the general-to-specific lag selection criterion concludes 1 lag is needed. The rank test does not reject the hypothesis that there is 1 cointegration relationship (p value = 0.385); therefore, Model 4 is estimated using $k = 1$ and $r = 1$.

The estimated integration order, $\hat{d}$, is 0.454 with standard error 0.084. Therefore, the p value of the Ljung-Box Q-test is 0.586, which implies that the residuals are not serially correlated. On the other hand, presidential approval decreases with inflation volatility in the long run (its coefficient is -190.999). In other words, a permanent increase in inflation volatility (uncertainty) of 100 basis points would lead to a decrease in presidential approval ratings (included in log-odds) from, for instance, 50% to 13%. Monetary theory states that high inflation volatility negatively affects the real economy through investment uncertainty. For instance, Judson and Orphanides (1999) present empirical evidence of the negative effects of inflation volatility on growth for a large cross-country dataset. Moreover, they show that this negative correlation also holds during low-inflation periods. Inflation uncertainty also incorporates media information about dramatic events. According to the theory of political equilibrium developed by Ostrom and Smith (1992), it is plausible that inflation volatility (uncertainty) holds a long-run relationship with political support because it is a good proxy for both ordinary and extraordinary events.

The hypothesis that $d = 1$ is strongly rejected (12.481 with p value = 0.000). Both variables enter the cointegration vector, since $\mathcal{H}_1^\beta$ and $\mathcal{H}_2^\beta$ are rejected with 95% confidence. $\mathcal{H}_1^\alpha$ and $\mathcal{H}_2^\alpha$ are also rejected, which means that the two variables respond to changes in political equilibrium. Finally, the coefficients for the speed of adjustment are $\hat{\alpha}_1 = -0.079$ and $\hat{\alpha}_2 = -0.005$, suggesting that both variables are error-correcting and presidential approval adjusts faster to deviations from the long-run equilibrium, since $|\hat{\alpha}_1| > |\hat{\alpha}_2|$. For example, consider a 100-basis-point increase in inflation volatility. This leads to an increase of $190.99 \times 0.01 = 1.901$ in $\hat{\nu}_t$. As a consequence, pa_t and σ_t^2 decrease by $1.901 \times -0.079 = -0.150$ and $1.901 \times -0.005 = -0.001$, respectively.

4.6 Model 5: Presidential Approval, Interest Rate, and Inflation Volatility (Uncertainty)

Table 6 contains the relevant estimation results for Model 5. The selection criteria indicate that the system requires one lag. Next, the hypothesis $r = 1$ cannot be rejected, since the p value of the rank test is 0.937. The estimated integration order is $\hat{d} = 0.732$. Furthermore, the residuals are not serially correlated, as the Ljung-Box Q test's p value = 0.364. From the estimated cointegration vector, presidential approval increases with the interest rate (0.539) and decreases with inflation volatility (-221.049). As asserted by previous models, a high interest rate is associated with positive demand shocks in an economy under inflation targeting, whereas inflation volatility is a proxy for uncertainty.

The hypothesis that $d = 1$ is rejected (10.239 with p value = 0.001). Additionally, the tests on the long-run coefficients reject the null hypotheses, $\mathcal{H}_1^\beta$ and $\mathcal{H}_3^\beta$, with p-values 0.014 and 0.000, respectively; and the interest rate enters the relationship with a p-value of 0.023. The interest rate's significance loss is to be expected, since its movements are mainly due to changes in inflation dynamics. If the interest rate is taken as a proxy for inflation, the results are similar to Judson and Orphanides (1993). They show that both inflation and inflation volatility are important for growth, but the coefficient for inflation loses significance when they are jointly estimated.

Next, the weak exogeneity tests fail to reject $\mathcal{H}_2^\alpha$ (p value = 0.948), again because of discretionary monetary policy. $\mathcal{H}_1^\alpha$ and $\mathcal{H}_3^\alpha$ are rejected with p-values of 0.009 and 0.000, respectively. A correspondent restricted model is estimated with $\alpha_2 = 0.000$. The restricted model estimation results in $\hat{d} = 0.732$. First, residuals are serially uncorrelated ($Q_{\hat{\varepsilon}}(12) = 112.617$ with p value = 0.361). Additionally, the estimated long-run coefficients are barely distinct from the unrestricted model. Finally, presidential approval and inflation volatility are error-correcting, with the same speed of adjustment as in the unrestricted model. As in Model 4, the faster adjustment in political support implies that exogenous shocks have a greater impact on political support than on inflation volatility.

4.7 Model 6: Presidential Approval and Principal Components

This paper further considers a large dataset of macroeconomic variables containing information about Peru's overall economic performance. Pérez, Ghurra, and Grandez (2017) construct a leading indicator for the Peruvian economy by extracting a common factor from six variables: electricity production, cement consumption, sales taxes, chicken sales, mining production, and monthly real GDP. Moreover, since the leading indicator has a correlation of 0.853 with annual GDP growth, the authors highlight its usefulness for forecasting purposes. Since this study is more concerned about the current rather than the future state of the economy, a PCA is performed.

First, the variables used by Pérez, Ghurra, and Grandez (2017) are added to the ones from our previous models: interest and nominal exchange rates, terms of trade, and manufacturing employment. Next, the total variance of the bigger dataset is decomposed in 11 dimensions or components (one per variable). An eigenvalue greater than 1 means that the factor in question provides more information than the individual series, so components 1 and 2 (denoted by ψ_t^1 and ψ_t^2) are kept. Figure 4 shows the five highest eigenvalues. Moreover, components 1 and 2 represent 85.2% of the total variance. In other terms, it is possible to summarize Peru's economic performance in just two variables. Additionally, the share of each variable's variance in the first two components is as follows: cement consumption, sales tax, chicken sales, electricity consumption, and real GDP contribute around 12% each to dimension 1. Additionally, nominal exchange rates, mineral production, and the terms of trade contribute 40%, 25%, and 11% to dimension 2. Finally, Figure 5 plots components 1 and 2. Component 1 (solid red) contains information about domestic demand indicators and follows the pattern of manufacturing employment, whereas the variables in component 2 (dashed blue) are influenced by external factors, mainly the terms of trade.

This paper further tests the economic voting hypothesis by estimating Model 6, which includes our political variable and components 1 and 2. If economic performance holds a long-run relationship with presidential approval, the political variable must cointegrate with the principal components of Peru's economy. The estimation results for Model 6 are presented in Table 7. The lag selection criterion suggests $k = 3$. The first hypothesis, $r = 0$, cannot be rejected at the common confidence levels, since it has a p value $= 0.157$. However, as pointed out by Jones, Nielsen and Popiel (2014), rank tests might have low power in small samples. Thus, it is possible to reject $r = 0$ and continue with the estimation procedure. The model is estimated with $k = 3$ and $r = 1$, given that the second rank test has a p value $= 0.997$. The estimated integration order is 1.191 as a result of the first component's upward trend. Moreover, the Ljung-Box Q test on the residuals concludes that there is no serial autocorrelation (p value $= 0.943$).

Specifically, both long-run coefficients are positive. Nevertheless, the coefficient for the second component, $\hat{\beta}_3 = 0.272$, is almost three times the first one ($\hat{\beta}_2 = 0.092$). Moreover, the coefficients for the speed of adjustment are all negative. The integration order is different from 1 (p-value of 0.010). It is possible to reject $\mathcal{H}_1^\beta$, $\mathcal{H}_2^\beta$, $\mathcal{H}_3^\beta$, and $\mathcal{H}_1^\alpha$ with 99% confidence. Thus, the error-correcting behavior of presidential approval is verified. However, $\mathcal{H}_2^\alpha$ and $\mathcal{H}_3^\alpha$ (p-values of 0.956 and 0.924, respectively) are not rejected, implying that both principal components are weakly exogenous to the system. This result should not raise concerns: Beck (1992) claims that cointegration analysis must consider political science arguments in favor of economic exogeneity, as political preferences respond to changes in the economy, but not vice versa. Moreover, since three lags are included, it is important to note that the system takes one quarter to go back to equilibrium after a shock. Estimation of the restricted model yields very similar results.

The inclusion of principal components in the system enriches the analysis of the long-run relationship, as they summarize the overall state of the economy. As mentioned before, component 1 is obtained from the variance of domestic demand indicators; and component 2 is based on nominal exchange rates, mineral production, and the terms of trade, which are indicators of external shocks on Peru's economy; see Dancourt (2017). The exercise shows that the external component may have a greater impact in moving presidential approval away from its long-run equilibrium than the domestic one. Following Campello and Zucco (2016), this finding suggests that Peruvian voters do not compare economic outcomes with those in foreign countries, which is consistent with the lack of political accountability and ambiguity explained in Model 2.

Kayser and Peress (2012) attribute political benchmarking to the media's efforts to disseminate domestic economic news in comparison with developed countries. Thus, a possible explanation of the absence of benchmarking in Peru is media preferences: they give greater coverage to exchange rate movements and their inflationary consequences than to domestic demand issues, such as the job market. As asserted by Ostrom and Smith (1992), media coverage defines the relevance of news in public sentiment. Additionally, Carlin, Love, and Martinez-Gallardo (2014) show that political accountability is conditioned by the economy: media scandals are more relevant when the economy is not going well. Thus, Peru's exposure to volatile international markets could enhance political instability. As an example, past corruption scandals emerged only under unfavorable international contexts; i.e., when there is a need for political consensus and new economic policies. More research is needed on the topic of information asymmetries through media coverage in Peru and its implications for the political system.

4.8 Model 7: Introducing Corruption

Lastly, a corruption index is added to the variables in Model 2. Concerning the corruption indicator, Google Trends is used to search for the word "corruption" in Peru from January 2004 to February 2018. Google calculates relative popularity by dividing individual data by the total number of searches in a location over a given period. These numbers are then scaled between 0 (the lowest interest recorded) and 100 (the highest). Thus, the higher the number of search results including the word "corruption", the higher the corruption perception and vice versa.

Table 8 shows the estimation results for Model 7. The lag selection criterion indicates $k = 1$. Next, the cointegration rank test strongly rejects both $r = 0$ and $r = 1$, with p values 0.000 and 0.090, respectively. The model is estimated specifying $k = 1$ and two cointegrating vectors, $r = 2$. Furthermore, the Ljung-Box test concludes that the residuals are uncorrelated. The first cointegrating vector excludes the terms of trade and the second one excludes the interest rate. As expected, both cointegrating vectors prove that there is a strong negative long-run relationship between presidential approval and corruption ($\hat{\beta}_2$ is -2.020 in the first vector and -1.868 in the second). At the same time, the estimated coefficients for the interest rate and the terms of trade are compatible with the magnitudes found in the former models: $\hat{\beta}_3 = 0.297$ and $\hat{\beta}_4 = 1.443$. Therefore, they are given the same interpretation. In this case, the tests on the long run-coefficients, β, test the hypothesis of the variable being excluded from the two cointegrating vectors. The hypothesis of exclusion is then rejected for all variables; see the p-values for $\mathcal{H}_1^\beta$, $\mathcal{H}_2^\beta$, $\mathcal{H}_3^\beta$ and $\mathcal{H}_4^\beta$ in Table 8.

Analogously, the hypothesis of the variable being weakly exogenous in the two vectors is assessed by testing α. $\mathcal{H}_1^\alpha$, $\mathcal{H}_2^\alpha$ and $\mathcal{H}_3^\alpha$ (p-values of 0.001, 0.078 and 0.000, respectively) are rejected. However, it is not possible to discard $\mathcal{H}_4^\alpha$, since it has a p value $= 0.603$. As explained before, exogeneity of the terms of trade ($\alpha_4 = 0$) is plausible, since they are determined by the international markets. Presidential approval is error-correcting in both vectors, $\hat{\alpha}_1 = -0.095$ and -0.102. The adjustment of the other variables is ambiguous. In the first vector, corruption and the interest rate deviate the system from equilibrium, while they are error-correcting in the second.

Corruption perception does respond to changes in both the economy and presidential approval. For example, corruption scandals might go unperceived if the economy is doing well and the president's popularity is high. It can also be argued that there is more space in the media for political and corruption scandals during good economic times. Thus, the response to corruption perception pushes the system back or away from its equilibrium, depending on the context. Additionally, the interest rate responds to uncertainty control in the same way as corruption perception. More importantly, the estimation results make clear that economic variables do not lose significance in the cointegrating vector once corruption is included. Finally, these results are compatible with the theory developed by Ostrom and Smith (1992). Corruption scandals enter the cointegrating vector because they have a permanent effect on presidential approval.

5 Other Estimation Results

As mentioned, there is no agreement on which variables enter voters' utility function. Therefore, different macroeconomic variables are combined as a proxy for overall wellbeing indicators. The search of long-run relationships between the economy and presidential approval implies estimating several models. Nevertheless, the cointegration relationship is not verified in all of them. The estimation results in this section are available upon request.

First, monthly real GDP, both in levels and growth rates, is not significant when explaining presidential approval dynamics. As a possible explanation, these series may not be informative about the general state of the economy due to publication delays; see Perez, Ghurra and Grandez (2017). Specifically, measuring aggregate production (GDP) is too complex (or debatable) a task to be completed in a month without noticeable errors.

On the other hand, contrary to expectations, it was not possible to find any cointegrating vector including presidential approval and inflation. The estimation results suggest that the model residuals are serially correlated whenever inflation is included. The same occurs with the inflation expectations of both individuals and firms. However, Models 4 and 5 show that inflation volatility is significant for presidential approval, indicating that inflation uncertainty is the relevant variable.

The rank test indicates that there is a cointegrating vector including presidential approval, the interest rate, and unemployment in Lima. Even though the estimated long-run unemployment coefficient is negative, as expected, it is not significant. Unemployment in Peru is poorly estimated as a consequence of high informal employment (about 70% of labor force). Thus, it is not a good indicator of overall economic performance.

Finally, the HP filter is applied to obtain the secular component of the series used. As in the case of CPI inflation, it is not possible to find any cointegrating vector and the estimated residuals present strong serial correlation.

6 Conclusions

This paper aims to show that presidential approval ratings hold a long-run relationship with macroeconomic variables in Peru. To that end, seven FCVAR models containing both types of series are estimated. Models 1 through 6 clearly show that the economic voting hypothesis holds for Peru in 2001-2018. Interest and nominal exchange rates, terms of trade, manufacturing employment, inflation volatility and principal components are cointegrated with presidential approval ratings. In the long run, presidential approval increases with the monetary policy interest rate, the terms of trade, manufacturing employment, and principal components. Conversely, approval ratings decrease as nominal exchange rates and inflation volatility increase. Moreover, corruption perception shows a negative relationship with presidential approval in Model 7. Additionally, economic variables do not lose significance when controlling for corruption perception: although there is a common belief that presidential approval in Peru is driven only by corruption and media scandals, this evidence is consistent with the economic voting theory. Lastly, the error-correcting behavior of the political variable in all the models implies that presidential approval responds to changes in economic performance.

The economic voting theory also states that the relationship between political and economic variables works as a democratic mechanism. Voters are able to communicate their conformity or dissatisfaction with the government's economic policies through presidential approval ratings; see

Ratto (2013). However, this paper finds that, as a consequence of external economic dependence, presidential approval ratings in Peru are very responsive to changes in the international context.

Therefore, it is not possible to discard the existence of political ambiguity (i.e., high political support due only to favorable external shocks) in Peru, as pointed out by Navia and Osorio (2015). It is important to highlight that, according to Campello and Zucco (2016), lack of political accountability encourages rent-seeking, one of the most important political issues in Latin America. More research is needed on information asymmetries through media coverage and its implications for Peru's political system.

To the authors' best knowledge, this is the first application of fractional cointegration analysis to political economy in Latin America. They hope the results of this research motivates political science to use the FCVAR approach to explain the intricate dynamics of political support in the region.

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Table 1. Univariate Analysis

	GPH Estimates			FAR(k) Estimates					
	$m = T^{0.5}$	$m = T^{0.6}$	$m = T^{0.7}$	$k = 0$		$k = 1$		$k = 2$	
	$\hat{d}$	$\hat{d}$	$\hat{d}$	$\hat{d}$	$Q_{\hat{\epsilon}}$	$\hat{d}$	$Q_{\hat{\epsilon}}$	$\hat{d}$	$Q_{\hat{\epsilon}}$
pa_t	0.863	0.611	0.900	0.941	21.473	0.793	18.484	0.795	18.949
	(0.129)	(0.129)	(0.141)	(0.329)	(0.044)	(0.329)	(**0.102**)	(0.220)	(**0.090**)
l_t	1.288	1.264	1.209	1.181	18.336	0.582	18.522	0.693	13.392
	(0.073)	(0.061)	(0.048)	(0.047)	(0.**106**)	(0.029)	(0.**101**)	(0.046)	(0.**807**)
i_t	0.473	0.665	0.891	1.272	19.357	0.748	20.490	0.786	20.496
	(0.245)	(0.148)	(0.120)	(0.066)	(0.008)	(0.082)	(**0.058**)	(0.151)	(**0.059**)
e_t	1.103	1.210	1.219	1.272	32.955	1.106	26.415	1.238	18.697
	(0.225)	(0.184)	(0.133)	(0.066)	(0.001)	(0.125)	(0.009)	(0.103)	(**0.096**)
tot_t	0.879	0.890	0.999	1.120	9.972	1.100	10.189	0.392	10.325
	(0.192)	(0.128)	(0.086)	(0.058)	(**0.322**)	(0.092)	(**0.599**)	(0.083)	(**0.296**)
σ_t^2	0.039	0.073	0.052	0.076	7.253	0.058	7.208	0.125	3.805
	(0.181)	(0.139)	(0.107)	(0.064)	(**0.840**)	(0.113)	(**0.844**)	(0.158)	(**0.987**)
ψ_t^1	1.037	1.053	1.007	0.852	46.282	1.021	28.677	1.118	18.493
	(0.045)	(0.028)	(0.021)	(0.043)	(0.000)	(0.044)	(0.004)	(0.045)	(**0.102**)
ψ_t^2	1.174	1.192	1.089	1.082	26.073	1.188	19.399	1.208	12.103
	(0.239)	(0.163)	(0.112)	(0.055)	(0.010)	(0.067)	(0.**079**)	(0.075)	(**0.437**)
c_t	0.898	0.691	0.294	0.378	30.514	0.396	30.827	0.587	27.530
	(0.202)	(0.173)	(0.144)	(0.060)	(0.002)	(0.136)	(0.002)	(0.082)	(0.006)

GPH denotes Geweke and Porter-Hudak and FAR denotes fractional AR model. Inflation volatility (σ_t^2) does not have long memory. Corruption perception (c_t) estimates are based on a smaller sample and its residuals are uncorrelated with $k = 3$.

Table 2. FCVAR Results for Model 1
Presidential Approval Ratings, Interest and Nominal Exchange Rates

Rank Tests:

Rank	$\hat{d}$	Log-Likelihood	LR Statistic	P value
0	1.090	517.800	38.154	0.035
1	1.160	532.955	7.843	0.919
2	1.183	536.333	1.088	0.978
3	1.187	536.877	-	-

Unrestricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ i_t \\ e_t \end{bmatrix} - \begin{bmatrix} -0.118 \\ 4.373 \\ 1.249 \end{bmatrix} \right) = L_{\hat{d}} \begin{bmatrix} -0.210 \\ 0.077 \\ -0.005 \end{bmatrix} \nu_t + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.042)}{1.160}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.107)}{126.544}, \quad \log(\mathcal{L}) = 532.955$$

Equilibrium Relation:

$$pa_t = 5.429 + 0.132 i_t - 4.801 e_t + \nu_t.$$

Hypothesis Tests:

	$\mathcal{H}_1^d$	$\mathcal{H}_1^\beta$	$\mathcal{H}_2^\beta$	$\mathcal{H}_3^\beta$	$\mathcal{H}_1^\alpha$	$\mathcal{H}_2^\alpha$	$\mathcal{H}_3^\alpha$
df	1	1	1	1	1	1	1
LR	16.367	24.721	6.929	19.205	19.107	0.507	4.055
P value	0.000	0.000	0.008	0.000	0.000	**0.476**	0.044

Restricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ i_t \\ e_t \end{bmatrix} - \begin{bmatrix} -0.169 \\ 4.381 \\ 1.248 \end{bmatrix} \right) = L_d \begin{bmatrix} -0.218 \\ 0.000 \\ -0.005 \end{bmatrix} \nu_t + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.042)}{1.158}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.117)}{125.750}, \quad \log(\mathcal{L}) = 532.702$$

Equilibrium Relation:

$$pa_t = 5.222 + 0.112 i_t - 4.713 e_t + \nu_t.$$

Table 3. FCVAR Results for Model 2
Presidential Approval Ratings, Interest Rate and Terms of Trade

Rank Tests:

Rank	$\hat{d}$	Log-Likelihood	LR Statistic	P value
0	1.117	426.470	35.697	0.072
1	1.160	438.626	11.385	0.616
2	1.162	444.201	0.235	1.000
3	1.164	444.319	-	-

Unrestricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ i_t \\ tot_t \end{bmatrix} - \begin{bmatrix} -0.282 \\ 5.887 \\ 4.050 \end{bmatrix} \right) = L_{\hat{d}} \begin{bmatrix} -0.194 \\ 0.005 \\ 0.005 \end{bmatrix} \nu_t + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.041)}{1.160}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.117)}{125.730}, \qquad \log(\mathcal{L}) = 438.626$$

Equilibrium Relation:

$$pa_t = -7.757 + 0.180 i_t + 1.584 tot_t + \nu_t.$$

Hypothesis Tests:

	$\mathcal{H}_1^d$	$\mathcal{H}_1^\beta$	$\mathcal{H}_2^\beta$	$\mathcal{H}_3^\beta$	$\mathcal{H}_1^\alpha$	$\mathcal{H}_2^\alpha$	$\mathcal{H}_3^\alpha$
df	1	1	1	1	1	1	1
LR	16.659	12.335	5.272	10.474	11.838	0.003	1.040
P value	0.000	0.000	0.022	0.001	0.001	**0.954**	**0.308**

Restricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ i_t \\ tot_t \end{bmatrix} - \begin{bmatrix} -0.273 \\ 5.865 \\ 4.049 \end{bmatrix} \right) = L_d \begin{bmatrix} -0.196 \\ 0.000 \\ 0.000 \end{bmatrix} \nu_t + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.039)}{1.161}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.086)}{128.629}, \qquad \log(\mathcal{L}) = 438.106$$

Equilibrium Relation:

$$pa_t = -7.007 + 0.152 i_t + 1.443 tot_t + \nu_t.$$

Table 4. FCVAR Results for Model 3
Presidential Approval Ratings, Interest Rate and Industrial Employment

Rank Tests:

Rank	$\hat{d}$	Log-Likelihood	LR Statistic	P value
0	1.128	702.905	43.045	0.011
1	1.152	717.929	12.996	0.457
2	1.195	724.389	0.077	1.000
3	1.197	724.428	-	-

Unrestricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ i_t \\ l_t \end{bmatrix} - \begin{bmatrix} -0.054 \\ 5.681 \\ 4.311 \end{bmatrix} \right) = L_{\hat{d}} \begin{bmatrix} -0.098 \\ 0.070 \\ 0.003 \end{bmatrix} \nu_t + \hat{\Gamma} \Delta^{\hat{d}} L_{\hat{d}} (X_t - \hat{\mu}) + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.056)}{1.152}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.081)}{129.169}, \qquad \log(\mathcal{L}) = 717.929$$

Equilibrium Relation:

$$pa_t = -16.004 + 0.339 i_t + 3.253 l_t + \nu_t.$$

Hypothesis Tests:

	$\mathcal{H}_1^d$	$\mathcal{H}_1^\beta$	$\mathcal{H}_2^\beta$	$\mathcal{H}_3^\beta$	$\mathcal{H}_1^\alpha$	$\mathcal{H}_2^\alpha$	$\mathcal{H}_3^\alpha$
df	1	1	1	1	1	1	1
LR	5.342	11.297	10.179	16.135	7.146	1.822	7.818
P value	0.021	0.001	0.000	0.000	0.008	**0.177**	0.005

Restricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ i_t \\ l_t \end{bmatrix} - \begin{bmatrix} -0.213 \\ 5.733 \\ 4.310 \end{bmatrix} \right) = L_d \begin{bmatrix} -0.131 \\ 0.000 \\ 0.004 \end{bmatrix} \nu_t + \hat{\Gamma} \Delta^{\hat{d}} L_{\hat{d}} (X_t - \hat{\mu}) + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.058)}{1.134}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.084)}{128.811}, \qquad \log(\mathcal{L}) = 717.018$$

Equilibrium Relation:

$$pa_t = -15.517 + 0.239 i_t + 3.233 l_t + \nu_t.$$

Table 5. FCVAR Results for Model 4
Presidential Approval Ratings and Inflation Volatility

Rank Tests:

Rank	$\hat{d}$	Log-Likelihood	LR Statistic	P value
0	0.456	687.278	26.042	0.000
1	0.454	699.922	0.753	0.385
2	0.401	700.299	-	-

Unrestricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ \sigma_t^2 \end{bmatrix} - \begin{bmatrix} -0.181 \\ 0.084 \end{bmatrix} \right) = L_{\hat{d}} \begin{bmatrix} -0.079 \\ -0.005 \end{bmatrix} \nu_t + \hat{\Gamma}\Delta^{\hat{d}} L_{\hat{d}}(X_t - \hat{\mu}) + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.084)}{0.454}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.586)}{45.246}, \qquad \log(\mathcal{L}) = 699.922$$

Equilibrium Relation:

$$pa_t = 15.863 - 190.999\sigma_t^2 + \nu_t.$$

Hypothesis Tests:

	$\mathcal{H}_1^d$	$\mathcal{H}_1^\beta$	$\mathcal{H}_2^\beta$	$\mathcal{H}_1^\alpha$	$\mathcal{H}_2^\alpha$
df	1	1	1	1	1
LR	12.481	5.940	22.884	8.835	15.184
P value	0.000	0.015	0.000	0.003	0.000

Table 6. FCVAR Results for Model 5
Presidential Approval Ratings, Interest Rate and Inflation Volatility

Rank Tests:

Rank	$\hat{d}$	Log-Likelihood	LR Statistic	P value
0	0.662	640.419	41.826	0.000
1	0.732	659.762	3.140	0.937
2	0.715	661.224	0.217	0.946
3	0.728	661.332	-	-

Unrestricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ i_t \\ \sigma_t^2 \end{bmatrix} - \begin{bmatrix} -0.138 \\ 5.762 \\ 0.088 \end{bmatrix} \right) = L_{\hat{d}} \begin{bmatrix} -0.030 \\ -0.001 \\ -0.003 \end{bmatrix} \nu_t + \hat{\Gamma}\Delta^{\hat{d}} L_{\hat{d}}(X_t - \hat{\mu}) + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.061)}{0.732}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.364)}{112.528}, \qquad \log(\mathcal{L}) = 659.762$$

Equilibrium Relation:

$$pa_t = 16.201 + 0.539 i_t - 221.049 \sigma_t^2 + \nu_t.$$

Hypothesis Tests:

	$\mathcal{H}_1^d$	$\mathcal{H}_1^\beta$	$\mathcal{H}_2^\beta$	$\mathcal{H}_3^\beta$	$\mathcal{H}_1^\alpha$	$\mathcal{H}_2^\alpha$	$\mathcal{H}_3^\alpha$
df	1	1	1	1	1	1	1
LR	10.239	6.047	5.152	34.033	6.855	0.004	21.942
P value	0.001	0.014	0.023	0.000	0.009	**0.948**	0.000

Restricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ i_t \\ \sigma_t^2 \end{bmatrix} - \begin{bmatrix} -0.138 \\ 5.763 \\ 0.088 \end{bmatrix} \right) = L_d \begin{bmatrix} -0.030 \\ 0.000 \\ -0.003 \end{bmatrix} \nu_t + \hat{\Gamma}\Delta^{\hat{d}} L_{\hat{d}}(X_t - \hat{\mu}) + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.061)}{0.732}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.361)}{112.617}, \qquad \log(\mathcal{L}) = 659.760$$

Equilibrium Relation:

$$pa_t = 16.119 + 0.547 i_t - 220.556 \sigma_t^2 + \nu_t.$$

Table 7. FCVAR Results for Model 6
Presidential Approval Ratings, Components 1 and 2

Rank Tests:

Rank	$\hat{d}$	Log-Likelihood	LR Statistic	P value
0	1.151	173.366	32.616	0.157
1	1.191	187.054	5.241	0.997
2	1.147	189.052	1.244	0.956
3	1.154	189.674	-	-

Unrestricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ \psi_t^1 \\ \psi_t^2 \end{bmatrix} - \begin{bmatrix} -0.360 \\ -5.111 \\ 0.290 \end{bmatrix} \right) = L_{\hat{d}} \begin{bmatrix} -0.264 \\ -0.003 \\ -0.004 \end{bmatrix} \nu_t + \sum_{i=1}^{3} \Gamma_i \Delta^d L_d^i (X_t - \hat{\mu}) + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.040)}{1.191}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.943)}{85.794}, \qquad \log(\mathcal{L}) = 187.054$$

Equilibrium Relation:

$$pa_t = 0.031 + 0.092\psi_t^1 + 0.272\psi_t^2 + \nu_t.$$

Hypothesis Tests:

	$\mathcal{H}_1^d$	$\mathcal{H}_1^\beta$	$\mathcal{H}_2^\beta$	$\mathcal{H}_3^\beta$	$\mathcal{H}_1^\alpha$	$\mathcal{H}_2^\alpha$	$\mathcal{H}_3^\alpha$
df	1	1	1	1	1	1	1
LR	6.606	19.723	9.698	13.770	10.241	0.003	0.009
P value	0.010	0.000	0.002	0.000	0.001	**0.956**	**0.924**

Restricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ \psi_t^1 \\ \psi_t^2 \end{bmatrix} - \begin{bmatrix} -0.364 \\ -5.111 \\ 0.291 \end{bmatrix} \right) = L_d \begin{bmatrix} -0.264 \\ 0.000 \\ 0.000 \end{bmatrix} \nu_t + \sum_{i=1}^{3} \Gamma_i \Delta^d L_d^i (X_t - \hat{\mu}) + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.040)}{1.191}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.944)}{85.730}, \qquad \log(\mathcal{L}) = 187.047$$

Equilibrium Relation:

$$pa_t = 0.027 + 0.092\psi_t^1 + 0.273\psi_t^2 + \nu_t.$$

Table 8. FCVAR Results for Model 7
Presidential Approval Ratings, Google Corruption Index, Interest Rate and Terms of Trade

Rank Tests:

Rank	$\hat{d}$	Log-Likelihood	LR Statistic	P value
0	0.733	477.029	70.199	0.000
1	0.631	501.023	22.211	0.090
2	0.653	511.166	1.925	0.964
3	0.676	512.118	0.020	0.994
4	0.678	512.128	-	-

Hypothesis Tests:

	$\mathcal{H}_1^d$	$\mathcal{H}_1^\beta$	$\mathcal{H}_2^\beta$	$\mathcal{H}_3^\beta$	$\mathcal{H}_4^\beta$	$\mathcal{H}_1^\alpha$	$\mathcal{H}_2^\alpha$	$\mathcal{H}_3^\alpha$	$\mathcal{H}_4^\alpha$
df	1	2	2	2	2	2	2	2	2
LR	29.389	16.557	26.666	26.414	14.773	13.284	5.092	31.922	1.013
P value	0.000	0.000	0.000	0.000	0.001	0.001	0.078	0.000	**0.603**

Restricted Model:

$$\Delta^d \left(\begin{bmatrix} pa_t \\ c_t \\ i_t \\ tot_t \end{bmatrix} - \begin{bmatrix} -0.918 \\ -0.236 \\ 2.599 \\ 4.252 \end{bmatrix} \right) = L_d \begin{bmatrix} -0.095 & -0.102 \\ 0.041 & -0.237 \\ 0.354 & -0.529 \\ 0.000 & 0.000 \end{bmatrix} \begin{bmatrix} v_{1t} \\ v_{2t} \end{bmatrix} + \hat{\Gamma} \Delta^{\hat{d}} L_{\hat{d}} (X_t - \hat{\mu}) + \hat{\epsilon}_t$$

$$\hat{d} = \underset{(0.043)}{0.645}, \quad Q_{\hat{\varepsilon}}(12) = \underset{(0.299)}{201.810}, \qquad \log(\mathcal{L}) = 510.659$$

Equilibrium Relation:

$$pa_t = -2.167 - 2.020 c_t + 0.297 i_t + \nu_{1t}.$$
$$pa_t = -7.495 - 1.868 c_t + 1.443 tot_t + \nu_{2t}.$$

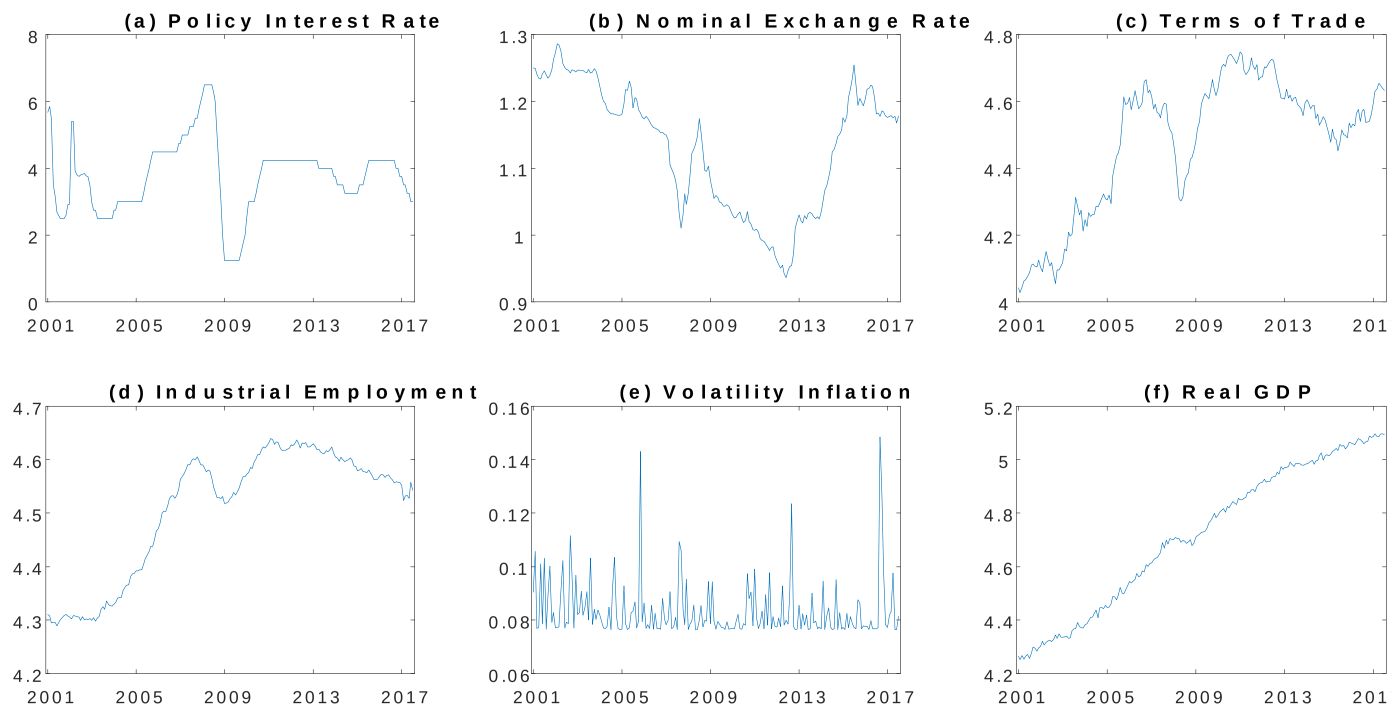

Figure 1. Variables in Levels

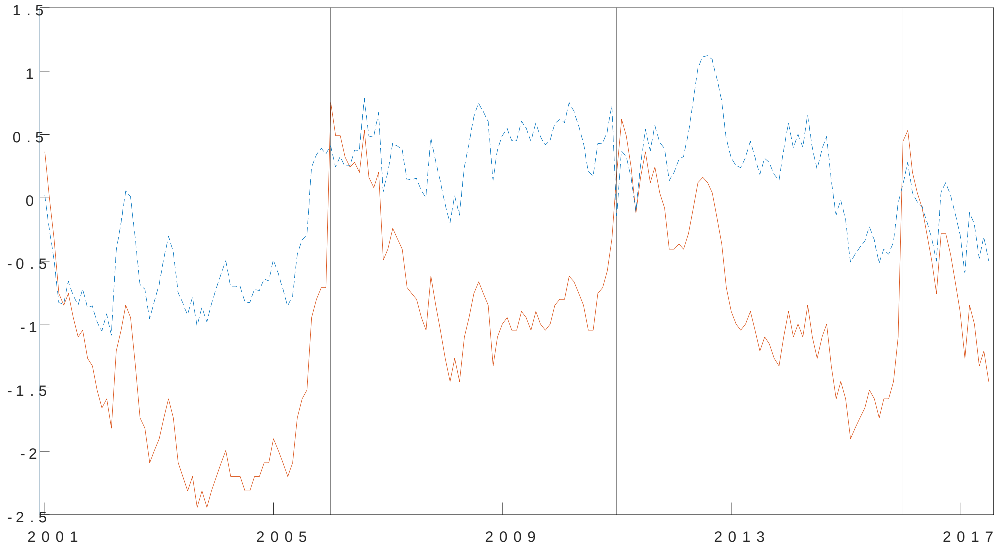

Figure 2a. President of Peru's Approval Ratings (solid-red) and Filtered Series (dashed-blue)

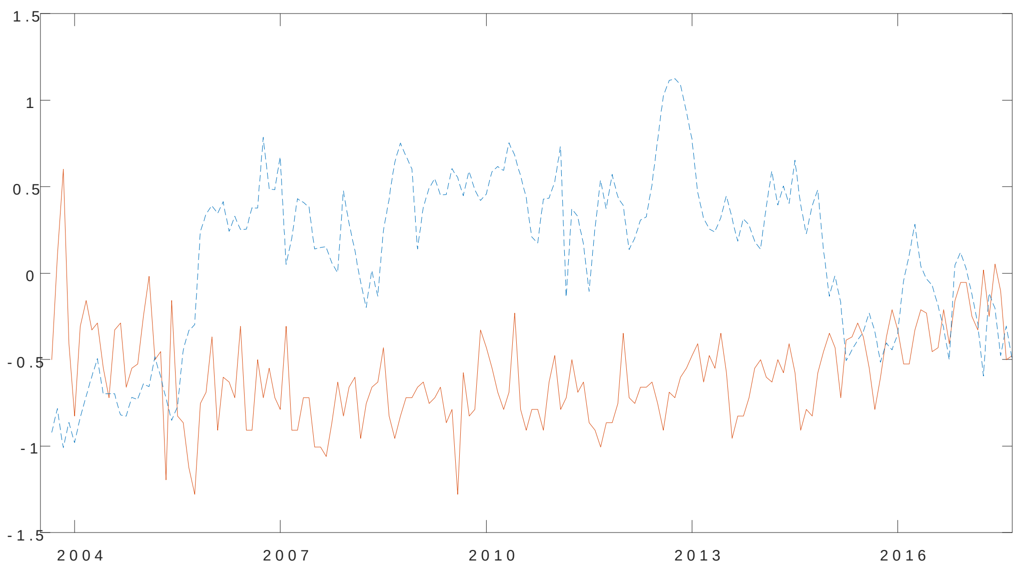

Figure 2b. Corruption Index (solid-red) and Approval Ratings (dashed-blue).

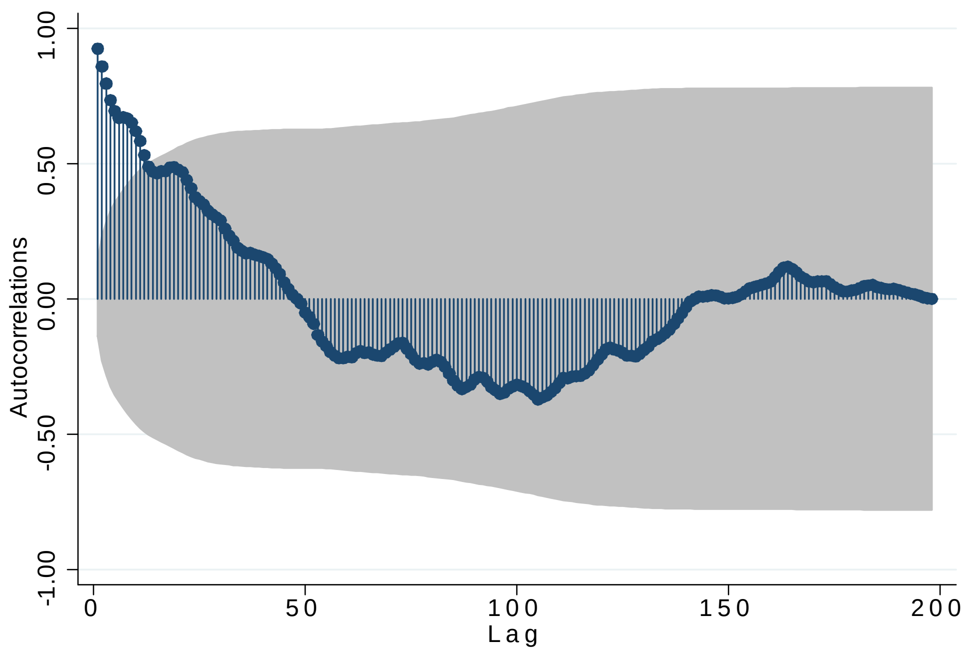

Figure 3a. Autocorrelation Function for President of Peru's Approval Ratings (Bartlett's formula for MA(q) 95% confidence bands)

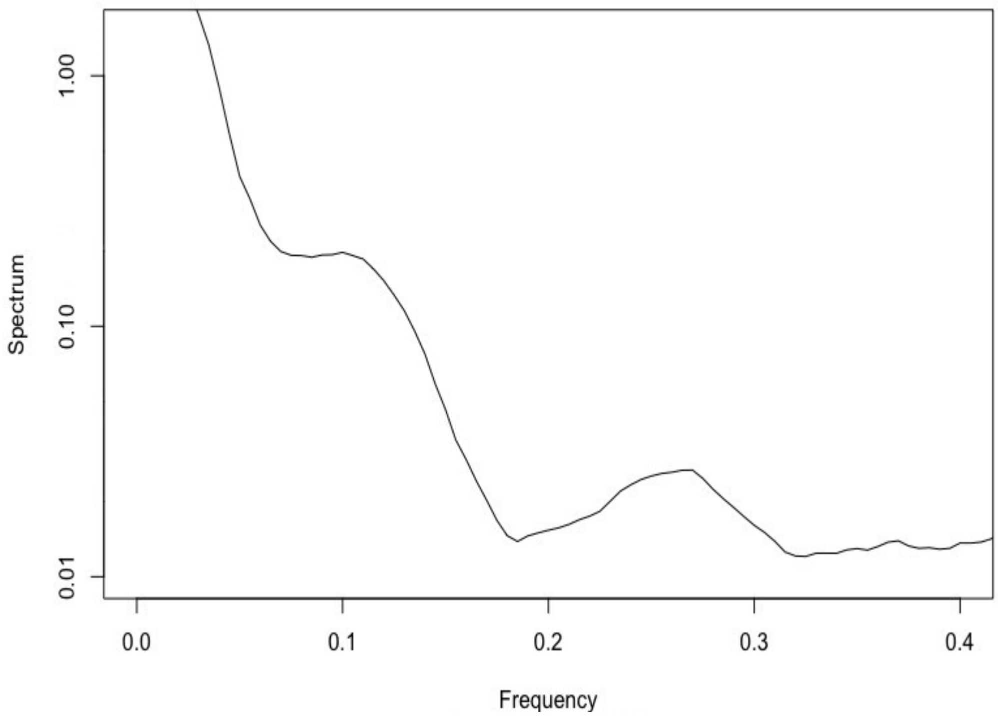

Figure 3b. Spectral density for President of Peru's Approval Ratings (bandwith=0.0183)

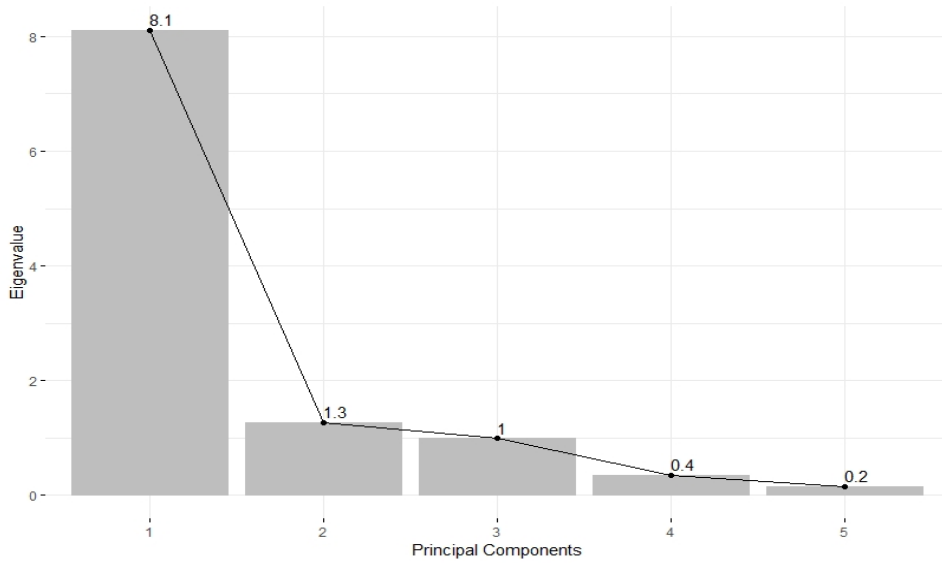

Figure 4. Eigenvalues of Principal Components

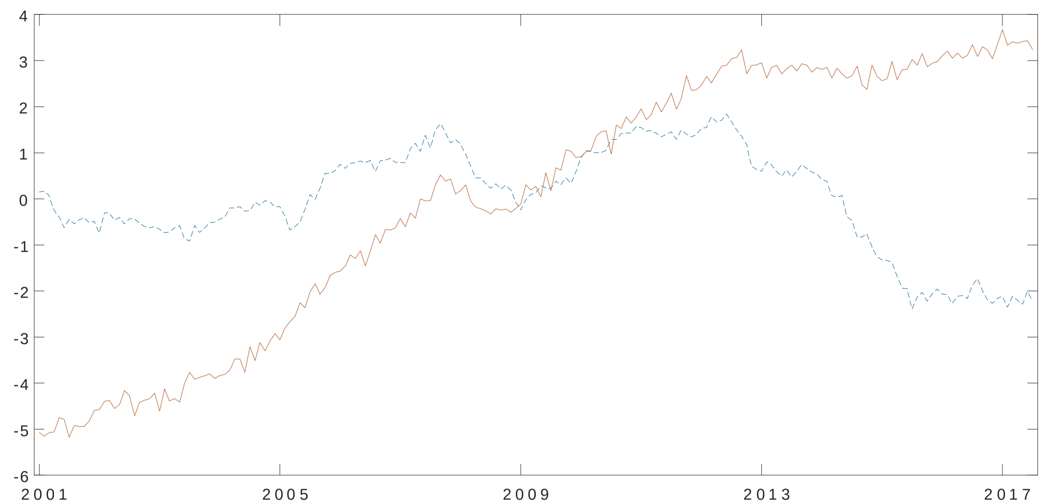

Figure 5. Principipal Components 1 (solid-red) and 2 (dashed-blue)

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