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UNRAVELING THE COMPLEXITY OF U.S. PRESIDENTIAL APPROVAL
—
A MULTI-DIMENSIONAL SEMI-PARAMETRIC APPROACH*

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Abstract

In this paper we show that findings of an apparently instable popularity function of U.S. presidents, as reported in the previous literature, are likely the consequence of the common use of linear estimation techniques. Employing Penalized Spline Smoothing in the context of Additive Mixed Models we allow for a-priori unspecified non-linear effects of possible economic determinants of presidential popularity. We find strong evidence for non-linear and negative effects of unemployment, inflation and government consumption on presidential approval and present empirical evidence in favor of the hypothesis of the existence of interaction effects between the economic variables. Additionally we give supporting evidence for the existence of honeymoon and nostalgia effects as well as general decline of support over time.

Keywords: Presidential Popularity, Approval, Penalized Splines, Mixed Models

JEL classification: C14, C32, E02, H11

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1 Introduction

In the aftermath of the pioneer work of Mueller [1970] a quickly growing empirical literature on the determinants of the popularity of U.S. presidents evolved. While it is more or less uncontroversial in this literature that political events as well as foreign conflicts exert significant effects on presidential popularity, the influence of the stance of the economy on presidential approval is less clear. Although there is a widespread belief that voters hold governments accountable for economic outcomes (see e.g. Norpoth (1984), Kiefer (1997), Gronke and Brehm (2002)) the empirical evidence summarized in Berlemann and Enkelmann [2012] indicates that the effects of economic variables on presidential popularity vary enormously between the existing studies. While proxies of inflation turned out to be significant in roughly 60% of all reported estimations, unemployment variables performed even poorer and delivered significant and plausible coefficients in only every second study where they were used. Berlemann and Enkelmann [2012] also show that the coefficients of economic variables turn out to be highly unstable when running the typically employed linear regression models for randomly chosen sub-periods. A possible reason for this finding is the inadequacy of the typically employed linear regression approach. Whenever the "true" relationship between presidential popularity and variables mirroring the stance of the economy is non-linear, the findings of linear regressions will strongly depend on the chosen sample period.

Up to now, non-linear estimation approaches have been used very rarely to estimate popularity functions. A few studies (e.g. Mueller [1970]) employ asymmetrically defined economic variables to account for possible non-linearities. A few others (see e.g. Smyth et al. [1991, 1994, 1995, 1999], Smyth and Taylor [2003]) use squared economic variables in their linear regressions thereby assuming the detrimental effects of unemployment and inflation to be smaller at lower levels of inflation and unemployment. However, while these approaches are first steps into the direction of a more general analysis of the effects of economic variables on presidential popularity, they nevertheless again make quite special and thus highly controversial assumptions on the exact type of non-linearity we possibly deal with.

In this paper we go beyond the existing approaches and employ non-linear estimation techniques to study the relation between U.S. presidential popularity and (possibly) important economic variables such as inflation, unemployment and government consumption. In order to do so we estimate semi-parametric Additive Mixed Models. In this approach the economic variables enter the estimation equation as a-priori unspecified functions to be estimated from the data. As a consequence, the relation between presidential popularity and economic variables might take any functional form. While we start out with studying additive effects we also investigate possible high-dimensional interaction effects between the economic covariates. While our focus is on the economic determinants of presidential popularity, we also allow for a non-linear time-in-office effect and non-linear effects of large wars on presidential popularity. Political events as well as smaller foreign conflicts enter the estimation equation as usual in the form of binary-coded variables (see Newman and Forcehimes [2010]).

We find significant non-linear effects for unemployment and inflation while government consumption (as a percentage of GDP) turns out to have an almost linear negative effect on presidential popularity. While the unemployment rate has little effect on presidential popularity on low unemployment levels, increases in the unemployment rate exceeding a threshold level of about 4% turns out to be detrimental to presidential popularity. Rising inflation rates in general tend to be harmful for the president. However, the strength of the effect depends strongly on the level of inflation. For inflation rates in between 4% and 7% presidential popularity tends to be unaffected by the economy's inflation performance. We also find supporting evidence for a general decline of presidential approval over time which is - however - reversed about one year before a presidential election. Moreover, we detect significant *honeymoon* effects.

Finally we find strong evidence for the existence of interaction effects between the economic covariates. Allowing for interaction effects between the economic covariates increases the explanatory power of the estimation results significantly. We might take this as an indication that the usually assumed additivity or even linearity of the effects of economic variables on presidential popularity is highly implausible. In reality, voters seem to make judgements on the economic situation and the government's undertaken measures as a whole.

The paper is organized as follows: The second section delivers a brief overview on the related literature. The third section introduces the applied non-linear estimation technique and section 4 describes the employed dataset. The estimation results are presented and discussed in section 5. In section 6, the analysis is extended to interaction effects between inflation, unemployment and government consumption. Section 7 summarizes and draws conclusions.

2 Review of the Related Literature

Over the last four decades, an extensive body of literature on vote and popularity functions evolved.[1] Vote and popularity functions have been estimated for numerous countries and differing sample periods. The existing empirical studies differ considerably in the used economic and political variables as well as the employed estimation techniques and model specifications. We make no attempt to summarize this literature here in length.[2] Instead, we focus the review on the aspect whether and how the existing studies control for possible non-linearities between presidential popularity and economic variables.

The vast majority of empirical popularity studies assumes a linear relationship between presidential approval and its determinants. In these studies, the employed popularity measure is typically

[1] It should be noted that *approval* and *popularity* are "conceptionally distinct" (Stimson [1976]). However, we will neglect this semantic subtlety and use the terms *popularity*, *approval* or *political support* interchangeably, as it is common in the field.

[2] For recent overviews on the literature see, for example, Paldam [2008], Bellucci and Lewis-Beck [2011] or Berlemann and Enkelmann [2012].

regressed on the level of a number of political and economic variables using the ordinary least squares estimator (OLS). This standard approach has rarely been questioned or discussed.

Interestingly enough, it was the pioneer study by MUELLER [1970] which already experimented with a (quasi) non-linear relationship between economic conditions and presidential approval. MUELLER's [1970] *economic slump* variable is generally defined as the difference between the unemployment rate at the beginning of the president's term and the unemployment rate at the time of the poll, but it is set to zero whenever this difference turned out to be negative. For a particular presidency, this transformation implies a kinked – hence, non-linear – relationship between government popularity and the unemployment rate. The introduction of this sort of non-linearity, however, was not motivated by economic or political theory, but chosen to make the data "come out right".

SMYTH, WASHBURN and DUA [1989] argue that "[t]he linearity assumption is inconsistent with standard utility theory" since such popularity functions produce a "preference map [that] consists of linear indifference curves". As a consequence, SMYTH and his co-authors in a number of papers addressed the issue of non-linearity by using squared terms of unemployment and inflation in popularity functions (SMYTH ET AL.[1991, 1994, 1995, 1999], SMYTH and TAYLOR [2003]).[3] A negative coefficient of a squared variable implies that the detrimental effects of these variables increase in the level of the referring variable. In general, SMYTH and his co-authors finds supporting evidence for squared economic variables such as inflation and unemployment. However, from a statistical perspective the choice of squared terms is as arbitrary as using a linear specification. Moreover, the authors include only a quadratic but no linear term of unemployment and inflation, thereby forcing the quadratic function to be symmetrical to the ordinate axis. Doing so is questionable, too.

BELLUCCI and LEWIS-BECK [2011] use interaction terms to model non-linear effects. In their cross-national study, the authors interact the perception of the state of the economy with a *clarity of responsibility* variable. The rationale behind this approach is that a worsening of the economic conditions has a stronger effect on approval ratings when this development can clearly be assigned to a single party or person. In times of divided governments, the impact of the economy on approval rates should be smaller. However, BELLUCCI and LEWIS-BECK find no significant effect for the interaction term. Hence, they find the relationship between presidential approval and the economic conditions to be again linear.

Non-linearities have also been studied for non-economic determinants of presidential approval. Repeatedly, the impact of time has been modeled in a non-linear fashion to capture the dynamics of presidential approval over the electoral cycle (e.g. BELLUCCI and LEWIS-BECK [2011], STIMSON [1976]). Often it is assumed that government popularity follows a U-shaped pattern between two elections – a honeymoon period followed by a phase of decline and, finally, a pre-election rebound due to campaign or farewell effects. In their early study of UK government popularity, GOODHART and BHANSALI [1970] call this pattern the "natural path of government popularity". BELLUCCI

[3] To our knowledge, the first to include squared economic variables into a popularity function was YANTEK [1988].

and Lewis-Beck [2011] include both a time-in-office variable and its squared counterpart to study the existence of a U-shaped time-pattern. However, the use of a linear-quadratic functional form is again quite arbitrary.

Summing up, one might conclude that the issue of non-linearities has only rarely been touched upon within the vote and popularity function literature. While linear estimation approaches dominate the literature, the concrete functional form of the covariates has almost always been chosen in an arbitrary and rather pre-defined way. One might suspect that the shortcoming to allow for more complex, possibly non-linear relationships between popularity and its determinants has contributed to the inconclusive findings of the existing literature on the (economic) determinants of presidential approval.

3 Estimation Approach

As outlined in the previous section, the predominant approach of estimating popularity functions follows the idea that the response or endogenous variable y depends on some covariates $x_1, \ldots, x_p$ in a linear fashion

$$y_i = \beta_0 + x_{i1}\beta_1 + \cdots + x_{ip}\beta_p + \epsilon_i, \tag{1}$$

with $\epsilon \sim N(0, \sigma^2)$ for $i = 1, \ldots, n$. Note, that defining $x_{p+1} = x_p^2$ and adding this quadratic component to (1) still yields a model which is linear *in the effects.*

Although the linear approach is both computationally efficient and easy to interpret, it might be too simplistic for the purpose of estimating a presidential popularity function. We therefore rely on a more general approach and employ a semi-parametric Additive Mixed Model, which was introduced in the statistical literature for instance by Ruppert et al. [2003], Wood [2006] and Zuur et al. [2008]. In the following, we outline the employed estimation approach in some more detail.

The standard model (1) is a special case of

$$y_i \; = \; f(x_{i1}, \ldots, x_{ip}) + \epsilon_i, \tag{2}$$

with $f(\cdot)$ being an unknown function quantifying the relationship of the p covariates on the response y_i. In a first step, we impose the assumption of additivity on (2) and therefore ease the rather strong assumptions of linearity in (1) and replace the structure by a functional and additive form

$$y_i = \beta_0 + f_1(z_{i1}) + \ldots + f_q(z_{iq}) + x_{i1}\beta_1 + \ldots + x_{ip}\beta_p + \epsilon_i, \tag{3}$$

with $Z = (z_{ij}), i = 1, \ldots, n; j = 1, \ldots, q$ consisting of q metrically scaled covariates and $X = (x_{im}), i = 1, \ldots, n; m = 1, \ldots, p$ consisting of p binary-coded indicator covariates. In (3), $f_r(z_{ir}) \,\forall\, r \in \{1, \ldots, q\}$ are assumed to be sufficiently smooth but a-priori otherwise unspecified functions in the corresponding ranges of the covariates to be estimated from the data. Models of the form (3) have been coined Additive Models by HASTIE and TIBSHIRANI [1990] and are extensively discussed in WOOD [2006]. Following RUPPERT ET AL. [2003] and FAHRMEIR ET AL. [2009], model (3) is a *semi-parametric* Additive Model due to binary-coded covariates in X (and the intercept β_0), entering the model in a linear way.

As FAHRMEIR ET AL. [2009] discuss, model (3) can not be identified without an additional constraint: any offset or other additional constant could simultaneously be added to $f_m(\cdot)$ and be subtracted from $f_o(\cdot)$ $(m \neq o)$, without changing the model's prediction. We therefore need to define the level or the height of each a-priori unspecified function. The most common way is to impose the constraint

$$\sum_{i=1}^{n} f_1(z_{i1}) \;=\; \ldots \;=\sum_{i=1}^{n} f_q(z_{iq}) \;=\; 0 \tag{4}$$

which centers each function around zero.

Fitting model (3) and therefore estimating the functional relationship of the metrically scaled covariates in Z is carried out using penalized spline smoothing. The underlying idea is to replace each function $f_j(x_j)\,\forall\, j \in \{1,\ldots,q\}$ in a first step by some high-dimensional basis representation

$$f_j(z_j) = B_j(z_j)b_j, \tag{5}$$

where $B(\cdot)$ can be constructed using a popular cubic smoothing spline (see WAHBA [1978] and DEBOOR 1978). Note that since basis $B(\cdot)$ is high-dimensional, the resulting fit employing OLS will be poor unless we use the coefficient vector b_j to control the relative weight to be given to the conflicting goals of matching the data appropriately and producing a sufficient smooth function f. A sophisticated way to achieve this goal is to impose a penalty on b_j by using quadratic penalties in the form $\lambda_j b_j^T D_j b_j$. In the latter, D_j is the penalty matrix (see WOOD [2006] for more details) and λ_j is the tunable penalty parameter steering the amount of smoothness of the function. The resulting penalized least-squares criterion for one single functional effect

$$\sum_{i=1}^{n} (y_i - f(z_i))^2 \;+\; \lambda \int f''(z_i)^2 dz \;\rightarrow\; \text{min!} \tag{6}$$

is connected to the curvature of the function by penalizing the integrated squared derivative of second order using the quadratic form of penalization.

The statistical literature discusses extensively the possibility to represent penalized regression in the context of mixed models in order to get an optimal value of λ for (6) data driven, see WAHBA [1978], WONG and KOHN [1996] or WOOD [2003]. By interpreting the quadratic penalty as a (Bayesian) prior on the spline coefficient vectors b_j, model (3) changes to

$$y_i \;=\; \beta_0 + \sum_{j=1}^{q} B_j(z_j)b_j + \sum_{u=1}^{p} x_u \beta_u + \epsilon_i \tag{7}$$

with $b_j \sim N(0, \sigma_b^2)$ and $\epsilon_i \sim N(0, \sigma_\epsilon^2)$. Although $B_j(\cdot)$ is high-dimensional, the Bayesian formulation in (7) is a well-known Linear Mixed Model (LMM), as described by SEARLE ET AL. [1992], MC-

Culloch and Searle [2001] and Zuur et al. [2008]. By following the corresponding criterion for the best linear unbiased predictors in LMM, Ruppert et al. [2003] show that $b_j \sim N(0, \lambda^{-1} D_j^{-1})$. The penalization parameter is therefore a variance component of the random effect and it follows that an optimal parameter steering the amount of smoothness is found by

$$\hat{\lambda} = \frac{\hat{\sigma_\epsilon^2}}{\hat{\sigma_b^2}}. \tag{8}$$

The estimation of the two variance components in (8) can either be carried out using Maximum Likelihood (ML) or Restricted Maximum Likelihood (REML) technique. Due to the large numbers of fixed parameters in the model (see Section 4), the ML based estimations of σ_ϵ^2 and σ_b^2 tend to be "badly biased" (see Wood [2006]). We therefore employ the REML technique, which integrates out the fixed parameters of the joint likelihood and allows for a more reliable estimation. For the technical details about REML we refer to Ruppert et al. [2003] and Fahrmeir et al. [2009].

For the aspired analysis with the data at hand we have to modify and amend the above motivated models with respect to two further aspects: First, presidential approval, as our response variable, is compiled on a monthly basis and therefore likely to be affected by unobserved effects in the short run. It is reasonable to assume that these effects occur randomly. We therefore supplement model (3) / (7) by a latent monthly-specific effect:

$$y_i \;=\; \beta_0 + \sum_{j=1}^{q} B_j(z_j) b_j + \sum_{u=1}^{p} x_u \beta_u + t_{i0} + \epsilon_i \tag{9}$$

with $t_{i0} \sim N(0, \sigma_t^2)$ and all of the above mentioned assumptions. t_{i0} allows for random monthly deviations from β_0 and controls additionally for serial correlation in the dataset. Note, that the latter is likely since the underlying data is a multivariate time series. The fitting can be carried out employing the same technique as motivated above since (9) is only a minor extension of (7) with respect to the parameters and is therefore again a Linear Mixed Model being estimated numerically with REML.

As second aspect, we have to supplement model (9) by the inclusion of interaction terms, capturing an interaction of metrically and binary-scaled covariates. In our case, we have to guarantee that well-defined metrically scaled covariates are only be estimated for the time period a certain binary-coded variable is set to 1. (9) therefore changes to

$$y_i \;=\; \beta_0 + \sum_{j=1}^{q} B_j(z_j) b_j + \sum_{k=1}^{w} \left(B_k(z_k) b_k\right) \tilde{x}_{ik} + \sum_{u=1}^{p} x_u \beta_u + t_{i0} + \epsilon_i \tag{10}$$

with $\tilde{x}_{ik}$ being binary-coded covariate k taking the value of either 0 or 1. If we do not capture the parametric effect itself ($\tilde{x}_{ik}\beta_k$) on the presidential approval, the corresponding smooth functions $f_1(z_{i1}), \ldots, f_w(z_{iw})$ do not have to be centred around zero (see Fahrmeir et al. [2009] for details). The estimation strategy as being motivated above is not affected by these two amendments and is straightforward.

The described estimation technique is implemented in R, see PINHEIRO and BATES [2000] and R DEVELOPMENT CORE TEAM [2010]. We make use of the R-package `mgcv` (see WOOD [2011]), which allows for a computationally stable and reliable estimation.

4 Data and Estimation Equation

Our empirical analysis bases on monthly observations dating from January 1953 to December 2006. Thus, the sample data covers the 10 U.S. presidents from Dwight D. Eisenhower to George W. Bush jun.

The endogenous variable, presidential approval, is defined as the average share of survey respondents answering positively to the following Gallup question: Do you approve or disapprove of the way *[name of the president]* is handling his job as president? Approval ratings are limited to an [0,100] interval, although actual values range between 23 (Watergate) and 88 (9/11 attacks). Since there are 45 months in which Gallup has not conducted the relevant surveys, our sample size is reduced from 648 to 603 months.

variable	**obs**	**mean [%]**	**std. dev.**	**min [%]**	**max [%]**
approval	603	55.39	12.01	23	88
unemployment	648	5.734	1.463	2.548	10.85
inflation	648	3.864	2.933	-0.854	14.59
gov.consumption	648	13.192	1.493	8.9	15.6

Table 1: Summary statistics, 1953–2006

In this study, three economic determinants of presidential approval are considered: (a) the seasonally adjusted, civilian unemployment rate, (b) the seasonally adjusted inflation rate, defined as the percentage change of the CPI over the previous 12 months and (c) government consumption as a percentage of GDP.[4] Two of them, unemployment and inflation, are very frequently used in the vote and popularity function literature. As a third variable we add government consumption as a percentage of GDP in order to control for the governments' fiscal policies. Tables 1 and 2 display descriptive statistics for the sample period and for each presidency.

Starting with MUELLER [1970], it has repeatedly been shown that presidential popularity tends to decline over time, probably in consequence of the so-called *cost of ruling* (see PALDAM [2008]) – although some authors criticize the use of time as an explanatory variable (e.g. KERNELL [1978]). Additionally, the literature suggests that presidents enjoy a *honeymoon* period during the first months as well as rebound effects (*nostalgia* effect) in the end of their presidencies (see e.g. GEYS [2010]). To capture all these potential influences, a *time-in-office* variable is included which takes on a value of zero in the first month of each presidency and increases linearly with every additional

[4] All series are taken from the FRED data base. The employed series are *UNRATE* (unemployment), *CPIAUCSL* (inflation) as well as *GOV.CON* government consumption as a share of GDP. Since the government consumption ratio is only available on a quarterly basis, we use the same quarterly value for each month during that quarter.

president	**approval** [%]	**unemployment** [%]	**inflation** [%]	**gov.consumption** [%]
Eisenhower	64.32	4.897	1.362	10.122
Kennedy	70.03	5.970	1.154	11.829
Johnson	54.01	4.193	2.567	12.816
Nixon	48.92	5.006	5.579	14.073
Ford	46.72	7.759	8.229	15.186
Carter	45.51	6.527	9.708	14.188
Reagan	52.49	7.541	4.659	13.628
Bush I	59.52	6.304	4.370	13.937
Clinton	54.62	5.205	2.594	13.669
Bush II	55.12	5.289	2.663	14.263
Total	55.39	5.734	3.864	13.192

Table 2: Summary statistics (mean values), by president, 1953–2006

month in office. Since we allow the president's time in office variable to enter the estimation equation in a non-linear fashion, a single variable is able to account for all three time-related effects. Different from the existing literature, we thus have not to impose any arbitrary assumption on the magnitude and duration of these effects.

Finally, we add a number of non-economic control variables to our model. Following Newman and Forcehimes [2010], we include binary coded variables to capture the effect of politically relevant events like the Cuban missile crisis, the Iran hostage crisis or the fall of Baghdad in April 2003. Altogether, we control for 120 events which are grouped together in eight variables. The variables include positive as well as negative events in the four categories *personal*, *domestic*, *international* and *diplomatic*.[5] Additionally, due to their extraordinary impact, two separate dummy variables are included for the Watergate affair and the 9/11 terror attacks, respectively. To account for the effect of major military conflicts, a dummy variable for the short Gulf War (Operation *Desert Storm*) as well as monthly casualty figures for the wars in Vietnam, Afghanistan and Iraq are included.[6] Moreover, a dummy variable for divided governments is employed to control for the clarity of responsibility of the reigning president (Powell and Whitten [1993]). Binary coded variables for each president are included to control for unobserved, president-specific effects. We chose Bill Clinton as the reference point since none of the major political events or wars fell into his period of office lasting 96 months.

The equation to be estimated is thus

[5] Among other criteria, extensive front page coverage in the *New York Times* is a necessary requirement for the inclusion of events. Since there was no event in the *negative diplomatic* category we end up with seven binary event variables. For a complete event list and details about the selection method, see Newman and Forcehimes [2010].

[6] The Gulf War dummy is 1 from January 1991 to September 1991. Casualty figures are taken from the National Archives (Vietnam) and the Department of Defense.

$$
\begin{aligned}
approval_i &= \beta_0 + f_1(inflation_i) + f_2(unemployment_i) \\
&+ f_3(gov.consumption_i) + f_4(time.in.office_i) \\
&+ f_5(vietnam.casualties_i) vietnam.war_i \\
&+ f_6(afghanistan.casualties_i) afghanistan.war_i \\
&+ f_7(iraq.casualties_i) iraq.war_i \\
&+ X\beta + t_{io} + \epsilon_i \qquad (11)
\end{aligned}
$$

5 Results

Before turning to the (potentially) non-linear determinants of presidential approval, we will shortly discuss the results for the binary covariates (matrix X) in model (11), which are displayed in Table 3.

First, we find highly significant effects of the three major events we control for. While the Watergate Affair exerted a negative effect on presidential popularity, Operation Desert Storm and 9/11 led to a positive rally effect.

Second, the controls for positive political events turn out to be significant for all four categories (domestic, foreign, diplomatic and personal) and the estimated coefficients all deliver the expected positive sign. Positive domestic political events turn out to influence presidential popularity by far the most. Among the group of negative political events only the negative domestic events turn out to be significant, however with the expected negative sign. While negative personal events also deliver a slightly negative coefficient, the coefficient is insignificant. The estimated coefficient of negative foreign events is positive, very small and highly insignificant. Altogether, the effects of the political event dummies turn out to be highly plausible.

Third, we find a significantly positive effect of divided governments on presidential popularity. During these times, the costs of ruling seem to be less important since the political and economic outcomes can not be attributed to the president alone.

Fourth, after having controlled for all other political, personal and economic effects, we find the presidents Kennedy, Johnson, Carter and Reagan to have reached significantly higher popularity scores compared to the Clinton administration. President Bush jr. is the only president who performed significantly worse in terms of popularity than president Clinton.[7]

The estimation results for the variables, entering the estimation equation as a-priori unspecified functions, are visualized in Figure 1. The displayed diagrams show the functional relation between presidential popularity and the referring covariates. The shaded areas depict the 95% pointwise confidence bands.

[7] This result should be interpreted with some caution, since some of the political events or affairs are directly linked to the person of the ruling president, such as the Watergate affair.

covariate	$\hat{\beta}_j$	p-value
(Intercept)	0.43	< 0.01
$watergate$	-0.13	< 0.01
$desert.storm$	0.26	< 0.01
$nine.eleven$	0.13	< 0.01
$neg.domestic$	-0.02	0.04
$neg.foreign$	0.003	0.91
$neg.personal$	-0.009	0.51
$pos.domestic$	0.67	0.01
$pos.foreign$	0.05	< 0.01
$pos.diplomatic$	0.02	0.09
$pos.personal$	0.04	0.04
$divided.gov$	0.15	< 0.01
$Eisenhower$	-0.03	0.5
$Kennedy$	0.13	< 0.01
$Johnson$	0.09	< 0.01
$Nixon$	-0.02	0.92
$Ford$	0.02	0.51
$Carter$	0.11	< 0.01
$Reagan$	0.05	< 0.01
$BushSr$	-0.01	0.52
$BushJr$	-0.14	< 0.01

Table 3: Parametric estimation results of model (11)

Inflation For inflation rates below 10%, we find an (almost) monotonic, negative relationship between inflation and popularity. However, this relationship turns out to be highly non-linear. As inflation increases from very low levels to roughly 4%, presidential popularity decreases significantly. For example, a rise in inflation form 0% to 4% decreases presidential popularity by almost 10 percentage points. In the range in between 4% and 7%, further increases in the inflation rate remain without any effect on popularity. However, for inflation rates larger than 7% higher inflation rates end up in further eroding presidential popularity. A rise in inflation from 7% to 11% leads to a decrease of presidential popularity of another 10 percentage points.

Somewhat surprisingly, we find a positive effect of further increases of inflation on presidential popularity for very high inflation rates. However, inflation rates above 10% have been observed in the U.S. only in the wake of the oil crises of 1973 and 1979. Apparently, the incumbent presidents during that times were not blamed for this inflation, thereby confirming the responsibility hypothesis of public support.[8] In fact, the 1973 oil embargo could rather be seen as an alien aggression. In consequence of this aggression the American voters aligned with their political leaders.

At first sight, one might be surprised that the electorate holds the U.S. president responsible for inflation although it is primarily the Federal Reserve Bank which is responsible for controlling inflation and the Federal Reserve Bank is formally independent from the U.S. government. However, the voter seems not to (be able to) distinguish between the role of different governmental institutions. Moreover, the U.S. president has the formal right to nominee the Chairman of the Board of Governors and thus might be held responsible for a suboptimal performance of U.S. monetary policy in controlling inflation.

Unemployment The relationship between unemployment and presidential popularity shows a strong non-linear pattern, too. For very low unemployment rates of up to 4% the effect of increasing unemployment is almost zero or at least very small. Changes in the unemployment rate in between 4% and 5.5% have moderately negative effects on popularity. An increase from 4% to 5.5% decreases popularity by 4 percentage points. In the range in between 5.5% and 7% increases of unemployment do not affect popularity. However, whenever the unemployment rate exceeds 7%, more unemployment turns out to be detrimental to presidential popularity. The strongest negative effect of unemployment on presidential popularity can be found around 8%. Unemployment rates exceeding 7% were reported in 118 of 603 months in our sample, mainly during the presidencies of Ford, Carter, Reagan and Bush Sen.

Government Consumption While inflation and unemployment have quite regularly been included in studies on the determinants of presidential popularity, much less evidence is yet available for the voters' perceptions of fiscal measures. When using government consumption (not including defence expenditures) as a proxy for a government's fiscal activity we find a significant negative,

[8] With respect to the responsibility argument it should be noted that the first oil crisis was part of a political conflict in the Middle East in which the U.S. played an important role. The results clearly depict that the U.S. President was not held accountable for the extraordinarily high inflation rates at that times.

but almost linear effect on popularity. Thus, U.S. voters in general seem (ceteris paribus) not to be interested in fiscal stabilization programmes. However, one might suspect that the voters' perceptions of stabilization programmes might depend on the overall state of the economy. We will return to this aspect in the next section where we allow for interaction effects between the three economic variables.

Time in office As discussed earlier, several hypotheses about the development of presidential approval over the term of office have been tested in previous empirical studies (costs of ruling, honeymoon and nostalgia effects). All these hypotheses are quite controversial in the literature since the empirical results depend very much on the way, how the expected pattern is modeled (typically by time-varying dummy variables). Our approach of modeling time-effects via a time-in-office variable and allowing for a non-linear relationship avoids this sort of criticism and is thus useful to shed light on the factual existence of these effects.

Interestingly enough, our results indicate that all three mentioned effects in fact exist. In general, presidential popularity tends to decrease over the first term of office of a president. This finding is in line with the cost-of-ruling argument. As predicted by the honeymoon hypothesis, presidents (ceteris paribus) turn out to have their highest approval ratings in the beginning of their first term. And finally we also find evidence in favor of the nostalgia effect. During the last year of the electoral term, popularity tends to increase again (at least slightly. However, as it is well known from the literature on political budget cycles, this effect might result from the incentives to implement favorable but costly policies to improve their short-term re-election chances.

Remarkably, the popularity pattern of presidents in their second term of office shows a somewhat similar pattern as the first one. Again, presidential popularity turns out to be maximal throughout the first months of the (additional) term of office, although the popularity level turns out to be much lower than in the first period. As time goes by, presidential popularity again erodes (although to a much lesser extend) but recovers roughly one and a half year before the second term of office ends. The fact that even in the second term of office a recovery of popularity occurs in the end of the final term of office might be taken as an indication that this effect is not primarily driven by re-election concerns. Especially presidents in their second term of office seem to profit from a strong nostalgia effect.

War casualties Since World War II, the U.S. were involved into three major wars of considerable length: Vietnam, Iraq and Afghanistan. While short external conflicts are well known to cause a rally around the flag effect boosting presidential popularity, the effect of longer lasting conflicts is dependent on the public's perception of the conflict. One might expect this perception to be related to the number of US soldiers killed in action. For all three wars (Vietnam, Iraq and Afghanistan) we in fact find a negative impact of casualty figures on presidential approval. While the negative effect is linear and small for the war in Afghanistan (a war with comparatively low monthly casualty numbers), we find stronger and non-linear effects in Vietnam and Iraq. In both cases, our approach implies some sort of adaptation effect. At a certain stage, voters seem not to distinguish between "high" and "very high" monthly casualty figures.

Summing up, we might conclude that is strong evidence for significant and non-linear effects of the stance of the economy on presidential popularity after controlling for the effects of political events. While the effect of government consumption turns out to be linear, unemployment and inflation seem to exert a strongly non-linear effect on presidential popularity. Moreover, we find non-linear effects of presidents' time in office, thereby delivering supporting evidence for the honeymoon-, the nostalgia- and the cost-of-ruling-hypothesis. Finally, we find negative effects of war-casualties on presidential approval. However, the detrimental effect of casualty figures on popularity seems to decrease in the level of the death-toll to be paid.

In Figure 2 we show a comparison of the popularity time series and the fitted values. Apparently, our estimation approach delivers a good fit for the time series to be explained.

6 Interaction Models

In the previously employed estimation approach we assumed additive effects of economic variables on presidential popularity. However, in reality this assumption must not necessarily hold true. The existence of interaction effects between the considered economic variables is well possible if not likely, but is a yet unexplored field of research.

In order to study the existence and relevance of interaction effects we relax the restriction of separate effects of the economic variables on presidential popularity in the following. More precisely, we allow for non-linear interaction effects between the economic variables while leaving the rest of model structure unchanged. The model to be estimated is then given by

$$\begin{aligned} approval_i &= \beta_0 + f_{1|2|3}(inflation_i, unemployment_i, gov.consumption_i) \\ &+ f_4(time.in.office_i) \\ &+ f_5(vietnam.casualties_i)vietnam.war_i \\ &+ f_6(afghanistan.casualties_i)afghanistan.war_i \\ &+ f_7(iraq.casualties_i)iraq.war_i \\ &+ X\beta + t_{io} + \epsilon_i \end{aligned} \tag{12}$$

with $f_{1|2|3}(\cdot)$ being a smooth but a-priori unspecified function of three metrically scaled covariates. The assumption of additivity is therefore eased for the effects of these economic covariates. With respect to the assumptions of the functional form, model (12) therefore takes a position between the models (2) and (3) with respect to a-priori assumptions on the functional form.

The common technique of estimating interaction effects in classical OLS models by creating products of the underlying data vectors can easily be transformed to get a new basis function and a resulting $\hat{f}_{1|2|3}(\cdot)$. The tensor product basis matrix is gained by using all possible interactions of the

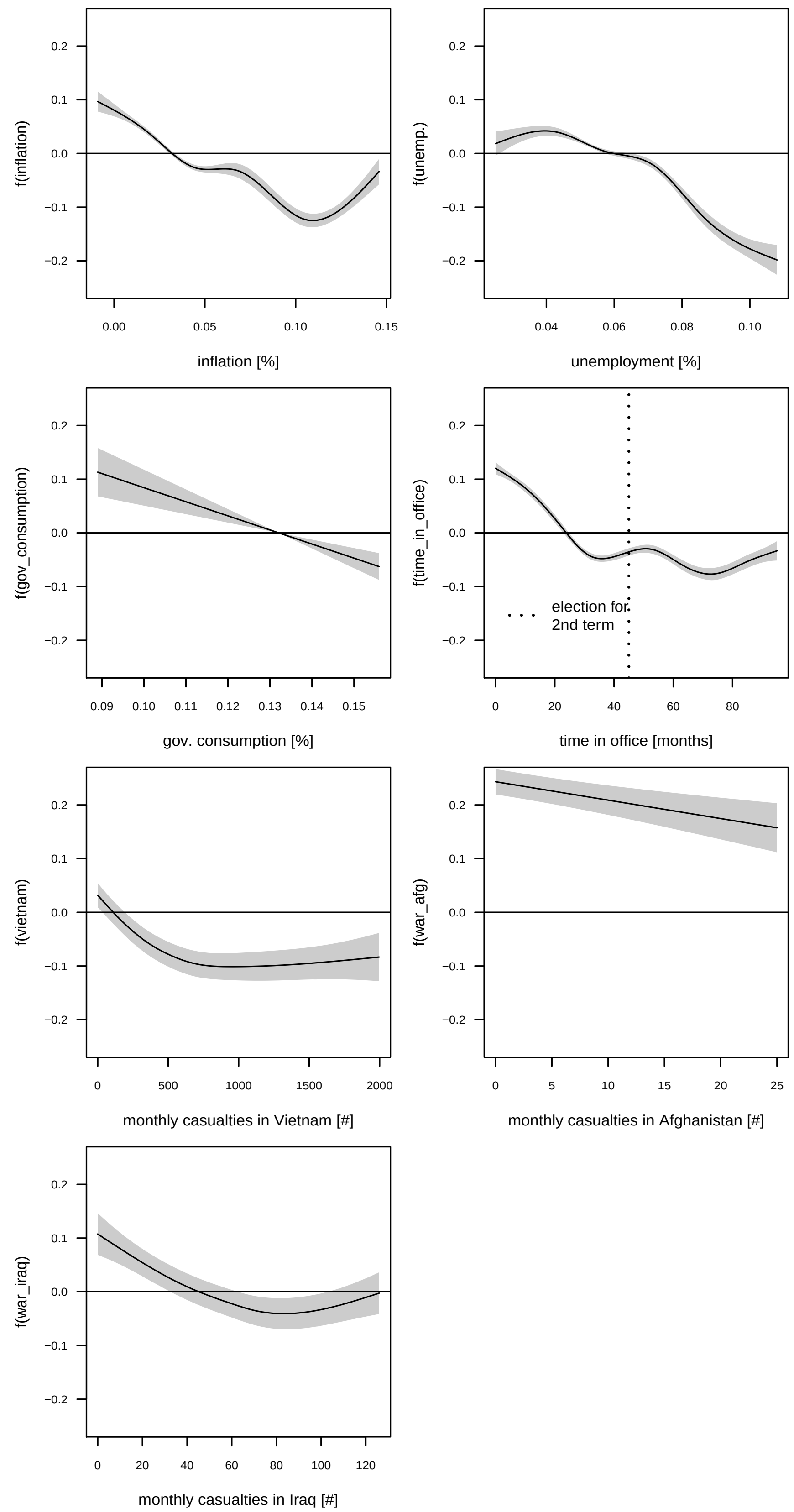

Figure 1: Fitted smooth effects of model (11) with 95% pointwise confidence intervals

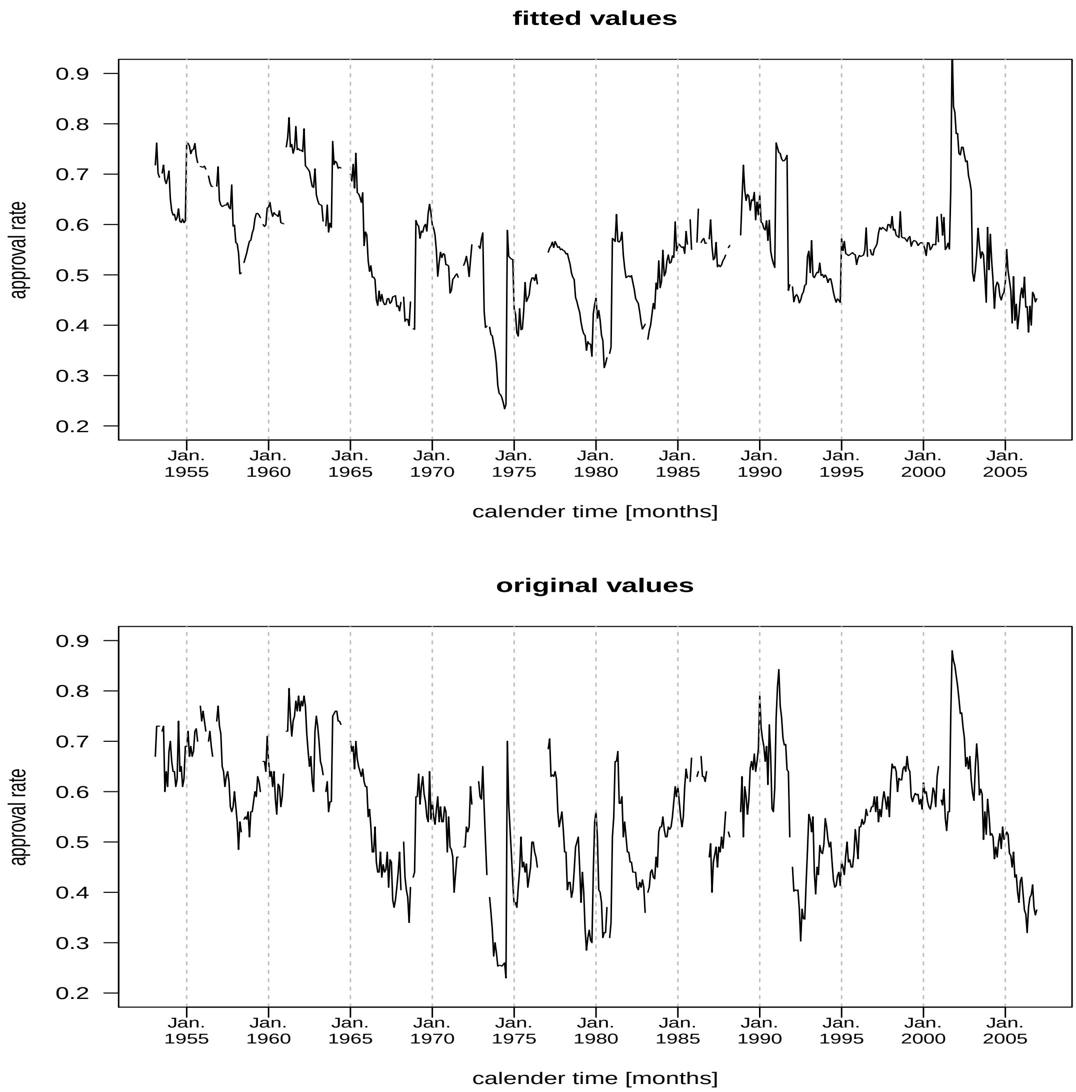

Figure 2: Fitted values and original values of model (11)

univariate splines for z_1, z_2 and z_3. The necessary penalization matrices for achieving sufficiently smooth three-dimensional functional effects are obtained by the Kronecker product

$$I_{d_3} \otimes I_{d_2} \otimes D_1 \,+\, I_{d_3} \otimes D_2 \otimes I_{d_1} \,+\, D_3 \otimes I_{d_2} \otimes I_{d_1}, \tag{13}$$

with $D_{\cdot}$ being the (univariate) penalization matrices and $I_{d_{\cdot}}$ the corresponding identity matrices in the matching dimensions. Further explanations can be found in Fahrmeir et al. [2009]. The smoothing technique, employing penalized splines, is therefore built upon the high-dimensional tensor products, which are invariant to any linear rescaling of the covariates and are computationally cheap (see Wood [2006]).
By introducing a three-dimensional interaction effect, model (10) changes to

$$y_i \,=\, \beta_0 + \tilde{B}(z_1, z_2, z_3)\tilde{b} + \sum_{j=3}^{q} B_j(z_j)b_j + \sum_{k=1}^{w} \left(B_j(z_j)b_j\right)\tilde{x}_{ij} + \sum_{u=1}^{p} x_u\beta_u + t_{i0} + \epsilon_i \tag{14}$$

with $\tilde{B}(\cdot)$ being the high-dimensional tensor product basis function and $\tilde{b}$ the corresponding (random) coefficient vector. For consistency, $\tilde{B}(\cdot)$ is again constructed by employing cubic smoothing splines. Note that

$$\sum_{i=1}^{n} f_{1|2|3}(z_{i1}, z_{i2}, z_{i3}) \,=\, 0 \tag{15}$$

additionally guarantees identifiability of the new model.
While an interaction of two metrically scaled covariates can be visualized by so-called *interaction surfaces*, a graphical analysis of a three- dimensional function, leading to a four-dimensional visualization, is not straightforward. However, by holding one of the three covariates of the joint effect constant at well-defined values, the joint effect can be visualized by an array of interaction surfaces. For numerical and graphical details see Wood [2011].
Fitting itself can again being carried out with REML estimation, thereby making use of the notation as a mixed model (see Fahrmeir et al. [2009] for details). Note, that the computational burden can be enormous since every d-dimensional basis representation of the splines leads to d^3 parameters for a three-dimensional functional relationship. The resulting d^3-dimensional system of linear equations has to be solved numerically, although penalization reduces the effective amount of parameters in the resulting model.
Again, we start out discussing the estimation results with the parametric effects displayed in Table 4. The effects of the major events remain qualitatively unchanged. However, the effect of 9/11 is now even more pronounced. The estimation results for the political events change only slightly. The only remarkable difference is that negative personal events now deliver a significantly negative coefficient. The effect of divided governments remains significant, but is numerically lower now. While all these results are very similar to the case without interaction effects, the binary coded covariates for each president do not remain unaffected by the change in the estimation approach. Only president Reagan is now judged to be significantly more popular than president Clinton after

controlling for all additional influences. However, as pointed out earlier, the presidential-specific effects should be interpreted cautiously, since their values obviously depend on the political and major events we control for explicitly.

In a next step we turn to the various non-economic variables for which we allow non-linear effects (see Figure 5). The time-in-office variable behaves quite similar to the non-interaction case. However, the effects in the second term of office are now more pronounced than before. In general, the war variables tend to lose some explanatory power. While the effects for Vietnam and Afghanistan remain at least similar, the casualties in the Iraq war lose their explanatory power completely.

In Figure 4 we show a visualization of the interaction effects between the economic variables. We decided to show the interaction effects between unemployment and government consumption as interaction surface for various levels of inflation. For the sake of clarity, statistical significance is not displayed in Figure 4. For all five displayed interaction surfaces one should be cautious in interpreting the results at extreme values, since statistical significance can not be guaranteed with a low number of observations.

We start out interpreting Figure 4 by focussing on the case of very low inflation rates (or moderate deflation). In times of very low inflation, increasing government consumption for almost any given rate of unemployment is harmful for the president. Especially when unemployment is very low, increasing government consumption leads to strongly decreasing popularity. One might suspect this to be due to the feeling that government spending programs are unnecessary during times of high employment. In times of high unemployment, voters do not punish increased spending as soon as a certain spending level of roughly 12% is already reached.

When moving to higher but still moderate inflation regimes (2.5%, 6%) the basic picture remains similar, but the described effects flatten significantly. In general, the interaction surface tends to tilt backward relatively to the imaginary axis between low government consumption and high unemployment. High government consumption at low levels of unemployment becomes less problematic while low government consumption at high levels of unemployment are now perceived worse than before.

Interestingly enough, in times of very high inflation, the situation turns completely around. Under high inflation, increased spending is not perceived as an evil any more. While government spending increases popularity only slightly in times of low unemployment, spending programs tend to be quite popular under high unemployment regimes.

It is an intriguing question why rising inflation has such a strong effect on the perceivement of government spending programs under varying labor market conditions. Often spending programs are financed via deficit spending. Excessive deficits contribute to a high level of public debt. However, in times of high inflation the public deficit erodes quickly in real terms at the burden of domestic and foreign creditors. This is especially true when large parts of the public deficit is not indexed such as in the United States.[9] Thus, voters might find the financing burden of spending

[9] As Aizenman and Marion (2009) argue, the U.S. government might have a strong incentive to inflate away the burden of the enormously risen public debt in consequence of the recent financial crisis.

covariate	$\hat{\beta}_j$	p-value
(Intercept)	0.54	< 0.01
$watergate$	-0.14	< 0.01
$desert.storm$	0.20	< 0.01
$nine.eleven$	0.17	0.01
$neg.domestic$	-0.02	0.02
$neg.foreign$	-0.01	0.68
$neg.personal$	-0.03	< 0.01
$pos.domestic$	0.06	< 0.01
$pos.foreign$	0.05	< 0.01
$pos.diplomatic$	0.02	0.04
$pos.personal$	0.03	0.08
$divided.gov$	0.10	< 0.01
$Eisenhower$	-0.24	< 0.01
$Kennedy$	-0.12	< 0.01
$Johnson$	-0.07	0.10
$Nixon$	-0.07	0.08
$Ford$	-0.06	0.16
$Carter$	-0.02	0.56
$Reagan$	0.06	0.03
$BushSr$	0.004	0.86
$BushJr$	-0.11	< 0.01

Table 4: Parametric estimation results of model (12)

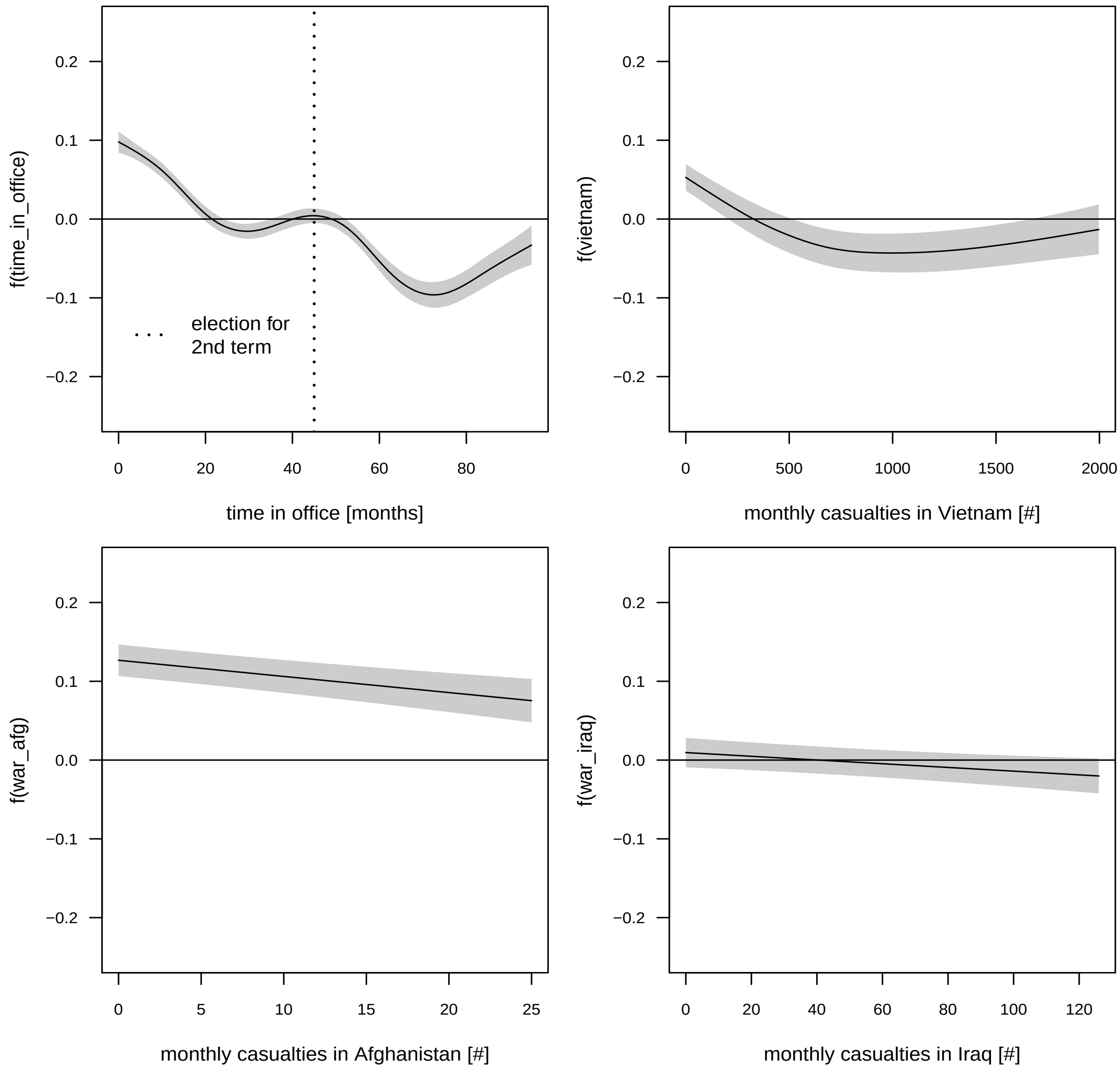

Figure 3: additional fitted smooth effects of model (12) with 95% pointwise confidence intervals

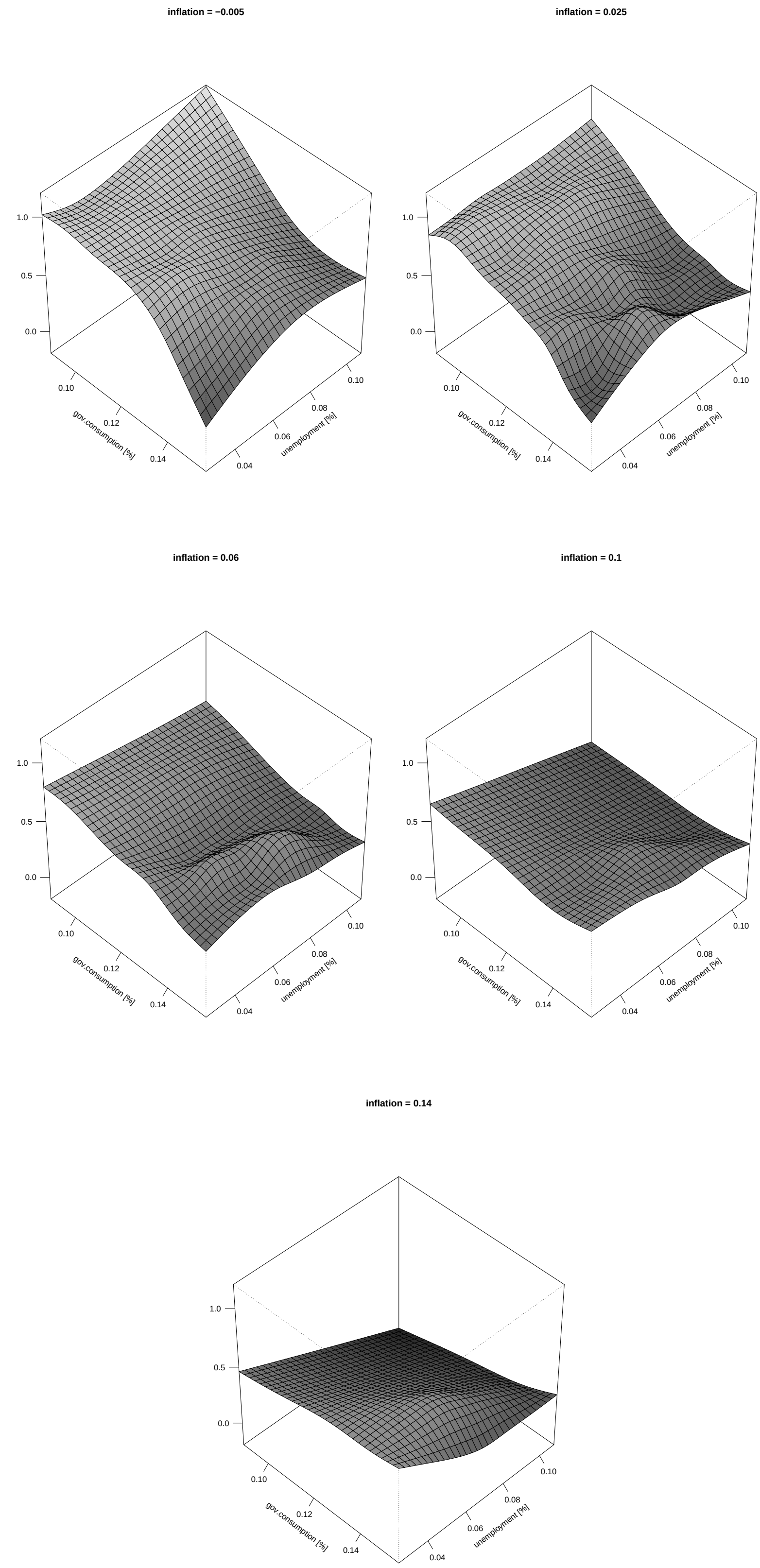

Figure 4: Fitted smooth interaction effect of inflation, unemployment and governmental consumption of model (12), holding inflation constant for selected values

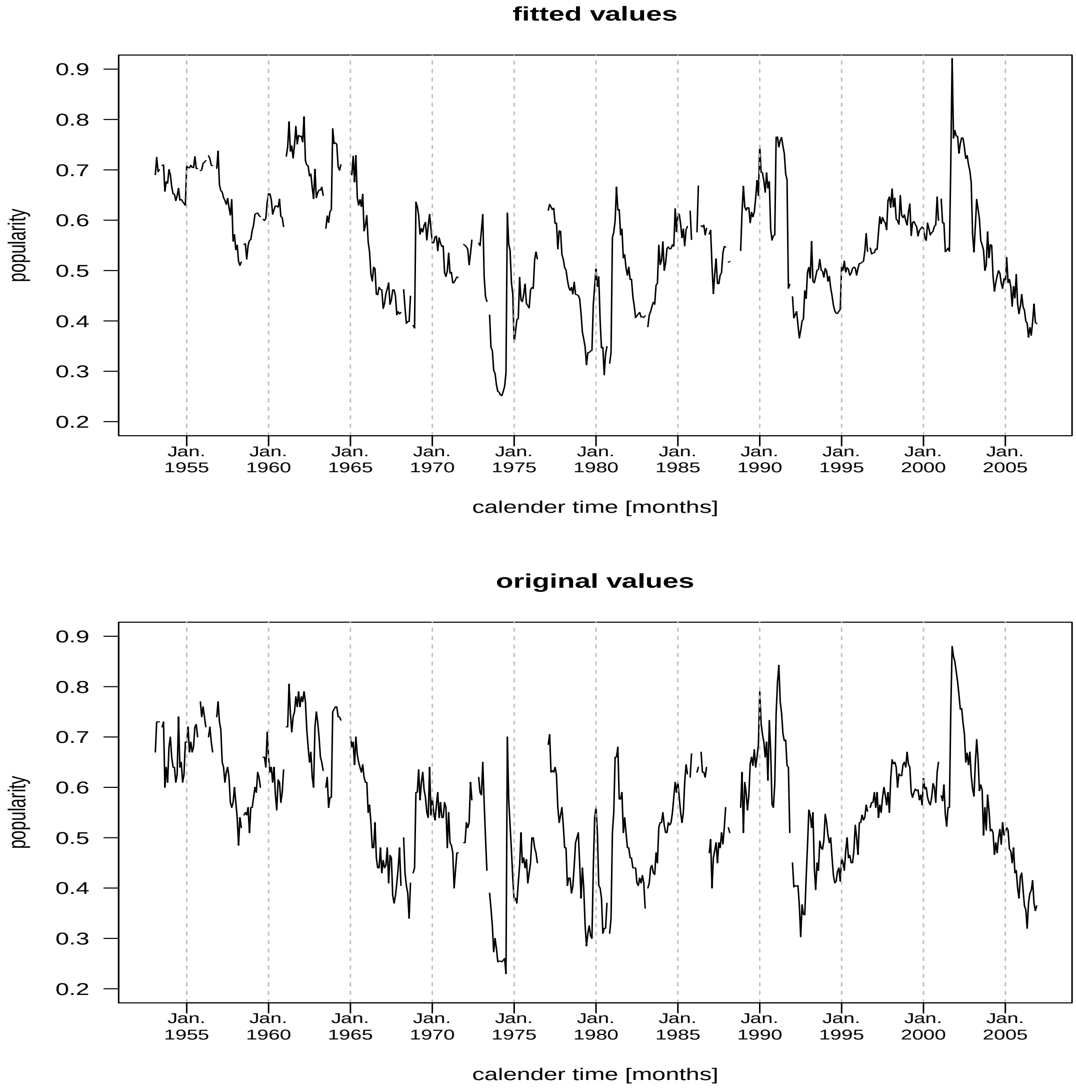

Figure 5: Fitted values and original values of model (12)

programs less problematic in episodes of high inflation. Especially in times of high unemployment presidents may thus profit from an increase in government consumption figures.

While the presented findings point into the direction that interaction effects between the economic variables are important and taking them into account is inevitable, it seems necessary to compare both presented modeling approaches with respect to econometric criteria. As shown in Figure 5 the model fit of the interaction model is even better than for the additive model. The models can be compared on the basis of the AIC and the BIC criteria.[10] As depicted in Table 5 both criteria clearly favor model (12).

[10] For a detailed discussion about the employed AIC and BIC measures see and their corresponding definition in the context of non-parametric estimation techniques, see Wood [2006] and Fahrmeir et al. [2009].

	model (11)	model (12)
AIC	-1491.74	-1648.79
BIC	-1317.77	-1457.72

Table 5: Model selection criteria

7 Conclusions

In this paper, we estimate popularity functions for the United States using a modern semi-parametric estimation approach. We deviate from the existing literature by not assuming any specific (and arbitrarily chosen) functional form, but allowing for a more data-driven and a-priori unspecified linkage between presidential approval and its likely determinants. Our results indicate that the most commonly used linearity assumption is inconsistent with the rather complex relationship between presidential approval rates and its determinants. The shortcoming to allow for non-linearities might have contributed to the fact that "[I]n spite of considerable efforts very little is 'cut and dried' in this field, and again and again discussions flare up when this or that result is found to be lacking in stability." (Paldam (1991)). The use of non-linear estimation techniques might therefore contribute much to deepening the understanding of the true determinants of presidential popularity.

We do not only find strong evidence for non-linearities in the relationship between economic variables and presidential popularity. We also show that strong interaction effects between these economic variables seem to exist. It is thus not useful to rely on purely additive (or even linear) effects when trying to uncover the determinants of presidential approval. Whenever these interaction effects exist, the perceivement of certain policies strongly depend on the macroeconomic situation. Presidents caring about their popularity among voters will have to take these effects into account when deciding on their policy measures. For example, spending programs might be perceived very differently under varying inflationary and employment regimes.

We also find strong evidence for the hypothesis that presidential approval rates on average follows a typical time pattern. While the literature up to now operated with quite specific assumptions about the exact form of this pattern, our non-parametric approach allows us to model the time in office as a possibly non-linear effect thereby allowing for any possible time pattern. In general, we find presidential popularity to decrease over the term of office, a finding which is in line with the cost-of-ruling argument. Moreover, we find supporting evidence for the hypotheses of a honeymoon and a nostalgia effect.

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