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Reconciling empirics on the political economy of the resource curse hypothesis. Evidence from long-run relationships between resource dependence, democracy and economic growth in Iran

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ABSTRACT

This paper analyzes the long-run relationships between resource dependence, democracy and per capita economic growth in Iran using the ARDL approach of cointegration (Pesaran et al. 2001). We find that resource dependence and democracy have positive, negative, or no significant effect on the long-term growth in Iranian GDP over the period 1970–2017. This multifaceted result is an additional reason to explain the lack of consensus on the empirics of the political economy of the resource curse hypothesis. We show as the interpretation of statistical tests only based on the average effect of resource dependence on economic growth may be misleading.

1. Introduction

A vast body of literature has focused on the relationship between natural resource abundance and economic growth. Auty (2001) coined the term of 'natural resource curse' to describe this surprising feature of economic life that resource-poor economies often outperform resource-rich economies in economic growth (Sachs and Warner, 1995). Several surveys point out both economic[1] and political[2] arguments to rationalize a negative correlation between resource abundance and economic growth. However, some econometric studies find a positive or also statistically not significant effect of resource abundance on GDP growth. This lack of consensus in the empirical literature emerges clearly in Havranek et al.'s (2016) meta-analysis. He found as 40% of empirical studies estimate a negative effect, 40% find no effect, and 20% of the analyzed studies showed that natural resources richness positively affects long-term economic growth.

In this research, we focus on one of the most influential hypotheses of this literature that argues as institutional quality plays a fundamental role in determining the effect of the abundance of natural resources in the national economy (Mehlum et al., 2006a, 2006b; Rosser, 2006; Robinson et al., 2006; Brunnschweiler, 2008; Deacon, 2011; Busse and Gröning, 2013; Vahabi, 2018). According to this hypothesis - also known as the *political economy of the resource curse hypothesis* - institutions are pivotal to reverse the "curse" into a "blessing" through several channels.[3] Good institutions can prevent rent-seeking activities (Auty, 2001), reduce corruption (Isham et al., 2005; Robinson et al., 2006), lower the risk of violent civil conflict (Collier and Hoeffler, 2005) and accelerate efficient resource allocation (Atkinson and Hamilton, 2003). Among these potential "channels", this research considers a specific type of institutional setting - that according to Collier and Hoeffler (2009: 294)

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[1] E.g. appreciation in the exchange rate - also known as "Dutch diseases"-, inefficiency of market economy because natural resources tend to be owned by firms with significant degrees of monopoly and monopsony power, less incentive for the rich-economy to diversify into different industries, etc.

[2] E.g. higher opportunities for rent-seeking activities, wars for ownership of resources, lower democracy, etc.

[3] These arguments solve some apparent empirical puzzles for purely economic explanations of resource curse hypothesis. Indeed, political economy resource curse hypothesis - including the role of institutions - is able to explain why for some resource-rich countries (e.g. Nigeria, Zambia, Venezuela) the resource abundance has been a "curse" while for other resource-rich countries (e.g. Norway, Botswana, Canada) it has been a blessing.

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"*has been the main recent institutional innovation in resource-rich countries*": the democracy.

Collier and Hoeffler (2009) distinguish between two dimensions of democracy: "electoral competition" and "checks and balances". They are two sides of the same coin but they have a different magnitude in mature and young democracies. Precisely, while, on the one hand, electoral competition is easy to establish in both mature and young democracies as there are strong incentives for participation, on the other hand, the system of checks and balances are generally lacking at the early stages of the democratization process.

This paper aims to contribute to the debate on the political economy of the resource curse hypothesis in several ways.

First, due to our focus on a specific type of institutional setting such as democratization, we use the index proposed by Vanhanen (2019) as a proxy of democracy. It specifically accounts for the aspects of democratization that, following Collier and Hoeffler (2009), may cause the resource curse, i.e. political participation and competition.

Second, following Bhattacharyya and Hodler (2010) and Arezki and Gylfason (2013) who find as resource rents lead to corruption only if the quality of democratic institutions is below a certain level, we estimate the value of this threshold for the Iranian economy. In particular, we calculate both the level of democratization and the percentage of resource rents on GDP that determine when the effects of resource dependence and democratization reverse from harmful to beneficial in economic growth.[4]

Third, taking into account Wright and Czelusta's (2004) and Brunnschweiler and Bulte's (2008) criticism on the empirical analyses of the resource curse hypothesis when resource dependence is used as a proxy of resource wealth,[5] we apply the autoregressive distributed lag (ARDL) bounds testing approach to identify the long-run relationship and to control for endogeneity issues.[6] In particular, to account for this issue, we test if the hypothesis of (weak) exogeneity holds in our model, so that efficient inference about the long-run coefficients can be conducted through the bounds testing approach to cointegration proposed by Pesaran and Shin (1999).

Forth, following Greene (2010), we provide graphical presentations of the overall marginal effects of resource dependence and democracy on long-term economic growth. This contribution may be noteworthy because it highlights as the lack of consensus in this literature may be due to a misleading interpretation of the statistical testing on the overall marginal effect of resource dependence on economic growth when an interaction term is included in the regression.

The paper provides in the next section an overview of the institutional view of the resource curse hypothesis. Section 3 describes the statistical approach, tests on the appropriateness of the proposed econometric approach, and converts the theoretical proposition of the political economy of the resource curse in two testable empirical hypotheses. Section 4 describes the empirical results and discusses the findings. Conclusions are drawn in section 5. Three appendixes provide details on the dataset and econometric analysis.

2. Literature review

While a thorough overview of the literature on the resource curse hypothesis is beyond the scope of this paper,[7] we intend to provide a brief review of a particular type of institution analyzed by the so-called political economy approach to the resource curse hypothesis[8]: the democracy.

According to Ross (2015), the most common argument in the political economy of the resource curse hypothesis is the "rentier effect" (Mahdavy, 1970; Crystal, 1990; Ross, 2001). It relies on the hypothesis that tax revenues (e.g. taxes on income, consumption, property) and non-tax revenues (e.g. non-renewable natural resources revenues) have different effects on authoritarian stability. When governments impose heavier taxes on citizens, thus taxpayers' dissatisfaction amplifies citizens' demand for greater accountability and a more inclusive political process. On the contrary, when undemocratic regimes "*gain higher nontax revenues, they are able to reduce taxation and hence attenuate these demands (*Bates and Lien 1985; Ross 2004; Brautigam et al. 2008*)*" (Ross, 2015: 246). This "political" curse for countries rich in nontax revenues generates also an "economic" curse because autocratic regimes have worse economic performances than democracies due to their higher levels of clientelism, corruption, income inequality, lower investment in education and market competition (Acemoglu et al., 2019). Recently, Prichard et al. (2018) find empirical support to this hypothesis by analyzing a panel data of 188 countries and over the period 1990–2012. They find that the natural resource curse is driven primarily by changes in the composition of government revenue.

A further explanation of the *political economy of the resource curse hypothesis* deals with the effect of resource wealth on economic growth as mediated by corruption. Bhattacharyya and Hodler (2010) argue that resource abundance hinders economic growth only if resource abundance causes more corruption. In particular, by analyzing a panel data of 124 countries over the period 1980–2004, they conclude that resource rents lead to an increase in corruption if the quality of the democratic institutions is relatively poor, but not otherwise.

Particularly relevant for our research is Collier and Hoeffler's (2009) argument. They hypostasize that resource-rich autocracies and mature democracies might out-perform equivalent young democracies since the latter suffer from a significant negative side effect in terms of misallocation of resources due to rent-seeking activities. Democracies may have also a second negative effect on the economic growth in resource-rich countries because of the bias that electoral competition generates on decision making for public investments. Accordingly, Collier and Hoeffler (2009) deduce that at the early stages of democratization, democracy may be harmful in terms of economic growth for under-provision of public investment.[9]

Ahmadov and Guliyev (2016) suggest a similar mechanism to explain why democracy may be detrimental to growth. They argue that in young democracies there are more opportunities for rent-seeking activities and corruption than in autocracies because they involve a

[4] Sarmidi et al. (2014) provide a similar analysis for a cross section of countries.

[5] Brunnschweiler and Bulte (2008) claim that find that, resource curse hypothesis may be a red herring. The most important source of bias for these empirical analyses is that resource abundance and economic growth suffer from endogeneity problems. Specifically, they find that resource abundance positively affects growth and institutional quality.

[6] Olayungbo and Adediran (2017) apply the same approach to test the resource curse hypothesis in Nigeria.

[7] Among the recent general surveys, see: Brunnschweiler and Bulte (2008), Van Der Ploeg (2011), Havranek et al. (2016), Venables (2016), Van Der Ploeg and Poelhekke (2017), Badeeb et al. (2017), Vahabi (2018), Shahbaz et al. (2019), Barbier (2019).

[8] Several studies review this strand of literature: Ross (1999, 2015), Bulte et al. (2005), Isham et al. (2005), Mehlum et al. (2006a, 2006b), Rosser (2006), Collier and Hoeffler (2009), Luong and Weinthal (2010), Haber and Menaldo (2011), Deacon (2011), Bhattacharyya and Hodler (2010), Brückner et al. (2012), El Anshasy and Katsaiti (2013), Brollo et al. (2013), Ahmadov and Guliyev (2016), Antonakakis et al. (2017), Abdulahi et al. (2019), Asif et al. (2020).

[9] By assuming that public investments have a crucial role for the growth in resource-rich countries, Collier and Hoeffler's (2009) hypothesis that democracy hinders economic growth is supported by Tavares and Wacziarg's (2001) evidence that democracies have lower investment than autocracies.

larger number of stakeholders. Indeed, while in established democracies there are institutional "barriers" against resource misallocation, these safeguards are lacking in young democracies. Accordingly, these rent-seeking activities, by wasting public and private resources, hamper the growth.

Some skeptic views on the role of democratic versus autocratic government in the resource curse hypothesis also exist. For instance, Haber and Menaldo (2011) claim that the existing literature may suffer from some serious sources of biases, the most serious issues are the omitted variable bias induced by unobserved country-specific and time-invariant heterogeneity and the use of datasets with a limited time dimension. By using a panel data from 1800 to 2006 to test if there is a long-run relationship between natural resource dependence and autocratic regime, Haber and Menaldo (2011: 25) concludes that "*oil and mineral reliance does not promote dictatorship over the long run. If anything, the opposite is true.*" A few years later, Andersen and Ross (2014) reanalyzed Haber and Menaldo's dataset and found that their results can be overturned by simply adding to the models a dummy variable for the post-1980 period, which is when oil began generating enormous rents and state-owned oil companies came into prominence in the market (Boutilier, 2017). Wright et al. (2015) confirm Andersen and Ross's (2014) finding, by reporting that higher levels of oil wealth deterred democratic transitions between 1980 and 2007. Moreover, Wright et al. (2015) criticize Haber and Menaldo's analysis also because they used a sample that excluded 51 autocratic countries (Ross, 2015).

A growing body of empirical studies deals with the issue of endogeneity as a potential source of unreliable inference. Among these researches, Antonakakis et al. (2017) analyze the role of political institutions in the resource curse hypothesis through a panel vector autoregressive approach. They conclude that while in economies with high quality of political institutions, having resources can be a blessing, in those countries with low-quality institutions, resources can become a curse for economic growth. According to Antonakakis et al. (2017: 163), "*all depends on how a country is prepared to take advantage of its natural resources, and quality of institutions is a key determinant of how each of the countries will use its resources to promote economic growth.*" In particular, they conclude that the resource curse hypothesis is mainly driven by the constraints imposed on the executives, such that the resource curse hypothesis holds for autocracies with limited constraints to the executive.

Ahmadov (2014) analyzes 29 econometric studies dealing with the resource curse literature by a statistical meta-analysis. He points out as the existing findings are inconclusive if analyzed separately[10], but, after removing the effects of differences on data coverage, model specifications and econometric approaches, then a negative association between oil and democracy across the globe emerges from his meta-analysis.

3. Empirical analysis

3.1. Data and econometric approach

To investigate the relationships between natural resource dependence, democracy, and economic growth, we consider the following baseline equation:

$$lnGDP_t = \alpha + \beta Res_t + \gamma Dem_t + \delta(Res_t * Dem_t) + \theta lnK_t + \rho FinDev_t + \mu_t \quad (1)$$

Where[11]:

- ln*GDP* stands for the natural log of Iranian real Gross Domestic Production in US dollars divided by the working-age population (people younger than 65 and older than 14 years old);
- *Res* indicates the proxy of resource dependence. It is measured as natural resource rents as a percentage of GDP.
- *Dem* is the proxy of democratization based on the index of Democracy (van_t) proposed by Vanhanen (2019). This index combines the degree of electoral competition - calculated by subtracting from 100 the percentage of votes won by the party that wins most votes in parliamentary elections or by the party of the successful candidate in presidential elections - and the degree of electoral participation - calculated as the percentage of the total population who voted in the election-. Specifically, it is formed by multiplying the degrees of electoral competition and electoral participation and then dividing the outcome by 100. In order to make the graphical interpretation easier, we rescale the original index on a scale from zero to ten by the following formula:

$$Dem_t = 10 * [van_t - min(van_t)] / [max(van_t) - min(van_t)].$$

- *Res#Dem* accounts for the interaction between the index of democratization and natural resource dependence. The inclusion of this interaction term plays a fundamental role in our research and follows the original intuition of Mehlum et al. (2006a, 2006b).

In order to avoid any potential omitted variable bias, we also control for variables typically used in the endogenous growth theory[12]:

- The proxy of physical investment per worker (*lnK)* is measured as the natural log of gross fixed capital formation divided by the working-age population. The inclusion of gross fixed capital formation per worker is aimed to approximate the effect of capital productivity factors, as in an "*Ak*" model of production function (e.g., Antonakakis et al., 2017).
- According to Yuxiang and Chen (2011), financial development plays an essential role in the relationship between resource wealth and long-run growth.[13] This proxy aims to control also for some purely economic explanations of the resource curse hypothesis. For instance, according to Van Der Ploeg and Poelhekke (2009), the volatility of the world price of natural resources is the mechanism of transmission between resource abundance and economic growth. In their hypothesis, a sound financial system can cope with large and sudden fluctuations in resource income and, accordingly, deactivates the resource curse. For Liu et al. (2015), financial development has a threshold-effect for economic growth: when the financial development is "over" the threshold, it can effectively alleviate the resource curse and promotes economic growth. Moradbeigi and Law (2017) deduce from their empirical analysis that a sound capital market helps to reduce the negative effect of oil abundance on the country's economic growth. Rongwei and Xiaoying (2020), find that financial

[10] 68% of studies report a negative correlation, 11% find a positive link, 21% do not find any statistically significant relationship.

[11] The data frequency is annual and spans over the period 1970–2017. Details on the database are provided in the appendix 1.

[12] We do not include as control other potential variable that may affect economic growth as trade openness. The economic rationale for this omission is that, for a country like Iran where the oil revenues weigh between half and three-quarters of the total exportation (and ¼ of GDP), resource dependence and international trade represent two faces of the same coin. Therefore, it may cause a relevant source of endogeneity. Moreover, given that the correlation between these two variables is significantly high (83%), its inclusion is also a source of multicollinearity without providing significant new information to explain economic growth.

[13] Bhattacharyya and Hodler (2014) investigate the opposite causal relationship i.e. from resource dependence to financial development. They find that resource-rich countries tends to be financially underdeveloped. They argue that "*the ruling elite has less incentive to foster contract enforcement while getting large natural resource rents, and because the financial sector cannot prosper without strong contract enforcement.*"

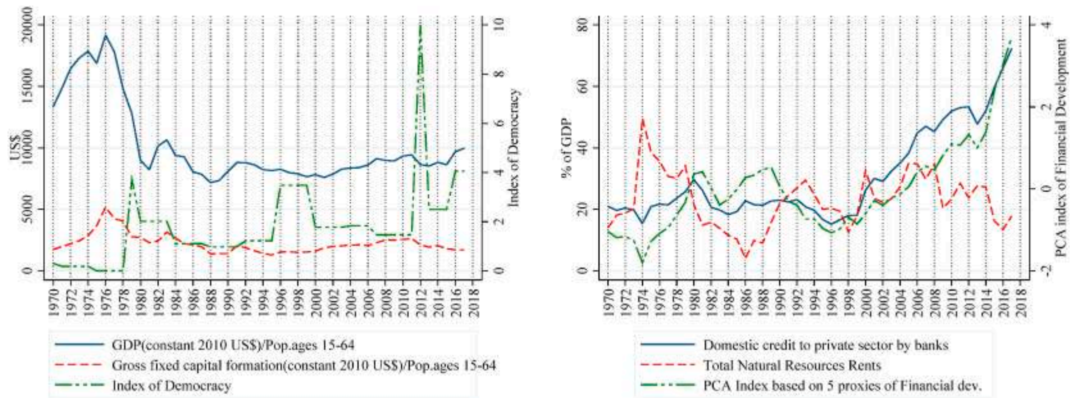

Fig. 1. Trends of the *variables*.

development has alleviated the "resource curse" in China trough two channels: on the short run, developed financial market contributes to reducing macroeconomic fluctuations (i.e., Van der Ploeg and Poelhekke's hypothesis); on the long-run, it provides financial support for education investment, human capital accumulation, and promotes scientific and technological progress. According to this theory, and following similar empirical research (e.g. Asif et al., 2020), we account for the role of financial development on economic growth by including the ratio between domestic credit to the private sector and gross domestic product (*FinDev)*. The numerator of this proxy refers to financial resources provided to the private sector by financial corporations, such as loans, purchases of nonequity securities, and trade credits and other accounts receivable, that establish a claim for repayment.[14] As a robustness check, we estimate an alternative index of financial development (*FinDev2*) based on five proxies of financial development extracted by the Global Financial Development Database and WDI.[15]

- Lastly, μ_t is the residual term and it is assumed to be normally distributed, not correlated, and with homoskedastic variance.

Fig. 1 reports the trends of *Real GDP* per capita*, Index of Democracy, Resource dependence, Gross fixed capital* per capita*, and two indexes of financial development.*

Fig. 1 highlights a significant decrease in GDP per capita between 1975 and 1981 and signals of nonstationarity and comovements between real GDP and physical capital. This graphical evidence suggests checking for the existence of a long-run equilibrium relation (i.e., cointegration). Accordingly, we apply the ARDL approach of cointegration originally proposed by Pesaran and Shin (1999) and furtherly developed by Pesaran et al. (2001) - (hereinafter PSS bounds test). This approach offers at least three advantages over the conventional cointegration testing: (i) it can be used with a mixture of stationary and integrated of order one series (see Table 1); (ii) it has a flexible specification in terms of lags to include for each variable; (iii) it can be easily rewritten as an Error Correction Model (ECM) that makes simpler testing the existence of long-run equilibrium and theoretical hypotheses on the structural relationships among variables.

3.2. Testing appropriateness of the PSS bond tests

Four main steps are required for the validity of PSS bond tests: (i) the variables should be purely I(0), purely I(1) or mutually cointegrated; (ii) there exists at most one conditional level relationship between the dependent variable (i.e. economic growth) and its regressors; (iii) the independent variables should be weakly exogenous for the parameters of interest (i.e. the coefficients of long-run equilibrium); (iv) the residuals of ARDL model should be normally distributed, homoskedastic, serially uncorrelated, and their coefficients should be stable over time.

The first preliminary test for the appropriateness of the ARDL model consists to reject the hypothesis that none of the variables is integrated of order two. Table 1 reports the order of integration of each series by applying Augmented Dickey-Fuller (ADF) and Phillips–Perron (PP) unit root tests.

Unit root analysis suggests a mixture of stationary and integrated of order one series, therefore, the existence of a cointegration vector is possible and an ARDL model may be beneficial to estimate the relationships among the variables in the long-run equilibrium.

The second test concerns the appropriateness of applying a single-equation approach instead of a system of equations - e.g. by a vector error-correction model (VECM) - without loss of relevant information. Specifically, we apply the Johansen cointegration procedure to determine the number of cointegrating equations in a VECM.[16] Table 2 reports Johansen's "*trace statistic*" (λ_{trace}), Schwarz's Bayesian (*SB*) and Hannan and Quinn (*HQ*) information criteria based on Johansen's maximum likelihood estimator of three[17] VECM specifications.

Statistical tests based on *trace statistics*, *SB* and *HQ* for all the three specifications converge to reject the null hypothesis that there is more than one cointegrating vector. Accordingly, we can rewrite the baseline equation (1) into an ARDL form:

$$
\begin{aligned}
lnGDP_t = {} & \alpha_0 + \sum_{i=1}^{p} \alpha_i lnGDP_{t-i} + \sum_{i=0}^{q} \beta_i Res_{t-i} + \sum_{i=0}^{r} \gamma_i Dem_{t-i} \\
& + \sum_{i=0}^{s} \delta_i (Res_{t-i} \# Dem_{t-i}) + + \sum_{i=0}^{v} \theta_i lnK_{t-i} + \sum_{i=0}^{z} \rho_i FinDev_{t-i} \\
& + \sigma D_{75-81} + \tau(Trend) + \mu_t
\end{aligned}
\tag{2}
$$

[14] In the last decade (2008–2017), the Iranian domestic credit to the private sector has been about 95% of the total domestic credit provided by the financial sector. It has shown a strong increase compared to the previous period (1970–2007) where the domestic credit provided to the private sector was only 55% of all the domestic credit to various sectors.

[15] Appendix 2 provides details on how this overall index has been estimated by applying principal component analysis.

[16] We apply the SB and the HQ information criteria to determine the optimal lag-order in a VAR that include all the variables of equation (1). The optimal lag-order for both criteria is one.

[17] I.e. a model that includes unrestricted constant, a second model including restricted constant and a third VECM with unrestricted linear trend.

Table 1
Unit Root Analysis (Null hypothesis: Unit Root).

Variable	ADF (C)	ADF (C + T)	PP (C)	PP (C + T)		ADF (C)	ADF (C + T)	PP (C)	PP (C + T)	Order
lnY_t	−2.109	−1.818	−4.301	−5.015	*F.Diff*	−4.800***	−4.906***	−28.78***	−28.27***	I(1)
Res_t	−3.019**	−2.964	−16.51**	−16.27	*F.Diff*	−7.709***	−7.648***	−49.47***	−49.73***	I(0/1)
Dem_t	−4.517***	−5.923***	−30.18***	−41.36***	*F.Diff*	−10.77***	−10.65***	−57.35***	−57.39***	I(0)
$Res_t\#Dem_t$	−4.790***	−6.197***	−32.05***	−40.83***	*F.Diff*	−10.56***	−10.44***	−56.44***	−56.43***	I(0)
lnK_t	−2.070	−2.425	−8.063	−10.09	*F.Diff*	−5.573***	−5.526***	−36.33***	−36.72***	I(1)
$FinDev_t$	2.116	−0.009	3.140	−0.505	*F.Diff*	−4.753***	−5.328***	−32.15***	−35.64***	I(1)
FinDev2	1.932	0.572	3.377	−0.284	*F.Diff*	−4.364***	−4.663***	−29.33***	−32.10***	I(1)

Note: *,**, *** indicate significance at 10%; 5% and 1%, respectively. "C" means only constant; "C + T" indicates that we include both constant and time trend. For ADF tests, we include zero or one lag according to their statistical significance; for PP tests, the number of Newey–West lags used in calculating the standard error is three.

Table 2
Johansen tests for cointegration.

H_0:	Unrestricted constant			Restricted constant			Restricted trend		
	λ_{trace}	*SB*	*HQ*	λ_{trace}	*SB*	*HQ*	λ_{trace}	*SB*	*HQ*
None (r = 0)	114.47	20.40	20.25	120.63	20.04	20.04	128.74	20.40	20.25
r ≤ 1	65.23*	20.25*	19.84*	70.82*	19.96*	19.67*	78.58*	20.32*	19.88*
r ≤ 2	42.43	20.51	19.87	47.76	20.29	19.75	54.62	20.63	19.94
r ≤ 3	24.64	20.70	19.89	29.71	20.56	19.83	36.33	20.89	20.01

Note: * indicates significance at 1%; Null hypothesis: there are no more than r cointegrating relations; Lag = 1; 47 Observations.

The third condition for the appropriateness of a single-equation approach to describe a VECM consists to test if the regressors of economic growth are weakly exogenous.[18] In economic terms, it implies that the deviations to long-run equilibrium cannot affect the regressors of economic growth but they may still react to lagged changes of the dependent variable (i.e. $\Delta lnGDP_{t-1}$).

In the following, we reparameterize the ARDL model (eq. (2)) into an ECM. In this form (eq. (3)), it provides short-run dynamics and long-run relationships of the determinants of the growth rate of real GDP per worker.

$$\begin{aligned}\Delta lnGDP_t = {} & c_0 - \phi(lnGDP_{t-1} - \omega_1 Res_{t-1} - \omega_2 Dem_{t-1} - \omega_3(Res_{t-1}\#Dem_{t-1}) \\ & - \omega_4 lnK_{t-1} - \omega_5 FinDev_{t-1}) + \sum_{i=1}^{p-1}\alpha_i^{sr}\Delta lnGDP_{t-i} + \sum_{i=0}^{q-1}\beta_i^{sr}\Delta Res_{t-i} \\ & + \sum_{i=0}^{r-1}\gamma_i^{sr}\Delta Dem_{t-i} + \sum_{i=0}^{s-1}\delta_i^{sr}(Res_{t-i} * Dem_{t-i}) + \sum_{i=0}^{v-1}\theta_i^{sr}\Delta lnK_{t-i} \\ & + \sum_{i=0}^{z-1}\rho_i^{sr}\Delta FinDev_{t-i} + \sigma D_{75-81} + \tau(Trend) + \mu_t\end{aligned} \tag{3}$$

Where:

- $\phi = 1 - \sum_{i=1}^{p}\alpha_i$ is the speed-of-adjustment coefficient and measures how strongly the GDP per capita reacts to a deviation from the equilibrium relationship in one period and can be interpreted as a measure of the speed at with the economy returns to the long-run equilibrium after a shock.
- $\omega_1 = \frac{\sum_{i=0}^{q}\beta_i}{\phi}$, $\omega_2 = \frac{\sum_{i=0}^{r}\gamma_i}{\phi}$, $\omega_3 = \frac{\sum_{i=0}^{s}\delta_i}{\phi}$, $\omega_4 = \frac{\sum_{i=0}^{v}\theta_i}{\phi}$ and $\omega_5 = \frac{\sum_{i=0}^{z}\rho_i}{\phi}$ measure the long-run coefficients and represent the equilibrium effects of the (weakly) exogenous variables on the dependent variable ($\Delta lnGDP_t$);
- α_i^{sr}, β_i^{sr}, γ_i^{sr}, δ_i^{sr}, θ_i^{sr}, ρ_i^{sr} are the short-run coefficients. They account for short-run fluctuations not due to deviations from the long-run equilibrium and the negative speed-of-adjustment coefficient.

The fourth step to appropriately apply the PSS bounds test consists in verifying whether the residuals of regression (2) are normally distributed, homoskedastic, serially uncorrelated, as well as whether there are stable coefficients over time. Table 3 reports these statistical tests on residuals. According to these statistical tests, we cannot reject the hypothesis that residuals are not significantly different from white noise. As a consequence of these preliminary tests, we can appropriately check the existence of structural relationships by the PSS bounds test. If the PSS bound test supports the existence of a cointegration vector, then we can use the long-run coefficients of ECM specification to test the key hypotheses of the political economy resource curse.

3.3. *Testing theoretical propositions on long-run effects of the resource curse hypothesis*

In this section, we describe how we apply the ARDL cointegration approach to answer to some of the main theoretical questions of this literature. Specifically, (Hp. 1) Is the political competition/democracy pivotal to reverse the resource dependence from "curse" into a "blessing"?, (Hp. 2) Does the resource curse emerge only if the political competition/democracy is sufficiently high? And, by looking at the same issue from a different perspective, this second question may be reformulated as follows: Is the political competition/democracy detrimental for economic growth if resources dependence is sufficiently high? These hypotheses are empirically tested as follows:

[18] In the context of cointegration, a variable is weakly exogenous if it "*does not respond to the discrepancy from the long-run equilibrium relationship*" (Enders, 2010: 407). See appendix 3 for these tests.

Table 3
ECM specifications - Dep. Variable: $\Delta lnGDP_t$.

Long-run c.	ECM 1	ECM 2	ECM 3(4)	ECM 4	Short-run c.	ECM 1	ECM 2	ECM 3	ECM 4
Res_{t-1}	0.033***	0.024***	0.038***	0.044***	$\Delta lnGDP_{t-1}$	−0.106	–	–	–
	(5.66)	(5.72)	(5.78)	(5.78)		(-0.92)			
Dem_{t-1}	0.451***	0.290***	0.344***	0.350***	$\Delta lnGDP_{t-2}$	−0.338***	–	–	–
	(5.41)	(4.69)	(5.48)	(5.57)		(-3.40)			
$Res\#Dem_{t-1}$	−0.021***	−0.014***	−0.015***	−0.016***	ΔRes_t	0.011***	0.010***	0.008***	0.009***
	(-5.74)	(-5.20)	(-5.59)	(-5.70)		(-6.02)	(5.59)	(3.87)	(4.76)
lnK_{t-1}	0.322***	0.405***	0.204*	0.296***	ΔRes_{t-1}	−0.001	–	−0.004***	−0.004***
	(3.84)	(4.84)	(1.94)	(3.36)		(-0.36)		(-2.80)	(-2.96)
$FinDev_{t-1}$	0.009***	0.002	0.011***	–	ΔRes_{t-2}	0.003**	–	–	–
	(3.57)	(0.88)	(3.20)			(2.26)			
$FinDev2_{t-1}$	–	–	–	0.114***	ΔRes_{t-3}	0.004**	–	–	–
				(3.38)		(2.51)			
Speed-of-adjust					ΔDem_t	0.077***	0.079***	0.528**	0.057**
$lnGDP_{t-1}$	−0.442***	−0.445***	−0.469***	−0.461***		(3.22)	(3.06)	(2.16)	(2.39)
	(-6.12)	(-6.06)	(-5.43)	(-5.56)	ΔDem_{t-1}	−0.060**	–	–	–
PSS Bound Test						(-2.12)			
F-stat§	11.190	10.289	10.795	11.595	ΔDem_{t-2}	−0.015*	–	–	–
I(0)	[0.000]	[0.000]	[0.000]	[0.000]		(-1.73)			
I(1)	[0.000]	[0.000]	[0.000]	[0.000]	ΔDem_{t-3}	−0.013**	–	–	–
t-stat§	−6.123	−6.062	−5.435	−5.555		(-2.14)			
I(0)	[0.000]	[0.000]	[0.000]	[0.000]	$\Delta Res\#Dem_t$	−0.003***	-.004***	−0.002**	−0.003**
I(1)	[0.001]	[0.001]	[0.010]	[0.008]		(-3.59)	(-3.47)	(-2.38)	(-2.60)
Residuals Tests					$\Delta Res\#Dem_{t-1}$	0.002*	–	–	–
H_0: *no autoc.*[a]	0.496	0.049	0.191	0.243		(1.96)			
H_0: *Const.var*[b]	0.711	0.660	0.750	0.892	Δk_t	−0.143***	0.180***	0.096*	0.137***
H_0: *homosked.*[c]	0.738	0.311	0.694	0.654		(-2.85)	(-2.25)	(1.76)	(2.78)
H_0: *Normal dis.*[d]	0.873	0.000	0.009	0.000	$\Delta FinDev_t$	−0.007**	0.001	0.005***	–
H_0: *White noise* [e]	0.797	0.546	0.900	0.946		(-2.08)	(0.93)	(3.16)	
H_0: *No str.break* [f]	0.308	0.554	0.686	0.598	$\Delta FinDev_{t-1}$	−0.007**	–	–	–
Observations	44	44	44	44		(-2.14)			
Sample	'74-'17	'74-'17	'74-'17	'74-'17	$\Delta FinDev_{t-2}$	−0.011***	–	–	–
Adjusted R^2	0.770	0.576	0.679	0.694		(-3.17)			
R^2	0.882	0.655	0.769	0.780	$\Delta FinDev2_t$	–	–	–	.053***
AIC	−137.164	−115.856	−125.485	−127.636					(3.47)
BIC	−97.912	−99.799	−102.290	−104.441	*Constant*	2.520***	2.431***	22.34***	17.319***
Weak Exogeneity	yes	Yes	no	no	*Dum75-81*	–	–	-0.202***	-0.207***
# Cointegr. rank	1	1	1	1	*Trend*	–	–	-0.010***	-0.007***

Notes: t-statistics in parentheses; ***p-value<0.01, ** p-value<0.05, * p-value<0.1; § we report estimated statistics and approximate p-values based on Kripfganz and Schneider (2018) critical values in squared parentheses; [a] lowest p-values of the Durbin's alternative test for autocorrelation up to 4 lags; [b]Breusch-Pagan/Cook-Weisberg test for heteroskedasticity; [c]Cameron and Trivedi test for overall heteroskedasticity, skewness, and kurtosis; [d] D'Agostino, Belanger, and D'Agostino Skewness/Kurtosis tests for Normality; [e]Portmanteau's test for white noise; [f] Cumulative sum test for parameter stability at 1% (test statistics). For this test, we cannot reject the null hypothesis of a constant mean at the 1% level because the test statistic values are always lower than the 1% critical level of 1.143.

Hypothesis 1. - "the political economy of the resource curse hypothesis"

Resource dependence is harmful to economic growth only if the level of democracy is higher than the following threshold: $-\omega_1/\omega_3$. This threshold is calculated through the overall marginal effect of resource dependence on economic growth in the long-run equilibrium as estimated in eq. (3):

$$\frac{\partial \Delta lnGDP_t}{\partial Res_{t-1}} = \begin{cases} \omega_1 + \omega_3 Dem_{t-1} < 0 & if\ Dem_{t-1} > -\omega_1/\omega_3 \\ \omega_1 + \omega_3 Dem_{t-1} > 0 & if\ Dem_{t-1} < -\omega_1/\omega_3 \end{cases}.$$

Hypothesis 2. - "Democratization is harmful to economic growth in resource-rich countries"

Democratization is harmful to economic growth only if the level of resource dependence is higher than the following threshold: $-\omega_2/\omega_3$. As the previous hypothesis, the threshold is estimated through the overall marginal effect of democracy on economic growth in the long-run equilibrium:

$$\frac{\partial \Delta lnGDP_t}{\partial Dem_{t-1}} = \begin{cases} \omega_2 + \omega_3 Res_{t-1} < 0 & if\ Res_{t-1} > -\omega_2/\omega_3 \\ \omega_2 + \omega_3 Res_{t-1} > 0 & if\ Res_{t-1} < -\omega_2/\omega_3 \end{cases}.$$

4. Empirical results

Table 3 reports the estimated coefficients, residuals diagnostics, goodness-of-fit tests, and the outputs of the PSS bounds tests for cointegration based on four alternative ECM specifications (i.e. with and without a dichotomous variable and time trend, and alternative optimal lag-order structures).[19]

The residuals tests indicate to not reject the null hypotheses of no serial correlation, homoskedasticity, white noise, and absence of structural break. However, on the basis of D'Agostino, Belanger, and D'Agostino Skewness/Kurtosis tests, we cannot reject the hypothesis that residuals are normally distributed only for the ECM 1.

As the PSS bounds test for cointegration concerns, it uses two separate statistics: an *F-test* on the joint null hypothesis that the coefficients on the level variables are jointly equal to zero and a *t-test* on the lagged level dependent variable. Given that for all the estimated models, the *F-*

[19] Specifically the lag structures for the ARDL(*p,q,r,s,v,z*) specifications are based on the optimal lag lengths with the Akaike information criterion (AIC) for Model 1 i.e. ARDL(3,4,4,2,0,3) while for model 2 ARDL(1,0,1,1,0,0), for model 3 ARDL(1,2,1,1,0,0)and for model 4 ARDL(1,2,1,1,0,0) the optimal lag lengths are selected with the Bayesian information criterion (BIC).

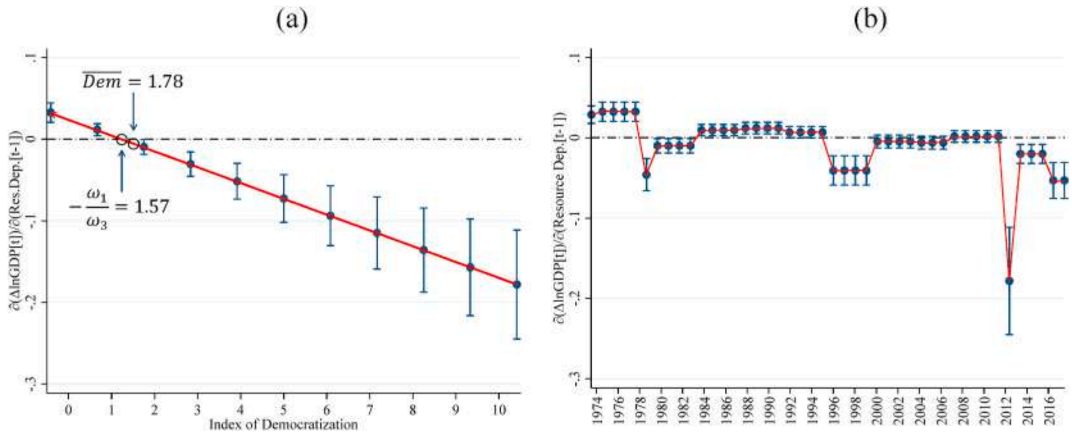

Fig. 2. Overall marginal effects of resource dependence on economic growth.

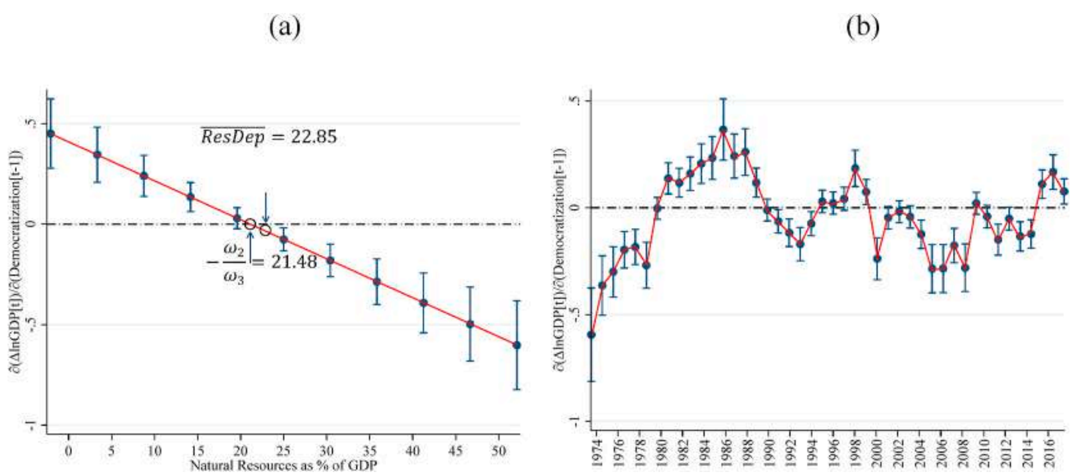

Fig. 3. Overall marginal effects of democracy on economic growth.

statistics and *t*-statistics exceed their respective upper critical values at 5%, we can reject the null hypothesis of no cointegration. Then, we conclude that there is evidence of a long-run relationship among the growth of GDP and the other covariates. These results are robust to alternative ARDL specifications. Taking into account that the marginal models of ARDL specifications 1 and 2 confirm the hypothesis of weak exogenous regressors,[20] we conclude the ARDL model 1 is the best specification because it has the highest adjusted R^2, the lowest AIC statistics and satisfies all the hypotheses request by Pesaran et al. (2001) for their approach of cointegration.

As the Hypothesis 1 - "the *political economy of the resource curse hypothesis*" concerns, we find that the index of democratization is on average higher ($\overline{Dem} = 1.78$) than, but not statistically different from, the threshold ($-\omega_1/\omega_3 = 1.57$), therefore, in the long-run, the hypothesis of the resource curse is not validated in Iran.[21] However, this result may be misleading. In order to make this statement clear, we follow Greene's (2010) suggestion to provide a graphical representation of the overall marginal effect.

Fig. 2(a) shows the estimated overall marginal effects of resource dependence on economic growth ($\partial \Delta lnGDP_t / \partial Res_{t-1}$) in the long-run equilibrium and their confidence intervals at 95% level over the range of democratization index in our sample (0–10). Fig. 2(b) displays how these effects change over time.[22]

These graphics highlight why, the common practice of testing the *political economy of the resource curse hypothesis* - i.e. checking if the (average) value of the institutional index is higher than the estimated threshold - may be deceiving to infer whether the resource curse exists. Indeed, if we consider the annual estimates of the overall effect instead of the average, we realize, as, for most of the observed periods, natural resources dependence has been detrimental for the Iranian economic development. Precisely, we find that resource dependence has been a "curse" for 15 years (1979–1983, 1996–1999 and 2012–2017), a "blessing" for 13 years (1974–1978 and 1984–1991) and it did not have a statistically significant effect (at 5% level of significance) on the economic growth for 16 years (1992–1995, and 2000–2011).

As the Hypothesis 2 - "*Democratization is harmful to economic growth*

[20] Appendix 3 provides details on the variable addition test proposed by Johanson (1992) to establish if weak exogeneity hypothesis holds for ARDL specifications.

[21] The average level of democracy over the sample is 1.78, at this level $\partial \Delta lnGDP_t / \partial Res_{t-1} = -0.005$ with a *t-statistics* $= -1.28$ and confidence intervals at 95% equal to -0.013 and 0.003.

[22] We report the estimates over the period 1974–2017 because the ECM 1 specification has q = 4, therefore, the first four years of the sample are lost.

for resource-rich countries" concerns, we find that the hypothesis that democratization is detrimental for long-term growth should be rejected because the level of resource dependence has been, on average, higher ($\overline{ResDep} = 22.85$) than, but not statistically different from, the threshold ($-\ \omega_2/\omega_3 =\ 21.48$).[23] Once again, this result may be misleading because, as Fig. 3(a) and 3(b) show, the sign of the overall marginal effect of democratization on growth significantly fluctuates around the threshold over the period.

Specifically, democratization has been beneficial for growth when resource dependence has been lower than 21.48% of GDP - it has occurred for 14 years (1981–1989; 1998–2000 and 2015–2017) - the opposite effect has occurred for 19 years (1974–1979, 1991–1994, 2004–2008, 2011 and 2013–2014) while democratization did not have effects on GDP for 11 years (1980, 1990, 1995–1997, 2001–2003, 2009–2010 and 2012). Fig. 3 shows these findings.

5. Conclusion

This research aims to contribute to the political economy of the resource curse hypothesis literature. We show as testing the hypothesis that resource curse is conditional on institutional quality - i.e. by checking the overall marginal effect of a model that includes an interaction term between proxies of resource wealth and institutional quality - may be misleading when based on the standard practice to use the average level of the variables. This study contributes to the current debate in several ways.

From a methodological viewpoint, we investigate the long-run relationships between economic growth, democratization, and resource dependence based on the ARDL approach. Following Green's (2010: 291) suggestion "*that the common practice of testing hypotheses about partial effects is less informative than one might hope, and could usefully be omitted from empirical analyses*", we propose a graphical representation of the overall marginal impacts of resource dependence and democracy on the long-run growth of GDP. This analysis has shown that, at least for one of the world's resource-dependent countries as Iran, is not always possible to answer the question whether resource dependence is a curse or blessing only using "yes" or "no" answer. A sounder scientific query should also concern "when" curse or blessing has arisen.

Our findings have revealed that, although the overall marginal effect of resource dependence on growth depends critically on the value of democracy, as expected by the theory of the political economy of the resource curse, this effect changes over time and clear-cut answers, straightforwardly based on a single value (e.g. the average), may be misleading.

From a positive viewpoint, a further result deals with the consequence of democratization for young democracies with richness in natural resources. In the same way as for the resource curse hypothesis, we have observed that the sign of the overall marginal effect of democracy on economic development significantly changes over time, therefore straightforward conclusions are still not possible.

From a theoretical viewpoint, we find empirical evidence of Collier and Hoefflier's (2009) hypothesis, that an increase of electoral competition may hamper economic growth. The rationale is that, at the early stages of democracy, the system of checks and balances are still insufficient to prevent the negative side effect in terms of misallocation of resources due to rent-seeking activities. Our results may also support the so-called "rentier effect" hypothesis if we assume that the natural resources dependence is a proxy of government reliance on non-tax revenue.[24] According to the "rentier effect" hypothesis, an abundant flow of natural resources revenues by reducing citizens' demand for greater accountability slows down the development of the system of checks and balances that is one of the most relevant aspects of the democratization process.

The results presented in this article can offer some opportunities for future research. The proxy used in our model captures just one type of institutional setting (i.e. democratization) but other features can affect the relationship between resource dependence and economic outcome (e.g. corruption, rule of law, etc.). It may be interesting to test if our results are robust to other measures of institutional quality and control variables. A further possible extension of this analysis consists in using a proxy of resource abundance instead of resource dependence.[25] Finally yet importantly, as soon as the data availability allows examining long-run relationships between resource dependence, democracy and growth for a sample of resource-rich countries and with a sufficiently long period, an analysis based on a panel ARDL approach may be a valuable contribution to this literature.

Declaration of competing interest

I declare that I have no conflict of interest.

[23] The average of resource dependence is 22.85%, therefore we get $\partial \Delta lnGDP_t / \partial Dem_{t-1} = -0.031$ with a *t-statistics* $= -1.23$ and confidence intervals at 95% equal to -0.082 and 0.021. This test suggests not rejecting the null hypothesis that democracy has no effects on long-term growth of GDP per capita.

[24] At least in the last decade this may be a reliable hypothesis because oil and natural gas revenues represent around the 55% of the Iranian government revenues (Central Bank of Iran, 2018).

[25] See Kropf (2010) and Shahbaz et al. (2019) for an overview on how empirical findings on resource curse hypothesis depend critically on how natural resource wealth is measured.

Appendix 1. Dataset (1970–2017)

Variable	Definitions of Manifest Variables (Indicators)	Sources (code series)	Mean	Min	Max
Real GDP per worker *(GDP)*	GDP (constant 2010 US$)/(Population ages 15–64, total)	WDI (NY.GDP.MKTP.KD/ SP.POP.1564.TO)	10,167.6	7,175.6	19,181.4
Resource Dependence (*Res*)	Total natural resources rents (% of GDP). Missing values in 1991 and 1992 are replaced by linear interpolation with the 1990 and 1993 values.	WDI (NY.GDP.TOTL.RT. ZS)	22.85	4.01	49.62
Democratization Index (*Dem*)	We standardize the index of democracy proposed by Vanhanen (2019; van_index) on a scale 0–10. (0 = No Democracy, 10 = maximum level of democracy over the period 1970–2017 in Iran)	QoG (van_index)	1.78	0.00	10
Fixed Capital per worker (*K*)	Gross fixed capital formation (constant 2010 US$)$)/(Population ages 15–64, total)	WDI (NE.GDI.FTOT.KD/ SP.POP.1564.TO)	2,222.7	1,280.3	5,182.8
Financial Development (*FinDev*)	Domestic credit to the private sector by banks (% of GDP). Missing value in 1978, and 2017 are replaced by linear interpolation.	WDI (FD.AST.PRVT.GD. ZS)	30.41	15.18	72.35
Financial Development2 (*FinDev2*)	Own calculation based on PCA using *FinDev; Fin_var2; Fin_var3; Fin_var4 and Fin_var5.*	See Appendix 2	0.00	−1.81	3.71
Fin_var2	Domestic credit provided by the financial sector (% of GDP). 4 Missing values replaced by linear interpolation.	WDI (FS.AST.DOMS.GD. ZS)	48.87	16.63	86.75
Fin_var 3	Index of Financial "Depth": Deposit money bank assets to deposit money bank assets and central bank assets (%). 3 Missing values replaced by linear interpolation.	Global Financial Development (GFDD. DI.04)	69.94	36.82	105.79
Fin_var 4	Bank deposits to GDP (%). 6 Missing values replaced by linear interpolation.	Global Financial Development (GFDD. OI.02)	37.13	15.78	87.98
Fin_var 5	Index of Financial "Stability": Bank credit to bank deposits (%). 3 Missing values replaced by linear interpolation.	Global Financial Development (GFDD. SI.04)	74.92	43.16	107.45

Note: "WDI" stands for World Development Indicators, version May 2020, published by World Bank (2020); "QoG" stands for The Quality of Government Basic Dataset, version Jan 2020, published by Dahlberg et al. (2020).

Appendix 2. Alternative proxies of Financial Development

Rongwei and Xiaoying (2020) state that the most used indicators to measure financial development are the ratio of private credit to GDP, the ratio of current liabilities to GDP, and the ratio of deposits and loans to GDP. However, with the exclusion of the ratio of private credit to GDP, the other two proxies are not available for Iran over the period 1970–2017. Consequently, once extracting all the indicators related to financial development available over the time span of our analysis from the Global Financial Development Database and World Development Indicators. We provide a set of robustness checks with alternative indexes of financial development.

Specifically, we use both a single-proxy approach - i.e. "*Domestic credit to the private sector by banks as a percentage of GDP*" (*FinDev*) and a multivariate statistical technique (i.e., based on Principal component analysis - PCA) to reduce all the correlated proxies of the financial system into an overall index. In the latter approach, the five correlated indicators of financial development include two indexes of the "size" of domestic credit as a percentage of GDP (i.e. "*Domestic credit to the private sector by banks*" and "*Domestic credit provided by financial sector"*); an index of financial "Depth" (i. e. "*Deposit money bank assets as a percentage of deposit money bank assets and central bank assets*"); an index of financial "Stability" (i.e. "*Bank credit as a percentage of bank deposits*"; and index of the size of bank deposit as a percentage of GDP (i.e. "*Bank deposits to GDP")*.

Table A.1 shows the five uncorrelated components, where each component is a linear weighted combination of the initial (correlated) variables. The components are ordered so that the first component (PC1) explains the largest possible amount of variation in the original data, subject to the constraint that the sum of the squared weights formula is equal to one. To control if the results depend on the statistical technique for data reduction, we also estimate an overall index of "financial development" using the Factor Analysis (FA) approach based on the method of principal factors.

Table A.1
Eigenvalues of the correlation matrix - PCA and FA.

Components	Eigenvalue	Difference	Proportion	Cumulative
Compon. 1	2.865	0.946	0.573	**0.573**
Compon. 2	1.920	1.755	0.384	**0.957**
Compon. 3	0.165	0.125	0.033	**0.990**
Compon. 4	0.040	0.030	0.008	**0.998**
Compon. 5	0.010	–	0.002	**1.000**
Factors				
Factor 1	2.819	0.973	0.599	**0.599**
Factor 2	1.847	1.755	0.392	**0.990**
Factor 3	0.092	0.105	0.020	**1.010**
Factor 4	−0.013	0.021	−0.003	**1.007**
Factor 5	−0.034	–	−0.007	**1.000**

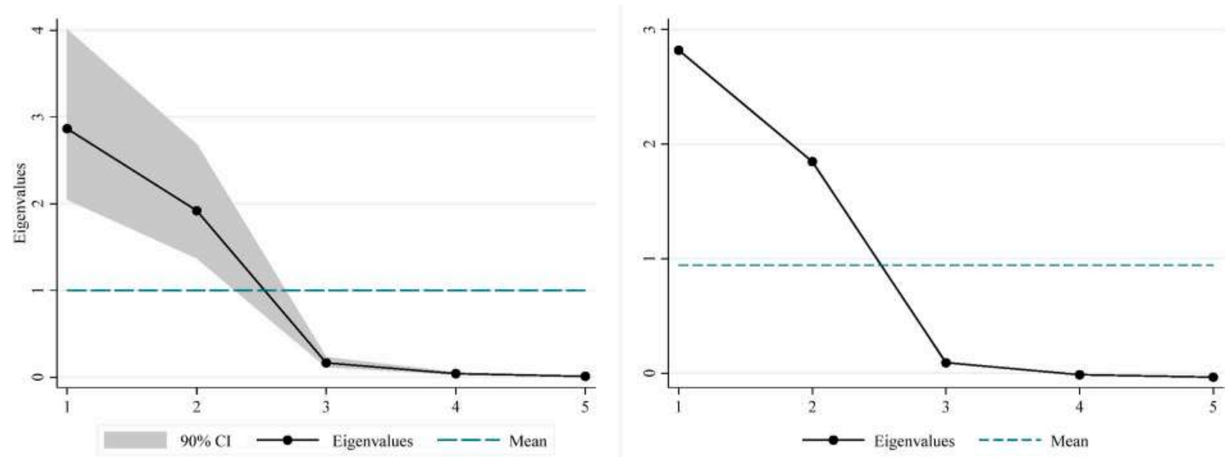

Fig. A.1. Scree plots of the eigenvalues of correlation matrix - PCA (left) and FA(right).

The scree plots of the eigenvalues of correlation matrixes suggest to consider the first two PCs and Factors (Fig. A.1). Table A.1 shows that the first two PCs (Factors) explain more than 95% of the variance. Table A.2 reports the eigenvectors and the unexplained variance of variables by using only 2 PCs and Factor loadings.

Table A.2
Eigenvectors of the first two PCs and Factors.

	Principal Component Analysis			Factor Loadings		
Variables	PC1	PC 2	*Unexplained*	Factor 1	Factor 2	*Uniqueness*
Fin_Dev	0.586	0.024	***0.014***	0.994	0.027	***0.012***
Fin_dep2	0.270	0.619	***0.056***	0.448	0.824	***0.120***
Fin_dep3	0.483	−0.385	***0.048***	0.799	−0.525	***0.086***
Fin_dep4	0.474	0.416	***0.024***	0.805	0.576	***0.020***
Fin_dep5	0.354	−0.544	***0.073***	0.588	−0.749	***0.095***

The indexes of financial development - $FinDev(PCA)$ and $FinDev(FA)$ - are calculated as a weighted average of the first two PCs and factors, respectively. Specifically, the weights are fixed equal to the proportion of explained variance (see Table A.1) as follows: $FinDev(PCA) = 0.5731 \cdot PC1 + 0.3840 \cdot PC2$ and $FinDev(FA) = 0.5985 \cdot Factor1 + 0.3920 \cdot Factor2$.

Fig. A.2 shows each overall index compared to the first two PCs and Factors.

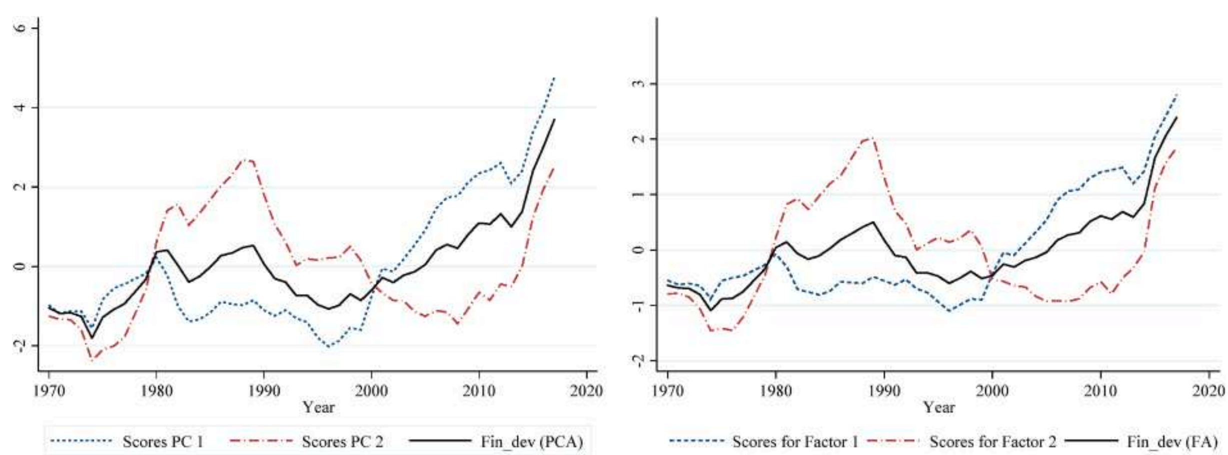

Fig. A.2. PC scores; Factor scores; *FinDev(PCA)* and *FinDev(FA)*.

Fig. A.3 shows the trends of the three indexes of financial development used in this empirical research.

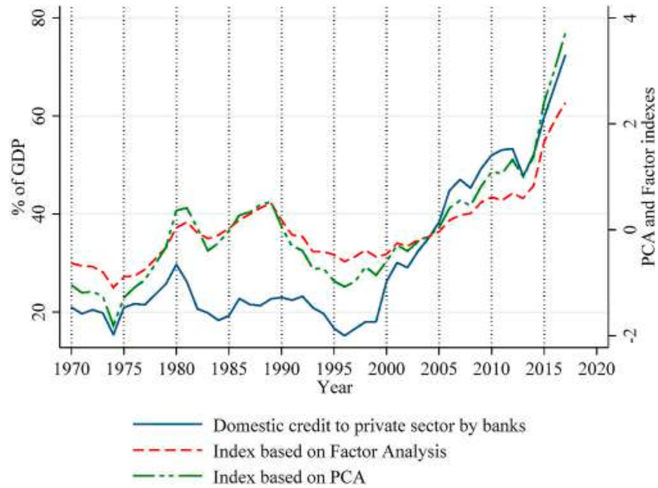

Fig. A.3. Financial Development indexes.

In the main text of the article, for the sake of brevity, we only report the empirical results based on *FinDev* and *FinDev2*, however, empirical outcomes based on factor analysis - i.e. *FinDev(FA)* - or using only the first component *(PC1)* or the first factor *(Factor1)* are qualitatively the same of reported findings of the model ECM 4 in Table 3. Table A.3 reports the correlations among these five proxies.

Table A.3
Correlation Matrix of the Financial Development Indexes.

	FinDev	FinDev2	PC1	FinDev(FA)	Factor1
FinDev	1.000				
FinDev2	0.886	1.000			
PC1	0.993	0.877	1.000		
FinDev(FA)	0.849	0.991	0.831	1.000	
Factor1	0.998	0.877	0.996	0.840	1.000

Appendix 3. Test for weak exogeneity

The objective of this appendix is to examine if the hypothesis of weak exogeneity holds in our model, in this case, efficient inference about the cointegration parameters can be conducted in a single equation framework. As Pesaran et al. (2001 - assumption 4) demonstrate, the weak exogeneity is one of the necessary conditions for the validity of the ARDL approach to test for the existence of long-run relationships among variables. Johansen (1992) provides a general test for weak exogeneity. By rewriting the (conditional) ARDL model in an error correct form:

$$\Delta y_t = c_0 - \phi(y_{t-1} - \omega X_{t-1}) + \sum_{i=1}^{p-1} \alpha_i^{sr} \Delta y_{t-i} + \sum_{i=0}^{q-1} \omega_i^{sr} \Delta X_{t-i} + \tag{a.1}$$

$$+\sigma D_{75-81} + \tau(Trend) + \varepsilon_t$$

Where: y_t is the natural log of GDP per capita, the vector X includes both the variables of interest (i.e. the indexes of democratization, natural resource dependence, and their interactions) and the control variables (i.e., *lnK; FinDev; FinDev2*) of our baseline regression; D_{75-81} is the dichotomous variable (from 1975 to 1981) and *Trend* is the deterministic linear trend.

We define that X is weakly exogenous for the parameters of (a.1). if the parameters of interest (i.e., the cointegration coefficients) are a function of the parameters in the conditional model (eq. a.1) and (ii) if the parameters in the conditional model and marginal model (eq. a.2) are variation-free, so they do not have any joint restrictions.

Accordingly, a feasible test for weak exogeneity implies to estimate the marginal models for each of the variables included in the vector X using a variable addition test to assess the statistical significance of the error correction term (ECT) estimated by equation (a.1).[26] Specifically, the marginal models are as follows:

$$\Delta x_t = c_0 + \delta \widehat{ECT}_{t-1} + \sum_{i=1}^{2} \alpha_i \Delta y_{t-i} + \sum_{i=1}^{2} \beta_i \Delta X_{t-i} + \sigma D_{75-81} + \tau(Trend) + \varepsilon_t \tag{a.2}$$

Where: x_t indicates each element of the vector X ; $\widehat{ECT}_{t-1} = (y_{t-1} - \widehat{\omega} X_{t-1})$ is the estimated long-run equilibrium error correction term of the conditional model; the dummy and/or the deterministic linear trend are included in the marginal regression only if their t-tests are statistically significant. The null hypothesis of weak exogeneity involves testing $\delta = 0$ - both as Likelihood-Ratio tests on each x variable and as

[26] See Calderòn et al. (2015) and Gemmel et al. (2016) for applications of this test on panel data.

a Wald test on the δ s that all the potentially endogenous variables are jointly equal to zero-. Taking into account that weak exogeneity tests are carried out with reference to a specific set of parameters of interest, we preliminary determine the optimal lag structure and the inclusion of dummy and deterministic trend of the marginal models by SB, AI and *adjusted R^2*.

Table A.4
Weak exogeneity tests.

Dependent var.	**ARDL 1**	**ARDL 2**	**ARDL 3**	**ARDL 4**
Δ*Res*	1.18	1.56	8.99***	12.08***
	(0.28)	(0.21)	(0.00)	(0.00)
Δ*Dem*	0.36	0.40	0.00	1.06
	(0.55)	(0.53)	(0.95)	(0.30)
Δ*Res#Dem*	1.09	0.11	0.10	0.66
	(0.30)	(0.74)	(0.76)	(0.42)
Δ*lnK*	2.55	1.68	0.73	0.60
	(0.11)	(0.74)	(0.39)	(0.44)
Δ*FinDev*	0.13	0.01	4.32**	–
	(0.72)	(0.94)	(0.04)	
Δ*FinDev*2	–	–	–	4.55**
				(0.03)
H_0: $\forall\delta = 0$	8.48	3.98	18.96***	19.91***
	(0.13)	(0.55)	(0.00)	(0.00)

Note: we report the χ^2 distributed with 1 degree of freedom for LR test for the tests on single coefficient and the Wald test with 5 degrees of freedom for joint tests on δ_s; p-values are in parentheses. *, **, *** indicate significant at 10%, 5%, 1% level. Further details are available upon request.

Table A.4 shows that we cannot reject the null hypothesis of weak exogeneity at 1% level for the ARDL models 1 and 2. This result suggests that for these two model specifications the *X*s affect economic growth but not the other way round in the long-run. In economic terms, this means that there is no need to model full mechanisms of independent (i.e. resource dependence, institutional and control) variables and economic growth jointly - e.g. by a VECM - as these variables are separable in the way that, holding economic growth fixed, we can infer all the effects from the conditional model (a.1).

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