Corruption and Growth: Long-run Historical Evidence (1790-2010)

This version: July 2021

Abstract
We employ a newly assembled indicator of corruption from Varieties of Democracy (V-Dem) to examine the effects of corruption on economic growth. The V-Dem indicator is coded for almost all contemporary and historical polities since the year 1900 and, for some countries, since the French Revolution. This global dataset allows us to exploit long-run, slow-moving variation within countries for identification, circumventing many of the difficulties faced by previous studies based on cross-section data or short panels. We present robust evidence of a negative effect of corruption on economic growth. Yet, we find that corruption interacts with political regime type, giving rise to heterogeneous effects. In particular, corruption is found to be significantly more deleterious for growth in democracies than in autocracies. Since corruption tends to be decentralised in democracies and centralised in autocracies, these findings are in line with theories of the industrial organisation of corruption. We find little to no evidence that other features of the institutional environment (state capacity, regulatory quality, property rights protection) exert a moderating influence on the magnitude of the corruption effect, casting doubt on the 'grease the wheels' hypothesis. Our findings provide a rational to target anti-corruption efforts and resources to young democracies.

Keywords
Corruption, economic growth, democracy

JEL Codes
D73, O43, P48

Highlights

- New expert-coded corruption indicator covering 186 countries over two centuries
- Long-run, slow-moving variation in corruption levels within countries used for identification
- Corruption has a negative effect on steady-state growth
- The negative effect is large in democracies, but significantly smaller in autocracies

1. Introduction

Since the early 1990s, corruption – the abuse of public office for private or sectional gain – has been regarded as one of the main obstacles to economic growth and development. Accordingly, national governments and international agencies have devoted an increasing share of their budgets to fighting graft, both at home and abroad (Marquette, 2003). But does corruption actually harm economic performance? If so, under what circumstances is corruption most harmful? Can corruption ever be compatible with (or even conducive to) economic dynamism?

The available evidence is suggestive of a negative effect of corruption on economic growth (see Ugur, 2013 for a meta-analysis). Yet, existing analyses are based primarily on cross-country regressions. Meanwhile, the few recent studies based on panel data rely on a very limited amount of over-time variation for identification. Due to these limitations, all existing empirical studies face a number of critical threats to causal inference, leading Aidt (2009: 288) to conclude that 'the search for a negative effect of corruption on the average growth rate of GDP per capita has failed to produce convincing and robust evidence'.

Here, we analyse a newly assembled, indicator of corruption from the Varieties of Democracy (V-Dem) dataset (Coppedge et al., 2020a). This data provides reliable, expert-coded information on the prevalence of corruption at the country-year level, covering almost all contemporary and historical polities since the year 1900 and, for some countries, since the French Revolution (1789). As such, V-Dem offers by far the most comprehensive data source on corruption available to date, improving dramatically upon all existing indicators, including Transparency International's Corruption Perceptions Index (1997-2019) and the International Country Risk Guide (ICRG) index (1982-2019). Although now at its 10th edition, the V-Dem indicator of corruption has not been used to shed light on the relationship between corruption and economic growth. By combining the V-Dem data with long-run historical information on GDP and population from Farris et al. (2017), this paper fills this gap, overcoming the severe data limitation problems that plague previous studies.

The identification of the causal impact of corruption on economic growth faces a number of critical challenges, which we address in the present study. *First*, cross-sectional regression estimates, as those in Mauro (1995), Mo (2001) and Pellegrini and Gerlagh (2004), are likely biased by the omission of country-level unobservables. In particular, cultural characteristics and historical legacies, which are idiosyncratic and difficult to measure, may exert a joint influence on corruption and economic performance. A related problem is that the earliest indicators of corruption were coded in the mid-1980s, which means that corruption is typically measured in the middle or even at the end of the growth period examined (e.g. 1970-2000). Thus, cross-sectional estimates are valid under the assumption that the relative prevalence of corruption across countries is stable over time. While plausible in the very short run, this assumption becomes much stronger when growth periods of 20-

30 years are considered.[1] Here, we address these two problems by using panel-data estimates with country fixed effects.

Second, existing panel-data estimates, such as those in D'Agostino et al. (2016a) and Cieślik and Goczek (2018), are only valid under the assumption that corruption indicators track the 'true' level of corruption precisely enough to permit a meaningful identification of causal effects from short-run variation. Unfortunately, the nature of existing corruption indices complicates 'any meaningful analysis of their variation over time, at least until longer time series are available' (Méon and Sekkat, 2005: 80). For one thing, existing indices are prone to small 'jumps' resulting from random measurement error (Treisman, 2007; Standaert, 2015). For another, the structural trend around which they fluctuate is subject to strong inertia. To circumvent these problems, we turn to the unprecedented time coverage provided by the V-Dem dataset, exploiting long-run, slow-moving variation in corruption levels within countries for identification. This variation is unlikely to result primarily from white noise, mitigating concerns of attenuation bias and increasing the precision of standard estimators.

Third, the 'dynamics of the relationship between corruption, governance, and economic performance […] are not yet fully understood' (Méon and Weill, 2010: 254). For one thing, the true effects of corruption on economic performance may be confounded by other time-varying features of the institutional environment that correlate with corruption. For another, the institutional environment may exacerbate or attenuate the economic consequences of corruption, giving rise to interaction effects. To make progress on these fronts, we first condition our fixed-effects estimates on other measures of institutional quality – state capacity, regulatory quality, property rights protection and the quality of democracy. Next, we add multiplicative terms to explore in a panel-date framework how corruption interacts with various dimensions of the institutional environment. In contrast to previous studies, which typically focus on one or two institutional dimensions as potential effect-modifiers, we examine a number of possible interactions in a systematic framework.

Lastly, our analysis tackles other potential threats to identification. For starters, we control for a number of observable determinants of economic growth that vary over time. Since the 'true' effect of corruption may also be confounded by the influence of time-varying unobservables, we present alternative specifications in which we model explicitly the dynamic adjustment of GDP, as in recent work by Acemoglu et al. (2019). In robustness tests, we employ two strategies to further mitigate concerns of simultaneity and omitted variable bias. First, we use instrumental variables to

[1] Existing attempts to address this threat to identification by means of (time-invariant) instrumental variables – e.g. ethno-linguistic fractionalisation, the share of Protestant adherents (e.g. Mauro, 1995, Lambsdorff, 2003) – are subject to the usual concerns regarding instrument validity.

isolate an exogenous component of variation in corruption levels over time. Second, we consider estimates that net out the influence of country- or region-specific time trends.

We find robust evidence of a negative causal effect of corruption on economic growth. This effect holds on average and, being identified from long-run time-series variation, it can be interpreted as an equilibrium effect. Yet, we find that the absolute magnitude of the corruption effect varies substantially depending on the quality of democratic institutions. Across a variety of model specifications, corruption is significantly less detrimental for economic performance in more autocratic regimes. An explanation for this finding, for which we find suggestive evidence in the data, is that the type of corruption that typically prevails in autocracies (centralised) is less harmful for growth than the form of corruption (decentralised) that is commonly found in democracies (Shleifer and Vishny, 1993; Ehrlich and Lui, 1999).

We find little to no evidence that other features of the institutional environment – including state capacity, regulatory quality, property rights protection – exert a moderating influence on the corruption-growth relationship. This finding runs counter to the so-called 'grease the wheels' hypothesis, which claims that corruption may help firms and households circumvent pre-existing institutional dysfunctions, thus promoting (or at least not harming) economic performance in environments with weak institutions.

These findings add to a large empirical literature on corruption and growth, which is reviewed systematically in Appendix A. Virtually all the studies reviewed share the limitations identified earlier, and the validity of their results is questionable. The closest paper to ours is Gründler and Potrafke (2019), which examines the 'grease the wheels' hypothesis in a panel-data framework. Yet, their data cover as few as seven years, and they use *forecasts* of real GDP per capita to populate the most recent years of the dataset (2019: 4n).

The paper proceeds as follows. To guide the empirical analysis, we first present a theoretical framework. We then discuss our data, focusing specifically on the V-Dem index of Political Corruption (section 3). After examining empirically the 'average' effect of corruption on economic growth (section 4), we explore whether these effects are heterogeneous across different institutional environments (section 5). In section 6, we investigate the mechanisms and present additional results. Section 7 concludes.

2. Theoretical Framework: A Review of the Arguments

Theoretically, the impact of corruption on economic performance has long divided scholars. Most economists maintain that corruption has a 'sanding' effect on economic activity. An alternative (and older) view is that corruption may, under certain circumstances, 'grease the wheels' of the state apparatus and facilitate economic growth.

2.1 The 'Sand the Wheels' Thesis:

Several arguments imply a 'sanding' effect. As an additional 'tax' on earnings, bribes reduce the private marginal product of capital, discouraging investment (Mauro, 1995). Compared to taxation, corruption may even be more 'taxing', as the proceeds of corruption are typically squandered on luxury consumption, siphoned off to offshore accounts, or laundered into illicit activities, rather than used to provide public goods. Furthermore, the fact that corruption is illegal creates transaction costs. Corrupt officials have to exert effort to find corruptible partners while avoiding detection and punishment (Shleifer and Vishny, 1993). Worse still, the terms of a corrupt agreement cannot be upheld in the courts. Thus, bribe-paying firms must sometimes resort to costly non-legal means (e.g. hiring hoodlums) to enforce corrupt deals (Bardhan, 1997: 1324).

Holding constant the rate of investment, corruption may also harm economic performance by preventing an efficient allocation of resources. First, politicians and firms may bias the composition of investment towards lower-productivity sectors (e.g. defence, non-tradables) in which it is easier to collect bribes undetected (Mauro, 1998; Henisz, 2000).[2] Second, corruption may distort the outcome of competitive auctions, conferring an advantage to high-cost firms with the willingness to compromise on quality (Rose-Ackerman, 1997).[3] Third, corruption is subject to increasing returns, partly because the probability of detection and punishment decreases in the number of corrupt individuals (Murphy et al., 1993). Thus, an increase in corruption provides an incentive for entrepreneurs to become rent-seekers (Murphy et al., 1991).[4] The resulting misallocation of talent starves the productive sector of productivity-enhancing human capital.[5]

In the empirical analysis, we attempt to distinguish these two broad mechanisms through which corruption may affect economic growth – the rate of capital investment and the productivity of installed capital.

2.2 The 'Grease the Wheels' Thesis:

The view that corruption is bad for growth is not without its challengers. The proponents of the so-called 'grease the wheels' hypothesis argue that corruption can provide a second-best solution to

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[2] When corruption is high, businesses may also choose to go 'underground' to evade bribe payments, increasing the size of the unofficial economy relative to GDP (Johnson et al., 1998).

[3] In the model analysed by Lien (1986), bribery reproduces the welfare outcome of a competitive bidding procedure. This result, however, rests on a set of strong assumptions (Bardhan, 1997; 1322).

[4] Acemoglu and Verdier (1998) present a model in which corruption *control* induces a misallocation of talent.

[5] Corruption may also discourage productivity-enhancing 'innovation by outsiders if expanding the ranks of the elite can expose existing corruption practices' (Shleifer and Vishny, 1993: 616).

coordination problems, correcting or alleviating pre-existing institutional weaknesses. There are at least *three* versions of this argument, which we distinguish in the empirical analysis. Samuel Huntington (1968) famously noted that bribes are akin to piece-rate payments. As such, they may help firms and households speed up the decisions of a sclerotic administration. In Lui's (1985) 'queuing' model, bureaucrats can price-discriminate between firms with different opportunity costs of time. Since those willing to pay the highest bribe are served first, 'speed money' minimises the average time cost of the queue. Second, corruption may function as a 'hedge' against 'bad' policies, helping firms pay their way around market-unfriendly regulations such as entry restrictions and trade barriers (Leff, 1964; Leys, 1965). Third, in countries with a weak rule of law, the possibility of making (corrupt) side-payments and 'political contributions' may improve bargaining outcomes between politicians and firms, enhancing the protection of property rights in a second-best world (Shleifer and Vishny, 1994). In the absence of bribery, firms would be exposed to a much greater risk of violence and expropriation by powerful politicians and well-connected competitors (Khan and Jomo, 2000; North et al., 2012).

If these (micro-level) mechanisms were operative, corruption would reduce macro-economic performance in countries with strong and well-functioning institutions, but would have no effect on (and may even promote) economic growth in countries with a weak and dysfunctional governance. In the former case, corruption only generates a social cost as highlighted by the advocates of the 'sand the wheels' thesis. In the latter, countervailing social benefits arise because firms can resort to bribery to mitigate or eliminate the negative effects of inefficient public administrations, anti-business regulations and property rights insecurity. Here, corruption functions as a 'substitute' for good institutions.

To fix ideas, denote the social cost of corruption as γ. The net effect of corruption on aggregate output may be written as $-\gamma + \alpha W$, where α is a positive parameter and W is an index of institutional weakness. When $W = 0$, corruption reduces growth by γ. When W is high, the benefits of corruption (αW) may offset (or even outweigh) its costs, leading to a zero (positive) net effect. The prediction that $\alpha W > \gamma$ (for a high W) is sometimes referred to as the 'strong form' of the 'grease the wheels' hypothesis, whereas the prediction that, at best, $\alpha W = \gamma$ (for a high W) as the 'weak form' (Méon and Weill, 2010).

By contrast, the proponents of the 'sand the wheels' hypothesis maintain that corruption can never mitigate the adverse consequences of institutional failures – that is, $\alpha = 0$. In fact, it may even *reinforce* them ($\alpha < 0$) (see, for instance, Ades and Di Tella, 1997). Furthermore, it is well known that bureaucrats and politicians have an interest in erecting institutional barriers *precisely* because they allow them to solicit bribes (Myrdal, 1968). To fix ideas, let the expected growth rate be a function of an index of corruption C, an index of institutional dysfunctions W, and their interaction: $g - \gamma C +$

$\alpha(C \cdot W) - \beta W$ (with $g > 0$ and $\beta > 0$). If bad institutions are endogenous to corruption ($W = \omega C$, with $\omega > 0$), the growth rate can also be written, in expectation, as:

$$g - (\gamma + \beta\omega)C + \alpha\omega C^2 \qquad (1)$$

and the total effect of corruption as $-(\gamma + \beta\omega) + 2\alpha\omega C$. For some values of the parameters[6], this effect is negative (but declining in C) even though the effect of corruption *conditional on* institutional weaknesses ($= -\gamma + \alpha W$) is positive.

2.3 The Industrial Organisation of Corruption:

Economists have also examined the 'industrial organisation of corruption' (Shleifer and Vishny, 1993; Celentani and Ganuza, 2002; Waller et al., 2007). Suppose that firms need to obtain *two* complementary licenses to enter an economic activity. The officials in charge of licensing enjoy a monopoly position, which they use to create scarcity rents and collect bribes. In doing so, they can either act as independent monopolists (decentralised corruption) or collude to form an organised syndicate (centralised corruption). The centralised case has been shown to be less distortionary.

To illustrate[7], let $Q(p)$ denote the total demand for licenses, and $p = p_1 + p_2$ the total bribe paid by firms to obtain license 1 and license 2. In the decentralised case, each monopolist chooses the bribe level p_i that maximises their own revenues $p_i Q(p)/2$, with $i \in \{1,2\}$, taking the other monopolist's pricing decision as given. In the centralised case, by contrast, the monopolists act as a cartel: they choose the total bribe p that maximises their *combined* revenues $pQ(p)$ and split the proceeds. In practice, the bribes may be paid to high-level politicians, who then share the spoils with their underlings, leading to a prevalence of 'grand corruption' in the centralised case. In the decentralised case, by contrast, 'petty corruption' is rife as even street-level bureaucrats enjoy monopoly powers.

For simplicity, let $Q(p)$ have the linear form $-ap + c$ (with $a > 0$, $c > 0$). When the monopolists act independently, the quantity of licenses issued in equilibrium is $c/3$. Collusion under a centralised authority, however, increases equilibrium output to $c/2$. Thus, the supply of licenses – and, hence, economic activity – is closer to the level ($= c$) that obtains in the absence of corruption ($p = 0$). Yet, corruption is 'high' in *both* cases.[8] In this example, firms must obtain only *two* licenses to enter the market. An increase in the number of complementary government inputs (i.e. the number

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[6] That is, for $\gamma < \omega/\alpha C$ and $\alpha < \beta/C$.
[7] This is a simplified version of the argument as presented by Aidt (2003: 644).
[8] The per-unit bribe price is higher in the decentralised ($c/3a$) than in the centralised case ($c/4a$). Yet, the total bribe transacted is lower in the decentralised ($c^2/9a$) than in the centralised case ($c^2/8a$). The data do not allow us to distinguish between bribe prices and total revenues from corruption.

of bureaucrats that must be bribed) drives the sale of licenses, and thus economic activity, further below $c/3$.

The models examined by Ehrlich and Lui (1999) and Blackburn and Forgues-Puccio (2009) place this argument in an explicitly macro-economic framework, showing that an economy in which bureaucrats coordinate their rent-seeking activities grows faster than an economy affected by disorganised corruption. Khan and Jomo (2000) and Kelsall (2013) highlight additional benefits of centralised rent management. A low dispersion of bribes makes it easier for the state to discipline the recipients of industrial policy rents, providing and withholding support based on economic rather than political calculations. The centralisation of corruption also makes bribe payments relatively more 'predictable', in the sense that firms can be sure to acquire full property rights to the license after paying a bribe to the joint monopoly (Campos et al., 1999). These arguments have been used to explain the paradox of high corruption and fast growth in the East Asian newly-industrialised countries (Wedeman, 1997; Rock and Bonnett, 2004; Huang, 2016) and, more recently, in China (Sun, 1999).

Ehrlich and Lui (1999) link the industrial organisation of corruption to political regime type. An '"organised corruption" scheme that simulates an efficient monopoly solution' may be achieved when 'bureaucrats are endowed […] with a tightly controlling political organisation – an autocratic regime – in which a powerful, but rational, leadership is capable of imposing its will on its members' (1999: 282). Communist Russia is an oft-cited example of an autocracy with centralised corruption (Shleifer and Vishny, 1993) In democracies, by contrast, the executive has to confront multiple centres of organised political power, and cannot easily prevent lower-level bureaucrats from running their own independent corruption rackets (Khan and Jomo, 2000). Post-communist Russia is a case in point.

On this argument, the quality of democratic institutions affects the degree of bribe centralisation. Thus, holding everything else constant, democracy amplifies the negative effects of corruption on growth. Aidt et al. (2008) arrive at a similar prediction. In the model they examine, there are multiple growth regimes characterised by political institutions. The corruption-growth relationship is regime-specific. When political accountability is high (as in democracies), growth reduces and is reduced by corruption. In the low-accountability regime (autocracy), corruption is at its maximum and its relationship with economic growth breaks down.

3. Data

3.1 The V-Dem dataset

Varieties of Democracy (V-Dem) is an international data-collection project based at the University of Gothenburg, Sweden (Coppedge et al., 2020a). The management team consists of about 50 social scientists from various academic institutions around the world, supported by over 3200 country experts and an international advisory board. Covering various aspects of democratic quality, the V-Dem indicators are based on an electronic expert survey filled in by V-Dem's country experts. The dataset provides country-year observations covering most polities around the world since year 1900, and starting with V-Dem's Version 7 (v.7), since 1789. The historical portion of the dataset also provides information on a number of (mostly European) state entities that are now defunct (e.g. the German Democratic Republic, the Grand Duchy of Tuscany). We employ the v.10 version of the V-Dem dataset, published in April 2020.

Each country-year observation relies on independent information provided by at least five coders, who can be either domestic experts coming from or residing in the country they code, or international experts with substantial in-country experience. Country experts are recruited by V-Dem based on criteria of expertise, focusing specifically on the subject area they are being asked to provide answers for, and impartiality. 'This expertise is usually signified by an advanced degree in the social sciences, law, or history; a record of publications; or positions in outside political society that establish their expertise in the chosen area' (Coppedge et al., 2018: 20). The 'historical' (pre-1900) segment of the V-Dem time series is typically coded by 1-2 independent experts per country-year, in light of the smaller pool of available experts on 'niche' historical subjects, e.g. 19th century Bavarian political history (Knutsen et al., 2019).

The expert coders provide answers to each survey question by choosing from a set of response categories, each of which is accompanied by a detailed rubric. A Bayesian ordinal Item Response Theory (IRT) model is then used to aggregate the ordinal ratings and estimate a latent trait variable, taking coder characteristics, biases[9], and cross-coder reliability into account (Pemstein et al. 2020). The measurement model is designed to improve cross-country and inter-temporal comparability, leading to continuous indicators with a standard normal distribution.

[Table 1]

3.2 The corruption indicator

We focus on V-Dem's political corruption index (*v2x_corr*), which measures the extent to which political corruption is 'pervasive', 'tap[ping] into several [distinct] types of corruption: both "petty" and "grand"; both bribery and theft; both corruption aimed [at] influencing law-making and that affecting implementation' (Coppedge et al., 2020a: 279). The index is arrived at by taking an unweighted average of four separate variables measuring corruption in the bureaucracy, the

[9] The main source of bias is 'differential item functioning', that is, different experts having different thresholds for how to map perceptions onto the ordinal answer categories provided by the survey.

executive branch, the legislature[10] and the judiciary (see Table 1).[11] To ensure comparability, all these measures were rescaled to run from a theoretical minimum value of 0 (signifying no corruption) to a theoretical maximum of 10 (signifying most corruption).[12]

Extending all the way back to 1900, or even 1789 for some polities, the V-Dem index has by far the best coverage along the time dimension of any extant indicator of corruption, including Transparency International's Corruption Perception Index (CPI) and the World Bank's Control of Corruption index (CC). A potential downside is that it is constructed retrospectively by country experts, rather than coded year by year, and may therefore contain a larger measurement error component than either CPI or WB, particularly when it comes to 'historical' data points. Furthermore, *v2x_corr* is entirely expert-coded, while CPI and WB are 'polls of polls' aggregating indicators compiled by analysts, business leaders and citizens, as well as experts. The biases inherent in these categories of respondents may thus cancel each other out, generating a more reliable indicator than a purely expert-coded measure.

On the upside, country experts (many of them, academics) may be more reliable estimators of corruption than either business leaders or citizens, who may themselves have been involved in episodes of corruption. Furthermore, *v2x_corr* is constructed hierarchically by aggregating lower-level indicators based on very precise questions about corruption in four different sectors of the state (see Table 1). All else equal, this procedure is likely to generate a more reliable measure than one obtained from simple questions about corruption in general. Ultimately, however, there may be an inevitable trade-off between coverage (V-Dem's undisputed strength) and reliability. Thus, analysing this indicator provides a valuable complement to existing studies based on CPI and CC.

[Figure 1]

[10] In countries with no legislature, *v2x_corr* is arrived at by taking the average of bureaucratic, executive and judicial corruption only.

[11] In turn, the *v2x_pupcorr* and *v2x_execorr* indices are obtained by taking an unweighted average of two indicators measuring corrupt exchange and theft/embezzlement in the public sector (*v2excrptp* and *v2exthftps*) and the executive (*v2exbribe* and *v2exembez*), respectively. For instance, for executive theft/embezzlement, the country coders were asked to answer the question "How often do members of the executive (the head of state, the head of government, and cabinet ministers), or their agents, steal, embezzle, or misappropriate public funds or other state resources for personal or family use?", and to provide an answer on an ordinal response scale ranging from "0: Constantly. Members of the executive act as though all public resources were their personal or family property" to "4: Never, or hardly ever. Members of the executive are almost always responsible stewards of public resources and keep them separate from personal or family property".

[12] Since the continuous indicators produced by the IRT model resemble *z*-scores, the rescaled variables were defined as $\Phi(z)$, where Φ represents the standard-normal cumulative distribution function.

Reassuringly, *v2x_corr* is very highly correlated with both CPI and CC (see Figure 1 for illustration). Across 3,496 country-year observations during 1996-2018, the correlation coefficient with the World Bank's CC is 0.9. In a sample including 1,371 observations during 2012-2019[13], the correlation coefficient between *v2x_corr* and CPI is also very high, at 0.89.[14] In addition, *v2x_corr* is highly correlated (0.81, N = 5,254, 1984-2017) with Standaert's (2015) Bayesian Corruption Indicator, which exploits the strong time-dependence of its component measures to filter out some of their random 'jumps', leading to an indicator that is more stable over time than either CPI or CC.[15]

More formal data-validation exercises give us additional confidence in the validity of the V-Dem measures. Focusing on *v2x_corr* (amongst others), Coppedge et al. (2020b, ch. 6) show that most coder traits (e.g. age, employment type, etc.) and ideological inclinations (e.g. self-reported degree of support for free markets) do not systematically predict the deviation between *v2x_corr* and either CPI or CC.[16] Moreover, intercoder disagreement is fairly low and randomly distributed, in the sense that coder traits and preferences do not predict their ratings (McMann et al., 2016).[17]

For illustration, Figure 2 displays the evolution of *v2x_corr* in four countries. Although both corruption and corruption perceptions are inherently subject to inertia, the V-Dem measure displays substantial, albeit slow-moving, variation over time. In the full sample, which contains information for 199 countries over 129 years on average, the within standard deviation is 1.38, which implies that there is more than *half* as much (57 percent) variation over time in the V-Dem indicator as there is between countries (2.41). By contrast, the within standard deviation of CC (0.39) is equal to less than *one-fifth* of its between standard deviation (1.99). Thus, the V-Dem index provides substantially more over-time variability than CC.[18] Compared to the CC index, *v2x_corr* is also relatively more stable in the short run. Controlling for country fixed-effects, the first partial autocorrelation coefficient (in a regression with four lags covering the period 2006-2018) is 0.905 for *v2x_corr* and 0.747 for the CC index. A visual comparison confirms the lower 'jumpiness' of the V-Dem measure (Figure 3).

[13] This is the period during which the CPI measures are comparable over time.

[14] The correlation, of course, is still less than 1. For instance, TI and CC rate Qatar (Trinidad and Tobago) as much less (more) corrupt than the V-Dem indicator.

[15] As with CC and CPI, a lot of the difference between the Bayesian indicator and *v2x_corr* has to do with the rating of corruption in oil-rich economies such as Qatar, Bahrein and Azerbaijan.

[16] The most robust predictor of the absolute residual between *v2x_corr* and either CPI and CC is the degree of disagreement between coders, 'a finding more indicative of stochastic error than systematic bias (Coppedge et al., 2020b: 154).

[17] Intercoder disagreement, however, does vary systematically with the difficulty of the coding task, with a greater rating dispersion observed when the availability of information on corruption is lower, as is likely to be the case in countries with less freedom of expression and for observations pertaining to earlier years. The relationship between year and coder disagreement is not statistically significant, however (Coppedge et al., 2020b: 163).

[18] The ability of country-coders to track variation in corruption levels in the distant past is lower than for more recent years. This might explain why movements in *v2x_corr* become somewhat more sluggish as we move further back in time: prior to year 1900, the within-country standard deviation is 0.89, as against 1.45 after this date.

Unsurprisingly, given these differences, the *within* correlation between the V-Dem and the CC index is only 0.34 (N = 3,496).

[Figures 2 and 3]

3.3 Institutional variables

We also employ four indicators of institutional quality from V-Dem, which are described in detail in Table 2. First, we follow Cornell et al. (2020) in constructing an indicator of 'Weberian' bureaucratic capacity. This is done by taking an unweighted average between a measure of meritocratic recruitment and promotion in the state administration, and an index of impartial bureaucracy. The resulting indicator (*statecap*) correlates very highly with the World Bank's index of Government Effectiveness (0.83, N = 3,423), a widely accepted measure of state capacity (Kaufmann et al., 2010). It is also highly correlated with *v2x_corr* (-0.69).

We also use a measure of the degree of state ownership of the economy[19], which we interpret as a proxy for the incidence of market-unfriendly policies and/or excessive market regulation. The assumption is that governments that promote or maintain state ownership of productive assets are also more likely to implement interventionist economic policies, imposing a high regulatory burden on firms, than governments that preside over privatisation programmes and/or refrain from nationalising private assets. Indeed, the state ownership indicator correlates positively (0.51) with the World Bank's measure of Regulatory Quality (Kaufmann et al., 2010). At -0.20, the correlation with *v2x_corr* is decidedly lower.

[Table 2]

Next, we consider a measure of property rights protection that gauges the extent to which property rights security is universally enjoyed by a country's resident population.[20] This measure correlates moderately highly with the World Bank's Rule of Law's indicator (0.61), which captures the quality of contract enforcement and the likelihood of crime and violence in addition to property rights protection (Kaufmann et al., 2010). At – 0.31, the correlation with v2x_corr is somewhat lower.

[19] The possible answer categories to this question range between 0 and 4 and include, for example: "0: Virtually all valuable capital belongs to the state or is directly controlled by the state. Private property may be officially prohibited"; "2: Many sectors of the economy either belong to the state or are directly controlled by the state, but others remain relatively free of direct state control."

[20] The possible answer categories to this question range between 0 and 4 and include, for example: "0: Virtually no citizen enjoy private property rights of any kind."; "3: More than half of men enjoy most private property rights, yet a smaller share of men have much more restricted rights.".

[Table 3]

Lastly, to capture the quality of democracy, we use the main V-Dem democracy index (or 'polyarchy'). This indicator has been widely issued in the political science literature and is described in detail in Knutsen et al. (2019). The correlation between *v2x_corr* and the index of democracy (-0.39) is not very high.

The variables presented in Table 2 are intended to measure conceptually distinct, although inevitably inter-related, dimensions of institutional quality. As such, with the exception of democracy and property rights protection, all these measures are moderately, but far from highly correlated with each other (Table 3).

3.4 Other variables

Our dependent variable is the percentage growth rate of GDP per capita. The income data comes from Fariss et al. (2017), who employ a dynamic latent-trait model on historical GDP and population sources to produce estimates that are less error-prone than other existing data sources (e.g. the Maddison Project data). Furthermore, the Fariss et al. (2017) model imputes missing values, mitigating potential sample selection bias arising from poor, highly corrupt countries being more prone to missingness. We use their time series benchmarked to the PPP-adjusted, Maddison times series, which are expressed in constant 2011 US$. The data cover the period 1789-2014.

In the analysis, we also consider a number of control variables that are standard in the growth literature. In selecting appropriate controls, the main constraint is data availability, since we need variables that are measured with sufficient frequency over two centuries. We use a measure of population growth from Fariss et al. (2017), and a measure of life expectancy from V-Dem (Coppedge et al., 2020a: 342), which we interpret as a proxy for human capital.[21] To proxy for political instability, we use V-Dem information to construct a count of all the general elections held in a given time period. Lastly, we define an indicator that takes the value one if a country experiences (at least) one civil or inter-state conflict in a given time period. The underlying data was compiled by Brecke (2001) and sourced from the Clio-Infra database.

[21] The life expectancy variable is compiled by V-Dem from various sources to ensure the broadest possible coverage.

4. Corruption and growth: Estimating the 'average' effects of corruption

4.1 Empirical specification

We first examine the 'average' impact of corruption on the growth of aggregate output, disregarding the potential effect-modifying influence of background institutions. To smooth out the influence of business cycle fluctuations, we divide the dataset into 5-year intervals – a standard approach in the empirical growth literature. Following Islam (1995), we specify a growth equation in panel-data form:

$$\Delta \ln y_{it} = \sum_{n=1}^{2} \rho_n \ln y_{it-1}^n + kCorr_{it-1} + \varphi X_{it-h} + \sigma_i + \tau_t + \varepsilon_{it} \quad (2)$$

$\Delta \ln y_{it}$ is the 5-year growth rate of GDP per capita, annualised using a geometric mean formula.[22] i indexes countries and t indexes 5-year periods. $Corr_{it-1}$ denotes the value of corruption, as measured by *v2x_corr*, in the first year of period t.[23] By conditioning the estimates on the start-of-period level of GDP per capita ($\ln y_{it-1}^n$), equation (2) allows countries to be out of their steady state and therefore experience convergence dynamics. In our preferred specification, the economy's out-of-steady-state behaviour is allowed to have a non-linear form (Fiaschi and Lavezzi, 2007). In particular, we enter $\ln y_{it-1}$ as a polynomial function of degree $n = 2$. An additional rationale for a quadratic specification is that the relationship between corruption and economic development is non-linear: corruption first increases weakly and then declines strongly as a country develops economically, producing an inverted-U or J-shaped relationship with per-capita income (Saha and Gounder, 2013; Paldam, 2020). If omitted, the non-linear component of $\ln y_{it-1}^n$ (which is correlated with $Corr_{it-1}$) would enter as part of the error term, potentially biasing the least-squares estimate of k.

σ_i denotes a full set of country fixed effects, which absorb the impact of any time-invariant or slow-moving determinants of the steady state (e.g. culture, history and geography). Since many of these factors may also be drivers of corruption, omitting σ_i might lead to biased estimates of k. τ_t are time-period dummies that control for technological progress at the frontier, as well as cyclical effects in the global economy.

Changes in corruption could be driven by time-varying factors related to future economic performance. Thus, we condition the estimate of k on a vector X_{it-h}, $h = \{0,1\}$ of time-varying

[22] That is $100 \cdot [(y_{it}/y_{it-1})^{1/5} - 1]$. Similar results are obtained if we use $100 \cdot (\ln y_{it} - \ln y_{it-1})/5$ instead.
[23] We also tried alternative specifications with further lags of corruption. Yet, their estimated coefficients (not reported) are always individually and jointly insignificant and sum up to 0, while the estimated coefficient on $Corr_{it-1}$ does not change substantially. These findings lend support to our choice of lag structure.

observables. These include the average annual rate of population growth (P_t)[24], the total number of elections held in each 5-year period, life expectancy at the beginning of each 5-year period, and an indicator for the incidence of at least one violent conflict in the 5-year period.[25]

Even after controlling for $\ln y_{it-1}^{n}$, σ_i and X_{it-h}, the estimated coefficient on $Corr$ cannot be interpreted as the 'true' causal impact of corruption on economic performance if contemporaneous and past shocks to GDP growth have an impact on the incidence of corruption. To address this potential concern, we consider an alternative specification that models GDP dynamics explicitly by including lags of the dependent variable ($\Delta \ln y_{it-h}$, $h = \{1,2,3\}$) on the right-hand side of equation (2), as proposed by Acemoglu et al. (2019). This specification isolates the relationship between corruption and subsequent economic growth, treating countries as if they had experienced the same growth trajectory in the past. As such, it corrects for possible violations of the classic 'parallel trends' assumption.

Since equation (2) is a dynamic specification, the within estimates of the parameters are subject to 'dynamic panel bias' (Nickell, 1981). As is well known, however, this bias disappears asymptotically for $T \to \infty$. In our dataset, T is very large – on average, each country is observed 20.5 times over 102.5 years.[26] Thus, we use OLS as a natural starting point. In additional tests, we check the robustness of our results to correcting for Nickell bias.

We base inference on standard errors that are heteroskedasticity consistent and robust to very general violations of the assumption of independently distributed residuals. Contemporaneous spatial dependencies across panels are likely to be substantial, reflecting unobserved growth spill-overs across national borders. Being local, these cross-country correlations are unlikely to hold for every pair of cross-sectional units, so that they are not absorbed by the time-period dummies. If unaddressed, they might introduce bias in the standard errors and invalidate inference. Here, we use the nonparametric covariance matrix estimator proposed by Driscoll and Kraay (1998) and updated by Hoechle (2007). This estimator produces standard errors that are heteroskedasticity consistent and robust to very general forms of cross-sectional (spatial) and serial correlation in the residuals. As an alternative to OLS, we also use an FGLS (Prais-Winsten) estimator with panel-corrected standard errors (PCSE). PCSE produces consistent standard error estimates even if the disturbances are cross-sectionally dependent, heteroskedastic and AR(1) serially correlated. Compared to the Driscoll-Kraay

[24] As is standard in the growth literature we add 0.05 to the rate of population growth, expressed as a fraction, and take logs (see Islam, 1995: 1135).

[25] We regard elections as a potentially serious confounding factor. Based on panel-data evidence from 94 democracies, for instance, Potrafke (2019) documents an increase in corruption (perceptions) during election years, while Vadlammanti (2015) finds evidence consistent with incumbent politicians exerting greater effort to control corruption prior to elections.

[26] Furthermore, Monte Carlo simulations have found that this type of bias primarily affects the estimate of the coefficient on the lagged dependent variable (Acemoglu et al., 2019: 65).

estimator, FGLS-PCSE has the advantage of allowing the autocorrelation structure to be panel-specific (Beck and Katz, 1995).

In our least-squares estimators, inference is valid under the assumption that both GDP growth *and* corruption are stationary processes, conditional on the covariates. To test this assumption, we perform augmented Dickey-Fuller tests on each country's time series (with 5-year intervals), combining the results into an overall test statistic using Fisher-type meta-analysis (Choi, 2001).[27] The inverse logit L* statistic comfortably rejects the null that the panels' time series are nonstationary for both GDP growth and corruption at the 1 percent level. Similar results are obtained by testing for a unit-root process around a linear trend.

3.2 Main results

Table 4 presents our main results. For comparison, column 1 reports the parameter estimates of a pooled cross-section model that omits σ_i, while model 2 (and all the subsequent specifications) include σ_i. In both models, the estimated effect of corruption is negative and highly significant. Controlling for country-level unobservables (model 2), however, reduces the absolute magnitude of this effect by 21 percent, suggesting that previous estimates based on cross-sectional regressions may be subject to a meaningful upward bias.[28] Allowing for a richer (quadratic) specification of the convergence dynamics (full results not reported) leaves the estimated effect of corruption essentially unchanged.[29]

[Table 4]

Model 4 controls for population growth, life expectancy, political instability, and violent conflict. The estimated impact of corruption is only slightly smaller as compared to a benchmark model estimated on the same restricted sample (model 3). The control variables enter with the expected sign, although only the coefficient on the war dummy is statistically significant. A possibly serious concern is that $\hat{k}$ may be spuriously picking up the influence of other institutional features, which may affect or be affected by the prevalence of corruption. To properly distinguish the effects of corruption

[27] To prevent the aggregate test statistics to be distorted by cross-sectional correlations, we remove the cross-sectional means from each time series. Since each country's mean level of corruption is always non-zero by construction, we always also include a drift term.
[28] In line with standard results (Islam, 1995), flexibly controlling for country-level differences in the steady state by including fixed effects leads to an estimate of faster convergence, as shown by a 2.3-fold increase in the absolute magnitude of $\hat{\rho}$ (from -1.53 to -3.56).
[29] The coefficients on the income terms, however, are jointly significant (at 1 percent) but individually insignificant. Adding a cubic term in $\ln GDPpc_{it-1}$ yields an estimate of k (=-0.163) that is only slightly lower but still highly significant (full results not reported).

from those of other institutional failures, model 5 conditions the estimate of k on our four indices of institutional quality from V-Dem (Table 2). The estimate of k is negative and significant, and almost 40 percent *larger* in magnitude than in model 3.[30]

These estimates may still be subject to bias if ‘corrupt’ and ‘non-corrupt’ countries happen to be on systematically different growth trajectories (see Acemoglu et a., 2019). For instance, spikes in corruption may occur endogenously as a result of growth spurts or economic crises. To address this potential threat to causal inference, model 7 accounts for the dynamic adjustment of GDP by including three lags of GDP per capita growth, which enter as jointly significant. The estimated effect of corruption is reduced by about 20 percent, relative to our benchmark. Yet, it remains statistically significant.[31]

Lastly, we consider a ‘kitchen-sink’ specification that includes all the control variables and the three lags of GDP per capita growth simultaneously. This specification is estimated by OLS (column 7) and FGLS (column 8). In both cases, our main results remain qualitatively unaltered. In the OLS model, inference about the effects of corruption remains unchanged if we use panel-corrected (PCSE) standard errors (0.065), or even conventional robust errors clustered at the country level (0.092), instead of Driscoll-Kraay standard errors (0.063).[32]

Does corruption ‘sand’ or ‘grease’ the wheels of economic growth? These findings indicate that corruption is, on average, an unambiguous ‘sander’. The magnitude of this ‘sanding’ effect is meaningful economically. Consider a standard-deviation (2.89) improvement in the V-Dem corruption index, which corresponds approximately to the difference between Turkey (5.73) and Italy (2.86) in 2018, or the difference between the year 1800 (3.45) and 1860 (0.65) in Sweden. The most conservative point estimate (Table 4, column 7) implies that an improvement of this magnitude would increase a country’s annual growth rate by half a percentage point, which corresponds to half the average annual rate of growth recorded during 1810-2000 (1.1), or 18 percent of a standard deviation (3.1). This effect is smaller than that reported in the cross-country literature[33], but is still substantial. If sustained, it would lead to a 64 percent difference in the level of GDP per capita after a century, and a 170 percent difference after two centuries.

[30] The coefficient on state capacity is negative and significant, a counterintuitive result. This finding, however, is not robust (full results not shown).

[31] Similar results (not reported) are obtained by dropping the second and third lags of GDP growth, which enter as insignificant (both individually and jointly). We also consider the following version of equation (2): $\ln y_{it} = (1+\rho_1)y_{it-1} + \rho_2 y_{it-1}^2 + kCorr_{it-1} + \varphi X_{it-k} + \sigma_i + \tau_t + \varepsilon_{it}$. Adding y_{it-2} and y_{it-3} to this regression equation leads to a specification of the GDP dynamics that is closer to the one examined by Acemoglu et al. (2019). In this model, too, the estimate of k is consistent with the results reported in Table 4, column 6 (full results not reported).

[32] In OLS, PCSE standard errors are robust to heteroskedasticity and cross-sectional correlation across panels (but not to serial correlation within panels), while clustered standard errors are robust to heteroskedasticity and serial correlation within panels (but not to cross-sectional correlation).

[33] Mauro (1995: 701), for instance, reports a standardised effect of 0.8 percentage points.

3.3 Robustness analysis

In Table 5, we evaluate the robustness of these findings in various ways. First, we investigate the potential endogeneity of corruption. Modelling the GDP dynamics explicitly (Table 4, columns 6-8) may not completely purge the influence of unobserved shocks to GDP. Besides, corruption is measured with error, so that $\hat{k}$ may be subject to attenuation.

[Table 5]

Motivated by these considerations, the model reported in column 2 (Table 5) uses two (time-varying) instruments to isolate a plausibly exogenous component of variation in $Corr_{it-1}$, conditional on country fixed effects. The first instrument, suggested by Gründler and Potrafke (2019), is the jack-knifed average level of corruption in a country's geopolitical region.[34] Corruption tends to be spatially correlated within regions because of trade relations, political exchange, and the transmission of cultural attitudes through migration (Gründler and Potrafke, 2019). Accordingly, this instrument is quite highly correlated with $Corr_{it-1}$ (0.61, N = 4,722). A potential threat to instrument validity may arise if regional economic performance is affected negatively by regional corruption, while having a positive impact on a country's growth path. For this reason, we condition the second-stage estimates on the (jack-knifed) regional average of GDP per capita growth. As an additional instrument, we also use the second lag of corruption ($Corr_{it-2}$), which is very highly correlated with $Corr_{it-1}$ (0.97).

The instruments are highly relevant (as implied by Kleibergen-Paap's F-statistic being much higher than the Stock-Yogo critical value). In addition, Hansen's J-test cannot reject the validity of the instruments' exclusion restrictions. Conditional on country and time-period fixed effects and our full set of controls, the 2SLS estimate of k is negative, statistically significant (based on clustered standard errors), and only about 7 percent smaller than the corresponding OLS estimate, reported as a benchmark in column 1.[35]

A possible concern is that the corruption index may be correlated with time-varying unobservables that are responsible for generating country-specific time trends in growth performance. For instance, the rate of technological progress may not be constant across countries

[34] This average is jack-knifed in the sense that it excludes the own-country level of corruption from the regional mean. We use six broad geo-political regions, which are Eastern Europe and Central Asia, Latin America and the Caribbean, Middle East and North Africa, Sub-Saharan Africa, Western Europe and North America, Asia and Pacific. The results are qualitatively robust to using finer geopolitical categorizations (results available upon request).

[35] The estimated coefficient on corruption remains negative in just-identified specifications with regional corruption (-0.464, s.e.=0.283) and twice-lagged corruption (-0.218, s.e.=0.089) entered individually as instruments.

(Islam, 1995: 1149). Alternatively, business cycles may be staggered across different economies. If these omitted trends have a bearing on the incidence of corruption, their influence on economic performance may be spuriously picked up by $\hat{k}$. While an instrumental variable approach addresses this concern, it is subject to the usual concerns regarding instrument validity. An alternative way-forward is to condition the estimates on country-level trends.

We begin y including a full set of linear trend terms ($\sigma_i t$), in addition to country (σ_i) and time-period (τ_t) fixed effects.[36] At -0.258 (column 4), the estimated effect of corruption in this (rather demanding) specification is negative, large and highly significant. Allowing the time trends to have a quadratic functional form ($\sigma_i t + \sigma_i t^2$) leads to very similar results (column 5). In column 6, we investigate the potential confounding influence of cyclical growth patterns by including periodic trends with a 30-year frequency ($\sigma_i \sin(2\pi t/6) + \sigma_i \cos(2\pi t/6)$, where t denotes a 5-year time period). The results are again unchanged.

The stability of the corruption effect is remarkable, given the fact that the trend terms are powerful predictors of GDP per capita growth. The within R-squared of the regression is 0.35 in the linear trend model, 0.50 in the specification with quadratic trends and 0.36 in the trigonometric regression, up from 0.02 in a benchmark model (column 3) without trend terms and control variables. Following Altonji et al. (2005), we calculate the following ratios of coefficients: $\hat{k}/|\hat{k}^3 - \hat{k}|$, where $\hat{k}$ is the estimate based on a model with an extensive set of controls, and $\hat{k}^3$ is the estimate based on a restricted specification (column 3). The ratios for models 4, 5 and 6 are, respectively, 7.8, 6.7, and 2.5, indicating that the confounding influence of omitted variables would need to be up to 8 times greater than that of included variables to explain away the entire effect of corruption on economic growth. This makes it highly unlikely that the estimated effect can be fully attributed to omitted variable bias.

In models 4, 5 and 6, we imposed an arbitrary functional form on the time trends. To relax these assumptions, model 7 allows for fully flexible time trends at the level of geo-political regions by replacing τ_t in equation 2 with a full set of time-period × region fixed effects.[37] Of course, this specification reflects a trade-off, since we must assume that all the countries within a given geo-political region share the same growth trend. The estimated effect of corruption is now substantially smaller in magnitude but still significant at 5 percent. The Altonji ratio is also much smaller (1.2). Yet, the continent-level trends are also somewhat less powerful than the country-level trends, as evidenced by the R-squared of the regression (0.23) being lower than that of models 4, 5 and 6.

Next, we examine whether our dynamic panel-data estimates are sensitive to Nickell bias. The specification reported in column 8 uses a 'first-difference' GMM estimator that treats $\ln y_{it-1}$ as pre-

[36] The trend terms enter as jointly significant at the 1 percent level.
[37] The region categories are: Eastern Europe and Central Asia, Latin America and the Caribbean, Middle East and North Africa (MENA), Sub-Saharan Africa, Western Europe and North America, Asia and the Pacific.

determined and instruments for its difference using lagged levels (Arellano and Bond, 1991). Standard diagnostic tests cannot reject the hypothesis that the instruments are valid. The estimated corruption effect is still statistically significant and, indeed, larger in magnitude than in a corresponding OLS model (-0.214, s.e.=0.054).[38] In column 9, we also report estimates based on the Nickell bias-corrected least-squares dummy variable (LSDVC) estimator proposed by Kiviet (1999). An advantage of LSDVC, relative to GMM, is that it allows us to preserve our preferred quadratic specification of the convergence dynamics.[39] We also include the three lags of GDP per capita growth. The findings are very similar to those obtained from a comparable OLS model (column 7, Table 4), confirming that our dynamic specification is unlikely to be an important source of bias.

The estimated negative effect of corruption is extremely robust to a host of alternative specifications accounting for potential omitted confounders and Nickell bias. This finding strengthens our confidence that the estimates reflect at least partly the true causal effects of corruption on growth.

5. ‘Greasing’ or ‘sanding’ the wheels?

Our findings so far indicate that $-\gamma + \alpha W < 0$ (using the notation of section 2.2). Yet, if $\alpha \neq 0$, the estimates may only be interpreted as the effects of corruption in an economy with an average quality of institutions. Here, we consider specifications that allow us to estimate α and hence detect heterogeneous effects across different institutional contexts. To do so, we allow the parameter k in equation (2) to depend linearly on the four V-Dem measures of institutional quality (Table 2), which we reverse for ease of interpretation so that a higher value denotes a lower quality of institutions (W). This leads to a specification with interaction terms:

$$\Delta \ln y_{it} = \sum_{n=1}^{2} \rho_n \ln y_{it-1}^{n} + \gamma Corr_{it-1} + \sum_{m=1}^{4} \alpha_m Corr_{it-1} \times W_{it-1}^{m} + \sum_{m=1}^{4} \beta_m W_{it-1}^{m} + \varphi X_{it-k} + \sigma_i + \tau_t + \varepsilon_{it} \qquad (3)$$

Since $\partial \Delta \ln y_{it} / \partial Corr_{it-1} = \gamma + \sum_{m=1}^{4} \alpha_m W_{it-1}^{m}$, $\hat{\gamma}$ measures the effects of corruption in countries with good institutions ($W^m = 0, \forall m$). We expect $\hat{\gamma}$ to be negative and significant, and to provide an estimate of the ‘pure’ costs of corruption – that is, the costs of corruption in the absence of ‘greasing effects’. When W denotes (the inverse of) bureaucratic capacity, regulatory quality or property rights

[38] Full estimates not reported for brevity.
[39] In a Monte Carlo simulation, Bruno (2005) also shows that LSDVC outperforms ‘first-difference’ GMM when N is small.

protection, a positive and significant $\hat{\alpha}$ provides evidence for the 'grease the wheels' hypothesis: $\partial\Delta \ln y_{it}/\partial Corr_{it-1}$ is less negative (potentially even positive) in environments characterised by pre-existing institutional weaknesses. An insignificant, or negative and significant, $\hat{\alpha}$ is evidence *against* the thesis that a positive countervailing effect of corruption arises when institutions are dysfunctional. When W measures the quality of democracy, a *positive* and significant $\hat{\alpha}$ is interpreted as evidence that corruption is less harmful in autocracies, in line with theories of the industrial organisation of corruption (Shleifer and Vishny, 1993; Ehrlich and Lui, 1999). An insignificant, or negative and significant, $\hat{\alpha}$ is interpreted as evidence against these theories.

Note that in equation (3), all four interaction terms ($\alpha_m Corr_{it-1} \times W_{it-1}^m$) are included simultaneously in the regression. This systematic specification allows us fully to distinguish the effect-modifying influence of our four institutional dimensions, allaying the concern that the institutional features that shape the costs of corruption may become confounded.

5.1 Results

In Table 6, we present estimates of equation (3). Column 1 reports a pooled cross-section model without country fixed effects (σ_i). The model in column 2 (and all the subsequent ones) include σ_i in the regression. In the specification reported in column 3, we add our vector of time-varying controls and the three lags of GDP per capita growth.[40] For comparison, we also report models in which the interaction terms are entered one by one (columns 1-6). The results of our preferred, kitchen-sink specification (column 3) are qualitatively unchanged if we use an FGLS estimator with panel-corrected standard errors (not reported), instead of Driscoll and Kraay's (1998) OLS.

[Table 6]

Across models 1-6, the estimate of γ is significantly negative and up to three times as large in absolute magnitude as the 'average' corruption effect ($\hat{k}$) estimated in models with no interaction terms (see Tables 4 and 5). In economies with no institutional dysfunctions, corruption is substantially more detrimental to aggregate performance than in an economy with 'average' institutions. Regarding the interaction terms, the pattern in the data is clear. State capacity, regulatory quality and property rights security do not exert a modifying influence on the estimated magnitude of the corruption effect, as shown by the coefficients on their interactions with $Corr$ being always statistically indistinguishable from zero (with the exception of model 2).

[40] In this model, inference remains unchanged if we use clustered or panel-corrected standard errors instead of Driscoll-Kraay standard errors. Full results available upon request.

In contrast, the estimate of α for the Corruption × Democracy interaction is consistently *positive* and highly significant across specifications 1-3 and 7, implying an attenuated adverse effect of corruption in more autocratic regimes. This interaction effect doubles in magnitude when country fixed effects are included in the regression (compare model 2 to model 1), suggesting that cross-country estimates may severely underestimate the extent to which political regime type determines the costs of corruption. The corruption × democracy interaction increases further in magnitude when the potential effect-modifying influences of other dimensions of institutional quality are controlled for in the regression (compare model 7 to model 3), pointing to another potential source of downward bias in previous studies.

[Table 7]

The modifying influence of democratic quality is not only statistically but also economically significant, leading to a meaningful heterogeneity of corruption effects. In Italy, an advanced democracy with an inverted polyarchy index of 1.46 (in 2010), the growth effects of corruption (-0.598, s.e.=0.169, based on model 3) are more than twice as high as in a country with an 'average' quality of democracy (-0.231, s.e.=0.071), and almost *five* times as high as in China (-0.125, s.e.=0.052), an autocracy with an inverted polyarchy index of 9.04.

Lastly, we note that the estimated coefficients (β) on the institutional variables are mostly in line with expectations. According to models 4-7, property rights insecurity is associated with lower growth (as in Acemoglu and Johnson, 2005). In columns 1-3, the β coefficient on property rights (the effect when $Corr = 0$) is statistically insignificant, but the average effect for a country with a mean level of corruption (-0.171, s.e.=0.057, based on model 3) is significant at 1 percent. The effects of state capacity (inverted) are mixed, consistent with previous findings (Knutsen, 2013; Cornell et al., 2020).

5.2 Interpretation of the results

These findings indicate that the economic consequences of corruption are shaped by the broader institutional context. Yet, the only institutional dimension that is found to exert a modifying influence on the effects of corruption on growth is democratic quality. The aspects of the institutional environment held as relevant by the advocates of the 'grease the wheels' thesis (state capacity, regulatory quality and property rights protection) are not found to shape the economic consequences of graft. Thus, there is little evidence to conclude that a positive countervailing effect of corruption may arise in institutionally dysfunctional environments, partly compensating for (let alone overriding) the costs of bribery. Yet, the long-run historical evidence is consistent with corruption being less detrimental to growth (though never growth-enhancing) in more autocratic regimes. A plausible explanation of these findings is suggested by theories of the industrial organisation of corruption (e.g. Ehrlich and Lui, 1999).

Our findings partially contradict previous empirical tests of the ‘grease the wheels’ hypothesis (Méon and Weill, 2010; Hodge et al., 2011) which attribute a significant role to state capacity and regulatory quality in moderating the corruption-growth relationship. A possible explanation is that these studies do not estimate these interaction effects simultaneously with democracy and with each other. Our findings are also at odds with the conclusions of a recent study by Gründler and Potrafke (2019), who argue based on a short panel that bad institutions may actually *reinforce* the adverse consequences of corruption ($\alpha < 0$). Based on long-run, slow-moving variation in corruption levels within countries, we find that the effects of corruption are neither mitigated nor aggravated by the institutional dysfunctions typically highlighted by the proponents of the ‘grease the wheels’ thesis (regulatory quality being a partial but definitely non-robust exception). At the same time, our finding that corruption is less growth-reducing in autocratic regimes is consistent with previous findings in the empirical literature (Aidt et al., 2008; Aidt, 2009).

[Figure 4]

In equation (3), we can view democracy as the factor that modifies the effects of corruption on economic growth. Alternatively, we can also view corruption as the effect-modifier, focusing instead on the relationship between democratic development and growth (to do this, just differentiate the regression equation with respect to the democracy index). As shown in Figure 4, democracy is significantly growth-enhancing at low levels of corruption (the marginal effects of the inverted index is negative), but becomes a burden on macroeconomic performance when corrupt practices are widespread, plausibly because it has the effect of decentralising the way corruption networks are organised, making corruption more harmful. In economies with a corruption score between 3 and 7, the effects of democracy are statistically indistinguishable from zero. These findings echo the mixed results found in the literature on democracy and growth, with some studies reporting a negative effect (Barro, 1996; Tavares and Wacziarg, 2001) and others a positive effect of democracy (Gründler and Kriegler, 2013; Acemoglu et al., 2019). But they also suggest a possible way of making sense of these heterogeneous results: democracy can either support or harm growth depending on the prevailing level of corruption.

5.3 Robustness analysis

We examine the robustness of the results reported in Table 6 in a number of ways (Table 7). First, we consider variants of equation (3) that: condition the estimates on country-specific linear, quadratic and periodic time trends (columns 1-3); replace the time-period FE with a full set of time-period x region FE (column 4); estimate a 'first-difference' GMM model that treats lagged income as pre-determined (column 5); and use Kiviet's (1999) Nickell-bias corrected estimator (column 6). Our findings are qualitatively unaltered. In all but one case (model 2)[41], the coefficients on the corruption x democracy interaction terms are positive and significant at conventional levels, although less stable across specifications than the average corruption effects estimated earlier. By contrast, the interactions of corruption with the other institutional dimensions are always statistically insignificant. These findings strengthen our confidence that the heterogeneity of effects associated with democracy is unlikely to be driven primarily by omitted effect-modifiers. Rather, it is likely to reflect at least partly the true moderating influence of political institutions on the corruption-growth nexus. This influence is also unlikely to simply be an artefact of dynamic panel bias.

[Table 7]

The estimates presented so far assume that the marginal effects of corruption on economic growth are a *linear* function of democratic quality. Next, we relax this assumption. We divide the observations into five equal-sized subgroups based on the values of each institutional indicator; we then allow the estimated effects of corruption – parameter k in equation (2) – to be different across quintiles of the distributions of institutional quality. Within each subgroup of observations, there is substantial variation in levels of corruption to permit a balanced estimation.[42] The results, presented in Figure 5, confirm our previous findings. With one single exception (Q1, state capacity), the marginal effects of corruption are significantly negative over the distribution of state capacity, regulatory quality and property rights security, with no discernible patterns. Yet, the marginal effects become substantially *less* negative in higher quintiles of the distribution of democratic quality (inverted), and statistically indistinguishable from zero in the top quintile, which contains the most autocratic regimes (e.g. China prior to 1990).

[Figure 5]

[41] Here, the OLS coefficient on the corruption × democracy interaction is only significant at the 17 percent level. Yet, basing inference on panel-corrected (PCSE), instead of Driscoll-Kraay, standard errors, makes this estimate marginally significant at the 10 percent level (full results not reported).
[42] With the exception of only two sub-groups, the standard deviation of corruption is always greater than 2 (full descriptive statistics available upon request).

An additional concern with our estimates of equation (3) is that they could be driven by influential outliers – for instance, Italy, a slow-growing advanced democracy with high corruption perceptions. Yet, removing the 131 observations with a Cook's D influence statistic greater than $4/N$ (where $N = 2615$) leaves our results qualitatively unchanged. To further allay this concern, we allow the moderating influence of democracy to be different in magnitude across geo-political regions and time periods.[43] We define six broad regional country groups (as in previous tests) and three time-period categories (1795-1900, 1900-1960 and 1960-2010). Across regions and time periods, the estimated coefficient on the corruption × democracy interaction is invariably positive and (with the exception of the Western Europe and North America region) always statistically significant.[44] These results confirm that the estimated effect-modifying influence of democracy is not driven by particular regions or time periods but holds quite uniformly in the data.[45]

[Table 8]

Lastly, we check the sensitivity of our results to the choice of panel structure. As mentioned earlier, the parameter estimates may be contaminated by the influence of unobserved, country-specific cyclical trends. Using panels with five-year intervals (or even controlling explicitly for country-level cyclical trends) may not be sufficient to remove these potential confounders from the model residuals. An alternative approach is to use longer time intervals, so that the ε_{it}'s are further apart. Table 8 shows that the estimated coefficient on the corruption × democracy interaction term remains very stable across specifications using panels with 5- (column 2) or 10- (column 3) year intervals, or even if we pool all available cross-sections over consecutive years (column 1). In the specification with 10-year intervals, however, the estimated effects are smaller in absolute magnitude, and at very low levels of democracy (e.g. in China in 2010) they become statistically insignificant (0.011, s.e.=0.085).

6. Mechanisms and Extensions

In this section, we further elucidate the relationship between corruption, governance quality and economic performance.

[43] To do so, we interact both $Corr_{it-1}$ and $Corr_{it-1} \times Inst_{it-1}$ with a categorical variable for either geo-political regions or time periods. The full results are available upon request.

[44] In Western Europe and North America, the variation in the quality of democracy across countries is substantially lower than in other geo-political regions.

[45] If the coefficient on the interaction terms loses statistical significance at conventional levels (but remains positive) in the pre-1990 sub-period, this is likely to be because the within-country variation in democratic development prior to 1990 is considerably lower than after 1990 (when most episodes of democratisation occurred).

6.1 Forms of corruption, democracy and growth:

First, we investigate the relationship between the industrial organisation of corruption, political regime type and economic growth. A possible explanation of our findings is that corruption is less detrimental to growth in autocracies because it is more centralised (as in Ehrlich and Lui, 1999). To corroborate this interpretation, we present suggestive evidence that the effect-modifying influence of democracy is indeed related to the form that corruption takes – centralised or decentralised – across different political regimes.

As discussed in section 3.2, The V-Dem index of corruption (*v2x_corr*) is an unweighted average of four expert-coded indicators measuring corruption in the bureaucracy, the executive branch of the state, the legislature and the judicial system. These components are highly (albeit not perfectly) correlated, with correlation coefficients ranging between 0.70 and 0.88, so that their relative contribution to the total corruption effect cannot be distinguished precisely. Still, in Table 9, we replace *v2x_corr* with its four components, allowing each to exert a separate effect on growth.

In column 2, we omit the index of legislative corruption, which is only available for the much smaller sub-sample of countries with a working legislature. Columns 3 and 4 report OLS and FGLS estimates, respectively; columns 5 adds three lags of GDP growth and column 6 a full set of country-specific linear trends. The estimates indicate that bureaucratic (that is, petty) corruption, but also corruption in parliaments, are detrimental to growth performance. After accounting for bureaucratic and legislative corruption, executive (that is, grand) corruption and corruption in the judiciary do not exert any additional growth-reducing effect.

Next, we report suggestive evidence that executive (grand) corruption is more prevalent in autocracies than in democracies, while the relative incidence of the other three forms of corruption is unrelated to political regime type. In Table 10, we show the results of regressing the various corruption indicators on their first lag (to account for their tendency to persist over time) and the V-Dem index of democratic quality, controlling for country and time-period fixed effects. As shown in Panel A, the only component of the V-Dem Political Corruption index that is associated with democracy, is the indicator of executive (that is, high-level) corruption (see Table 1).

In the literature, the level of development has been identified as the most important and robust predictor of the magnitude of corruption (Treisman, 2007). Thus, in Panel B, we condition the estimates on lagged (per-capita) income and lagged income squared. The association between autocracy and grand (executive) corruption is now even stronger and more precisely estimated. Our findings remain unaltered if we use a 'first-difference' GMM specification that corrects for Nickell bias by instrumenting for the lagged dependent variable in GMM style (results not reported). It is also interesting to note that the income coefficients (not reported) imply that corruption traces a J-shaped pattern in the course of economic development, confirming the findings of previous cross-sectional studies by Saha and Gounder (2013) and Paldam (2020), and justifying our non-linear specification of convergence dynamics in equation (2).

Taken together, the regressions reported in Tables 9 and 10 corroborate our interpretation of previous results. A plausible reason why corruption is less harmful in countries with a low level of democratic development is that grand (executive) corruption is typically the form that corruption takes in autocratic regimes. Yet, in contrast to petty (bureaucratic) corruption, grand (executive) corruption is unrelated to economic growth, all else equal. On these grounds, we feel confident to interpret the estimated effect-modifying influence of political institutions as evidence that corruption may be more or less harmful depending on its 'industrial' organisation (Shleifer and Vishny, 1993; Ehrlich and Lui, 1999).

6.2 Non-linear effects:

Next, we consider the possibility that corruption may impact the quality of political institutions. Both the 'grease the wheels' thesis and the industrial organisation of corruption argument rely on the classical theory of the second best, in which the constraints to achieving the first-best equilibrium are treated as exogenous.[46] Yet, the institutional dysfunctions that corruption helps overcome (e.g. bureaucratic delays), or under which corruption is less harmful (autocracy), are very often introduced and maintained precisely *because of* corruption (Myrdal, 1968). If corruption affects the quality of democracy (say, linearly), the marginal effects of corruption on growth should be negative *and* diminishing, as in equation (1). The intuition is that at higher levels of corruption, the effects of a marginal increase in graft are attenuated by political institutions becoming more autocratic.

[Figure 6]

To check this, we first add $k_2 Corr_{it-1}^2$ to equation (2), conditioning on all the controls included in our 'kitchen sink' specification (Table 4, column 7). While the OLS estimate of k is -0.411 (s.e. = 0.157), $\hat{k}_2$ is positive (0.021, s.e. =0.012) and statistically significant at 10 percent (full results not reported). Both $\hat{k}$ and $\hat{k}_2$ are larger in absolute magnitude and more precisely estimated in a corresponding FGLS model, as well as in an OLS regression with linear time trends instead of GDP lags. Qualitatively similar (but more nuanced) findings are obtained by using Baltagi and Li's (2002) partially linear fixed-effects estimator, which allows corruption to enter the equation non-parametrically. The predicted partial relationship between GDP per-capita growth and corruption is shown in Figure 6.

[46] I would like to thank an anonymous referee for noting this point.

While growth slows down rapidly as corruption first increases above 0, this negative relationship tends to level out once corruption is more widespread. Note, in addition, that the quadratic relationship is concave but monotonic, in contrast to previous findings by Swaleheen (2011).[47]

These results are consistent with the proposition that corruption harms the quality of democracy. Not only do exogenous changes in political institutions affect the slope of the curve displayed in Figure 5 (as shown in the previous section), but the slope changes endogenously at different levels of corruption, as implied by Myrdal's (1968) well-known argument. The economic significance of this change, however, is ~~relatively more~~ modest.

6.3. Heterogeneity across levels of income:

A related question is whether the estimated effects of corruption on economic growth are heterogeneous at different levels of income. The claim that economic development leads to the establishment and consolidation of democratic institutions lies at the heart of the modernisation hypothesis. Several studies find that a positive relationship between income and democracy holds in panel regressions with country fixed effects (Heid et al., 2012; Benhabib et al., 2013; Cervellati et al., 2014). Focusing on within-country variation only, Moral-Benito and Bartolucci (2012) document a positive but *diminishing* effect of per-capita GDP on democracy. These findings imply that the impact of corruption on economic growth should be smaller in magnitude in lower-income – and therefore less democratic – economies.

[Figure 7]

To check this, while allowing for a non-linear effect of income on democracy, we include the following two interaction terms in equation (2): $\sum_{n=1}^{2} \eta_n \ln y_{it-1}^{n} \times Corr_{it-1}$. The marginal effects of corruption ($= k + \sum_{n=1}^{2} \eta_n \ln y_{it-1}^{n}$) are plotted in Figure 7 as a function of $\ln y_{it-1}$. Panel A reports results based on an OLS model with a full set of controls. Panel B displays FGLS estimates of the same specification, while the models underlying panels C and D condition on GDP lags and country-specific linear trends, respectively. Consistent with expectations, the findings indicate that the adverse effects of corruption on growth are smaller or even statistically indistinguishable from zero (depending on the specification) in lower-income economies. A plausible (though by no means the only) interpretation of this finding is that low-income economies are less likely to have well-functioning democratic institutions, and hence more likely to harbour centralised corruption networks.

[47] In addition, our findings run counter to those of Mendez and Sepulveda (2006), who find that the relationship between corruption and growth is non-monotonic and convex, and conclude that corruption is growth-maximising at low levels, as in the model examined by Acemoglu and Verdier (1998).

6.4 Corruption, Investment and growth:

Lastly, we examine whether corruption affects the economy's steady-state growth indirectly, i.e. by lowering the rate of capital accumulation, or directly, by shifting the aggregate production function. While early studies found that corruption had no additional impact on economic growth after controlling for the investment share in GDP (Mauro, 1995; Mo, 2001), more recent empirical analyses show that corruption may also exert a direct effect on aggregate productivity (e.g. Méon and Weill, 2010).

To separate these two channels, we estimate a growth regression with and without the share of gross capital formation in GDP (in logs)[48], and observe how the estimates of our coefficients of interest change. In performing this analysis, the main challenge is the lack of long time-series records on the rate of savings or investment. Thus, we can only make use of a severely restricted sample covering just over 100 countries during 1975-2000. Over this time period, the within-group standard deviation of the V-Dem indicator of corruption is only about 33 percent of the corresponding between-group variation (as against 57 percent in the full sample). Consequently, the model parameters are likely to be much less precisely estimated. They may also suffer from sample selection bias. For these reasons, the following analysis should be interpreted with caution.

[Table 11]

Given the short time period, we focus on the corruption × democracy interaction. The results are shown in Table 11. Column 1 reports the parameters of a standard specification. The estimated effect-modifying influence of democracy is positive and significant, although the marginal effects of corruption are less precisely estimated. This pattern of results remains unaltered when the rate of investment, denoted as $\ln(s_t)$, is held constant (column 2), and corruption is allowed to influence the growth of output only by shifting the production function. This finding is consistent with total-factor productivity being the main channel through which corruption affects economic performance. In line with theoretical expectations, the rate of investment has a positive and significant effect on steady-state growth. The coefficient on population growth is negative, as implied by the neo-classical framework, though not statistically significant.[49]

We also replicate this analysis, but taking advantage of the greater data availability over this shorter and more recent time period to make use of an alternative (and better) set of control variables.

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[48] The data are from the World Bank.

[49] As is common in the growth literature, population growth is interpreted as a proxy for labour force expansion.

In particular, we use a measure of (start-of-period) educational attainment from Barro and Lee (2013), instead of life expectancy as a measure of human capital. This variable measures the number of years spent in education by the average worker, and is available from 1950 on a 5-year basis. To capture political stability more accurately, we use the number of coups that took place in each 5-year period (Przeworski et al., 2013), instead of the number of elections held. We also control for the rate of inflation, using data compiled by Coppedge et al. (2020a: 339) from various sources, as well as total trade (imports plus exports, over GDP) and total income from natural resources (oil, gas, coal, minerals) as a share of GDP.[50] The results, reported in columns 3 and 4, confirm our previous findings.

To further investigate whether corruption has an effect on the mobilisation of factor inputs, we regress population growth and the investment rate on their first lag (to account for persistence), corruption, and corruption × democracy, conditional on country and time-period fixed effects and the same set of observables as in columns 3 and 4. The results of these regressions, shown in columns 5 and 6, rule out a direct impact of corruption on capital accumulation and fertility behaviour, confirming that corruption affects economic growth primarily by lowering the productivity of factors inputs. The arguments that corruption biases the composition of investment, distorts competition, and leads to a misallocation of human capital (see section 2.1) are all consistent with these findings. Our data, however, does not allow us to distinguish these particular sub-mechanisms.

7. Conclusion

We used a novel indicator of political corruption with unprecedented historical coverage to examine the causal effects of corruption on economic performance. Our data and design allow us to focus on long-run, slow-moving variation in corruption levels within countries over nearly two centuries. Unlike the short-run variation used for identification in previous studies, which may reflect white noise, variation over the long run is more likely to track actual improvements or deteriorations in the incidence of corruption on the ground, making causal inference more reliable. Accounting for the potential confounding influence of time-varying observables and unobservables, we also examined how corruption interacts with the broader institutional environment in a systematic framework, uncovering heterogeneous effects across different institutional contexts.

Our findings indicate that, on average, corruption slows growth, potentially by lowering productivity. We find little evidence for a 'greasing' effect of corruption in contexts of institutional dysfunctionality: corruption is not any less detrimental (let alone beneficial) for economic growth in economies characterised by slow bureaucracies, anti-business regulations or property rights

[50] Both the trade and natural resource data are from the World Bank.

insecurity. Yet, we find that corruption is substantially less growth-reducing in more autocratic regimes. In autocracies, high-level politicians can deploy political authority to centralise the collection of bribes, leading to a lesser impact of corruption on output compared to regimes that make it easier for bureaucrats and politicians to set and collect bribes independently (Shleifer and Vishny, 1993; Ehrlich and Lui, 1999). Accordingly, we find that grand (executive) corruption is more prevalent in less democratic regimes, and is unrelated to economic growth after accounting for the prevalence of petty (bureaucratic) corruption. The negative effects of corruption on growth are also less severe when corruption is already widespread and at lower levels of economic development – two situations in which political institutions are typically less democratic. By highlighting an interaction effect between corruption and democracy, we can also make sense of previous conflicting results in the democracy literature: democracy is growth-enhancing in the absence of corruption, but significantly growth-reducing where corruption is widespread.

Our findings point towards an explanation for the apparent paradox of rapid growth with corruption in autocracies such as China (Wedeman, 1997; Sun, 1999). They may also explain the frequent episodes of growth collapse observed in newly democratised countries with high corruption perceptions – for instance, Tunisia after the Arab Spring. Policy makers and 'good governance' practitioners should target anti-corruption efforts and resources to where corruption is most harmful – that is, young democracies. This is not only to reward democratisation efforts, but also because corruption control, if successful, may deliver more 'bang for the buck' in democracies than in autocracies.

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Table 1 – Corruption: V-Dem measures

Variable	Question/Clarification (Coppedge et al., 2020a)	Observations	Mean (s.d.)
Political corruption index (*v2x_corr)*	"How pervasive is political corruption?"	25,683	4.6 (2.8)
Public sector corruption index (*v2x_pubcorr*)	"To what extent do public sector employees grant favors in exchange for bribes, kickbacks, or other material inducements, and how often do they steal, embezzle, or misappropriate public funds or other state resources for personal or family use?"	26,143	4.5 (2.9)
Executive corruption index (*v2x_execorr)*	"How routinely do members of the executive, or their agents grant favors in exchange for bribes, kickbacks, or other material inducements, and how often do they steal, embezzle, or misappropriate public funds or other state resources for personal or family use?"	26,006	4.7 (3.0)
Legislature corrupt activities (*v2lgcrrpt*)	"Do members of the legislature abuse their position for financial gain?"; "This includes any of the following: (a) accepting bribes, (b) helping to obtain government contracts for firms that the legislator (or his/her family/friends/political supporters) own, (c) doing favors for firms in exchange for the opportunity of employment after leaving the legislature, (d) stealing money from the state or from campaign donations for personal use."	17,374	4.6 (3.6)
Judicial corruption decision (*v2jucorrdc*)	"How often do individuals or businesses make undocumented extra payments or bribes in order to speed up or delay the process, or to obtain a favorable judicial decision?"	26,506	4.5 (3.5)

Notes: all the variables are rescaled to run from 0 (no corruption) to 10 (most corruption)

Figure 1 – V-Dem corruption index vs. other indicators (2018)

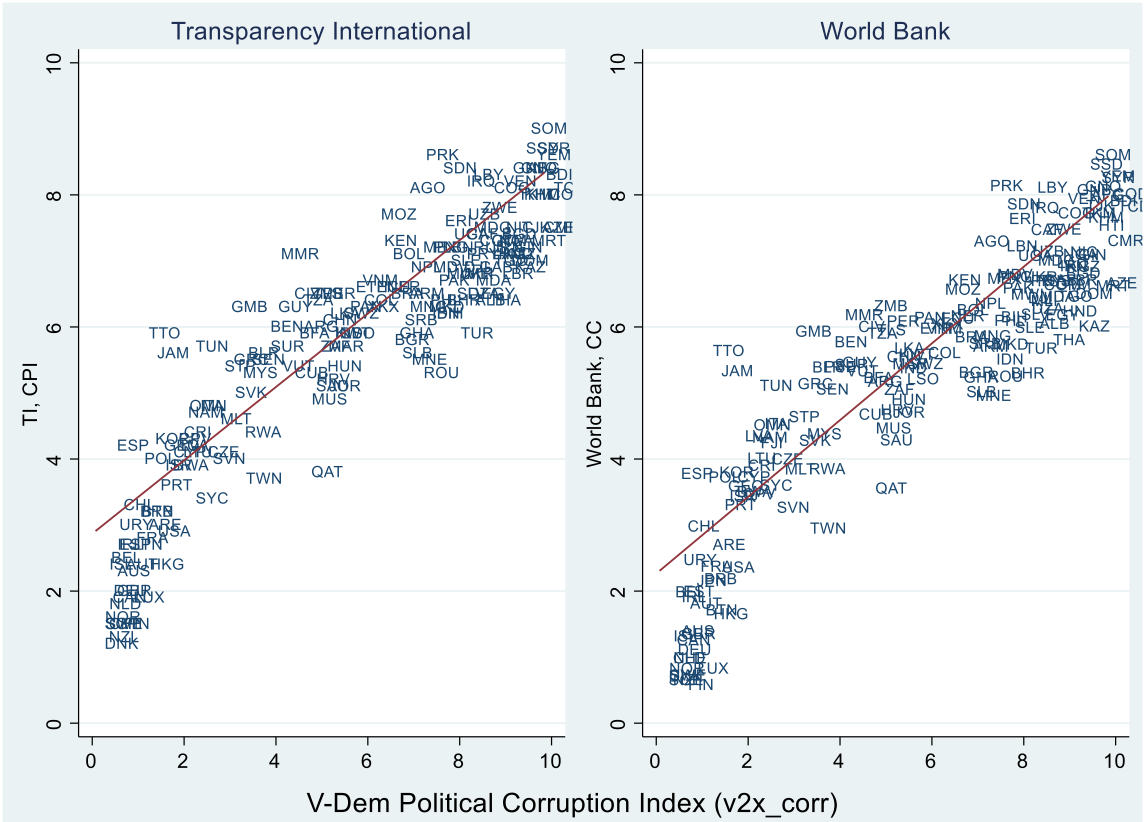


Figure 2 – V-Dem corruption index in four countries, 1789-2019

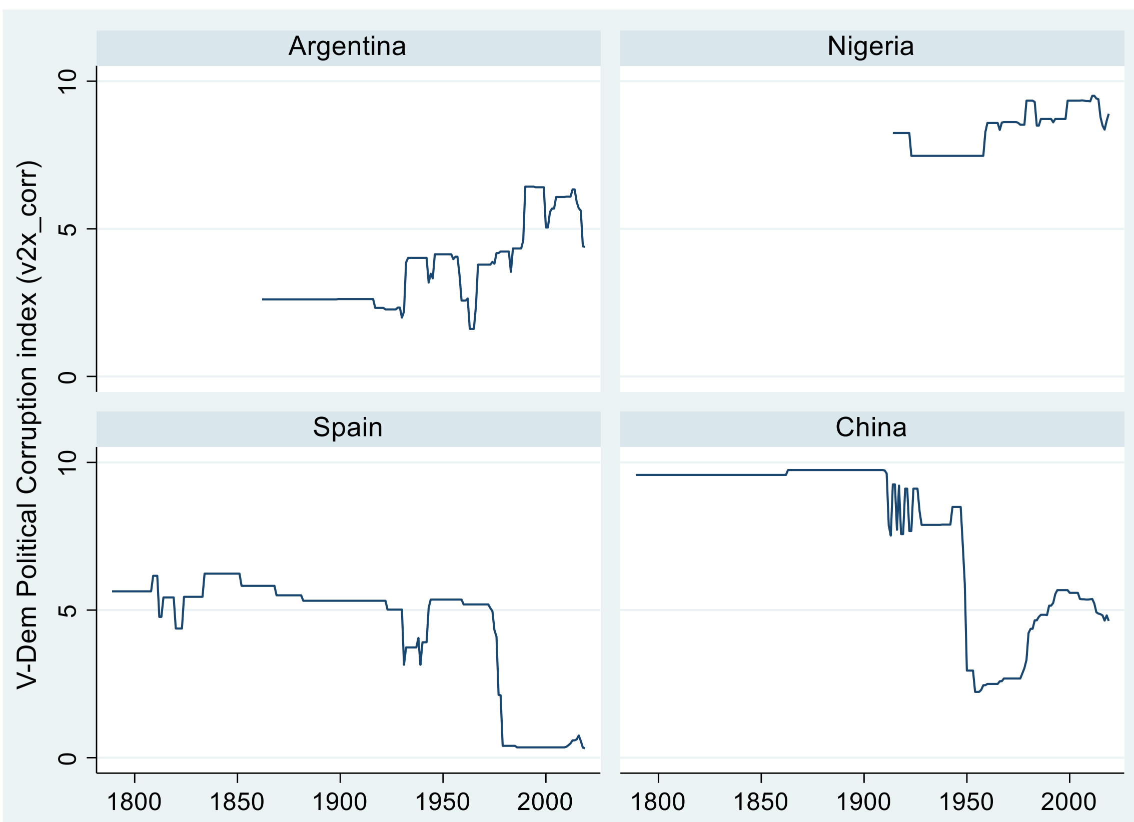


Figure 3 – V-Dem (blue solid) vs. World Bank 's CC (red dashed) indices of corruption, 1996-2019

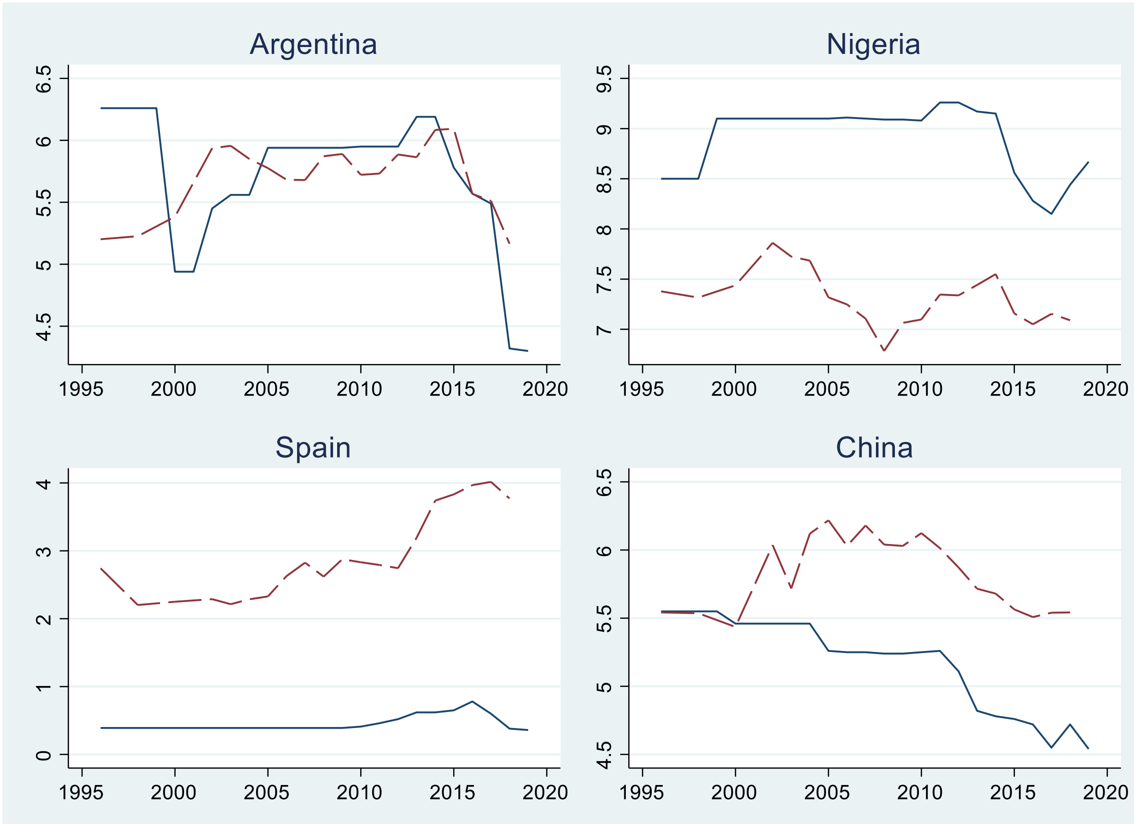


Table 2 - Background institutions: V-Dem measures

Variable	Question/Clarification (Coppedge et al., 2020a)	Observations	Mean (s.d.)
1. State capacity:			
Criteria for appointment decisions in the state administration (*v2stcritrecadm*)	"To what extent are appointment decisions in the state administration based on personal and political connections, as opposed to skills and merit?"; "Appointment decisions include hiring, firing and promotion in the state administration. Note that the question refers to the typical de facto (rather than de jure) situation obtaining in the state administration, excluding the armed forces. [...]"	24,870	4.69 (3.50)
Rigorous and impartial public administration (*v2clrspct*)	"Are public officials rigorous and impartial in the performance of their duties?"; "This question focuses on the extent to which public officials generally abide by the law and treat like cases alike, or conversely, the extent to which public administration is characterized by arbitrariness and biases (i.e., nepotism, cronyism, or discrimination). [...]"	26,416	4.55 (3.47)
2. Regulatory quality:			
State ownership of economy (*v2clstown*)	"Does the state own or directly control important sectors of the economy?"; "This question gauges the degree to which the state owns and controls capital (including land) in the industrial, agricultural, and service sectors. It does not measure the extent of government revenue and expenditure as a share of total output. [...]"	26,681	5.42 (3.50)
3. Property rights:			
Property rights protection (*v2xcl_prpty*)	"Do citizens enjoy the right to private property?"; "Private property includes the right to acquire, possess, inherit, and sell private property, including land. Limits on property rights may come from the state which may legally limit rights or fail to enforce them; customary laws and practices; or religious or social norms. This question concerns the right to private property, not actual ownership of property".	26,687	4.27 (2.84)

4. Quality of democracy:			
Electoral democracy index (*v2x_polyarchy*)	"To what extent is the ideal of electoral democracy in its fullest sense achieved?"; "The electoral principle of democracy seeks to embody the core value of making rulers responsive to citizens, achieved through electoral competition for the electorate’s approval under circumstances when suffrage is extensive; political and civil society organizations can operate freely; elections are clean and not marred by fraud or systematic irregularities; and elections affect the composition of the chief executive of the country. In between elections, there is freedom of expression and an independent media capable of presenting alternative views on matters of political relevance."	25,342	2.63 (2.61)

Notes: all the variables are rescaled to run from 0 (low-quality institutions) to 10 (high-quality institutions).

Table 3 - Background institutions and corruption: cross-correlations

	Corruption	State capacity	Regulatory quality	Property rights	Democracy
Corruption	1				
State capacity	-0.69	1			
Regulatory quality	-0.20	0.40	1		
Property rights	-0.31	0.55	0.51	1	
Democracy	-0.39	0.61	0.46	0.73	1

Table 4 – Average effects of corruption on economic growth, 1795-2010

	Full sample		Restricted sample					
Dependent variable: 5-year GDP per capita growth	Pooled (1)	FE (2)	Benchmark (3)	Controls (4)	Back. Inst. (5)	Lags gr. (6)	Kitchen sink (7)	FGLS (8)
ln GDP pc$_{t-1}$	-0.459***	-1.593***	3.765**	4.319**	3.189*	2.520	2.435	3.072
	(0.156)	(0.516)	(1.599)	(1.639)	(1.651)	(1.708)	(1.748)	(1.920)
ln GDP pc$^2_{t-1}$			-0.287***	-0.328***	-0.255**	-0.220**	-0.225**	-0.281**
			(0.099)	(0.100)	(0.104)	(0.105)	(0.108)	(0.117)
Corruption$_{t-1}$	-0.219***	-0.173***	-0.174***	-0.154*	-0.243***	-0.140**	-0.190***	-0.198***
	(0.044)	(0.060)	(0.055)	(0.060)	(0.065)	(0.052)	(0.063)	(0.072)
ln(P_t) (5 year)				-0.958			-1.082	-1.676***
				(0.778)			(0.773)	(0.585)
Life expectancy$_{t-1}$				0.006			0.010	0.014
				(0.016)			(0.013)	(0.013)
N. of elections held$_t$ (5 year)				-0.030			-0.046	-0.054
				(0.064)			(0.067)	(0.047)
War dummy (5 year)				-0.473***			-0.398***	-0.428***
				(0.094)			(0.097)	(0.113)
State capacity$_{t-1}$					-0.105***		-0.103***	-0.119***
					(0.035)		(0.032)	(0.044)
Regulatory Quality$_{t-1}$					0.029		0.031	0.031
					(0.031)		(0.025)	(0.027)
Property Rights$_{t-1}$					0.170***		0.163***	0.072
					(0.051)		(0.049)	(0.048)
Democracy$_{t-1}$					0.002		-0.015	0.121**
					(0.056)		(0.061)	(0.057)
Lags of Growth (1-3) [*p*-value][1]	No	No	No	No	No	Yes [0.001]	Yes [0.000]	Yes [0.017]
Country FE	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time period FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	3,819	3,819	2,615	2,615	2,615	2,615	2,615	2,615
N. of countries	186	186	141	141	141	141	141	141
Average N. of 5-year periods	20.5	20.5	18.5	18.5	18.5	18.5	18.5	18.5
Adjusted R-squared (within[2])	0.11	0.12	0.16	0.17	0.17	0.20	0.22	0.37

Notes: OLS regressions with Driscoll-Kraay standard errors in parentheses in columns 1-7; FGLS (Prais-Winsten) regression with panel-corrected standard errors in column 8; *** p<0.01, ** p<0.05, * p<0.1; [1] joint test of significance of the three lags of the dependent variable; [2] Overall R-squared in the pooled model (column 1) and in the FGLS model (column 8).

Table 5 – Average effects of corruption on economic growth: Alternative specifications.

Dependent variable: 5-year GDP per capita growth	OLS (1)	2SLS Overid. (2)	OLS Benchmark (3)	OLS Linear trends (4)	OLS Quadratic trends (5)	OLS Cyclical trends (6)	OLS Time x region FE (7)	'First-Difference' GMM (8)	Kiviet LSDVC (9)
ln GDP pc_{t-1}	4.065*	4.069*	4.487*	12.23**	26.39***	1.759	7.759*	-16.92***	2.370*
	(2.164)	(2.137)	(2.369)	(5.443)	(5.921)	(1.509)	(4.210)	(2.506)	(1.369)
ln GDP pc^2_{t-1}	-0.324**	-0.324**	-0.281*	-1.032***	-2.426***	-0.158*	-0.598**		-0.220**
	(0.147)	(0.145)	(0.148)	(0.380)	(0.426)	(0.091)	(0.288)		(0.075)
$Corruption_{t-1}$	-0.231***	-0.214**	-0.291***	-0.258***	-0.253***	-0.207***	-0.159**	-0.275***	-0.183***
	(0.083)	(0.092)	(0.055)	(0.073)	(0.090)	(0.051)	(0.062)	(0.102)	(0.067)
GDP pc growth (region av.)	0.273***	0.273***							
	(0.076)	(0.075)							
AR(1) test [*p*-value]								[0.020]	
AR(2) test [*p*-value]								[0.210]	
Hansen J test [*p*-value]		[0.302]						[0.714]	
Kleibergen-Paap F-stat		476							
Stock-Yogo 10% critical value		20							
N. of instruments		2						128	
Control variables	Yes	Yes	No	Yes	Yes	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time-period FE	Yes	Yes	No	Yes	Yes	Yes	No	Yes	Yes
Country-level time trends	No	No	No	Yes	Yes	Yes	No	No	No
Time-period x Continent FE	No	No	No	No	No	No	Yes	No	No
Observations	2,830	2,830	2,941	2,941	2,941	2,941	2,941	2,759	2,615
N. of countries	150	150	154	154	154	154	154	146	141
Adjusted R-squared (within)	0.17		0.02	0.35	0.50	0.36	0.23		

Notes: the standard errors, which are reported in parenthesis, are clustered within countries in columns 1, 2 and 8, Driscoll-Kraay in columns 3-7, and bootstrapped based on 50 iterations in column 9; *** p<0.01, ** p<0.05, * p<0.1.

The AR(1)/AR(2) tests are the Arellano-Bond tests of no residual AR(1)/AR(2) autocorrelation. AR(2) autocorrelation invalidates the instrument matrix (see Arellano and Bond, 1991). The null of Hansen's J test is that the instruments' exclusion restrictions are valid. In model (8), the lag range used to form moment conditions is (1 3). In column (2), the Kleibergen-Paap F-statistic tests the null that the instruments are not weakly correlated with the endogenous variable. An F-statistic larger than the Stock-Yogo critical value leads to a rejection of the null.

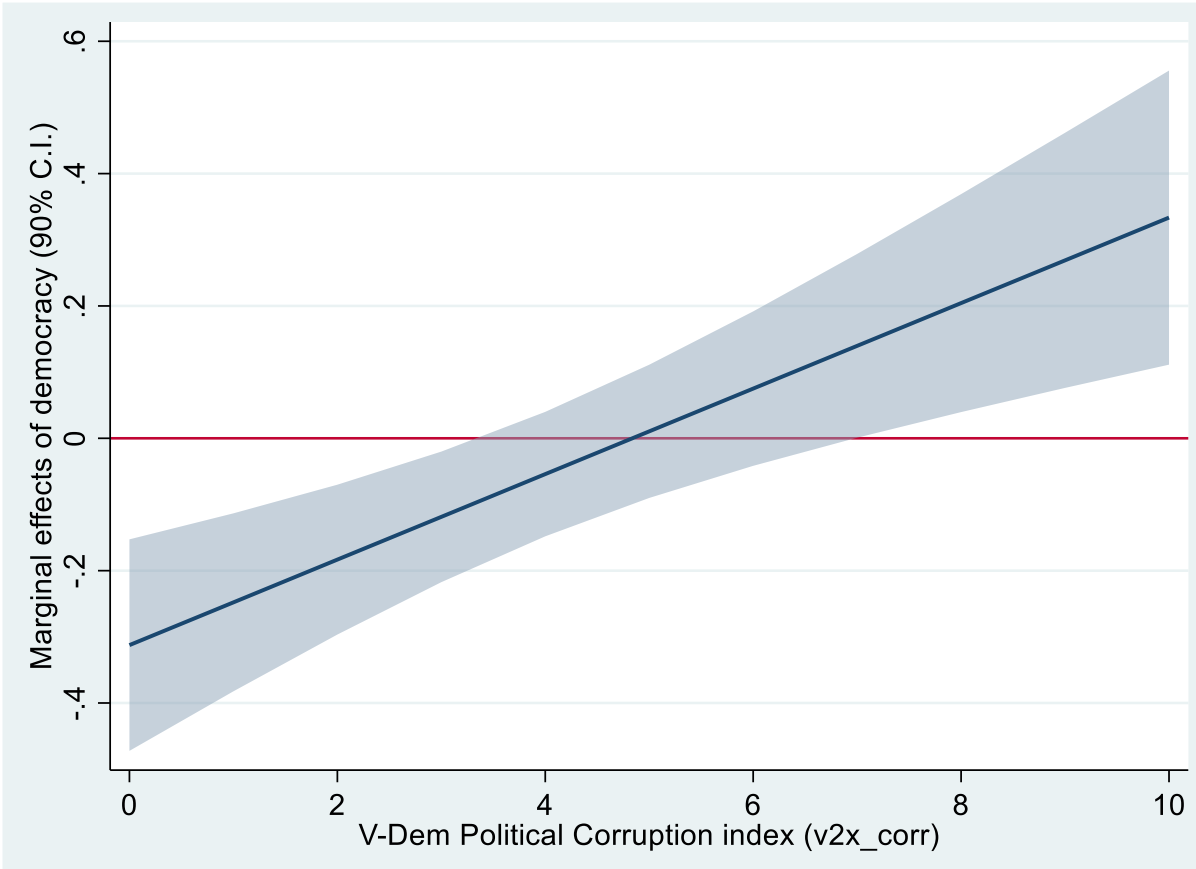
The control variables included are: the log of population growth, life expectancy, the n. of elections held during each 5-year period, the war dummy, and four measures of background institutions (state capacity, regulatory quality, property right protection, democracy). Model 7 also includes three lags of GDP per capita growth.

Table 6 – Corruption and background institutions: Main results, 1795-2010

	All interactions			Interactions one by one			
Dependent variable: av. y.o.y. GDP per capita growth	Pooled (1)	FE (2)	Controls (3)	State capacity (4)	Regulatory Quality (5)	Property rights (6)	Democratic quality (7)
ln GDP pc_{t-1}	-0.600**	-1.790***	3.891*	2.661	2.391	2.702	4.067*
	(0.251)	(0.651)	(2.054)	(1.723)	(1.735)	(1.918)	(2.026)
ln GDP pc^2_{t-1}			-0.325**	-0.239**	-0.222**	-0.243**	-0.335**
			(0.131)	(0.107)	(0.108)	(0.119)	(0.129)
$Corruption_{t-1}$ [γ]	-0.378**	-0.421**	-0.505***	-0.247***	-0.138*	-2.237**	-0.569***
	(0.151)	(0.195)	(0.145)	(0.082)	(0.171)	(0.942)	(0.154)
$Corruption_{t-1}$ x State $capacity_{t-1}$ [α]	-0.006	0.002	-0.005	0.009			
	(0.011)	(0.016)	(0.011)	(0.007)			
$Corruption_{t-1}$ x Regulatory $quality_{t-1}$ [α]	0.012	-0.015	-0.013		-0.010		
	(0.008)	(0.011)	(0.008)		(0.009)		
$Corruption_{t-1}$ x Property $rights_{t-1}$ [α]	-0.030	-0.041*	-0.018			0.007	
	(0.017)	(0.022)	(0.019)			(0.014)	
$Corruption_{t-1}$ x $Democracy_{t-1}$ [α]	0.044**	0.067**	0.063***				0.045***
	(0.021)	(0.031)	(0.019)				(0.013)
State $capacity_{t-1}$ [β], inverted	0.064	-0.031	0.093	0.057	0.096***	0.103***	0.086***
	(0.075)	(0.081)	(0.058)	(0.038)	(0.034)	(0.032)	(0.028)
Regulatory $Quality_{t-1}$ [β], inverted	-0.065	0.061	0.036	-0.029	0.026	-0.032	-0.040
	(0.053)	(0.084)	(0.063)	(0.024)	(0.064)	(0.025)	(0.026)
Property $Rights_{t-1}$ [β], inverted	-0.034	-0.109	-0.085	-0.159***	-0.168***	-0.192***	-0.151***
	(0.073)	(0.074)	(0.073)	(0.048)	(0.050)	(0.062)	(0.050)
$Democracy_{t-1}$ [β], inverted	-0.166	-0.168	-0.174	0.019	0.016	0.022	-0.106
	(0.155)	(0.170)	(0.123)	(0.060)	(0.061)	(0.066)	(0.090)
Control variables	No	No	Yes	Yes	Yes	Yes	Yes
Lags of Growth (1-3) [p-value][1]	No	No	Yes [0.000]	Yes [0.000]	Yes [0.000]	Yes [0.000]	Yes [0.000]
Country FE	No	Yes	Yes	Yes	Yes	Yes	Yes
Time period FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	3,536	3,536	2,615	2,615	2,615	2,615	2,615
N. of countries	183	183	141	141	141	141	141
Average N. of 5-year periods	19.3	19.3	18.5	18.5	18.5	18.5	18.5
Adjusted R-squared (within[2])	0.13	0.14	0.22	0.22	0.22	0.22	0.22

Notes: OLS regressions with Driscoll-Kraay standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1; [1] joint test of significance of the three lags of the dependent variable; [2] Overall R-squared in the pooled model (column 1) and the FGLS model (column 8).The control variables are: the log of population growth, life expectancy, the n. of elections held during each 5-year period, and the war dummy.

Figure 4 – Marginal effects of democracy (inverted) on economic growth, conditional on the level of corruption



Notes: the margins plot displays the marginal effects of corruption on GDP per capita growth as a linear function of the level of corruption (with 90 percent confidence intervals), holding the co-variates constant at the means. The marginal effects are based on FGLS estimates of the kitchen sink specification reported in column 3, Table 7.

Table 7 – Corruption and background institutions: alternative specifications

Dependent variable: Av. y.o.y. GDP per capita growth	OLS Linear trends (1)	OLS Quadratic trends (2)	OLS Cyclical trends (3)	OLS time x region FE (4)	'First-Difference' GMM (5)	Kiviet LSDVC (6)
$Corruption_{t-1}$	-0.716***	-0.645***	-0.216***	-0.384***	-0.596**	-0.367**
	(0.184)	(0.191)	(0.076)	(0.141)	(0.252)	(0.144)
$Corruption_{t-1}$ x $State\ capacity_{t-1}$	-0.003	-0.008	-0.001	-0.000	-0.004	-0.008
	(0.014)	(0.016)	(0.009)	(0.012)	(0.021)	(0.017)
$Corruption_{t-1}$ x $Regulatory\ quality_{t-1}$	-0.013	-0.001	-0.007	-0.005	0.005	-0.015
	(0.009)	(0.013)	(0.008)	(0.009)	(0.017)	(0.010)
$Corruption_{t-1}$ x $Property\ rights_{t-1}$	0.001	0.030	-0.009	-0.017	-0.011	-0.015
	(0.029)	(0.024)	(0.021)	(0.020)	(0.028)	(0.016)
$Corruption_{t-1}$ x $Democracy_{t-1}$	0.063***	0.032	0.040**	0.044***	0.047*	0.053***
	(0.019)	(0.023)	(0.015)	(0.014)	(0.027)	(0.019)
Observations	2,941	2,941	2,941	2,941	3,099	2,951
Number of countries	154	154	154	154	154	160
Adjusted R-squared (within)	0.35	0.50	0.36	0.24		

Notes: the standard errors, which are reported in parenthesis, are Driscoll-Kraay in columns 1-4, cluster-robust in column 5, and bootstrapped based on 50 iterations in column 6; *** p<0.01, ** p<0.05, * p<0.1.

All regressions control for country FE and (with the exception of model 4) time-period FE. They also control for lagged income, lagged income squared, population growth, life expectancy, the number of elections held, the war dummy, and the four indicators of background institutions. All institutional indicators are inverted, so that a higher number denotes greater institutional weaknesses along the relevant dimension. The model in column 6 also includes the first three lags of GDP per capita growth. The estimates of the other parameters are not reported to save space

Figure 5 – Marginal effects of corruption on economic growth by quintiles of institutional quality

Notes: the coefficient plots display the marginal effects of corruption on GDP per capita growth in different quintiles of the distribution of each institutional indicator (inverted), with 90 percent confidence intervals. Each quadrant corresponds to a different model. In each model, the other three institutional dimensions, and their interactions with corruption, are included in the regression, along with lagged income and lagged income squared, our vector of control variables, three lags of GDP per capita growth, and country and time-period FE.

Table 8 – Corruption and background institutions: 1- and 10-year panels (1800-2010)

Dependent variable: Av. y.o.y. GDP per capita growth	Consecutive years (1)	5-year panel (2)	10-year panel (3)
Corruption_{t-1}	-0.421	-0.508**	-0.343*
	(0.287)	(0.209)	(0.192)
Corruption_{t-1} x $\text{State capacity}_{t-1}$	-0.003	0.000	-0.017
	(0.019)	(0.016)	(0.021)
Corruption_{t-1} x $\text{Regulatory quality}_{t-1}$	-0.011	-0.017	-0.032*
	(0.012)	(0.012)	(0.016)
Corruption_{t-1} x $\text{Property rights}_{t-1}$	-0.059**	-0.037	-0.017
	(0.026)	(0.032)	(0.030)
Corruption_{t-1} x Democracy_{t-1}	0.083***	0.076***	0.076***
	(0.028)	(0.023)	(0.025)
Average marginal effects of corruption	-0.216***	-0.237***	-0.117*
	[0.083]	[0.056]	[0.068]
Observations	17,480	3,472	1,726
Number of countries	172	172	172
Average N. of 5-year periods	101.6	20.2	10.1

Notes: OLS regressions with Driscoll-Kraay standard errors in parenthesis; delta-method standard errors in brackets; *** p<0.01, ** p<0.05, * p<0.1.
All regressions control for country and time-period FE, lagged income, lagged income squared, and the four institutional indicators, which are inverted, so that a higher number denotes greater institutional weaknesses along the relevant dimension.

Table 9 – Forms of corruption and economic growth

Dependent variable: Av. y.o.y. GDP per capita growth	V-Dem index (1)	Index components (2)	Index components (3)	FGLS (4)	GDP Lags (5)	Linear Trends (6)
Corruption (*v2x_corr*)	-0.241***					
	(0.055)					
Types of corruption:						
Bureaucratic (v2x_pubcorr)		-0.197***	-0.189**	-0.099	-0.204*	-0.124*
		(0.069)	(0.088)	(0.077)	(0.113)	(0.077)
Executive (*v2x_execorr*)		0.010	-0.000	-0.032	0.058	0.070
		(0.057)	(0.061)	(0.056)	(0.070)	(0.060)
Legislative (*v2lgcrrpt*)			-0.120**	-0.173***	-0.099*	-0.137**
			(0.050)	(0.047)	(0.051)	(0.053)
Judicial (*v2jucorrdc*)		-0.030	0.052	0.080	0.015	-0.103
		(0.033)	(0.040)	(0.065)	(0.046)	(0.070)
Observations	2,941	2,941	2,106	2,106	1,903	2,106
Number of countries	154	154	153	153	141	153

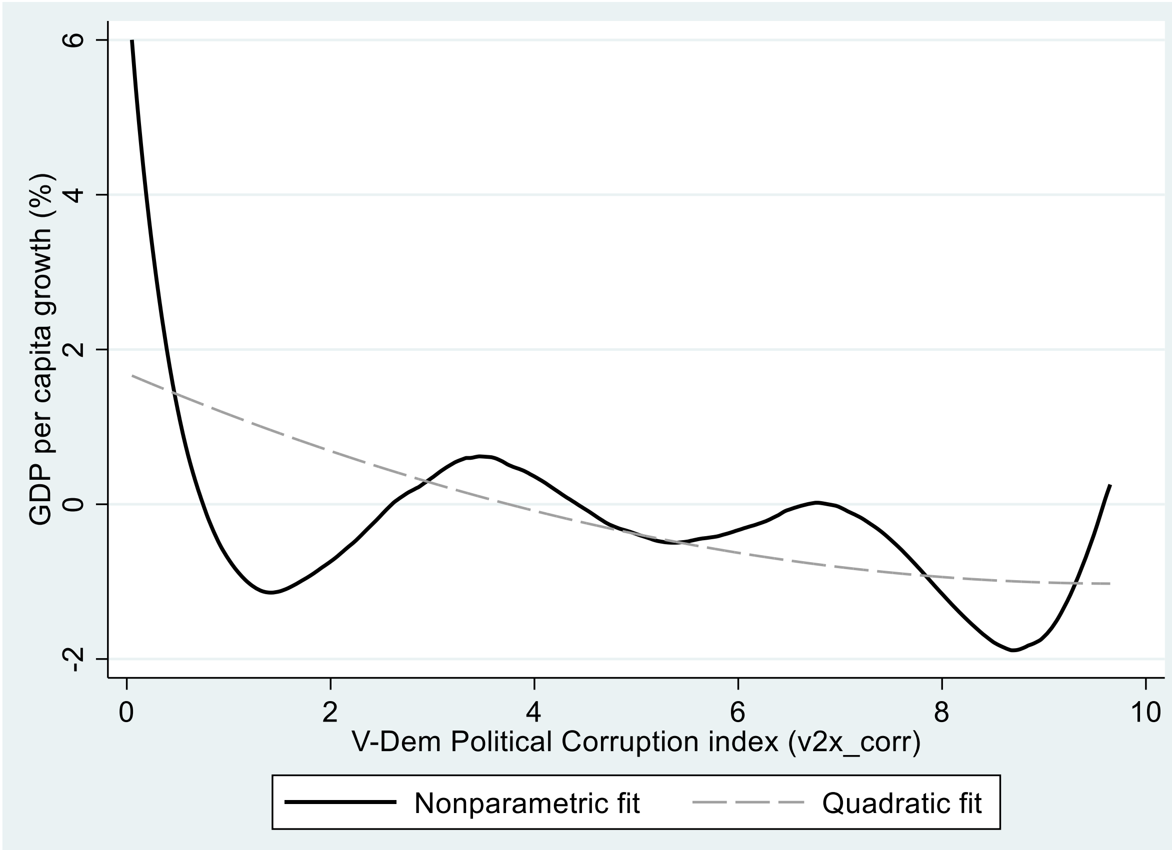
Notes: OLS regressions with Driscoll-Kraay standard errors in parentheses in columns 1-3 and 5-6; FGLS (Prais-Winsten) regression with panel-corrected standard errors in column 4*** p<0.01, ** p<0.05, * p<0.1. All models control for country and time-period FE, lagged income, lagged income squared, the usual vector of time-varying controls and the four institutional indicators from V-Dem.

Table 10 - Forms of corruption and democracy

	Dependent variable $_t$:				
	Corruption (*v2x_corr*) (1)	Bureaucratic (*v2x_pubcorr*) (2)	Executive (*v2x_execorr*) (3)	Legislative (*v2lgcrrpt*) (4)	Judicial (*v2jucorrdc*) (5)
Panel A: Regressions with country and time-period FE					
Dependent variable $_{t-1}$	0.876***	0.881***	0.832***	0.844***	0.835***
	(0.015)	(0.019)	(0.019)	(0.025)	(0.017)
Democracy $_{t-1}$	-0.003	0.000	-0.031**	0.007	0.012
	(0.008)	(0.010)	(0.011)	(0.013)	(0.008)
Observations	4,515	4,571	4,580	2,795	4,616
Number of countries	198	198	198	192	198
Panel B: Regressions with country FE, time-period FE, lagged income and lagged income squared					
Dependent variable $_{t-1}$	0.825***	0.839***	0.781***	0.820***	0.782***
	(0.023)	(0.025)	(0.024)	(0.034)	(0.025)
Democracy $_{t-1}$	-0.016*	-0.008	-0.051***	-0.006	0.004
	(0.009)	(0.010)	(0.013)	(0.015)	(0.009)
Observations	3,711	3,735	3,740	2,553	3,764
Number of countries	186	186	186	181	186

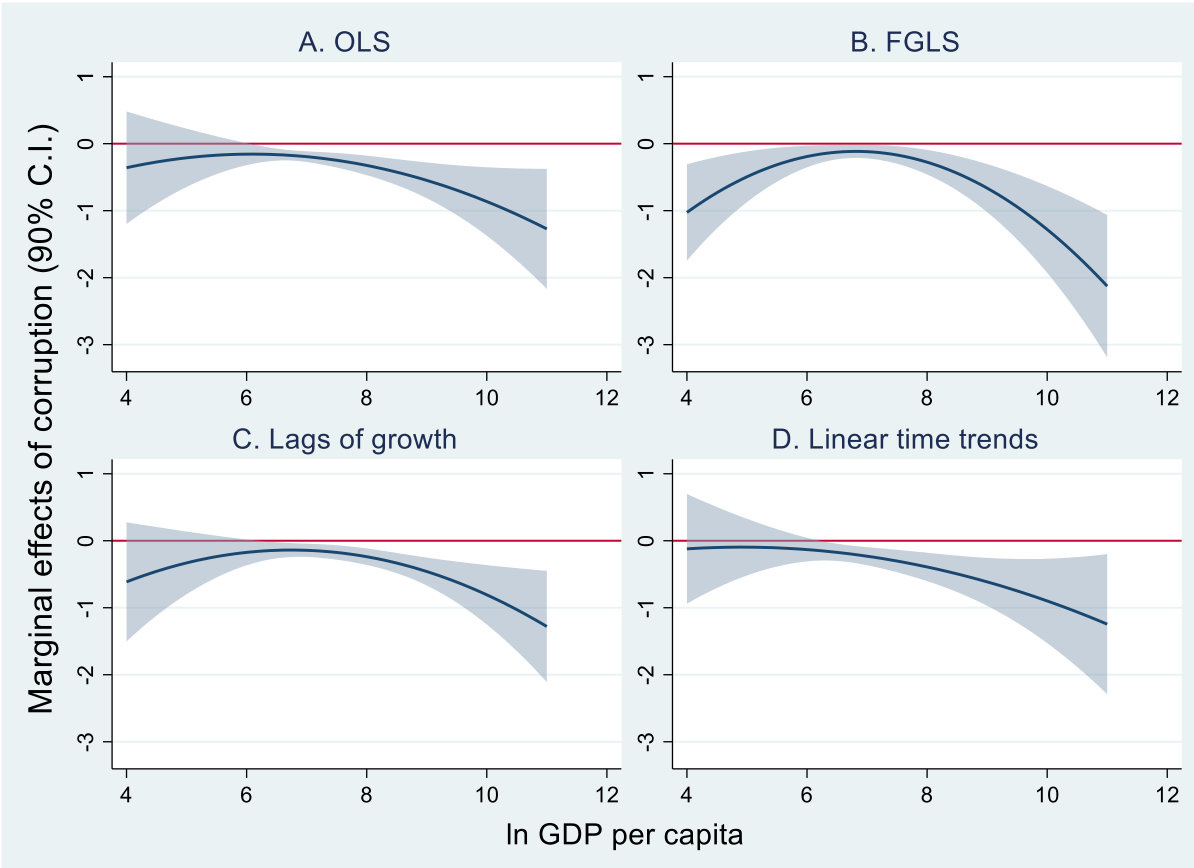
Notes: OLS regressions with Driscoll-Kraay standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1. The democracy indicator is *not* inverted here.

Figure 6 – Corruption and Economic Growth: Nonlinearities



Notes: the diagram displays the predicted values of GDP per capita growth, holding the other co-variates constant at the means, based on a 'kitchen sink' specification in which corruption enters the equation non-parametrically. The estimates are obtained using Baltagi and Li's (2002) series estimator of partially linear panel-data models with fixed effects, as implemented by Libois and Verardi (2013). The nonparametric part of the regression uses cubic splines to fit the partial relationship between GDP per capita growth and corruption.

Figure 7 – Corruption and Economic Growth across different levels of income



Notes: the diagram displays the conditional marginal effects of corruption on GDP per capita growth (with 90% confidence intervals), holding the other co-variates constant at the means.

Table 11 - Controlling for investment and other covariates, 1975-2005

Dependent variable:	$\Delta \ln y_t$	$\Delta \ln y_t$	$\Delta \ln y_t$	$\Delta \ln y_t$	$\ln(P_t)$	$\ln(s_t)$
	(1)	(2)	(3)	(4)	(5)	(6)
ln GDP pc_{t-1}	-4.199***	-4.689***	-5.295***	-5.412***	0.070**	0.050
	(0.808)	(0.852)	(0.826)	(0.820)	(0.035)	(0.051)
$\ln(P_t)$	-1.176	-1.335	-0.812	-0.945		
	(2.214)	(2.169)	(2.109)	(2.076)		
$\ln(s_t)$		2.414***		1.657***		
		(0.459)		(0.553)		
$\ln(P_{t-1})$					0.115***	
					(0.031)	
$\ln(s_{t-1})$						0.220***
						(0060)
Corruption_{t-1}	-0.223	-0.311	-0.406*	-0.453*	-0.005	0.009
	(0.184)	(0.219)	(0.217)	(0.226)	(0.005)	(0.021)
Corruption_{t-1} x Democracy_{t-1}	0.052***	0.069***	0.082***	0.089***	0.000	-0.002
	(0.019)	(0.015)	(0.023)	(0.021)	(0.001)	(0.003)
Democracy_{t-1} (inverted)	-0.112	-0.196*	-0.286**	-0.311**	0.004	0.009
	(0.129)	(0.114)	(0.130)	(0.125)	(0.005)	(0.010)
Control variables	Yes	Yes				
Alternative set of controls			Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Time-period FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	502	502	502	502	509	470
Number of countries	109	109	109	109	108	102
Adjusted R-squared	0.23	0.27	0.30	0.31	0.20	0.35

Notes: OLS regressions with Driscoll-Kraay standard errors in parentheses; *** p<0.01, ** p<0.05, * p<0.1. The control variables are life expectancy, the number of elections held and the war dummy. The alternative set of controls includes the average number of years of education, the number of coups, the log of natural resource rents, trade dependence, the inflation rate and the war dummy.

Appendix A - Empirical literature on Corruption and Economic Growth

Study	Data type	N. of countries (time-period)	Measure of corruption	Estimation (instrument)	Outcome variable	Grease the wheel (confirm)	Institutional interactions	Effect of corruption	Main channels
Mauro, 1995	Cross-section	58 (1960-85)	BI (1980-83)	IV (ethnic fractionalisation)	GDP pc growth, investment	No		-	Private investment
Ehrlich and Lui, 1999	Panel	68 (1981-92)	BI (1981)	OLS	GDP pc growth	No		0	
Olson et al., 2000	Cross-section	51 (1960-87)	ICRG (1982)	OLS	Productivity growth	No		-	
Mo, 2001	Cross-section	46 (1960-85)	TI (1980-85)	OLS	GDP growth, and others	No		-	Political stability, investment
Lambsdorff, 2003	Cross-section	69 (2000)	TI (2001)	IV (share of Protestants)	Productivity (GDP/capital stock)	No		-	
Pellegrini and Gerlagh, 2004	Cross-section	48 (1975-96)	TI (1980-85)	IV (legal origin)	GDP pc growth	No		-	Investment, trade openness
Rock and Bonnett, 2004	Cross-section	90 (1980-96)	BI (1980-83), TI (1988-92), WB (1994-1996), ICRG (1984-86)	OLS	GDP pc growth, investment	Yes (Yes)	Implicit (by country)	- (but + in East Asia)	Productivity (East Asia), investment (elsewhere)
Meon and Sekkat, 2005	Cross-section	67 (1970-1998)	TI (1999), WB (1997-98)	GLS	GDP pc growth	Yes (No)	Rule of law, government effectiveness, lack of violence	- (more - in countries with bad institutions)	Investment, productivity
Mendez and Sepulveda, 2006	Cross-section and Panel	85 (1960-2000, 1984-2000 for the panel)	ICRG (1982-2001), IMD (1990-2000), TI (1996-2003)	OLS, OLS with FE	GDP pc growth	Yes (Yes)	Democracy, implicit (non-linear effects)	- (but only in democracies), non-monotonic	Investment, productivity
Aidt et al., 2008	Cross-section	64 (1995-2000), 58 (1970-2000)	TI (1996-2002), WB (1996-2002)	IV (ethnic fractionalisation, n. years democracy)	GDP pc growth	Yes (Yes)	Democracy	- (but only in democracies)	Productivity

Aidt, 2009	Cross-section	60 (1970-2000)	TI, WBES (1999-2000)	IV (ethnic fractionaliation, democracy)	GDP pc growth	Yes (Yes)	Rule of law, democracy	+ (but less + in countries with rule of law), - in democracies	Productivity
Meon and Weill, 2010	Cross-sections (pooled)	92 (2000-03)	TI (2000-03), WB (2000-03)	Stochastic frontier (ML)	Output per worker	Yes (Yes)	Government effectiveness, regulatory quality	- (but less - in countries with bad institutions)	Productivity
Heckelman and Powell, 2010	Cross-section	82 (2000-2005)	TI (1995-2000)	OLS	GDP pc growth	Yes (Yes)	Economic freedom	+ (but less + in economically free countries)	Productivity
Swaleheen, 2011	Cross-section and Panel	122 (1984-2007)	ICRG (1984-2007), TI (1995-2007)	OLS, OLS-FE, GMM	GDP pc growth	Yes (Yes)	Implicit (non-linear effects)	- (but + in countries with high corruption)	Productivity
Hodge et al., 2011	Cross-sections (pooled)	81 (1984-2005)	ICRG (1984-2005)	3SLS	GDP pc growth	Yes (Yes)	Government effectiveness, regulatory quality	- (but less - in countries with bad institutions)	Investment, political stability
Huang, 2016	Panel	13 (1997-2013)	TI (1997-2013)	Bootstrap Granger causality testing	GDP pc growth	Yes (Yes)	Implicit (by country)	0 (but + in South Korea)	
D'Agostino et al., 2016a and 2016b	Panel	106 or 48 (1996-2010)	WB (1996-2010)	GMM (diff)	GDP pc growth	No		-	Productivity, public investment, military spending
Cieslik and Goczek, 2018	Panel	142 (1996-2014)	WB (1996-2014)	GMM (sys)	GDP pc growth	No		-	productivity, investment
Grundler and Potrakfe, 2019	Panel	175 (2012-18)	TI	IV-FE (regional corruption)	GDP pc growth	Yes (No)	Rule of law, government effectiveness, democracy	- (but + in countries with good institutions)	FDI, inflation