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# Who profits from windfalls in oil tax revenue? Inequality, protests, and the role of corruption

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# Abstract

We investigate the relationship between oil windfalls and income inequality using the subnational data of one of the resource-richest and most unequal countries in the world – Russia. While previous literature produced contradictory findings due to the use of an aggregate measure of oil rents mainly in cross-national settings, we focus exclusively on oil rents that accrue to the subnational governments across one country. Our estimation strategy takes advantage of the two specific features of Russian oil taxation: 1) the policy change when sharing oil extraction taxes with local budgets was discontinued; and 2) the oil tax formula being tied directly to the international oil prices making oil price shocks an exogenous measure of change in oil rents. When we look at the period with oil tax revenues shared with the regional governments, we find that oil windfalls had increased income inequality and benefited the wealthiest quintile of the population in regions with more intense rent-seeking. Further, positive oil price shocks combined with greater rent-seeking reduced the share of labor income but increased the income share from unidentified sources traditionally associated with corruption. These effects of oil windfalls disappeared after the Russian government discontinued oil tax revenue sharing with regional governments. Finally, we examine some political implications of rising inequality due to the appropriation of oil windfalls. We find a positive effect of rising inequality on the frequency of protests associated with grievances among the poor and disadvantaged social groups; this effect, however, exists only in relatively democratic regions.

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# 1 Introduction

Many countries heavily tax their oil industry on the premise that natural resources belong to their people, but do all the people benefit from oil tax revenues equally? This important question has eluded an unambiguous answer. Additional budget revenues from oil windfalls can improve incomes of the poor, and middle class through direct redistribution (as discussed in Parcero and Papyrakis, 2016) or increased social spending (Kim and Lin, 2018), but the very same revenues are often misappropriated by the wealthy and politically connected elite, thus increasing income inequality. Indeed, misappropriation of oil rents by political elites in the form of corruption has been well documented in the existing literature (Arezki and Brückner, 2012; Caselli and Michaels, 2013; Andersen et al., 2017; James and Rivera, 2021; Asher and Novosad 2022), but it has not been previously linked to rising income inequality. *We hypothesize that the intensity of current corruption significantly affects the impact of oil windfalls on income distribution in society, resulting in a positive association between oil price shocks and inequality in a highly corrupt environment.*

We test our hypothesis using the subnational data for Russia – one of the biggest producers of oil and, at the same time, one of the most unequal countries in the world (Novokmet et al., 2018). Studying inequality within one country circumvents a major problem of selective underreporting of data on income inequality by resource-rich nations that plagues cross-country studies (e.g., Ross, 2007; Parcero and Papyrakis, 2016). There are several additional advantages of using the Russian subnational data for our estimation strategy. First, Russian regions are relatively homogenous in their legal system, fiscal and monetary policies, but about a third of them are endowed with oil wealth, and their budgets had been directly benefiting from a degree of decentralization of oil extraction tax revenues until this regime was gradually discontinued in 2010-2012. The heterogeneity of oil endowments across regions and the change in the assignment of oil tax revenues over time due to plausibly exogenous policy change enables us to explore the relationship between oil tax revenues and income inequality. Second, the formula for calculating regional oil tax revenue included the international oil price, and, therefore, oil price shocks can serve as an exogenous measure of oil tax revenue fluctuations. We find no evidence that oil price shocks directly affect regional income inequality, whether regional budgets receive a share of oil tax revenue or not. To examine how this effect is mediated by the ability of regional elites to capture government rents,

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we interact the oil price shocks with the measures of regional intensity of corruption introduced in Schulze et al. (2016). We find that in a more corrupt environment, oil windfalls increase the share of the top quintile of income distribution and the Gini coefficient, but this effect appears only in the period when oil-rich regions were receiving a share of oil extraction taxes.[1] Our findings are robust to using an alternative, albeit less precise, regional measure of rent-seeking, the inclusion of control variables, and is confirmed by the placebo tests of substituting corruption incidents with overall crime or murder.

Further, we reinforce our main results with the evidence that oil windfalls in a corrupt environment affect the composition of income sources by increasing the share of income that can be plausibly linked to corruption. We employ regional estimates of income shares derived from employment, entrepreneurship, social transfers, property, and from "other sources" calculated as a residual from subtracting savings and expenditures from overall income. This latter category is the most useful for our case since "other income" includes revenues from the shadow economy and illegal activities, and Gorodnichenko and Sabirianova Peter (2007) propose using it as a measure of corruption income. Our data show that in the environment of increasing corruption, the share of "other income" rises consistently after a positive oil price shock while the income share from employment declines.

Finally, we look at the political implications of rising inequality due to the appropriation of oil windfalls. On the one hand, rising inequality exacerbates economic grievances leading to more frequent public protests (e.g., De Juan and Wegner, 2019, Kurer et al., 2019). On the other hand, oil rents have been previously found to decrease the likelihood of political protests (Smith, 2004; Vadlamannati and Soysa, 2020) by allowing the incumbent politicians to buy out the opposition or finance a sizable security apparatus (Basedau and Lay, 2009). We employ a newly assembled dataset on regional protests in Russia introduced by Lankina and Tertytchnaya (2020) that differentiates between various categories of protests.[2] We find a significant positive effect of rising inequality instrumented with oil price shocks under rent-seeking on the "social" protests – "*protests by, and specifically furthering the aims of, socially vulnerable groups of people*"

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[1] This result is consistent with Gupta et al. (2002), who used cross-country data to show that higher corruption is associated with greater inequality and poverty.
[2] The dataset is published as Lankina (2018).

(Lankina and Tertytchnaya, 2020: 34), but not the protests associated with other issues such as anti-government, cultural, environmental, or against specific legal or economic policies. However, this effect is driven solely by the subsample of relatively democratic regions. In contrast, no effect can be found among the relatively authoritarian regions, presumably because in these latter regions, the authorities are relatively more capable of suppressing protest activity, protesting is more costly to the participants, or there are too few activists, to begin with (Tertytchnaya and Lankina, 2020). To our knowledge, we are the first to show the positive relationship between the oil windfalls mediated by different corruption regimes and protest frequency.[3]

Our research advances the existing empirical literature on the effect of natural resources on inequality (Gylfason and Zoega, 2003; Ross, 2007; Buccellato and Mickiewicz, 2009; Goderis and Malone, 2011; Carmignani, 2013; Howie and Atakhanova, 2014; Parcero and Papyrakis, 2016; Kim and Lin, 2018) in several dimensions. First, this is the first study to show that this effect depends on the degree to which the institutional environment allows for rent-capture, while the previous literature often assumed the relationship between natural resources and inequality to be linear and direct. Second, we conduct our empirical analysis in a within-country setting that provides us with a better identification strategy and circumvents the usual drawbacks of cross-country studies. Third, contrary to the previous literature that has traditionally treated oil rents as a homogeneous aggregate, we differentiate among different types of oil rents as they often accrue to different actors and can serve different purposes. For example, the oil sector normally produces rents in the form of oil firms' profits, dividends to shareholders, higher wages of employees and management, export taxes, extraction taxes, and unofficial taxes (Gaddy and Ickes, 2005), and it is reasonable to assume that they might have disparate implications for income inequality.[4] The major novelty of this paper is thus the focus on the oil rent in the form of government tax revenue that is supposed to benefit the population as a whole, according to the official narrative. However, the

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[3] We acknowledge a recent paper by Ishak and Farzanegan (2021) that looks at the cross-country evidence and finds that negative oil price shocks (i.e., the decline in oil rents in oil exporters) increase the incidents of demonstrations, strikes, and riots but only in countries with a relatively small size of the shadow economy.

[4] For example, bonuses to the top management would obviously work towards increasing the overall inequality; the same pro-inequality effect but potentially at a lesser rate would apply to rents to the shareholders of oil companies; on the contrary, most of the tax-related rents should be more equally distributed via the provision of public goods and services, and especially social transfers, leading presumably to lower income inequality. Our claim in this paper is that tax windfalls accruing to local budgets are particularly prone to increasing inequality in a highly corrupt environment.

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existing literature suggests that revenue increases in local budgets do not necessarily guarantee an improvement of living conditions for the local population (Brollo et al., 2013), and, instead, might be misappropriated by officials in charge of managing the budgets (Nikolova and Marinov, 2017).[5] Therefore, we argue that oil rents accruing directly to the local budgets are particularly likely to exacerbate inequality.

Our findings also add to the ample literature on the effect of economic inequality on non-violent protests. Contrary to many studies that find inequality to depress protest activities (Dubrow et al. 2008; Solt, 2008, 2014), we show a positive effect of rising income inequality on the occurrence of social protest. This is, however, consistent with Kurer et al. (2019), who finds that not the levels but the changes in economic grievances (e.g., deterioration of economic prospects) tend to galvanize protest activity.

Finally, our paper contributes to the literature on corruption in Russia by introducing a novel dataset on the registered incidents of embezzlement acquired from the Ministry of Internal Affairs of the Russian Federation, which covers all of 83 Russian regions over 14 years.[6] These data offer an alternative proxy for the risk of rent capture for testing our hypothesis. In our estimation strategy, we rely only on the within-variation of the corruption and embezzlement measures using first-differencing and fixed effects specifications to account for region-specific time-invariant characteristics.

The paper is structured as follows. Section 2 reviews the existing literature and identifies our main contributions. Section 3 describes the data. Estimation strategy and empirical results are presented and discussed in Section 4. Section 5 provides robustness checks. Section 6 concludes with a summary of the findings. The Appendix contains descriptive statistics, comparisons of general economic trends between two periods under study, and additional regressions.

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[5] The appropriation of the windfalls is not exclusive for the subnational budgets: for example, a recent paper by Andersen et al. (2020) finds that the foreign aid provided to national governments is strongly associated with surging cash flows to the offshore financial centers.

[6] Most previous literature on regional corruption in Russia relies on perception-based corruption measures (e.g., Dininio and Orttung 2005; Detkova et al. 2021), which are often criticized for being biased (e.g., Jensen et al. 2010; Sharafutdinova 2010). Schulze and Zakharov (2018) provide a literature review on subnational corruption in Russia.

# 2 Literature review on oil rents and income inequality

Our paper contributes to the literature on the effect of natural resources on inequality by focusing on one particular type of oil rents within a single country. Previous research is represented mainly by cross-country comparisons that do not differentiate between the types of rents and produce mixed results (Sebri and Dachraoui, 2021). In the earliest investigation of the topic, Gylfason and Zoega (2003) find that natural resource dependence increases inequality. However, this result was obtained in a cross-section of countries using per capita GDP as a control variable. Alexeev and Conrad (2009a) argued that this approach is problematic. Nonetheless, even adjusting for the endogeneity of per capita GDP, the share of mining in GDP appears to lead to higher income inequality in a country, although this result does not hold for oil abundance (Alexeev and Conrad, 2005). Two more recent papers produced contradictory results. The first paper by Carmignani (2013) examines a cross-section of 80 countries and finds that countries rich in subsoil resources are more unequal on average. Goderis and Malone (2011) employ a more sophisticated identification strategy by using a dynamic panel model with an exogenous variation in international commodity prices and find that subsoil-mineral booms in developing countries lower inequality in the same year, but inequality returns to the pre-boom level in the long-run. The idea of a nonlinear relationship between natural resources and inequality is proposed by Fum and Hodler (2010), who find that natural resource abundance is mediated by ethnic polarization so that natural resources are associated with greater income inequality in more polarized countries. This finding is related to our hypothesis that the likelihood of rent-capture affects how oil windfalls are distributed since ethnic polarization has been linked to rent-seeking (e.g., Dincer, 2008). Hartwell et al. (2019) use the World Bank measure of natural resource rents and cross-country data to show that the impact of these rents on inequality is conditional on the level of democracy: natural resource rents in more (less) democratic countries result in lower (higher) inequality.[7] Most studies mentioned above employ a very broad measure of natural resources that include minerals, hydrocarbon fuels, land

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[7] Although Hartwell et al.'s argument is reminiscent of ours, their econometric approach is very different. In addition to using only cross-sectional data, their regressions include the interaction term between a measure of democracy and natural resource rents, but not the democracy measure on its own. This makes it unclear whether democracy mediates the relationship between resource rents and inequality or is by itself the main cause of lower inequality.

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wealth, forests, and raw agricultural materials. This approach is problematic since rents generated by different resources tend to differ in their characteristics and may affect inequality differently.[8]

Two papers that focus exclusively on oil rents are relatively recent and support a mostly negative relationship. The first one by Parcero and Papyrakis (2016) finds a U-shaped relationship such that oil wealth decreases income inequality in modestly oil-rich nations but increases it in the relatively few cases of extreme oil abundance. Importantly, Parcero and Papyrakis (2016) also show that underreporting the inequality data is associated with oil-richness making cross-country findings less reliable. Kim and Lin (2018) deal with another inherent problem of cross-country settings in the form of cross-sectional dependence and common correlated effects. They estimate the effect of oil production and rents on income inequality using traditional econometric techniques and find it significant and positive. However, accounting for the common shocks and cross-sectional dependence, the results become significantly negative, i.e., oil abundance and oil dependence reduce income inequality. Interestingly, they also show that the negative effect is persistent both in ethnically polarized and homogenous nations, contradicting the nonlinear effect claimed by Fum and Hodler (2010).

Most recently, Davis (2020) uses a large sample of countries and obtains results that are similar to Alexeev and Conrad (2005). Namely, mining production tends to be associated with higher income inequality, while oil output does not increase inequality and might even somewhat decrease it. While cross-country studies provide important insights, they often suffer from unobserved heterogeneity, omitted variable bias, and uneven data quality. Parcero and Papyrakis (2016) and Van der Ploeg (2011) argue that researchers should exploit the variations within single countries for more reliable findings.

To our knowledge, there are only two papers that attempt to test the relationship between oil rents and income inequality at the subnational level. Howie and Atakhanova (2014) estimate the uniform effect of global oil prices on the quarterly measures of inequality in all of 16 provinces of Kazakhstan and find a strong negative relationship. Note, however, that their estimation strategy does not account for regional differences in oil endowments across Kazakh provinces, meaning that their measure of the oil price shock is identical across the country. This implies that the authors

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[8] In particular, Boschini et al. (2007) argue that resources may differ in the degree to which rents associated with them can be captured.

are unable to disentangle the impact of changes in oil prices from other common time-varying shocks such as exchange rate fluctuations, interest rates, and inflation. Moreover, the regressions suffer from a potentially significant spatial correlation of errors. In effect, given the commonality of oil prices for all provinces, Howie and Atakhanova's (2014) estimates of the impact of oil price changes on inequality in Kazakhstan are based on the number of observations corresponding to the number of quarters (i.e., at most 36), although the standard errors in their regressions are apparently calculated using the total number (more than 400) data points. As a result, the confidence levels of oil price coefficient estimates are almost certainly greatly overestimated.

The paper by Buccellato and Mickiewicz (2009) is closer to our work as it looks at subnational inequality and oil in Russia. The authors find a positive relationship using oil production in tons per capita as a proxy for oil rents. Their measure is likely to be endogenous to the production decisions and can correlate with omitted variables. In addition, the production volume does not necessarily correspond to the size of rents, which depend on both physical output and prices. We argue that a better strategy for establishing causality is using exogenous variation in global oil prices weighted by the average share of oil output in the gross regional product (GRP) over a period of time (e.g., Goderis and Malone, 2011, Bruckner et al., 2012) as we do in our paper. In fact, using the time frame of Buccellato and Mickiewicz (2009), we show that the plausibly exogenous oil price shocks do not by themselves increase subnational inequality.

Our paper also contributes to a broader empirical literature on the consequences of oil abundance or oil dependence. Recent resource curse literature often takes a nuanced view arguing that whether natural resources are a curse or a blessing depends largely on the quality of institutions. For example, oil wealth tends to stifle media freedom in authoritarian countries (Egorov et al., 2009) and hinder financial development in countries with poor institutions (Bhattacharyya and Hodler, 2014). In a survey of the resource curse literature, Badeeb et al. (2017) note that "in some countries, the windfall of resource revenues increases the power of elites, who have the capacity to widen income inequalities," although they do not provide references to studies demonstrating this effect. We show that an increase in oil rents accruing to the government budget tends to exacerbate inequality in a corrupt environment.

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# 3 Data

## 3.1 Oil windfalls

Russia, one of the largest global oil producers, exports over two billion barrels of oil annually,[9] thus, even a small increment in international oil prices leads to an ample windfall in oil rents. Most of this windfall accrues to the budget system. Zakharova and Goldsworthy (2010) estimate that an additional dollar in oil exports results in up to 90 cents in tax revenue.[10] Alexeev and Conrad (2009b) provide a similar observation about the Russian fiscal policy, calculating that over 90 percent of the net present value of oil deposits is appropriated by the government, provided that oil is exported. The prevalence of oil tax revenues relative to the other rents associated with oil and their high volatility due to the exogenous changes in international oil prices are the two factors that make Russia an ideal setting for studying the effects of this type of oil rent.

Our identification strategy takes advantage of the geographical and intertemporal variation in oil tax rents in Russia. While almost a third of Russian regions produce oil, the quantities vary significantly, as Figure 1 illustrates by plotting the average subnational production of oil between 2005 and 2018, measured in tons per capita. Between 2005 and 2009, the decentralized tax revenue system assigned five percent of the oil extraction tax to the budgets of the regions where oil was extracted. This system was discontinued in 2010, with only a partial compensation of the foregone regional revenues by 60, 40, and 20 percent in 2010, 2011, and 2012 respectively. This policy change allows for comparing two distinct time intervals: with oil tax revenue sharing (2005-2012) and without (2013-2018) and constitutes an important part of our identification strategy. If changes in oil rents accruing to regional governments affect income inequality in the regions, we may find a relationship between oil rents and inequality during 2005-2012, but we expect to find no similar effect after the oil tax revenue sharing ended in 2012. This approach is valid because the two periods were otherwise comparable in terms of several economic parameters: both periods contain only one occurrence of economic downturns, resulting in similar increases in unemployment and inflation, as we show in Figures A1 and A2 in Appendix A. Note that the second period is slightly more economically stable. Also, although the second period includes foreign economic sanctions and

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[9] BP Statistical Review of World Energy July 2021 (http://www.bp.com/statisticalreview).
[10] This is conditional on oil prices being above 25 dollars per barrel.

Russian countersanctions that followed the annexation of Crimea and the Ukrainian conflict, the established literature typically finds their economic and political impact to be insubstantial (e.g., Ashford, 2016; Dreger et al., 2016).

Figure 1: Geographic distribution of oil extraction in Russia in 2005-2018



Source: Alexeev & Zakharov based on data from the Russian Federal State Statistics Service

Another crucial characteristic of the Russian oil extraction tax is its formula that calculates the oil tax revenue proportionally to the international oil price (Alexeev and Conrad, 2009b). This enables us to use international oil price shocks as an exogenous component of the subnational oil windfalls.

We follow the standard approach to compute the oil price shocks (e.g., Brückner et al., 2012; Collier and Goderis, 2012; Brückner and Gradstein, 2015). The regional oil price shock (*OPS*) is calculated as an annual log-difference of the international oil price converted to constant Russian rubles multiplied by the average ratio of oil output to GRP over the time span covered by our regressions:[11]

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[11] Using the average ratio of oil output to regional GDP over a period of time as a weight to calculate the size of the oil price shock addresses the potential endogeneity between the dependent variable in the regressions and oil production decisions in any given year. However, one may argue that such weight does not adequately reflect the current oil price shock, which is determined by both price and quantity changes.Our results remain qualitatively the same when we measure the oil price shock using the product of the log-change in oil price and the ratio of the current value of oil output to GRP. The results of regression using this measure of the oil price shock are available upon request.

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$$OPS_{it} = \Delta \ln OilPrice_t \times \frac{1}{T} \sum\nolimits_{t=1}^{T} \frac{OilProduced_{it} \times 100}{GRP_{it}}, \qquad (1)$$

where $i$ and $t$ index regions and years, respectively; $\Delta\ln$ indicates the log-change between years $t$ and $t-1$; $OilPrice$ is the price of a barrel of Brent Crude oil (UNCTAD database) converted to rubles using the average Dollar-Ruble exchange rate in the given year; $OilProduced$ is the value of oil extracted in region $i$ in year $t$ evaluated using $OilPrice$; $GRP$ is the gross regional product expressed in constant rubles in region $i$ in year $t$.[12] Because the foregone oil tax revenues were only partially compensated during three years after the tax-sharing was discontinued in 2010, we adjusted the OPS in these years by the factors 0.6, 0.4, and 0.2 for 2010, 2011, and 2012, respectively. Appendix A, Figure A3 illustrates the dynamics of international oil prices and the amplitude of oil price shocks by years over the studied period, demonstrating that the pattern of oil price shocks was substantially similar in 2005-2012 and 2013-2018.

Our data cover all the Russian regions, excluding the cities of Moscow and Saint Petersburg as potential outliers since they profit from hosting the headquarters of the major oil companies and the Chechnya Republic due to local military conflict and data availability.[13] Overall, 34 of the 79 regions in our sample have produced at least some amount of oil during the period 2005-2018.

## 3.2 Rent capture

Our central hypothesis implies that the rents accruing to the government are particularly vulnerable to capture by the politically connected elites. We proxy the risk of capture by the dynamics of corruption in the region: the more common is the misuse of public office for private gain, the more likely is the misappropriation of the oil windfalls.
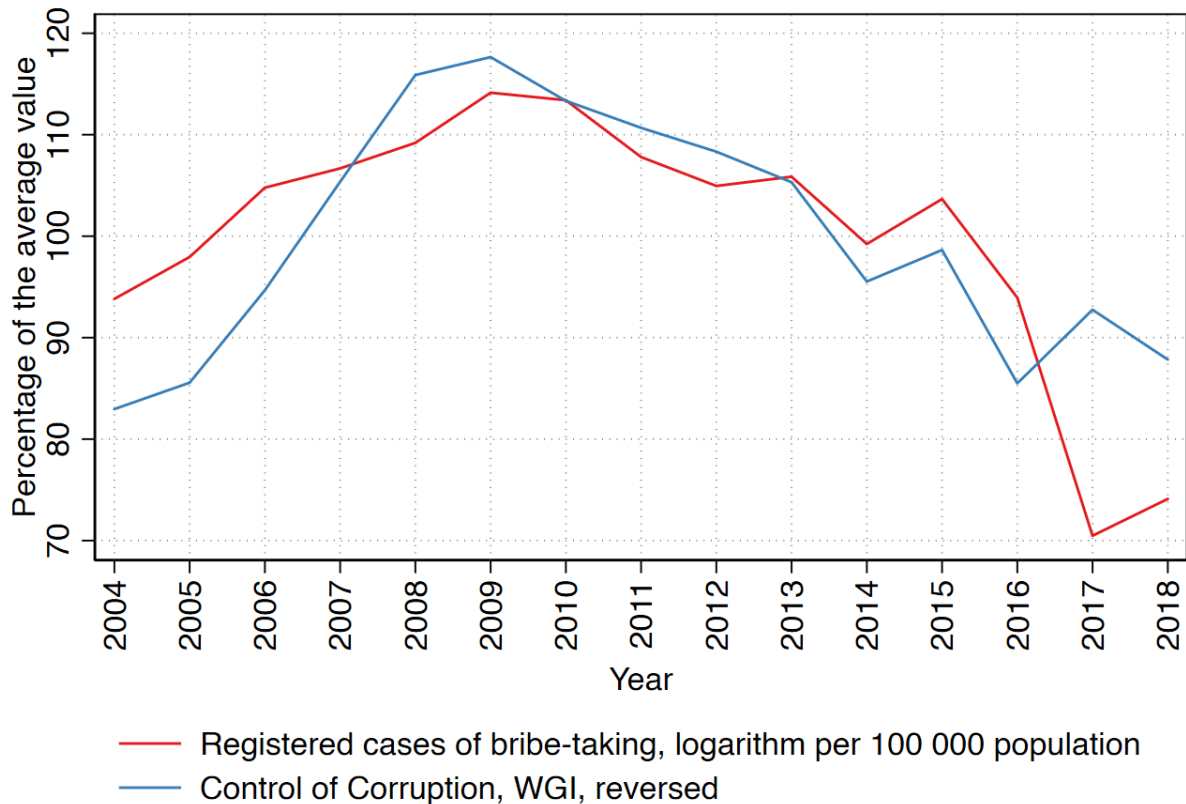
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[12] The first differencing of oil prices is necessary to satisfy the stationarity condition (e.g., Brückner et al., 2012). The GRP data are from Russia's Federal State Statistical Service (FSSS).

[13] The Nenets Autonomous Okrug is aggregated with the Arkhangelsk Oblast because of its minuscule population size. The Khanty-Mansi Autonomous Okrug and Yamalo-Nenets Autonomous Okrug are included in the dataset separately. We do not consider annexed Ukrainian regions – Sevastopol and the Crimea Republic as a part of Russia.

For our primary measure of corruption, we employ a recent dataset on the regional bribe-taking incidents registered by police authorities that was introduced by Schulze et al. (2016). Bribery is the most common type of corruption (Ryvkin and Serra 2012). Article 290 of the Russian Criminal Code, under which the bribe-taking is penalized, defines this crime as "acceptance of money, securities, or other valuables by a public official (personally or through an intermediary) for his or her performance (action or inaction) for the benefit of a giver or an affiliated person if such action implies that the public official exploits his or her position or authority or installs patronage". We use the registration of bribe-taking cases and not the convictions for three main reasons identified in Schulze et al. (2016). First, crime registration is the initial stage of criminal procedure that occurs within three days after the crime is reported or exposed, while a conviction takes significantly longer; therefore, registrations relate to the timing of the crime much more precisely. Second, if the juridical process is plagued with corruption, that would bias the conviction data, but not the data based on the registration by the police. Finally, corruption can still affect the registration process but only marginally since registration does not impose severe consequences on the suspects and can be further obstructed at any time during the process. Meanwhile, non-registration of a crime can be penalized by law if the victim reports to the higher law-enforcement authorities (Schulze et al., 2016).

We assume that the variation across time represents the actual dynamics of local corruption even if the registration of bribe-taking is affected by the region-specific characteristics (e.g., the quality of the police, law-enforcement resources, and priorities). Indeed, when we compare the overall time-trend in the registered cases of bribe-taking (a logarithm of the overall number of cases per 100 000 population) with a reversed control of corruption indicator from the World Bank, we observe a very similar pattern as shown in Figure 2. Therefore, we use first-differencing to construct our corruption variable ($\Delta$Bribes) as a log-difference of bribe-taking cases registered per 100 000 population.

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Figure 2: Dynamics of corruption in Russia



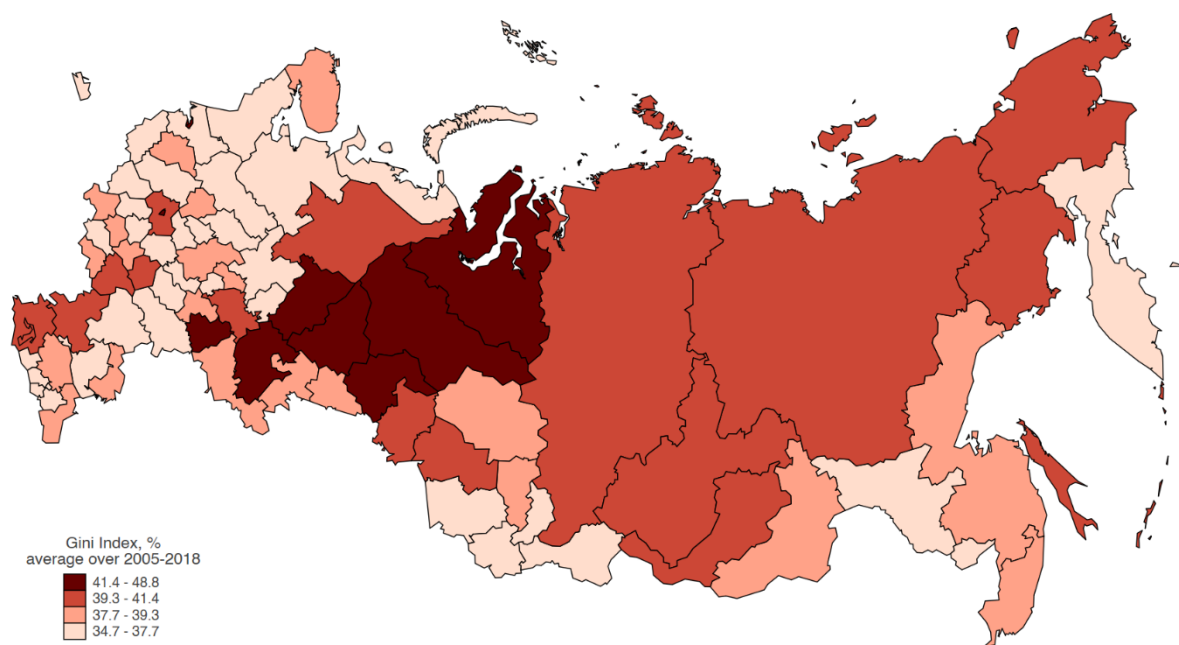Source: Alexeev & Zakharov based on data from the Russian Ministry of Internal Affairs and the World Bank

While the bribery data is probably the best available time-variant proxy for corruption in the Russian regions, no measure is perfect. Following the standard practice in corruption research, we introduce an alternative corruption measure to check the robustness of our results.

Our alternative measure is the annual number of embezzlement incidents registered by the law enforcement authorities obtained from the Ministry of Internal Affairs of the Russian Federation. It is similar to the bribery proxy but covers a different crime (Article 160 of the Russian Criminal Code), which also indicates the risk of misappropriation. The embezzlement cases are positively correlated with incidents of bribe-taking in Russian regions, but unlike bribe-taking, embezzlement crime is not committed exclusively by public officials and, therefore, it is not a corruption measure *per se*. The data on embezzlement also covers a shorter period of time: 2006-2018. The variable (ΔEmbezzlement) is constructed as a log-difference of registered embezzlement cases per 100 000 population.

## 3.3 Income inequality

Russia is one of the most unequal countries in the world (Novokmet et al., 2018), but inequality varies significantly across its subnational regions. Figure 3 maps the regional data on income inequality averaged over 2005-2018. Income inequality is measured by the Gini index in percentage published by Federal State Statistical Service (FSSS). One can observe that the oil-rich regions are also more unequal on average.

Figure 3: Income inequality in Russian regions, 2005-2018



Gini Index, %
average over 2005-2018
41.4 - 48.8
39.3 - 41.4
37.7 - 39.3
34.7 - 37.7

Source: Alexeev & Zakharov based on data from the Russian Federal State Statistics Service

Our regression analysis employs two sets of inequality measures provided by the FSSS: the Gini coefficient (%) and the share of income by quintile population groups. The methodology for calculating all of the inequality measures did not change during our period of investigation.

Additionally, we employ the data on the composition of income by sources: salary, entrepreneurship, social transfers, property income, and other income. These data are also available from the FSSS. The average composition of income in the regions over 2005-2018 is presented in Figure 4. The main income source for the Russian population is the official salary, with the second and third places shared by income from social transfers and other income, while income from self-employment and entrepreneurship and income from property come fourth and fifth, respectively. We are mostly interested in the category of other income calculated as a difference between overall

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income on the one hand and expenditures and savings on the other. This category includes income from sales of foreign currencies, incoming money transfers, and hidden income (e.g., legal but undeclared income or illegal income). Gorodnichenko and Sabirianova Peter (2007) link this income category to illegal income from corruption. Our hypothesis predicts that oil windfalls captured in a corrupt environment enrich the elites mostly through other income.

Figure 4: Composition of household income by sources, 2005-2018
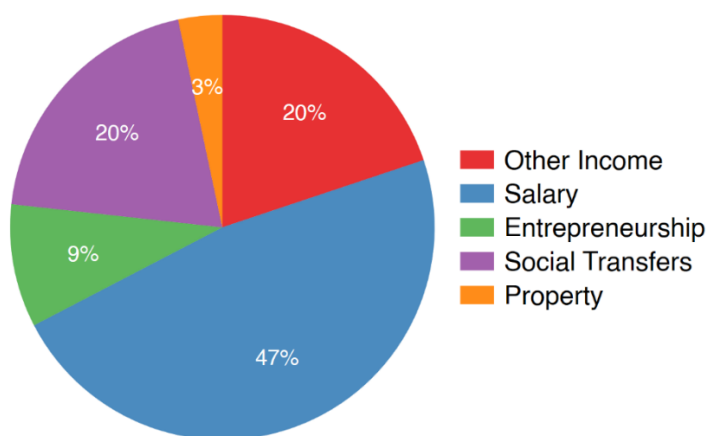


Source: Alexeev & Zakharov based on data from the Russian Federal State Statistics Service

## 3.4 Protests

The data on regional incidents of protests come from Lankina and Tertytchnaya (2020) and cover all Russian regions for 2007-2016. Lankina and Tertytchnaya (2020) assemble the dataset using the information from a crowd-sourced online platform (*namarsh.ru)* that aggregates reports on the various public forms of citizen protest collected by journalists, activists, and ordinary citizens. All entries were verified by human coders to include only events of public gathering (one or more participants), at the same time excluding petitions, entries that describe only intentions to protest, and political activities organized by the state authorities and pro-governmetal organizations (e.g., by pro-governmental youth movements or the ruling party, United Russia). The protest events are defined as "public expression of dissent or critique" and are recorded in the dataset at the level of an event-day-location. Therefore, for our analysis, we aggregate protests by region-year, divide by

millions of regional population and take the logarithm of this number plus one. The dataset contains almost six thousand incidents of protests, with 3549 (61%) of them held outside of Moscow or St. Petersburg (i.e., in the regions under our investigation).

The main advantage of the dataset for our purposes is the classification of the protests by agenda, as it allows us to look specifically at the inequality-related protests. Lankina and Tertytchnaya (2020) identify six categories of protests: 1) political protests include anti-government and anti-regime protests with broad agendas; 2) economic protests are against economic policies, such as worker-rights regulations or wage arrears; 3) legal protests are against unpopular legislation and illegal acts by state and private entities; 4) cultural protests mainly oppose the deconstruction of monuments and historically valuable sites; 5) environmental protests are about environmental issues; 6) social protests aim to further the interests of particularly vulnerable social groups, including the poor.[14] We are most interested in the latter category as the most relevant for the rising inequality: the economically disadvantaged groups of the society should be especially intolerant to the rising inequality from appropriating oil rents that could be instead allocated to redistribution. Interestingly, this is the second-largest category after political protests, and it accounts for nearly one in four protests. Figure 5 illustrates the compositions of protest activities in the Russian regions by category.

Figure 5: Composition of protest incidents by category, 2007-2016
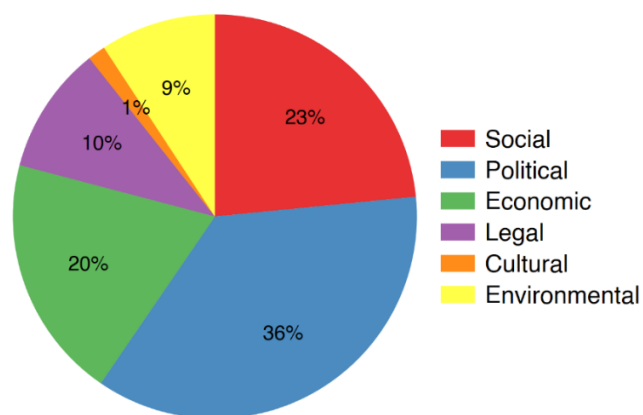


Source: Alexeev & Zakharov based on data from the Lankina and Tertytchnaya (2020)

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[14] The same protest event can be simultaneously assigned to several categories.

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Another worthy observation from the dataset is that the frequency of social protests is almost twice as high for oil-producing regions, where, as we observed earlier, both income inequality and average incomes are substantially higher than in non-oil-producing Figure 6 demonstrates the annual dynamics of social protests in regions with and without oil production. The illustration suggests that social protests were much more common when oil tax revenues were shared with regional budgets (2007-2012), and their frequency converged to the level of regions without oil production after 2011. This pattern indicates a potential link between protests and oil tax rents.

Figure 6: Dynamics of social protest in regions with and without oil production
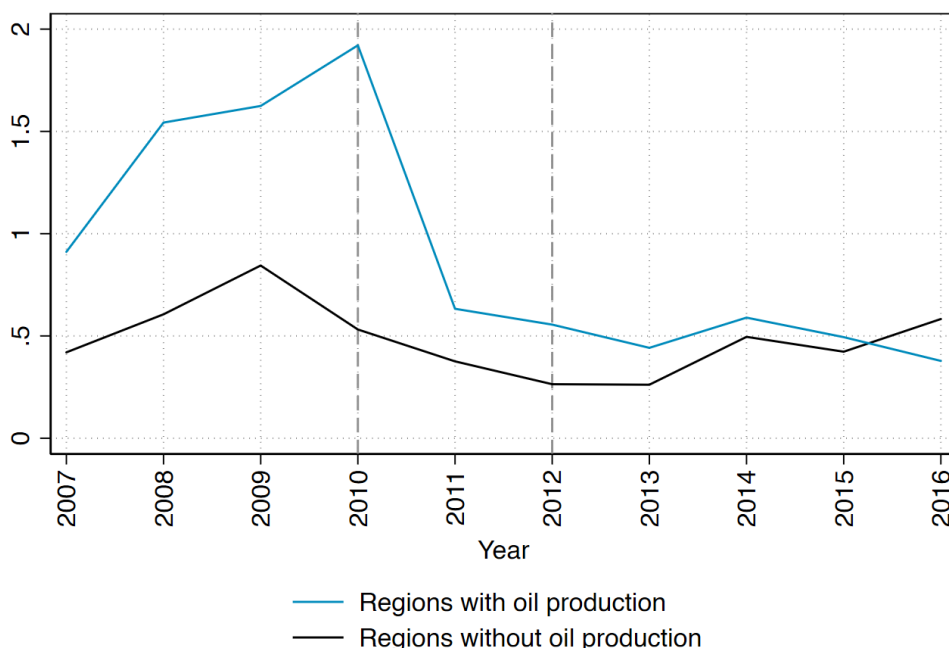


Source: Alexeev & Zakharov based on data from the Lankina and Tertytchnaya (2020)

While we expect an increase in the frequency of social protests due to rising inequality, we note that protests activities in non-democratic countries like Russia are generally costly to participants. However, due to the high heterogeneity in the regional degree of non-democracy in Russian regions, these costs vary depending on the local political regime.[15] Tertytchnaya and Lankina (2020), who studied the frequency of subnational protests in Russia incited by the exposure of electoral fraud at

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[15] The heterogeneity in subnational political regimes, particularly during the period under study, is a well-established feature of the Russian political system that has been studied widely in political science: Obydenkova and Libman (2012) provide a short summary.

national parliamentary elections in 2011, address this issue by employing a commonly used index of Russian subnational democracy developed by Petrov and Titkov (2013) on behalf of the Moscow Carnegie Center.[16] They show that protests due to electoral fraud occurred more often in relatively more democratic regions. We follow their approach and employ the democracy index from Petrov and Titkov (2013) to test for a heterogeneous effect of income inequality on protests. The index is based on expert opinions about ten dimensions of democracy in each region in the previous five years (election quality, regional political system, pluralism, quality of local elites, openness, economic liberalization, media independence, control of corruption, civil society, and local self-governance). Each dimension is graded on a scale of 1 to 5, where 5 is the most pro-democratic evaluation.[17] Hence, the overall index ranges from 10 to 50, with the higher value indicating a relatively more democratic region. We employ the regional democracy index for 2006 – the first year of the protest dataset – to avoid any potential interaction between oil rents and democracy, and then we split all regions in our sample into halves using the median value of democracy.[18] Unsurprisingly, we observe that, on average, protests are 80% more frequent in relatively democratic regions; social protests are almost 90% more frequent. We expect the effect of inequality on protests to be more pronounced in regions where public protests are a more common activity.

In addition to the data described above, we use some regional-level variables in the robustness checks. All of those variables are obtained from the FSSS. The descriptive statistics for the data used in our regressions are shown in Table A1 in Appendix A.

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[16] The Moscow Carnegie Center is a reputable think tank affiliated with the Carnegie Endowment for International Peace operating in Russia since 1994.

[17] The index has been widely used in economics (e.g., Libman, 2012; Berkowitz et al., 2014; Belmonte and Rochlitz, 2019) and comparative political science (e.g., Freinkman and Plekhanov, 2009; Beazer, 2012; Buckley and Reuter, 2019).

[18] The number of oil-producing regions is equal in both subgroups. The median value of the democracy index is 29.

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# 4 Estimation strategy and the results

## 4.1 The effect on income inequality

Our estimation strategy follows a standard approach in the literature (e.g., Brückner et al. 2012, Collier and Goderis 2012, Brückner and Gradstein 2015). We first estimate the linear effect of oil price shocks in a dynamic panel model using the system GMM as in the following equation:

$$\Delta y_{it} = \alpha_0 \times \Delta y_{it-1} + \alpha_1 \times OPS_{it} + \mu_i + \tau_t + \varepsilon_{it} , \qquad (2)$$

where $y_{it}$ is the value of the dependent variable in region $i$ and year $t$, $OPS_{it}$ is the oil price shock, and, $\mu_i$ and $\tau_t$ denote region and time fixed effects, respectively; $\Delta$ denotes the change of the variables between years $t$ and $t-1$. Note that the oil price shock is necessarily the first-differenced variable, and, therefore, it is appropriate to first-difference the other variables. The system GMM specification allows for the dynamic nature of the panel and can also account for potential endogeneities in the data.

To test the nonlinear effect of oil price shock under different corruption levels, we modify equation (2) to include the corruption variable and the interaction term:

$$\Delta y_{it} = \beta_0 \times \Delta y_{it-1} + \beta_1 \times OPS_{it} + \beta_2 \times \Delta Corr_{it} + \beta_3 \times OPS_{it} \times \Delta Corr_{it} + \mu_i + \tau_t$$
$$+ \varepsilon_{it} , \qquad (3)$$

where $Corr_{it}$ is our measure of corruption. While our baseline estimation is a system GMM, we also provide OLS FE estimates as a robustness check later in our analysis, as in Brückner et al. (2012).

We estimate both specifications for the period when oil extraction taxes were shared with local budgets (2005-2012), where we expect rent-seeking to mediate the effect of oil windfalls on inequality, and the period after the tax-sharing regime was discontinued (2013-2018), where the windfall effect should disappear.

Table 1 presents the results for inequality as measured by the Gini coefficient. Columns 1 and 2 show that oil price shocks have no impact on inequality under the oil tax sharing regime (2005-2012) whether we control for corruption environment or not. However, the interaction term in column 3 is positive and highly significant, confirming our hypothesis that the rent capture

mediates the inequality-raising effect of windfalls on the regional budgets.[19] The effect is robust to using our alternative measure of corruption – registered incidents of embezzlement (column 4) and OLS FE estimation instead of GMM (column 5). The magnitudes of the effect are very similar across all specifications. To illustrate them, we use the OLS FE estimation to compute the average marginal effect of oil price shocks conditional on the log-change in bribery incidents and present them in Figure 7. The marginal effects are positive and significant at a 5% level, starting slightly under 0.5 log change in corruption.[20]

Figure 7: Marginal effects of oil price shocks on the Gini coefficient (2005-2012)



Source: Alexeev & Zakharov, own calculations
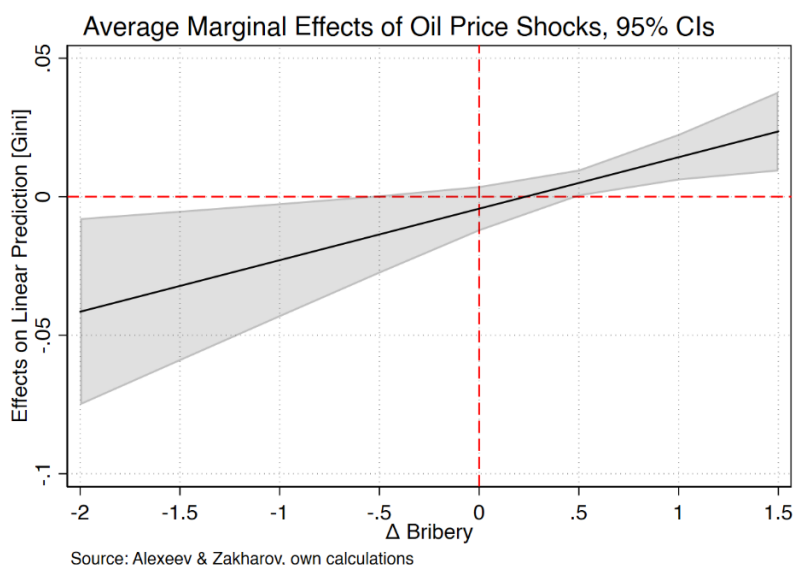
What happens when the oil tax revenue is no longer shared with regional budgets? We run the same regressions for the period of 2013-2018 and report the results in columns 6-8. While the linear term of oil price shocks is again not significant in estimation with and without the inclusion of the cor-

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[19] We also note that system GMM regressions satisfy the standard AR(2) and overidentification tests, and the number of instruments is comfortably below the number of groups (regions), implying the absence of instrument proliferation problem.

[20] Marginal effects and their standard errors are calculated using Stata's *margins* command. We estimate marginal effects only for OLS FE regressions because calculating marginal effects and their standard errors for system GMM specification is problematic due to the presence of the lagged dependent variable among the regressors.

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ruption variable (columns 6-7), the interaction term between oil price shocks and corruption becomes indistinguishable from zero (column 8).[21] This result supports our hypothesis that the non-linear effect is driven by the incremental oil tax revenue accrued to the local budgets.

[21] These results are also robust to the inclusion of the alternative measure of corruption and OLS FE as an alternative estimation technique, and they are available upon request.

Table 1: Oil price shocks, corruption, and income inequality (2002-2018)

(dependent variable: Δ Gini coefficient)

| Time period: | 2005-2012 | | | | | 2013-2018 | | | 2002-2004 |
|---|---|---|---|---|---|---|---|---|---|
| Estimation method | System GMM | | | | OLS FE | System GMM | | | |
|  | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Oil price shock | -0.0002 | -0.0008 | -0.0089** | -0.0011 | -0.0043 | 0.0009 | -0.0014 | 0.0001 | 0.0027 |
|  | (0.965) | (0.823) | (0.012) | (0.786) | (0.301) | (0.871) | (0.801) | (0.984) | (0.866) |
| ΔBribes |  | 0.0022 | 0.0044 |  | 0.0149 |  | 0.3081 | 0.0038 |  |
|  |  | (0.972) | (0.944) |  | (0.853) |  | (0.189) | (0.963) |  |
| **OPS× ΔBribes** |  |  | **0.0288\*\*\*** |  | **0.0186\*\*\*** |  |  | **-0.0050** |  |
|  |  |  | **(0.000)** |  | **(0.008)** |  |  | **(0.694)** |  |
| ΔEmbezzlement |  |  |  | -0.0704 |  |  |  |  |  |
|  |  |  |  | (0.270) |  |  |  |  |  |
| **OPS× ΔEmbezzlement** |  |  |  | **0.0370\*\*\*** |  |  |  |  |  |
|  |  |  |  | **(0.006)** |  |  |  |  |  |
| Lagged dep. variable | 0.0238 | -0.0466 | -0.0379 | -0.0631 |  | -0.6248*** | -0.6480*** | -0.5380*** | 0.1219 |
|  | (0.656) | (0.365) | (0.466) | (0.288) |  | (0.000) | (0.000) | (0.000) | (0.457) |
| Observations | 632 | 632 | 632 | 474 | 632 | 474 | 474 | 474 | 237 |
| Panels | 79 | 79 | 79 | 79 | 79 | 79 | 79 | 79 | 79 |
| AR(2) p-value | 0.79 | 0.87 | 0.71 | 0.71 |  | 0.11 | 0.27 | 0.45 | . |
| Hansen test (p-value) | 0.184 | 0.160 | 0.318 | 0.566 |  | 0.334 | 0.145 | 0.204 | 0.316 |
| Instruments | 26 | 38 | 52 | 38 |  | 20 | 32 | 44 | 8 |
| R squared |  |  |  |  | 0.33 |  |  |  |  |

Notes: p-values in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects. The regression for 2013-2018 in Column 6 includes a second lag of the dependent variable in order to make it pass the AR(2) test. The statistical significance of the coefficients is not affected by the addition of this variable.

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Statistical significance of the regression coefficients is important, but it does not directly convey the economic importance of the effect of oil price shocks on income distribution. This effect depends on the degree of corruption as well as on the role of oil in the regional economy. On average, in oil-producing regions, one standard deviation of the positive oil price shock combined with the largest observed increase in corruption would result in about a quarter of the standard deviation increase in the annual change of the Gini coefficient. In a region with a smaller change in corruption, the impact of the oil price shock on inequality would be somewhat weaker.

Note that the standard deviation of the oil price shock differs across oil-producing regions because it is proportional to the ratio of the mean value of oil output and regional GDP over 2005-2012. Therefore, given the same change in corruption, the effect of oil price changes on inequality would be greater in regions with a larger share of oil output in the economy. For example, the standard deviation of the oil price shock for the top five oil producers in our sample is twice that of the standard deviation of the price shock for all oil producers. Therefore, given our point estimates, the impact of oil price shocks and corruption on income inequality would also be twice as large in the top oil producers. It is also important to mention here that the "within" R-squared in all but one of our inequality regressions is about 0.3, which is perhaps surprisingly large given the small number of explanatory variables.

Finally, we look at the period under investigation by Buccellato and Mickiewicz (2009). In their estimation strategy, the authors rely exclusively on the physical measure of oil extraction that was likely to be endogenous to the production decisions and find a positive effect of oil on inequality. We replicate their regressions in column (9) using the exogenous variation in the oil windfalls due to the changes in international prices and find no significant effect. We are unable to investigate the nonlinear effect in this earlier period due to the lack of data on corruption.

Our main regressions do not include any control variables besides time and region fixed effects. In the robustness checks section, we add (first differences of) the unemployment rate, the share of the working-age population, the degree of urbanization, a measure of educational attainment of the population, and the logarithm of regional per capita GDP – variables that change over time and have been identified as potentially important in the literature on the factors influencing

income distribution.[22] In the same section, we also perform placebo tests by using overall crime and murder rates instead of measures of corruption as indicators of governance quality. We note that because we use the data on regions within the same country as well as time fixed effects, we do not include factors that are likely to affect all regions similarly, such as those related to monetary policy or international trade. As we will demonstrate later, both the robustness checks and placebo tests confirm our original results.

While our main finding is the mediating effect of corruption on the relationship between oil rents accruing to the government and income inequality, we are also interested in who gains the most from the captured rents. For this purpose, we employ the data on income distribution by quintiles, using the changes in their shares as the dependent variables.

As the results in Table 2 demonstrate (columns 1-5), positive oil price shocks combined with increased corruption are associated with declining shares of the first four income quintiles and a rising share of the fifth quintile. As before, this effect is present only during the period when oil tax revenues were shared with the regions, and it disappears in the post-2012 period (column 8). We present the graph of marginal effects of bribery on the share of the fifth income quintile calculated based on OLS FE regression (column 7) in Figure 8. Note that these marginal effects become positive and significant at the 5% level, almost exactly at the same value of bribery as in the Gini coefficient regression. As in the case of income inequality, the results for income quintiles based on the bribery measure are confirmed by the regression using the embezzlement variable (column 6). Finally, the interaction of oil price shock and corruption in a period without oil tax sharing (column 8) is no longer significant.[23]

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[22] See Kuznets (1955), Barro (2000), and more recently Hacibedel et al. (2019) for the effects of GDP. Nielsen and Alderson (1995, 1997), Higgins and Williamson (2002), and Bergh and Fink (2008) are examples of the studies that addressed the impact of urbanization, age structure, and education, respectively. Unemployment is one of the "robust" drivers of inequality identified by Furceri and Ostry (2019).

[23] While the direct term is statistically significant at the 10% level, its significance disappears in the linear estimation as in equation (2). Results are available upon request.

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Table 2: Oil price shocks and income quintiles (2005-2018)

| Time period: | 2005-2012 | | | | | | | 2013-2018 |
|---|---|---|---|---|---|---|---|---|
| Estimation method: | System GMM | | | | | | OLS FE | System GMM |
| Dependent variable:<br>Independent variables: | ΔShare of 1st income quintile | ΔShare of 2d income quintile | ΔShare of 3d income quintile | ΔShare of 4th income quintile | ΔShare of 5th income quintile | ΔShare of 5th income quintile | ΔShare of 5th income quintile | ΔShare of 5th income quintile |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Oil price shock | 0.0022** | 0.0023*** | 0.0021** | 0.0007 | -0.0076*** | -0.0008 | -0.0039 | 0.0068* |
| | (0.037) | (0.009) | (0.022) | (0.266) | (0.009) | (0.809) | (0.269) | (0.093) |
| ΔBribes | 0.0048 | 0.0043 | 0.0007 | -0.0028 | -0.0045 | | 0.0050 | -0.0240 |
| | (0.782) | (0.785) | (0.964) | (0.629) | (0.927) | | (0.934) | (0.679) |
| **OPS× ΔBribes** | **-0.0051***** | **-0.0080***** | **-0.0068***** | **-0.0046***** | **0.0247***** | | **0.0167***** | **0.0077** |
| | **(0.001)** | **(0.000)** | **(0.000)** | **(0.000)** | **(0.000)** | | **(0.004)** | **(0.315)** |
| ΔEmbezzlement | | | | | | -0.0756 | | |
| | | | | | | (0.136) | | |
| **OPS× ΔEmbezzlement** | | | | | | **0.0293***** | | |
| | | | | | | **(0.009)** | | |
| Lagged dep. variable | -0.0438 | -0.0357 | -0.0717 | -0.2641*** | -0.0410 | -0.0617 | | -0.0763 |
| | (0.396) | (0.425) | (0.138) | (0.000) | (0.408) | (0.313) | | (0.588) |
| Observations | 632 | 632 | 632 | 632 | 632 | 474 | 632 | 474 |
| Panels | 79 | 79 | 79 | 79 | 79 | 79 | 79 | 79 |
| AR(2) p-value | 0.43 | 0.37 | 0.77 | 0.35 | 0.76 | 0.61 | | 0.84 |
| Hansen test (p-value) | 0.430 | 0.395 | 0.484 | 0.492 | 0.437 | 0.526 | | 0.319 |
| Instruments | 52 | 52 | 52 | 52 | 52 | 38 | | 44 |
| R squared | | | | | | | 0.33 | |

Notes: p-values in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects. The regression for 2013-2018 in Column 8 includes a second lag of the dependent variable in order to make it pass the AR(2) test. The statistical significance of the coefficients is not affected by the addition of this variable.

Figure 8: Marginal effects of oil price shocks on the share of the 5th income quintile (2005-2012)



Source: Alexeev & Zakharov, own calculations
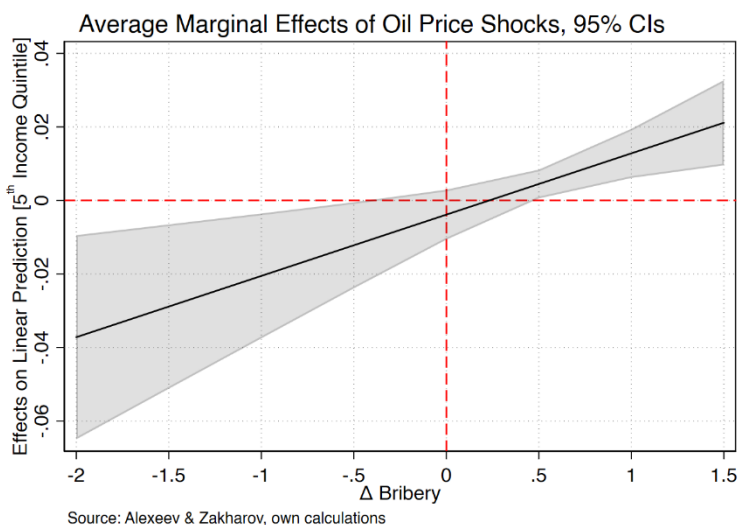
## 4.2. The effect on income composition

Having demonstrated the complementary impact of the changes in oil rents accruing to the regional budgets and corruption on income inequality, we conjecture that this impact might also be reflected in the changing shares of different types of income. Specifically, we expect that oil windfalls in the environment of rising corruption will be associated with an increasing share of "other income", which includes incomes derived from informal or illegal activities (Gorodnichenko and Sabirianova Peter 2007). To test this hypothesis, we estimate regression (3) using changes in the shares of different income sources as dependent variables. We find that the interaction between oil price shocks and changes in bribery is positively and statistically significantly associated with the change in the share of other income (see Table 3). The significance level is 1% in the system GMM specification (column 5) and 10% in the OLS FE regression (column 7). The interaction term between the oil price shock and embezzlement is also positive and statistically significant (column 6). As before, the effect cannot be detected after oil tax sharing was discontinued (column 8). In addition, the interaction term coefficient has a negative and significant sign in both the OLS FE and system GMM wage income regressions (column 1). The shares of income from entrepreneurship, social transfers, and property are not significantly affected by the oil rents fluctuations (columns 2).

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Table 3: Oil price shocks and sources of income (2005-2018)

| Time period: | 2005-2012 | | | | | | 2013-2018 | |
|---|---|---|---|---|---|---|---|---|
| Estimation method: | System GMM | | | | | | OLS FE | System GMM |
| Dependent variable: / Independent variables: | ΔShare of salaries | ΔShare of entrepre-neurial in-come | ΔShare of social transfers | ΔShare of income from prop-erty | ΔShare of other in-come | ΔShare of other in-come | ΔShare of other in-come | ΔShare of other in-come |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Oil price shock | 0.0095 | -0.0023 | 0.0017 | 0.0117 | -0.0111 | 0.0013 | -0.0000 | 0.0153 |
| | (0.676) | (0.806) | (0.801) | (0.166) | (0.637) | (0.967) | (1.000) | (0.398) |
| ΔBribes | -1.1593* | -0.1337 | -0.1955 | -0.1038 | 0.3813 | | 0.2803 | -0.1655 |
| | (0.091) | (0.330) | (0.182) | (0.522) | (0.307) | | (0.459) | (0.651) |
| **OPS× ΔBribes** | **-0.1351**** | **-0.0180** | **0.0046** | **0.0061** | **0.0861*****| | **0.0671*** | **0.0316** |
| | **(0.013)** | **(0.122)** | **(0.522)** | **(0.615)** | **(0.004)** | | **(0.072)** | **(0.655)** |
| ΔEmbezzlement | | | | | | -0.7299 | | |
| | | | | | | (0.168) | | |
| **OPS× ΔEmbezzlement** | | | | | | **0.1885**** | | |
| | | | | | | **(0.015)** | | |
| Lagged dep. variable | -0.0274 | 0.1443** | -0.0461 | -0.0211 | -0.0867 | -0.0245 | | -0.3018*** |
| | (0.881) | (0.049) | (0.491) | (0.761) | (0.117) | (0.682) | | (0.000) |
| Observations | 632 | 632 | 632 | 632 | 632 | 474 | 632 | 474 |
| Panels | 79 | 79 | 79 | 79 | 79 | 79 | 79 | 79 |
| AR(2) p-value | 0.24 | 0.77 | 0.24 | 0.26 | 0.19 | 0.92 | | 0.13 |
| Hansen test (p-value) | 0.144 | 0.371 | 0.184 | 0.332 | 0.185 | 0.299 | | 0.314 |
| Instruments | 48 | 71 | 52 | 52 | 52 | 38 | | 62 |
| R squared | | | | | | | 0.15 | |

Notes: p-values in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects. The regression for 2013-2018 in Column 8 includes a second lag of the dependent variable in order to make it pass the AR(2) test. The statistical significance of the coefficients is not affected by the addition of this variable.

Figure 9 shows the graph of marginal effects of the oil price shocks on the other income share conditional on corruption. Interestingly, these marginal effects become significant at the 5% level at almost exactly the same point as in Figures 7 and 8. We conclude that there is indirect evidence that rising corruption combined with increasing oil rents affects income inequality via increasing informal incomes.

Figure 9. Marginal effects of oil price shocks on the share of "other income" (2005-2012)



Source: Alexeev & Zakharov, own calculations
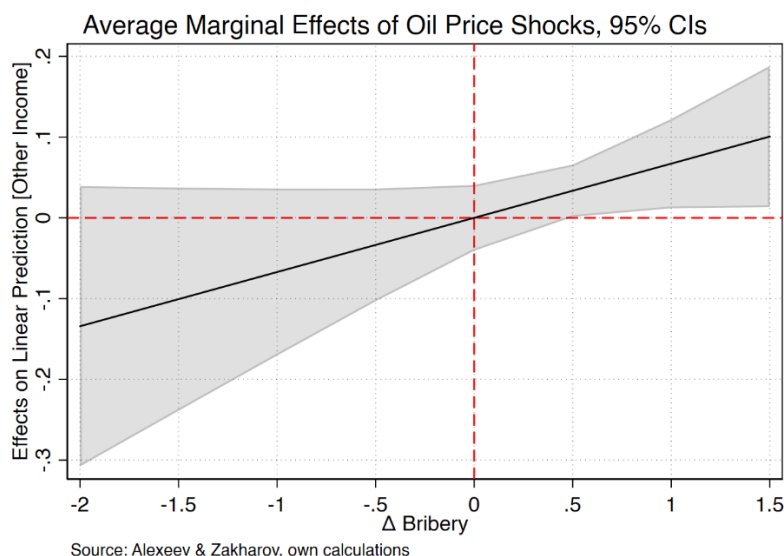
## 4.3 The effect on protests

This section investigates the political implications of rising inequality. We expect that income inequality caused by the capture of oil rents generates grievances among the population and potentially lead to more frequent protests – particularly the social protests aimed at furthering the interests of the poor. Earlier in Figure 6, we demonstrated that social protests took place in oil-producing regions significantly more often than in regions without oil endowment, but only in the period when the oil tax revenue was shared with the regional budget.

To determine the transmission mechanism from the impact of the oil price shocks and corruption on protest, we estimate a structural model using the instrumental variable (IV) approach. We instrument the $\Delta$Gini coefficient with oil price shocks, the bribery measure, and their interaction

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similarly to Table 1 where column 5 represents the first-stage regression.[24] Table 4 presents the results. First, we look at the overall frequency of protests and find a positive effect of increasing inequality at the 10% significance level only in the subgroup of relatively democratic regions (column 2), but not for the overall sample (column 1) and relatively undemocratic regions (column 3). The structural form allows for interpreting the magnitudes of the estimated local average effects in the relatively democratic regions: the change in inequality by one standard deviation of the first-differenced Gini explains 47 percent of the standard deviation in first-differenced protests per million population. According to our main hypothesis, the effect should be driven by the social protests, and, indeed, we find that instrumented inequality increases their frequency (column 4), and this effect is again driven exclusively by the relatively democratic regions (column 5). The magnitude of the effect in relatively democratic regions is very similar to the overall protest: one standard deviation change in ΔGini explains 46 percent of the standard deviation in social protests variable. Non-democratic regions might still accumulate grievances over the unfair redistribution of natural resources, but because protests are more costly for potential participants, we do not find the effect in this subgroup (column 6). Finally, Table 4 shows that inequality is not associated with protests in the period after oil tax revenue sharing was discontinued.[25] Table A2 in Appendix A demonstrates that this relationship cannot be found for other categories of protests.

For robustness purposes, Table A3 in Appendix A provides the reduced form estimation for our IV regressions as in Table 4. The reduced form confirms our main results for the social protests: the interaction of oil price shocks and corruption is a good predictor of social protests (at the 10% level), and this effect is larger and more statistically significant (at the 5% level) in the subgroup of relatively democratic regions, but absent among the relatively autocratic regions. As before, the period when the oil tax revenue was no longer shared with the regional budget shows no relationship between oil price shocks, corruption, and protests.

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[24] We obtain similar results when we instrument the change in income inequality only with the interaction term.

[25] Predictably, the F-statistics in the post-2012 period are very low. The results for the post-2012 period are presented only for democratic regions where the interaction effect was strongest previously, but we can report that no effect exists in non-democratic regions, and the full sample: results are available upon request.

Table 4: Oil windfalls and protest (IV regressions; second stage; 2007-2016)

| Time period: | 2007 - 2012 | | | | | | 2013 - 2016 | |
|---|---|---|---|---|---|---|---|---|
| Dependent variable: | ΔAll Protests | | | ΔSocial Protests | | | ΔAll Protests | ΔSocial Protests |
| | All regions | Democratic | Non-democratic | All regions | Democratic | Non-democratic | Democratic | Democratic |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| **ΔGini coefficient** | **1.9737** | **3.0648**[*] | **-0.4368** | **1.6987**[*] | **1.0972**[**] | **-0.5192** | **-0.9057** | **0.3720** |
| | **(0.229)** | **(0.060)** | **(0.752)** | **(0.093)** | **(0.046)** | **(0.506)** | **(0.917)** | **(0.913)** |
| Observations | 395 | 195 | 200 | 395 | 195 | 200 | 156 | 156 |
| Panels | 79 | 39 | 40 | 79 | 39 | 40 | 39 | 39 |
| F-statistic | 9.36 | 15.14 | 9.63 | 9.36 | 15.14 | 9.63 | 0.07 | 0.07 |
| Hansen Test (p-value) | 0.20 | 0.29 | 0.18 | 0.20 | 0.29 | 0.18 | 0.93 | 0.93 |

Notes: p-values in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects.

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# 5 Robustness checks

In this section, we perform additional robustness checks. First, we implement placebo tests by running regressions (3) using overall crime and murder rates instead of corruption measures as the conditioning variables. Crime rates and, particularly murders, can also serve as indicators of institutional quality (Cole and Gramajo, 2009;Popov, 2013), but they are not specific to rent-capturing. Therefore, we do not expect to have a consistently positive and statistically significant association between oil price shocks interacted with changes in institutional quality based either on overall crime rates or on murder rates. Second, we estimate regressions (3) with several time-variant regional control variables. In what follows, to save space, we show only the main "inequality" regressions, i.e., those for the 5th quintile and the Gini coefficient, as well as the system-GMM regressions for social protests.[26]

## 5.1. Placebo tests

Tables 5 and 6 present the results of the placebo tests where we use overall crime rates and murder rates instead of governance quality variables that are related to rent capture. None of the interaction terms between oil price shocks and overall crime or murder rates has a statistically significant coefficient, suggesting that it is precisely the rent-capturing property of corruption that drives our main results.

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[26] The other regressions (placebo tests for the first four income quintiles, placebo tests with additional control variables, regressions for sources of income besides "other", regressions for all protests with additional control variables, and the IV regressions for social protests with additional controls) confirm the robustness of our main results, although the IV regressions for social protests with additional control variables produce somewhat weaker statistical significance for the Gini coefficient for the more democratic regions. These regressions are available upon request.

Table 5: Placebo test with the number of overall registered crime (system GMM; 2005-2012)

| Dependent variable: | ΔShare of 5th income quintile | Δ Gini coefficient | ΔShare of other income | ΔSocial Protests (all regions) | ΔSocial Protests (democratic regions) | ΔSocial Protests (non-democratic regions) |
|---|---|---|---|---|---|---|
| Independent variables: | | | | | | |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Oil price shock | 0.0010 | 0.0010 | -0.0098 | -0.0075 | -0.0018 | 0.0406 |
| | (0.557) | (0.616) | (0.698) | (0.781) | (0.948) | (0.389) |
| ΔCrime | -0.3929 | -0.5246 | -1.8133 | -1.6907 | -0.7949 | -2.3617 |
| | (0.238) | (0.217) | (0.528) | (0.410) | (0.826) | (0.212) |
| **OPS× ΔCrime** | **-0.0086** | **-0.0125** | **-0.0739** | **0.0621** | **0.1698** | **-0.1650** |
| | **(0.820)** | **(0.786)** | **(0.705)** | **(0.874)** | **(0.694)** | **(0.743)** |
| Lagged dependent variable | 0.0108 | 0.0207 | -0.0517 | -0.0615 | 0.0588 | -0.3784*** |
| | (0.834) | (0.706) | (0.296) | (0.758) | (0.800) | (0.001) |
| Observations | 632 | 632 | 632 | 316 | 156 | 160 |
| Panels | 79 | 79 | 79 | 79 | 39 | 40 |
| AR(2) (p-value) | 0.94 | 0.92 | 0.17 | 0.20 | 0.36 | 0.16 |
| Hansen test (p-value) | 0.317 | 0.282 | 0.425 | 0.532 | 0.563 | 0.249 |
| Instruments | 58 | 58 | 58 | 25 | 25 | 25 |

Notes: p-values in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects.

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Table 6: Placebo test with the number of registered murders (system GMM; 2005-2012)

| Dependent variable: <br> Independent variables: | ΔShare of 5th income quintile | Δ Gini coefficient | ΔShare of other income | ΔSocial Protests (all regions) | ΔSocial Protests (democratic regions) | ΔSocial Protests (non-democratic regions) |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Oil price shock | -0.0013 | -0.0022 | -0.0202 | -0.0021 | -0.0065 | 0.0883[*] |
| | (0.394) | (0.253) | (0.506) | (0.895) | (0.742) | (0.053) |
| ΔMurder | 0.0143 | 0.0279 | 0.4614 | -0.3475 | -1.6370 | 0.0200 |
| | (0.902) | (0.854) | (0.583) | (0.577) | (0.315) | (0.958) |
| **OPS× ΔMurder** | **0.0138** | **0.0108** | **-0.1720** | **0.0828** | **0.0531** | **0.3287** |
| | **(0.384)** | **(0.557)** | **(0.251)** | **(0.369)** | **(0.462)** | **(0.225)** |
| Lagged dependent variable | 0.0348 | 0.0452 | -0.0596 | -0.0576 | 0.0994 | -0.3881[***] |
| | (0.490) | (0.401) | (0.210) | (0.777) | (0.672) | (0.001) |
| Observations | 632 | 632 | 632 | 316 | 156 | 160 |
| Panels | 79 | 79 | 79 | 79 | 39 | 40 |
| AR(2) (p-value) | 0.68 | 0.64 | 0.21 | 0.18 | 0.38 | 0.20 |
| Hansen test (p-value) | 0.148 | 0.157 | 0.737 | 0.551 | 0.187 | 0.314 |
| Instruments | 58 | 58 | 58 | 25 | 25 | 25 |

Notes: p-values in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects.

## 5.2. Additional controls

The regressions we have estimated so far did not have any control variables except for time and region fixed effects. This approach follows much of the literature on the effects of oil price shocks based on panel data (e.g., Arezki and Brückner 2012 and Brückner et al. 2012). However, although the oil price shocks are plausibly exogenous in the sense that they are not affected by the developments within regions, it might be appropriate to include other time-variant control variables that could be correlated with these price shocks and corruption while influencing inequality. Accordingly, we add the first-differenced unemployment rate, working-age population share, the degree of urbanization, educational attainment of the population, and the logarithm of regional per capita GDP in constant 2016 Russian rubles. We note, however, that with a possible exception of the unemployment rate, these variables change relatively slowly, and thus they can be problematic in first-differenced panel regressions. The resulting estimates are shown in Tables 7-8. The inclusion of these additional variables does not significantly affect our main results. The coefficients for the interaction term between oil price shocks and corruption have the same signs and similar or just slightly lower significance levels as in the corresponding regressions in Tables 2-4.

As expected, the unemployment variable is the only one among our control variables with a statistically significant coefficient, although it has different signs in the OLS FE and system-GMM regressions.[27] This inconsistency prevents us from making any firm conclusions about the association between unemployment and income inequality.

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[27] We do not report the OLS FE specifications with controls, but the results are available upon request.

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Table 7: Robustness check with additional control variables (system GMM; 2005-2012)

| Dependent variable: | ΔShare of 5th income quintile | ΔGini coeffi-cient | ΔShare of ot-her income | ΔSocial pro-tests (all regi-ons) | ΔSocial protests (democratic re-gions) | ΔSocial protests (non-demo-cratic regions) |
|---|---|---|---|---|---|---|
| Independent variables: | | | | | | |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Oil price shock | -0.0067** | -0.0073** | -0.0140 | -0.0337 | -0.0484* | 0.0876** |
| | (0.011) | (0.022) | (0.554) | (0.269) | (0.076) | (0.029) |
| ΔBribes | 0.0068 | 0.0274 | 0.1420 | -0.2258 | -0.4511 | 0.1807 |
| | (0.893) | (0.672) | (0.756) | (0.224) | (0.127) | (0.423) |
| **OPS× ΔBribes** | **0.0238***** | **0.0271***** | **0.1573***** | **0.0673** | **0.0979**** | **0.1065** |
| | **(0.001)** | **(0.002)** | **(0.004)** | **(0.185)** | **(0.018)** | **(0.586)** |
| ΔUrbanization | -0.0721 | -0.0738 | -0.8813* | -0.0173 | -0.4842 | -0.1526 |
| | (0.516) | (0.563) | (0.071) | (0.927) | (0.149) | (0.374) |
| ΔEducational achievement | 0.0005 | 0.0025 | -0.0558 | -0.0640 | -0.1030 | -0.0626 |
| | (0.943) | (0.750) | (0.372) | (0.361) | (0.399) | (0.180) |
| ΔUnemployment | -0.0111 | -0.0121 | 0.0147 | 0.0436 | -0.0615 | 0.1034 |
| | (0.424) | (0.488) | (0.907) | (0.625) | (0.723) | (0.130) |
| ΔWorking age population | -0.2206 | -0.2840 | -2.0506 | 0.7034 | -1.5064 | 1.5940** |
| | (0.498) | (0.476) | (0.350) | (0.605) | (0.376) | (0.042) |
| ΔLog(PC GRP) | -0.1396 | -0.3566 | -5.1559 | -2.8666 | -1.9163 | -0.7147 |
| | (0.745) | (0.506) | (0.115) | (0.166) | (0.616) | (0.562) |
| Lagged dep. variable | -0.0234 | -0.0272 | -0.0730 | -0.0937 | -0.0291 | -0.4170*** |
| | (0.637) | (0.589) | (0.217) | (0.644) | (0.899) | (0.000) |
| Observations | 631 | 631 | 631 | 315 | 155 | 160 |
| Panels | 79 | 79 | 79 | 79 | 39 | 40 |
| AR(2) (p-value) | 0.50 | 0.40 | 0.15 | 0.14 | 0.39 | 0.18 |
| Hansen test (p-value) | 0.346 | 0.406 | 0.121 | 0.263 | 0.238 | 0.908 |
| Instruments | 25 | 25 | 25 | 21 | 29 | 21 |

Notes: p-values in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects.

Table 8: Robustness checks for IV regressions with additional controls

| Time period: | 2007-2012 | | | | | | 2013-2016 | |
|---|---|---|---|---|---|---|---|---|
| Dependent variable: | ΔAll Protests | | | ΔSocial Protests | | | ΔAll Protests | ΔSocial Protests |
| | All regions | Democratic | Non-democratic | All regions | Democratic | Non-democratic | Democratic | Democratic |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| **ΔGini coefficient** | **1.9656** | **3.0541** | **-1.2471** | **1.6484*** | **1.1655*** | **-0.6390** | **-1.9757** | **-0.2807** |
| | **(0.286)** | **(0.115)** | **(0.421)** | **(0.081)** | **(0.077)** | **(0.452)** | **(0.761)** | **(0.903)** |
| ΔUrbanization | -1.0677 | -2.6084** | -0.6986 | -0.5563* | -1.0414*** | -0.4342* | -0.3240 | -0.1335 |
| | (0.140) | (0.015) | (0.261) | (0.095) | (0.001) | (0.083) | (0.540) | (0.281) |
| ΔEducational achievement | -0.1925 | -0.4697* | 0.0634 | -0.0945 | -0.1359 | -0.0172 | 0.1123 | 0.0289 |
| | (0.206) | (0.076) | (0.657) | (0.156) | (0.264) | (0.641) | (0.503) | (0.653) |
| ΔUnemployment | 0.4648* | 0.6860 | 0.1366 | 0.1542 | 0.2108 | -0.0344 | -0.2141 | -0.0240 |
| | (0.068) | (0.177) | (0.481) | (0.200) | (0.337) | (0.552) | (0.750) | (0.886) |
| ΔWorking age population | -2.2334 | -6.0727 | -0.1862 | -1.0042 | -2.8590* | 0.3618 | 3.1996 | -0.5979 |
| | (0.335) | (0.147) | (0.916) | (0.341) | (0.080) | (0.473) | (0.652) | (0.801) |
| ΔLog(PC GRP) | 0.4531 | 2.8624 | -2.1389 | -0.3581 | 0.3616 | -1.0535 | 5.2632 | 1.3020 |
| | (0.899) | (0.727) | (0.613) | (0.817) | (0.906) | (0.437) | (0.552) | (0.648) |
| | 1.9656 | 3.0541 | -1.2471 | 1.6484* | 1.1655* | -0.6390 | -1.9757 | -0.2807 |
| Observations | 394 | 194 | 200 | 394 | 194 | 200 | 154 | 154 |
| Panels | 79 | 39 | 40 | 79 | 39 | 40 | 39 | 39 |
| F-stat | 7.51 | 8.34 | 9.35 | 7.51 | 8.34 | 9.35 | 0.13 | 0.13 |
| Hansen test (p-value) | 0.22 | 0.33 | 0.17 | 0.22 | 0.33 | 0.17 | 0.86 | 0.86 |

Notes: p-values in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects.

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# 6 Conclusion

We study empirically how oil windfalls that accrue to subnational budgets affect inequality in a rent-seeking environment. Our estimation strategy takes advantage of the unique setting of oil tax policy in the Russian Federation. We find that unexpected growth in oil tax revenues – proxied by oil price shocks – increases income inequality, particularly the income share of the wealthiest quintile but only conditional on the level of corruption in the region. We interpret our findings as evidence of growing inequality due to the capture of the oil windfalls by politically connected elites. Additionally, we provide indirect evidence for our hypothesis by showing the impact of oil tax revenues under rent-seeking on income composition by sources: the interaction of oil price shocks and rent-seeking decreases the share of income from employment but increases the share of "other income" traditionally associated with corruption.

Finally, we examine the potential political implications of rising inequality. We confirm our prediction that redistribution of the oil rents through corruption is associated with higher protest activity, particularly protests aimed at fostering the interests of the poor and vulnerable groups of society. This effect is unsurprisingly driven by the subgroup of regions with relatively democratic institutions. This result identifies inequality as a potentially distinct transmission channel for the detrimental effect of oil rents on political stability and requires further empirical research.

Our focus on one specific category of oil rents is one of the main contributions to the literature that has previously treated all types of oil rents as a single aggregate. It also allows us to produce more specific policy implications. The Russian data suggest that discontinuing oil tax sharing with regional budgets is probably the most straightforward way of reducing inequality caused by oil windfalls. However, this might still be suboptimal: the reallocation of the oil tax revenues to the federal center may subject these funds to similar or even higher risks of misappropriation at the federal level. In this case, a policy alternative to the complete reallocation of the oil tax revenues to the central government is to provide a transparent mechanism of direct distribution of unexpected oil windfalls to the general population, as is the case in Alaska, USA, and the province of Alberta, Canada (Ross 2007) and has been advocated by Moss et al. (2015). This, however, would not be needed if the government could tackle the problem of rent-seeking in the first place.

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# Appendix A

Table A1

Panel A: Descriptive statistics based on the 2005-2012 period, first-differences

| Variables | Mean | St. Dev | Maximum | Minimum |
|---|---|---|---|---|
| 1st income quintile (%) | -0.09 | 0.17 | 0.50 | -0.80 |
| 2d income quintile (%) | -0.10 | 0.19 | 0.62 | -0.70 |
| 3d income quintile (%) | -0.08 | 0.16 | 0.50 | -0.92 |
| 4th income quintile (%) | -0.02 | 0.07 | 0.25 | -0.26 |
| 5th income quintile (%) | 0.29 | 0.55 | 2.52 | -1.84 |
| Gini coefficient (%) | 0.37 | 0.70 | 3.03 | -2.60 |
| Wage income (%) | -0.53 | 2.53 | 8.70 | -9.60 |
| Entrepreneurial income (%) | -0.29 | 1.35 | 6.30 | -7.80 |
| Social transfers (%) | 0.75 | 1.78 | 8.50 | -5.60 |
| Property income (%) | -0.22 | 1.39 | 7.10 | -8.80 |
| Other income (%) | 0.30 | 3.69 | 15.40 | -13.50 |
| Oil price shock | 0.42 | 5.59 | 53.86 | -69.06 |
| Oil price shock for oil producers | 0.99 | 8.49 | 53.86 | -69.06 |
| Bribe cases per 100,000 pop-n, log | 0.03 | 0.40 | 1.42 | -2.07 |
| Embezzl. cases per 100,000 pop-n, log | -0.11 | 0.42 | 1.59 | -1.63 |
| Crime per 100,000 pop-n, log | -0.03 | 0.13 | 0.41 | -0.35 |
| Murders per 100,000 pop-n, log | -0.10 | 0.18 | 1.07 | -0.94 |
| PC GDP (1,000s 2016 RR), log | 0.03 | 0.08 | 0.50 | -0.34 |
| Urban share (%) | 0.07 | 0.45 | 1.89 | -4.28 |
| Higher education (%) | 0.72 | 2.86 | 15.00 | -17.40 |
| Unemployment (%) | -0.34 | 1.93 | 19.60 | -11.20 |
| Working age pop-n (%) | -0.19 | 0.49 | 1.80 | -1.22 |
| All protests per mln pop-n, log | 3.70 | 4.78 | 38.39 | 0.00 |
| Social protests per mln pop-n, log | 0.80 | 1.53 | 16.00 | 0.00 |
| Political protests per mln pop-n, log | 1.47 | 2.08 | 11.66 | 0.00 |
| Economic protests per mln pop-n, log | 0.84 | 1.44 | 9.25 | 0.00 |
| Legal protests per mln pop-n, log | 0.42 | 0.87 | 6.79 | 0.00 |
| Cultural protests per mln pop-n, log | 0.06 | 0.26 | 2.56 | 0.00 |
| Environmental protests per mln pop-n, log | 0.34 | 0.96 | 9.62 | 0.00 |

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Table A1 continued...

Panel B: Descriptive statistics based on the 2013-2018 period, first-differences

| Variables | Mean | St. Dev | Maximum | Minimum |
|---|---|---|---|---|
| 1st income quintile (%) | 0.10 | 0.16 | 0.70 | -0.50 |
| 2d income quintile (%) | 0.11 | 0.19 | 1.00 | -0.60 |
| 3d income quintile (%) | 0.10 | 0.15 | 0.70 | -0.40 |
| 4th income quintile (%) | 0.03 | 0.07 | 0.30 | -0.21 |
| 5th income quintile (%) | -0.34 | 0.53 | 1.70 | -2.70 |
| Gini coefficient (%) | -0.39 | 0.75 | 2.10 | -3.50 |
| Wage income (%) | 2.32 | 5.52 | 24.30 | -7.40 |
| Entrepreneurial income (%) | -0.60 | 1.98 | 6.90 | -15.40 |
| Social transfers (%) | 0.18 | 1.13 | 3.50 | -3.00 |
| Property income (%) | 0.01 | 0.48 | 2.70 | -2.70 |
| Other income (%) | -1.91 | 5.72 | 12.20 | -24.20 |
| Oil price shock | -0.35 | 5.95 | 58.90 | -57.29 |
| Oil price shock for oil producers | 0.69 | 11.22 | -57.55 | 80.52 |
| Bribe cases per 100,000 pop-n, log | -0.09 | 0.49 | 1.68 | -2.02 |
| Embezzl. cases per 100,000 pop-n, log | -0.08 | 0.31 | 1.42 | -1.42 |
| Crime per 100,000 pop-n, log | -0.04 | 0.08 | 0.19 | -0.23 |
| Murders per 100,000 pop-n, log | -0.06 | 0.19 | 0.35 | -0.73 |
| PC GDP (1,000s 2016 RR), log | 0.02 | 0.06 | 0.33 | -0.41 |
| Urban share (%) | 0.17 | 0.51 | 7.06 | -2.03 |
| Higher education (%) | 0.77 | 1.77 | 7.70 | -4.60 |
| Unemployment (%) | -0.14 | 0.94 | 3.20 | -13.90 |
| Working age pop-n (%) | -0.84 | 0.21 | -0.03 | -2.64 |
| All protests per mln pop-n, log | -0.30 | 2.05 | 10.10 | -9.07 |
| Social protests per mln pop-n, log | 0.03 | 1.06 | 6.05 | -5.90 |
| Political protests per mln pop-n, log | -0.25 | 1.27 | 3.86 | -9.08 |
| Economic protests per mln pop-n, log | -0.03 | 0.90 | 3.75 | -6.52 |
| Legal protests per mln pop-n, log | -0.04 | 0.41 | 1.87 | -1.87 |
| Cultural protests per mln pop-n, log | -0.01 | 0.14 | 0.79 | -1.01 |
| Environmental protests per mln pop-n, log | -0.03 | 0.51 | 4.72 | -3.01 |

Figure A1: Change in the unemployment rate over the 2004-2018, regional average
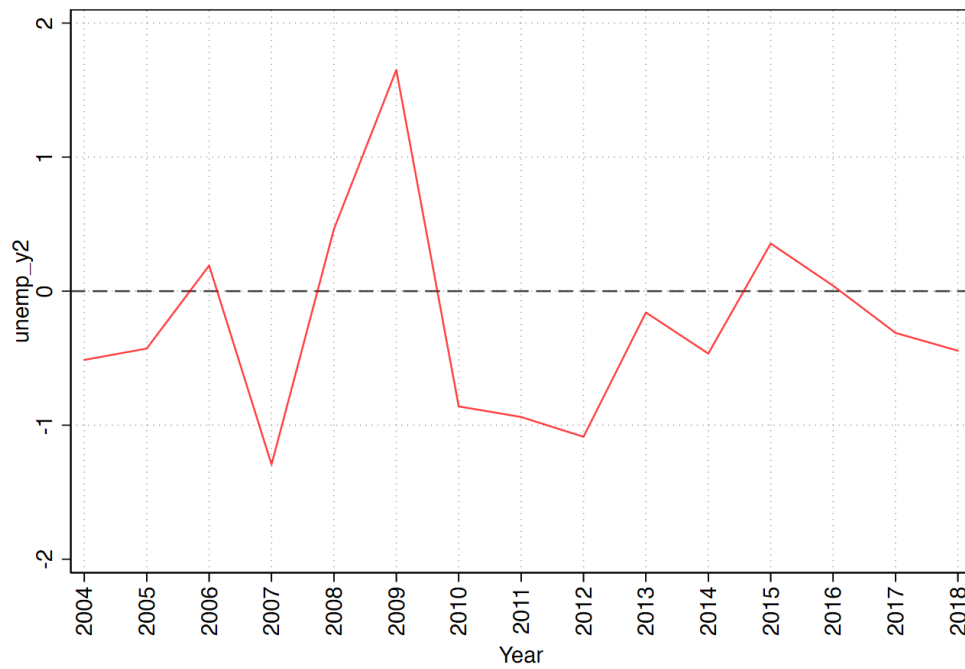


Source: Alexeev & Zakharov based on data from the Russian Federal State Statistics Service

Figure A2: Dynamics of a consumer price index over the 2004-2018, regional average



Source: Alexeev & Zakharov based on data from the Russian Federal State Statistics Service

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Figure A3: Dynamics of international oil prices and an oil price shock in an average oil-producing region by years
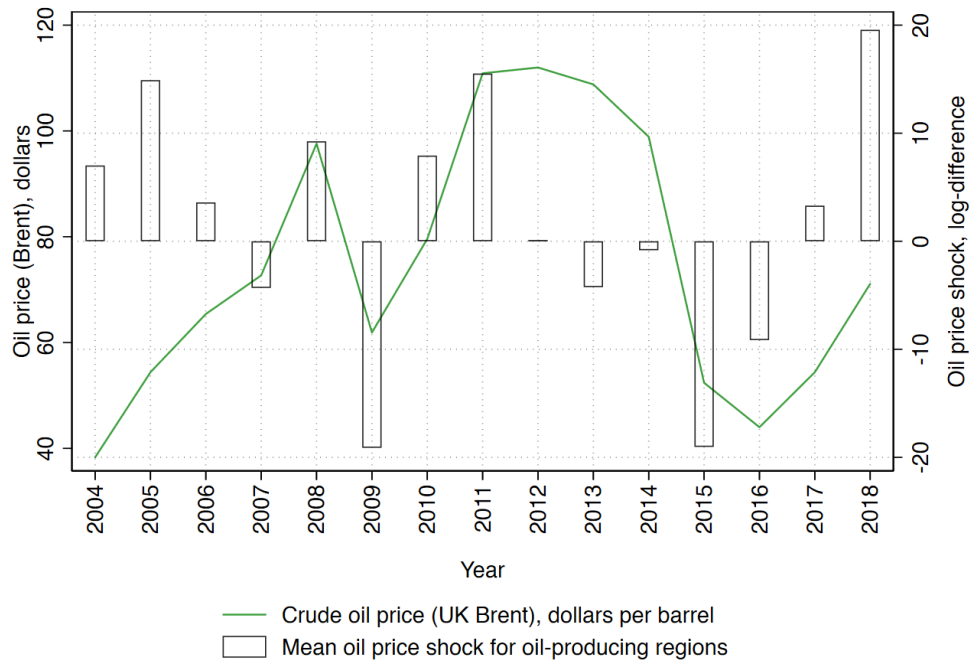


Source: Alexeev & Zakharov based on data from the Russian Federal State Statistics Service

Table A2: Oil windfalls and other protests (second stage IV regressions; 2007-2012)

| Dependent variable: | (1) ΔPolitical protest | (2) | (3) | (4) ΔEconomic protest | (5) | (6) | (7) ΔLegal protest | (8) | (9) | (10) ΔCultural protest | (11) | (12) | (13) ΔEnvironmental protest | (14) | (15) |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | All | D | ND | All | D | ND | All | D | ND | All | D | ND | All | D | ND |
| **ΔGini coefficient** | **-0.72** | **0.14** | **0.25** | **-0.41** | **1.05** | **-0.95** | **0.16** | **0.05** | **-0.19** | **0.20** | **0.13** | **0.05** | **0.24** | **0.08** | **0.10** |
| | (0.39) | (0.86) | (0.66) | (0.59) | (0.12) | (0.30) | (0.83) | (0.94) | (0.60) | (0.23) | (0.43) | (0.55) | (0.58) | (0.87) | (0.73) |
| Observations | 395 | 195 | 200 | 395 | 195 | 200 | 395 | 195 | 200 | 395 | 195 | 200 | 395 | 195 | 200 |
| Panels | 79 | 39 | 40 | 79 | 39 | 40 | 79 | 39 | 40 | 79 | 39 | 40 | 79 | 39 | 40 |
| F-stat | 9.36 | 15.14 | 9.63 | 9.36 | 15.14 | 9.63 | 9.36 | 15.14 | 9.63 | 9.36 | 15.14 | 9.63 | 9.36 | 15.14 | 9.63 |
| Hansen test (p-value) | 0.20 | 0.29 | 0.18 | 0.20 | 0.29 | 0.18 | 0.20 | 0.29 | 0.18 | 0.20 | 0.29 | 0.18 | 0.20 | 0.29 | 0.18 |

Notes: t statistics in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects; All – all regions; D – democratic regions; ND – non-democratic regions.

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Table A3: Oil windfalls and protest, reduced form (OLS FE)

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| Time period: | | | | 2007 - 2012 | | | | 2013 - 2018 |
| Dependent variable: | | ΔAll Protests | | | ΔSocial Protests | | ΔAll Protests | ΔSocial Protests |
| | All regions | Democratic | Non-democratic | All regions | Democratic | Non-democratic | | Democratic |
| OPS | 0.0005 | -0.0098 | 0.0890 | -0.0307 | -0.0393 | 0.0447 | 0.0199 | 0.0140 |
| | (0.991) | (0.859) | (0.198) | (0.216) | (0.151) | (0.249) | (0.466) | (0.398) |
| Δ Bribes | -0.1405 | 0.4942 | -0.6019 | -0.2465$^*$ | -0.3849 | -0.1408 | -0.2059 | -0.2187 |
| | (0.722) | (0.427) | (0.270) | (0.089) | (0.240) | (0.401) | (0.715) | (0.231) |
| **OPS× ΔBribes** | **0.0531** | **0.0809** | **0.2693** | **0.0457$^*$** | **0.0584$^{**}$** | **0.1831** | **-0.0296** | **-0.0263** |
| | **(0.356)** | **(0.238)** | **(0.487)** | **(0.080)** | **(0.044)** | **(0.561)** | **(0.552)** | **(0.186)** |
| Observations | 395 | 195 | 200 | 395 | 195 | 200 | 156 | 156 |
| Panels | 79 | 39 | 40 | 79 | 39 | 40 | 39 | 39 |
| R squared | 0.08 | 0.09 | 0.08 | 0.06 | 0.08 | 0.07 | 0.04 | 0.02 |

Notes: p-values in parentheses, * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Robust standard errors are clustered at the region level. All estimations include regional and year fixed effects.

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