Election Cycles and Mining Sector Governance in Post-conflict Kosovo

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ABSTRACT

Political business cycles are typically linked to the manipulation of fiscal or monetary policy instruments. In a recent paper, Imami, Lami and Uberti (ILU, 2018) argue that opportunistic politicians may also choose to manipulate non-fiscal/non-monetary policy instruments. Here, we extend ILU's study using time-series data on mining-sector licensing from post-conflict Kosovo (2001-2018). We find robust evidence that is consistent with electoral opportunism in the allocation of mining permits, despite the checks-and-balance mechanisms introduced by Kosovo's international administrators in an attempt to reduce the politicisation of licensing. That said, the cycle effect is only observed prior to scheduled, as opposed to early, elections. Disaggregating the data by license type, in addition, we find that the observed election cycle is driven primarily by the manipulation of licenses for the mining of construction materials. We argue that, in the context of post-conflict Kosovo, this is the category of licenses whose strategic manipulation offers the greatest pay-off to the incumbent. The results raise some questions about the feasibility of fighting political opportunism (and, relatedly, corruption) by establishing formal check-and-balance mechanisms.

KEYWORDS

Political business cycles; mining; construction; international administration; separation of powers; corruption

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1. Introduction

Political business cycles (PBCs) are macroeconomic fluctuations induced by the electoral cycle (Frey & Benz, 2002; Drazen, 2008a; Alesina, Roubini & Cohen, 1997). In traditional 'opportunistic' models, incumbent politicians stimulate the economy during election years in order to increase voter satisfaction and help their re-election prospects. By and large, the existing literature links PBCs to the manipulation of fiscal and monetary policy instruments such as government spending, tax rates and the money supply. Political business cycles, in other words, are driven by political *budget* cycles (Drazen, 2008b). Although the evidence for advanced countries is mixed, other contributions have found robust evidence of fiscal and monetary opportunism in developing and transition countries (Block, 2002; Khemani, 2004; Akhmedov & Zhuravskaya, 2004).[1]

In a recent paper, Imami, Lami & Uberti (2018, henceforth ILU) suggest that the logic of electoral competition may also lead incumbents to manipulate *non-fiscal/non-monetary* policy tools such as investment licenses. Electoral competition, in other words, may also induce *microeconomic* fluctuations. A number of scenarios may prompt the incumbent to diversify his/her electioneering strategy away from traditional expansionary measures: if a monetary and fiscal stimulus is unviable due to already-high inflation rates; if the government budget is subject to externally imposed consolidation measures (e.g. an IMF programme); or if the country lacks monetary sovereignty due to dollarization (or, as is the case in Kosovo, euroization). ILU (2018) explore these possibilities by examining the effect of elections on the issuing of mining licenses in post-socialist Albania. They find robust evidence of opportunistic effects in line with traditional PBC theory. In their argument, awarding additional licenses in the run-up to elections allows the incumbent to stimulate economic activity and/or reward client firms who provide campaign finance or pay bribes.

Here, we replicate and extend the ILU's (2018) study of Albania by using time-series data on mining-sector licensing from neighbouring Kosovo for the period 2001-2018. As noted by Hamermesh (2007, p. 715), replication is 'an activity that most economists applaud but few perform'. The most important justification for scientific replication is that 'one cannot expect econometric results produced for […] one country to carry over to another' (ibid., p. 727). In particular, whether election cycles in mining licensing are a recurring feature of transition economies or just a peculiarity of the Albanian context cannot be determined from a single country study, let alone a priori.

A potential solution is to employ panel data (e.g. Mosley & Chiripanhura, 2016). Multi-country studies of election cycles, however, face a number of difficulties. First, the underlying structural model may be different for different countries. Second, multi-country studies typically rely on annual data, which obscure the dynamics taking place *within* election years. Here, we elect to focus on a single country as it allows us to specify a model that is sensitive to the country context and to exploit the detailed monthly data that are available for Kosovo. As in the ILU (2018) study of Albania, of course, the limitation of this approach is that the cycle effects are identified from a limited number of elections. Accordingly, the conclusions may not immediately generalise beyond the context of Kosovo. That said, conducting a series of country-level replications allows us to accumulate evidence of a general pattern while also paying attention to country-level specificities.

[1] Brender & Drazen (2005) find that political business and budget cycles are a phenomenon specific to recently democratised countries, regardless of income levels. Thus, it is particularly relevant to investigate PBCs in the context of a post-socialist, post-conflict country such as Kosovo.

An important difference between Albania and Kosovo is that the regulatory institutions that preside over the mining sector are established differently. Whereas in Albania the agency responsible for granting licenses is established as a unit of a line ministry, Kosovo's mining regulator is an independent administrative agency. In theory, a separation of powers between the regulator and the executive should make it more difficult for incumbent politicians to influence the mining bureaucracy and manipulate the licensing process for opportunistic reasons. Accordingly, an election cycle effect, if observed at all, should be less prominent in Kosovo than in Albania. Yet, using a similar specification and estimation strategy to ILU (2018), we find robust evidence of an election cycle in the allocation of mining permits in Kosovo, too. In fact, the estimated pre-election effect is consistently *larger* in Kosovo than in Albania. We explain this finding by suggesting that incumbent politicians rely on informal networks of power to control the mining bureaucracy and sidestep formal rules. That said – and here is the second important difference between Kosovo and Albania – the election cycle is only observed prior to scheduled, as opposed to early elections, in line with previous findings (Khemani, 2004).

An additional innovation of this study is that it probes deeper into the logic of mining policy manipulation by using data disaggregated by license type. We examine whether the election cycle is driven primarily by a particular type of license. The data are consistent with *extraction* licenses for *construction* materials (e.g. sand, gravel, etc.) being the primary vehicle used by the incumbent to influence election outcomes. This finding accords well with the central role played by the construction industry in the political economy of post-conflict Kosovo. Finally, we perform a number of robustness checks that were neglected by ILU (2018).

Although further research is needed, we conclude by suggesting that election cycles are fundamentally a political phenomenon, and that formal institutional checks (e.g. a formally independent licensing authority) may be of limited use in constraining opportunistic behaviour by the incumbent. Before presenting our empirical analysis, we first provide some background on the mining sector in Kosovo and formulate testable hypotheses.

2. Mining Policy and Politics in Post-conflict Kosovo

Following the break-up of socialist Yugoslavia and a decade of inter-ethnic strife, the 1999 NATO military intervention brought Serbian direct rule in Kosovo to an end, leading to the establishment of a UN protectorate in the Albanian-majority province. Since then, Kosovo has been the site of unprecedented state- and institution-building efforts by the international community, becoming one of the world's largest recipients of development aid in per capita terms (Capussela, 2015, p. 186).

With Western backing, the local leadership proclaimed independence from Serbia in February 2008, paving the way for the end of the period of internationally 'supervised' independence in 2012.

[Table 1]

Kosovo offers a particularly suitable context to replicate ILU (2018). Both Albania and Kosovo are rich in mineral resources and used to be large primary commodity exporters during socialism (Palairet, 2003; World Bank, 2009).[2] In the early years of transition, the mining industry succumbed to underinvestment, cannibalisation and decay in both countries, but later experienced something of a revival in the wake of the global commodity super-cycle – albeit less in Kosovo than in Albania. Mining occupies a similar position in the economic structure of both countries, although it is relatively more important (as a share of value-added) in Kosovo's less diversified economy (Table 1). In addition, both Albania and Kosovo are lower-middle income economies with similar historical legacies (Ottoman and socialist rule) and a majority population that speaks the same language (Albanian) and shares to a large extent the same culture. There are, however, two key differences, which have important implications for model specification and the interpretation of the results.

Institutionally, the agencies responsible for issuing mining permits have a different legal status in the two countries. In Albania, the mining regulatory office[3] operates as an arm of the Ministry of Infrastructure and Energy.[4] As such, the bureaucratic apparatus is likely to be particularly exposed to political interference, potentially explaining the election-year manipulation of mining policy reported by ILU (2018). In post-1999 Kosovo, by contrast, the bodies responsible for mining licensing have enjoyed a high degree of legal and administrative independence.

During 2000-5, 'reserve powers' over mineral resources and mining policy were held by the Directorate of Mines and Minerals (DMM), a specialised unit of the United Nations Interim Administration Mission in Kosovo (UNMIK) that was led by European Union-appointed staff (Large et al. 2003: 6). At the same time, a constitutional framework unveiled in May 2001 paved the way for the establishment of Provisional Institutions of Self-Governance (PISGs), headed and staffed by Kosovar personnel. The framework outlined the progressive transfer of executive powers to the PISGs. The first elections to the new 120-seat legislative assembly were held in November 2001 (Bideleux & Jeffries, 2007, p. 565). In 2005, UNMIK went on to establish a politically independent regulatory body tasked with licensing and monitoring all mining activities in Kosovo (Uberti, 2014a).

[2] In 1989, minerals (including chrome, copper and nickel) accounted for nearly 80% of total Albanian exports by value (World Bank, 2009: 17). Throughout the 1970s, metals (including lead, zinc, bauxite, magnesium) accounted for nearly 46% of Kosovo's exports. By 1985, the share of metals in total exports had declined to 24%, however (Gashi and Pugh, 2015: 37).
[3] The *Agencia Kombëtare e Burimeve Natyrore* (National Agency for Natural Resources).
[4] Formerly, the Ministry of Energy and Industry.

Initially with a staff of 64, the Independent Commission on Mines and Minerals (KPMM)[5] inherited all the powers and functions previously carried out by UNMIK's DMM.[6]

UNMIK's institution-building template was overtly inspired by international best practice (Bastida, 2008). To reduce the risk of corruption and political interference, for instance, the World Bank advocates the establishment of government-independent regulatory bodies along the lines of independent central banks (World Bank, 1996, p. 45). Mining regulators are supposed to establish and extinguish mineral rights purely based on technical, financial and environmental criteria. In Kosovo, all licensing decisions are taken by the 5-member KPMM board within 3 months of a license application's submission date. The board members are appointed by Kosovo's legislative assembly for a term of four years (which is deliberately intended to be out of synch with the electoral cycle). They may not be active members of a political party or hold any other political positions in public institutions. By law, the decision to license a project must be based solely on the quality of the proposed mining or exploration plan and the applicant's financial and technical capabilities.

In theory, a separation of powers between politicians and regulators should make it difficult for politicians to influence licensing decisions, reducing the likelihood of policy manipulation during elections. Accordingly, the amplitude of the election cycle in Kosovo, where the regulator is political independent, should be expected to be smaller than in Albania, where licenses are issued by a line ministry. If the separation of powers completely eliminates political interference, the magnitude of the observed cycle effect should be zero ($H_0\colon cycle = 0$).

In practice, however, there are ways in which politicians may circumvent this separation of powers. For instance, KPMM board members may be linked to the ruling coalition through ties of personal, political or familial loyalty. The incumbent may rely on these informal networks to secure the board members' support for its electioneering strategy, eluding the formal system of checks and balances. If this is the case, we should expect to observe an election cycle effect in the allocation of mining licenses in Kosovo, too ($H_a\colon cycle \neq 0$).

Anecdotal evidence of corruption and informality in post-conflict Kosovo lends credence to this alternative hypothesis (Belloni & Strazzari, 2014; Danielsson, 2016). Capussela (2015, p. 188), for instance, argues that 'patronage in Kosovo primarily takes the form of granting employment in the public sector to groups connected to the factions of the elite, so as to link their future income to the maintenance of power by such factions'. Following an election, the winning coalition must 'divert to

[5] Komisioni i Pavarur për Miniera dhe Minerale
[6] According to UNMIK, introducing a comprehensive system of licensing was crucial to 'protecting resource assets' (Large et al., 2003: p. 15). Following NATO's 1999 intervention, leaders of the ethnic-Albanian guerrilla (KLA) had taken control of most of Kosovo's socially-owned mining operations (Grasten & Uberti, 2017). Consequently, UNMIK also saw the introduction of a licensing system as an urgent response to the chaos of 'informal' or 'spontaneous' privatization.

itself the loyalty of the existing civil servants, who had been employed by [previous] governments' (ibid.).

By securing the loyalty of the administration's rank-and-file, the ruling coalition can rely on the state apparatus to perform political functions, such as protecting or rewarding client firms. In an interview conducted by the authors, the owner of a ready-mix concrete company based in Prizren stated that their production site is surrounded by small-scale miners operating without a license. Although the quarries' owners may evade detection by bribing local KPMM inspectors, the interviewee also ventured that they may enjoy some degree of political protection higher up.[7] Confirming this perspective, the local representative of a foreign-owned lead and zinc miner observed that KPMM often refrains from adequately monitoring (or fining) private quarries in the case that they are owned by political figures or party bosses. On the other hand, he added, 'there are extensive cases of informal monitoring of [other] private enterprises'.[8] While we do not have direct evidence that politicians manipulate the allocation of licenses *for electioneering purposes,* the available evidence does suggest that KPMM may not be completely immune to pressure from incumbent politicians.

Whether a formal separation of powers is sufficient to insulate licensing decisions from political considerations, or whether the incumbent can use informal networks to influence licensing decisions, is ultimately an empirical question. This question can be examined by testing H_0 against H_a. Answering this question is important for various reasons. It allows us to determine whether the ability of the incumbent to act opportunistically (a key prediction of PBC theory) depends critically on the absence of institutional checks and balances (as suggested by Khemani, 2004, p. 127, amongst others). Relatedly, it allows us to determine whether establishing a separation of powers between politicians and regulators has the desired effect of curbing the discretionary power of the former over the latter.

The *second* important difference between Albania and Kosovo concerns their politics. Since political competition was introduced in December 1990, Albania has maintained a two-party system with a relatively high degree of political stability.[9] With the exception of the government that emerged out of the 1996 elections, all the post-socialist governments saw out their full four-year term in office, leading to a sequence of electoral contests that were fully anticipated by the incumbent and could therefore be planned in advance.[10]

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[7] Interview with non-metallic minerals company based in Prizren, 8 April 2015. The interviewee also noted that the KPMM is otherwise fairly 'clean': '[KPMM] might delay the granting of a certain permit [perhaps to solicit a bribe payment], but eventually they give it to you'.

[8] Interview with mining company based in Prishtina, 21 July 2015.

[9] At the level of parliamentary politics, at least.

[10] The government formed in 1996 was forced to resign following the outbreak of the 1997 civil unrest.

In Kosovo, by contrast, the political arena is considerably more fragmented, with governments typically formed out of loose and unstable coalitions of political parties. In addition, the 2008 post-independence constitution raised the length of parliamentary terms from three to four years. As a result, no post-independence government has ever survived for the full duration of its term, paving the way for early elections in three out of six electoral cycles. In particular, the November 2001, October 2004 and November 2007 elections took place according to the constitutionally mandated schedule. The contests of December 2010, June 2014 and June 2017, by contrast, were early elections called by the President of the Republic in the wake of a no-confidence vote against the government.

The distinction between early and scheduled elections was first highlighted by Khemani (2004), who finds patterns of electoral fiscal manipulation in the Indian states, but only prior to constitutionally scheduled elections. Snap elections are, by definition, sudden and unanticipated. As such, they force the incumbent to organise an election campaign on an extremely tight schedule. Indeed, in the case of government collapse, Kosovo's constitution mandates that new elections be held within 45 days of the dissolution of parliament. For this reason, the incumbent's ability to manipulate the allocation of licenses may be greatly diminished,[11] making it less likely for a political cycle to arise in relation to early elections.

As such, whether or not Kosovo's incumbent politicians are able to exert political influence on the KPMM, thereby generating an election cycle, we only expect them to be able to do so prior to scheduled elections. While theory produces conflicting predictions as to whether we should observe a cycle prior to *scheduled* elections, we expect to observe *no* cycle effect in the run-up to *early* elections. We now turn to presenting our empirical specification and estimation strategy.

3. Empirical Strategy

The outcome variable is the number of mining permits issued in month t. Our data, obtained from the Independent Commission of Mines and Minerals (KPMM), encompasses the full history of mining sector regulation in post-conflict Kosovo, for a total of 201 monthly observations. A time-series representation of the dependent variable (together with the timing of elections) is shown in Figure 1.

[Figure 1]

[11] This may be especially the case if the licensing authority is formally established as an independent agency (as in Kosovo). We cannot test whether an incumbent can manipulate a non-independent agency prior to early elections since (with the exception of the 1997 elections), all post-socialist elections in Albania were scheduled elections.

The outcome is a *non-negative*, integer-valued random variable y which we assume to be Poisson-distributed with population mean μ. The assumption is corroborated by an inspection of the data, which reveals a markedly right-skewed sample distribution (Figure 2). For this type of data, (event counts), the standard econometric approach is to estimate a Poisson model (Cameron & Trivedi, 2013). In this model, the conditional mean of the outcome $E(y_t|X_t) = \mu_t$ is parametrised as

$$\mu_t = \exp(\boldsymbol{\beta X}_t) \tag{1}$$

in order to ensure that μ_t is always greater than (or equal to) zero. $\boldsymbol{X}_t$ is a vector of regressors and $\boldsymbol{\beta}$ is a vector of parameters.[12]

[Figure 2]

Our specification of the vector of regressors adapts ILU's (2008) specification to the Kosovar context. Since KPMM is the 'monopoly supplier' of mining licenses, the equilibrium quantity of licenses issued depends solely on the KPMM's supply decisions (ILU, 2018). The decision to issue a license, in turn, may be influenced by the incumbent's opportunistic incentives (provided the incumbent is able to exert pressure on the KPMM) and other political and institutional factors. To model the influence of elections, we use dummies for pre-election years (A_{-1}), as is standard in the PBC literature. We also distinguish between early and scheduled elections. Thus, A_{-1}^{sched} (A_{-1}^{early}) takes the value 1 for all months belonging to the last year prior to a scheduled (early) election, and zero otherwise. Estimating a statistically significant coefficient on A_{-1}^{sched} implies a rejection of H_0.

The timing of scheduled elections is constitutionally set, so that it may be considered exogenous. On the assumption that the timing of early elections does not depend on mining policy decisions, the coefficients on the early election dummies may also be given a causal interpretation. This assumption is plausible. Early elections may be timed strategically by the incumbent to coincide with favourable macroeconomic outcomes (e.g. low unemployment, low inflation). Macroeconomic shocks may also pave the way for early elections. It is very unlikely, however, that the KPMM's licensing decisions may (lead the incumbent to) trigger early elections.

In addition to the election cycle dummies, we control for a number of potential confounders. First, we include a dummy for the post-independence (post-2008) period ($Indep$). This variable has a

[12] Even though the OLS estimator is asymptotically normal (regardless of the population distribution), a linear OLS model with form $\mu_t = \boldsymbol{\beta X}_t$ might produce negative predicted values, violating a fundamental feature of count data. That said, our baseline results (Model 2, Table 2) are qualitatively unchanged if the model is given this linear form and is estimated by OLS (with HAC standard errors). Results are available upon request.

particularly important control function since all scheduled (early) elections took place before (after) the February 2008 declaration of independence ($Indep$ correlates with A_{-1}^{sched} at -0.52 and with A_{-1}^{early} at 0.38). Since $Indep$ is also positively correlated with the outcome variable (0.42), omitting it may lead the estimated coefficient on A_{-1}^{sched} (A_{-1}^{early}) to be biased downwards (upwards). Secondly, we control for any influence from the legal regime governing licensing activities. During the period 2001-18, we distinguish between four legal regimes. Between May 2001 and August 2003, the DMM at UNMIK issued 139 licenses (mostly for quarrying) under the 1980 mining law inherited from socialist Yugoslavia. This law is treated as the reference category. In August 2003, UNMIK suspended all licensing operations pending a fully-fledged re-organization of the regulatory regime.[13] The new UNMIK-drafted mining law entered into force on 21 January 2005 and led to the establishment of the KPMM.[14] Following the full transfer of sovereignty to local institutions in 2008, the 2005 law was repealed and replaced by a new law on 27 August 2010.[15] Accordingly, we define the indicator variable $Suspend$, which takes the value 1 during all months between September 2003 and January 2005 (inclusive), and 0 otherwise; $Law2005$, which takes the value 1 between February 2005 and August 2010 (inclusive); and $Law2010$, which takes the value 1 from September 2010 onwards.

While the equilibrium quantity of licenses is supply-driven, we also allow the regulator's supply decisions to be influenced by firms' demand for mining (i.e. investment) licenses (here, we closely follow ILU, 2018, p. 103-4). The demand for licenses depends on firms' investment demand, which in turn depends on market conditions. When the market is buoyant, mining firms may lobby the regulator (or the incumbent) to relax the supply constraint. If, by chance, the lobbying efforts of firms coincide with the election cycle, the estimated coefficient on A_{-1} would spuriously pick up the influence of market forces. To obtain unbiased estimates of the election cycle, we therefore control for firms' investment demand.

Following standard microeconomic theory, we model investment demand as a function of exogenous (input and output) prices, measured using the log of a metal price index (MPI) compiled by the IMF and the 12-month LIBOR interest rate.[16] We assume it may take over a year for firms to

[13] The 2004 scheduled election took place during the suspension period.

[14] UNMIK Regulation 2005/3 'On Mines and Minerals in Kosovo'.

[15] Law No. 03/L-163 'On Mines and Minerals', 27.8.2010. This law transferred some regulatory powers away from the mining regulator (Uberti, 2014a). However, since these special executive powers were never exercised in practice, the 2010 reform did not effectively jeopardise the formal independence of the KPMM. The new law was also amended again in May 2013 (Law No. 04/L-158). Since the 2013 amendments were relatively minor, however, we do not distinguish between the 2010 law and the 2013 amendments in the analysis.

[16] Unlike in Albania, a substantial number of mining firms active in Kosovo (with the exception of firms owning small-scale operations) is foreign-owned. Since domestic banks provide little credit for large-scale industrial projects, firms typically rely on foreign banks for project financing. For this reason, we control for the rental cost of capital using the LIBOR rate, rather than the domestic interest rate (as ILU do). Premiums above the LIBOR rate typically take into account the risk of a mining project (which we assume to be randomly distributed) as well as the country risk (which we assume to be constant during 2001-18). A foreign-owned

respond to market signals, submit applications to KPMM and eventually obtain a permit. For both MPI and LIBOR, therefore, we include a vector of lags ranging from $N = -4$ to $N = -15$ (months). Our results are robust to omitting these (demand-side) variables from the model.[17]

The frequency of mining permit issued may also be subject to inter-temporal persistence. Information on mineral prospectivity (i.e. mining potential) may be subject to spill-over effects. Companies may also apply for permits as a way to emulate their competitors. Thus, we choose to specify a dynamic (autoregressive) model. In doing so, we follow the functional form recommended by Zeger & Qaqish (1988) and employed by ILU (2018). In the Zeger-Qaqish model, the lagged dependent variable is entered *in log form* to prevent the model from having an explosive behaviour. Zero values are rescaled to a constant $c = 0.5$ (so that $\ln y_t = \ln 0.5$ if $y_t = 0$). Once time dependence is accounted for, we find no evidence of trends or seasonality in the data. The dynamic specification of the conditional mean is thus:

$$\mu_t = \exp(\rho \ln y_{t-1} + \boldsymbol{\beta X}_t) \tag{2}$$

where ρ is a persistence parameter, and the population regression equation is as follows:

$$y_t = \exp\left(\rho \ln y_{t-1} + \beta_1 A_{-1t}^{sched} + \beta_2 A_{-1t}^{early} + \gamma_3 Indep_t + \gamma_4 Law2005_t + \gamma_5 Law2010_t \right. \\ \left. + \gamma_6 Suspend_t + \sum_{i=4}^{15} d_i MPI_{t-i} + \sum_{i=4}^{15} \delta_i LIBOR_{t-i}\right) \times \varepsilon_t \tag{3}$$

where the β's, γ's, δ's and d's are parameters to be estimated and ε_t is the model error. In an alternative specification (which is less common in the empirical PBC literature), we also include dummies denoting the first year after an (early or scheduled) election (A_{+1}) in order to control for any post-election effects.

Equation (3) may be used for non-linear regression analysis (NLS), but a more efficient option is to impose a distributional assumption (ε_t is i.i.d. Poisson) and estimate (3) by GLM or ML

firm indicated that, in their cash-flow calculations, they typically cost in an 8 percent interest rate, while return on capital ranges between 20-25 percent (Interview with mining company based in Prishtina, 9 September 2017). The MPI and LIBOR data are taken from the IMF (https://www.imf.org/external/np/res/commod/index.aspx) and the Federal Reserve Bank of St. Louis (https://fred.stlouisfed.org), respectively.

[17] Data on LIBOR and the metal price index for the last observations in the dataset (early 2018) were not yet available at the time of writing, leading to a loss of three observations when these variables are included.

(Cameron & Trivedi, 2013).[18] Consistency only requires a correct specification of the conditional mean. The GLM (or ML) t-statistics, however, are generally inflated if there is any distributional misspecification. Typically, autocorrelation leads to the sample variance being larger than the mean,[19] which violates the Poisson assumption of equidispersion. Indeed, the raw data are considerably over-dispersed, with the variance of y_t (34.8) being over five times the mean (6.9). In all the models reported below, we include as many lags of the dependent variable as is necessary fully to eliminate residual autocorrelation, as measured by the Ljung-Box (LB) portmanteau statistic (Cameron & Trivedi, 2013, p. 270).[20] The tables below report one or two lags of the dependent variable, depending on how many proved necessary to remove autocorrelation in the residual. Since some overdispersion may still persist even after the regressors (crucially, the lagged dependent variables) are included,[21] heteroskedasticity-robust standard errors are used throughout to support valid inference (Cameron & Trivedi, 2013, p. 282).

4. Main Results

Poisson GLM estimates of eq. (3) are reported in Table 2. Throughout specifications (1)-(5), the lagged dependent variable enters as positive and significant, confirming our choice of a dynamic specification. The estimated magnitude of the persistence effect (ρ) is very close to the estimates of ILU (2018; see their Table 1), suggesting that the same type of persistence mechanism may be operative in both Albania and Kosovo.

Pooling early and scheduled elections together (model 1), we find no evidence of election cycles in the distribution of mining licenses, in contrast to ILU's (2008) findings for Albania. In specification (2), which is our baseline, we distinguish between early and scheduled elections, as suggested by Khemani (2004). Consistent with expectations, the coefficient on A_{-1}^{early} is statistically insignificant and very close to zero. The model, however, estimates a statistically *significant* pre-election effect prior to *scheduled* elections, implying a rejection of H_0 (i.e. $cycle = 0$) at the 5 percent significance level. A Wald test rejects the equality of the coefficients on A_{-1}^{sched} and A_{-1}^{early} at the 10 percent significance level (p-value = 0.066).

[18] We use GLM. The ML and GLS estimators are identical when the conditional distribution is Poisson (Cameron & Trivedi, 2013, p. 76).
[19] This process known as 'true contagion' (Cameron & Trivedi, 2013, p. 161).
[20] The LB statistic is a weighted sum of the autocorrelation of the Pearson residuals. A higher number indicates higher residual autocorrelation.
[21] Residual over-dispersion may be a consequence of unobserved heterogeneity in the data (a process known as 'apparent contagion'). Although our data are generated by the same individual over time (the DMM and its successor agency, the KPMM), it is reasonable to assume that staff turnover at the agency may lead to different relative propensities to issue licenses over time.

In our view, the fact that the 2010, 2014 and 2017 contests were early, as opposed to scheduled, elections provides the best explanation for our failure to observe an opportunistic pre-election effect: early elections are by definition unforeseen, and incumbent politicians have very little time (only 45 days) to plan their opportunistic strategy and enlist the support of the bureaucracy for its execution. We considered, however, two additional explanations for the observed difference between the 2001/2004/2007 elections and the 2010/2014/2017 elections. However, we regard these explanations as implausible on a priori grounds. Firstly, it could be argued that Kosovo made substantial progress in democratic consolidation since 2008, limiting incumbent opportunism (Brender & Drazen, 2005). Yet, Kosovo has actually made very little progress in democratic consolidation beyond cosmetic changes (Cocozzelli, 2013; Schwandner-Sievers, 2013). Second, since all early elections took place in the post-independence period, the observed difference may be picking up a potential moderating effect of independence, with the transition to domestic rule supposedly alleviating electoral opportunism. This additional explanation cannot be tested explicitly due to perfect multicollinearity. Yet, it may be ruled out on a priori grounds. The end of international rule has, if anything, led to an *increase* in corruption and political discretion (Capussela, 2015). So, if the key difference was independence, rather than the early vs. scheduled distinction, then the 2010/2014/2017 early elections should have been associated with more (rather than less) political manipulation.

[Table 2]

Prior to scheduled elections, the extent of pre-electoral opportunism is substantial. The estimated coefficient (0.549) implies that in years preceding a scheduled election, the number of issued licenses increased on average by 73.1 percent relative to non-election years.[22] This corresponds to about 3.8 additional licenses per month, suggesting that, during 2001-18, electoral competition has led to as many as 46 new mining projects being licensed purely based on opportunistic political considerations. Interestingly, the pre-election effect in Kosovo (73.1 percent) is considerably *larger* than the effect estimated by ILU (2018) for Albania on the basis of similarly specified models (57.5 percent, from model 1 in their Table 1, or 37.3 percent from model 2 in Table 1).

To save on degrees of freedom, specification 3 omits the demand-side variables (the IMF metal price index and LIBOR lags), which entered as jointly insignificant in models 1 and 2 (see below). The coefficient on A_{-1}^{sched} is now less precisely estimated, but is still positive with a magnitude (0.433) that implies a 55.7 percent increase in the number of permits issued, compared

[22] $= 100 \cdot (e^{0.549} - 1)$

to a 29.2 percent increase estimated for Albania based on a similarly specified model with no demand-side variables (see model 3 in ILU, 2018: Table 1).

In specifications 1-3, the value of the rescaling constant ($c = 0.5$) is fixed. Model 4 allows the rescaling constant to be an additional parameter to be estimated – an alternative procedure recommended by Cameron & Trivedi (2013, p.282-3). The procedure consists of including an additional indicator variable for (lagged) zero counts, while estimating the autoregressive coefficient from non-zero counts only.[23] Although the estimated value of the rescaling constant (0.013) is statistically indistinguishable from zero (p-value=0.766) and much smaller than 0.5, the estimated coefficient on A_{-1}^{sched} is almost unchanged, indicating that the results do not depend critically on the handling of zero counts.

Lastly, column (5) reports a more extensively specified model that includes post-election year dummies, both of which prove to be insignificant. However, the estimated coefficient on A_{-1}^{sched} is now even larger and more precisely estimated, implying a 105.6 percent increase in the rate of licensing prior to scheduled elections in Kosovo, compared to an estimated increase of 46.9 percent in Albania based on a similarly specified model (ILU's model 5 in their Table 1). The coefficient on A_{-1}^{early}, by contrast, continues to be statistically insignificant. A Wald test rejects the equality of these two coefficients at the 1 percent level (p-value = 0.010). Figure 1 plots the fitted values from this model against the actual observations, whereas Figure 2 compares the fitted Poisson frequencies to the actual frequencies. With a pseudo-R^2 of 0.41, this model fits the data quite well.

The estimated coefficients on the control variables provide additional insights into the determinants of licensing decisions in Kosovo. Throughout specifications 1-5, we find no evidence that legal regime type (the specifics of the mining law in force) has any effect on the equilibrium quantity of permits issued (although the coefficient on the dummy for the suspension period is obviously negative and highly significant). This result runs counter to ILU's (2018) findings for Albania. Nevertheless, a very large (79-120 percent) increase in the monthly rate of licensing is observed, ceteris paribus, in the post-independence period. A possible explanation is that the settlement of Kosovo's status might have increased firms' propensity to invest in mining, pushing up demand for licenses and forcing the incumbent to relax the supply constraint. The 2008 declaration of independence was recognized by the US, Japan and most EU countries and is likely to have bolstered investor security.

Lastly, we computed linear combinations of the estimated coefficients on the lagged demand-side variables (not reported in Table 2 to save space). These linear combinations are always

[23] Equation (2) is re-written as: $\mu_t = \exp[\rho \ln \tilde{y}_{t-1} + (\rho \ln c)D_t + \boldsymbol{\beta X}_t]$, where $D_t = 0$ and $\tilde{y}_{t-1} = y_{t-1}$ if $y_{t-1} \neq 0$, and $D_t = 1$ and $\tilde{y}_{t-1} = 1$ if $y_{t-1} = 0$. Then, $c = \exp[(\rho \ln c)\,/\rho]$ is used to obtain an estimate of c.

statistically insignificant.[24] This result contrasts with ILU's findings for Albania – where output (input) prices have a positive (negative) effect on investment demand and the quantity of licenses issued, in line with theoretical predictions. A possible explanation is that the regulator in Kosovo is less responsive to firms' demand for licenses than in Albania. Alternatively, Kosovo-based firms may not respond to market signals to the same extent as Albania-based firms do.

[Figure 3]

To replicate fully the approach followed by ILU (2018), we also break the pre- and post-election years dummies (i.e. $A_{\pm 1}$) into their quarterly components and estimate a model including a dummy variable for each of the last four quarters before/after a scheduled/early election. The results are consistent with the findings reported in Table 2. A coefficient plot of the estimated election cycle parameters (the coefficients of the quarterly dummies) is shown in Figure 3.[25] The diagram illustrates pictorially the shape of the election cycle. It also suggests that the pre-election licensing spree is concentrated in the last six months prior to a scheduled election.[26]

To sum up, our findings are consistent with our expectation that, prior to early elections, the incumbent does not have enough time to enlist the help of the mining bureaucracy in the pursuit of its electioneering strategy. Prior to scheduled elections, by contrast, a statistically significant cycle effect (a rejection of H_a) indicates that the incumbent is indeed in a position to manipulate the allocation of mining licenses, in line with political business cycle theory. This result is particularly striking in the context of Kosovo, where the international community invested heavily in establishing a formal separation of powers. Yet, neither the formal independence of DMM and later KPMM, nor the large number of international staff employed at these agencies proved sufficient to prevent the opportunistic manipulation of mining policy. A possible explanation, we suggested, is that the members of the KPMM board may be linked to the ruling coalition through informal ties of political or personal loyalty.

5. Results by License Type

[24] Full results are available upon request.
[25] Full results are available upon request.
[26] The average value of the quarterly coefficients for the periods preceding *scheduled* elections (0.637, s.e. = 0.279, *p*-value = 0.023) is qualitatively consistent with the estimates obtained using annual pre-election dummies (see Table 1, model 5). Again, consistent with previous findings, the average value of the quarterly coefficients for the periods preceding *early* elections (-0.245, s.e. = 0.182, *p*-value = 0.179) is statistically indistinguishable from zero. We can reject the equality of these two averages at the 5% level (the *p*-value of the Wald test is 0.011).

Next, we investigate which particular types of license are favoured by politicians to influence election outcomes. To do so, we disaggregate the data by license type. The results are reported in Table 3. Columns (1) and (2) disaggregate the data by type of mineral. We distinguish between high-value added metallic and industrial minerals (e.g. lead, zinc, nickel) and low-value added construction materials (e.g. sand and gravel). Metallic and industrial minerals are typically associated with high capital and technology requirements and are often the preserve of multinational corporations. Construction materials, by contrast, are mined in hard-rock quarries and are less capital-intensive. Due to lower value-to-weight ratios, they are usually produced by local companies for domestic consumption. Columns (3) and (4), in addition, distinguish between *exploration* licenses, and *extraction* licenses. The former grant the licensee the right to prospect within a given license block, while the latter confer the right to begin production once a find has been confirmed and a production plan has been approved by the regulator.

[Table 3]

The results indicate that the incumbent's electioneering strategy (prior to scheduled elections) is based primarily on manipulating the allocation of *extraction* licenses for *construction* materials (see columns 2 and 4, Table 3). The explanation for this finding is two-fold. *First*, the construction industry has often occupied a pre-eminent position in post-socialist transition economies. In Kosovo, the construction sector gained further impetus from the post-war reconstruction efforts. During 2008-13, construction and real estate (which are typically labour-intensive sectors) accounted for some 16.2 percent of GDP.[27] As such, they provide the incumbent with an important lever to boost employment and improve its re-election prospects. In addition, the construction materials industry is composed exclusively of domestic firms and is known to be highly politicised.[28] Thus, it is plausible that firms involved in quarrying may mobilise local votes and influence for their political sponsors in exchange for licenses. Indeed, previous studies from Romania and Bulgaria have found evidence of electoral intimidation in the workplace (Mares et al., 2016). High-value added mining projects, by contrast, are typically foreign-owned (unlike in Albania) and relatively unimportant economically, as a large chunk of the former state-owned mining economy is yet to be privatised (Uberti, 2014b).[29] Ironically, perhaps, licenses for high-value added minerals are not particularly 'valuable' from the point of view of electioneering.

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[27] Authors' calculations based on data from *Statistical Agency of Kosovo,* National Accounts Statistics, 2015.

[28] It is not uncommon for politicians to own quarries, often located in their electoral strongholds. The trade in construction materials in Kosovo is also partly rooted in the legacy of war economies, with business opportunities arising from post-war reconstruction and supply contracts being distributed amongst former members of the same guerrilla factions.

[29] Notably, the Trepça mining and metallurgical complex in the northern city of Mitrovica.

Second, extraction (as opposed to exploration) licenses may be more valuable to mining firms due to the sunk costs incurred during exploration (i.e. the costs of drilling boreholes and ascertaining the quantity of exploitable minerals, and the costs of preparing a production plan). Consequently, the incumbent may have an incentive to create artificial scarcity in this particular license category, delaying the granting of a license until the upcoming election in order to extract rents from those mining firms that successfully hit upon a find.[30] Rents, in this context, may include campaign finance, local votes or other favours (e.g. jobs for party activists). We admit, however, that this second explanation is somewhat speculative.

Lastly, the results reported in Table 3 imply that *both* the 2008 declaration of independence *and* the introduction of the 2005 UNMIK law were followed by significant increases in exploration activities for high-value minerals. The estimated effect is economically large. On average, nearly three times as many licenses for high-value minerals (exploration and extraction) were issued in post-independence months as in the months preceding the 2008 declaration,[31] while the introduction of the 2005 law increased 13-fold the monthly frequency of licenses issued for high-value minerals.

Since these minerals are mined primarily by foreign companies, these findings suggest that the resolution of Kosovo's status made Kosovo more attractive as an investment destination for foreign mining companies. At the same time, the introduction of a regulatory regime based on international best practice (the 2005 law) had an independent salutary effect on investor security. This finding partly vindicates UNMIK's governance strategy for the mining sector. Although the institutions put in place by Kosovo's international administrators failed to eliminate political interference and electoral opportunism, they did succeed in 'providing incentives and security for foreign direct investment' (Large et al., 2003, p. 20).

6. Robustness Analysis

To check the robustness of our results, we also perform a simple falsification test based on regression discontinuity approaches. The results are reported in Table 4. For ease of reference, the first column reproduces column 5 of Table 2. In columns 2 and 3, the timing of elections (both early and scheduled) is spuriously shifted backwards by 6 and 12 months, respectively (e.g. the November 2001 election is now assumed to have taken place in May 2001). The results show that the estimated pre-election effect (scheduled elections) loses statistical significance (indeed, the sign is reversed), when a 'placebo treatment' is administered in lieu of an actually occurring election. Our test

[30] This interpretation of election cycles is proposed by ILU (2018).
[31] =exp(1.078) = 2.94

correctly rejects the hypothesis (that there is a pre-election effect) when the hypothesis is known a priori to be false, providing suggestive evidence of the power of the test.

[Table 4]

Lastly, we check the robustness of our results to alternative specifications of the conditional mean (eq. 2). The reason that this check is advisable is that the Zeger-Qaqish model $\mu_t = \exp(\rho \ln y_{t-1} + \boldsymbol{\beta X}_t)$ is only one of several alternatives and 'at this stage it is not clear which […] will become the dominant model for time series count data' (Cameron & Trivedi, 2013, p. 263). The tests are reported in Table 5. For ease of computation, all the models omit the demand-side variables (the lags of MPI and LIBOR) but include the two post-election dummies.

Column (1) reports a Zeger-Qaqish benchmark model estimated by Poisson GLM. The results are consistent with previous findings. Also, this model provides the best fit to the data (highest presudo-R^2). Model (2) handles residual over-dispersion by specifying the conditional distribution of the dependent variable to be negative-binomial (NB2) rather than Poisson (Cameron & Trivedi, 2013, p. 282).[32] Over-dispersion is modelled explicitly as a feature of the distribution rather than being 'corrected for' with heteroskedasticity-robust standard errors. Overall, the results are consistent, but the magnitude of the pre-election effect increases by about 16 percent, suggesting that our previous estimates may be conservative.

[Table 5]

Model (3) is obtained by logging both sides of equation (2), leading to $\ln \mu_t = \rho \ln y_{t-1} + \boldsymbol{\beta X}_t$. Since the dependent variable is now normally distributed, the parameters and standard errors may be estimated by OLS (even in small samples). This model, however, has been criticised for performing poorly (King, 1989) and has the disadvantage of producing predicted values that may be less than 0. Model (4) is the GLARMA (generalised linear autoregressive moving average) model proposed by Davis, Dunsmuir & Streett (2003). Serial correlation is handled by including the first lag of the Pearson residual from the corresponding static model: $\mu_t = \exp(\rho e_{t-1} + \boldsymbol{\beta X}_t)$, where e is the estimated Pearson residual. Lastly, model (5) is the INAR (integer-valued autoregressive) model proposed by Brannas (1995). The conditional mean is specified as $\mu_t = \rho y_{t-1} + \exp(\boldsymbol{\beta X}_t)$ and estimation is by Non-linear Least Squares (NLS). This model has the advantage of not requiring ad-

[32] In this model the over-dispersion parameter α is still large (0.404) and significantly different from zero (*p*-value = 0.000), even after the regressors (and the lagged dependent variable) are included. This indicates that over-dispersion does not result solely from 'true contagion' (i.e. serial correlation) but also from unobserved heterogeneity.

hoc adjustments for zero counts. It also avoids distributional assumptions. On the downside, NLS estimation is typically inefficient. Indeed, the estimated standard errors are uniformly larger than in any of the previous models.

In any case, our results are qualitatively unchanged throughout columns (3)-(5), suggesting that the estimated election cycle (prior to scheduled elections) is not an artefact of choosing the Zeger-Qaqish model. In fact, across all these alternative models, the estimated cycle effect prior to scheduled elections is *higher* than in the corresponding Zeger-Qaqish specification.

7. Conclusion

Using data from post-conflict Kosovo, we find evidence consistent with opportunistic election cycles in the allocation of mining licenses. We detect a statistical trace of behaviour that would otherwise be difficult to observe directly. Our interpretation of the evidence as the outcome of electoral manipulation is supported by interview evidence from related fieldwork research as well as by reference to published accounts of current developments in Kosovo, particularly with respect to the pervasive nature of clientelism as a deeply embedded mode of governance in Kosovo. Disaggregating the data by license type, we then find that the observed election cycle is driven primarily by the manipulation of extraction licenses for the mining of construction materials. As we argued, this is the category of licenses whose strategic manipulation has the potential to offer the greatest pay-off to incumbent politicians. The election cycle, however, is not observed prior to early elections, when time constraints make it difficult for the incumbent to enlist the support of the bureaucracy and tamper with the licensing process. We could not think of plausible factors that may be driving this effect heterogeneity other than the fact that the elections (not) preceded by a cycle effect were scheduled (early) elections. Our results are broadly in accordance with ILU's (2018) empirical findings from Albania, which is characterized by regular *scheduled* elections. In addition, our results support Khemani's (2004) earlier suggestion that empirical PBC studies should take care to distinguish between early and scheduled elections.

Our findings are particularly important, since Kosovo's licensing authority (unlike Albania's) was designed to be an independent, non-political agency. As such, it should have been immune to political interference. Yet, not only do we observe an election cycle in Kosovo too, but the estimated amplitude of Kosovo's cycle is systematically larger than Albania's. To explain this finding, we cited anecdotal evidence of patronage and politicisation in Kosovo's civil service. If recruitment at KPMM is also based on political loyalties, the incumbent may be able to rely on political insiders to manipulate the allocation of licenses and advance its opportunistic objectives. Clearly, the incumbent's ability to act opportunistically does not depend critically on the absence of institutional checks and balances (as is the case in Albania). These findings raise some questions about the feasibility of fighting political opportunism (and, relatedly, corruption) by establishing a formal separation of powers between elected politicians and regulatory agencies – a key element of the institutional reform agenda promoted by the donor community in mineral-rich countries (World Bank, 1996; Bastida, 2008).

A remaining issue concerns the interpretation of election cycles in mining licensing. What is the incumbent's motivation in fuelling an election-year licensing spree? How is this largesse expected to increase its chance of political survival? As mentioned in the introduction, ILU (2018) propose what may be termed an 'ecumenical' interpretation of election-year opportunism. As in traditional PBC theory, the incumbent may be seeking to engineer an economic expansion, which may in turn improve its popularity amongst voters. Alternatively, there may be an incentive to allocate mining rights to political supporters and party clients in exchange for bribes, votes, campaign contributions and/or other favours (e.g. jobs for party activists). This explanation, first proposed by Khemani (2004) in a PBC context, draws on the vast literature on rent-seeking in political economy (Khan & Jomo, 2000; North et al., 2009).

A 'rent-seeking' interpretation of PBCs may be particularly relevant in the following contexts: (1) If policy manipulation cannot lead to output and employment growth in the short-run (e.g. due to the long gestation periods of mining projects, or if firms apply for permits with a view to selling them on to third parties or to use them at a later stage); and/or (2) if relatively few voters stand to benefit from sector-level economic growth (e.g. due to capital intensity). If either or both of these conditions prevail, the pay-off to the incumbent from policy manipulation cannot come in the form of economic expansion and voter satisfaction. More plausibly, the electoral benefits of manipulation are likely to accrue to the incumbent more *directly* – in the form of bribes, political contributions and/or other favours.

An observable implication of the 'rent-seeking' theory (which is inconsistent with classic PBC theory) is that there should be an election cycles in the policy variable (e.g. the frequency of mining permits issued) with no corresponding cycle in the level of economic activity (e.g. in mining-sector investment, employment or output). While it is not possible to adjudicate between the traditional PBC and 'rent-seeking' interpretations based on our data, examining election cycles in related macroeconomic variables (e.g. mining value-added) may provide additional insights in future research. In addition, future research should explicitly investigate the effect of licensing policy manipulation on the incumbent's likelihood of re-election (as in Brender & Drazen, 2008). After all, only if policy manipulation does indeed improve the incumbent's re-election prospects can the researcher interpret the effect of elections on licensing activities as an *equilibrium* effect over a period of repeated interactions between rational politicians and voters (Khemani, 2004, p. 126).

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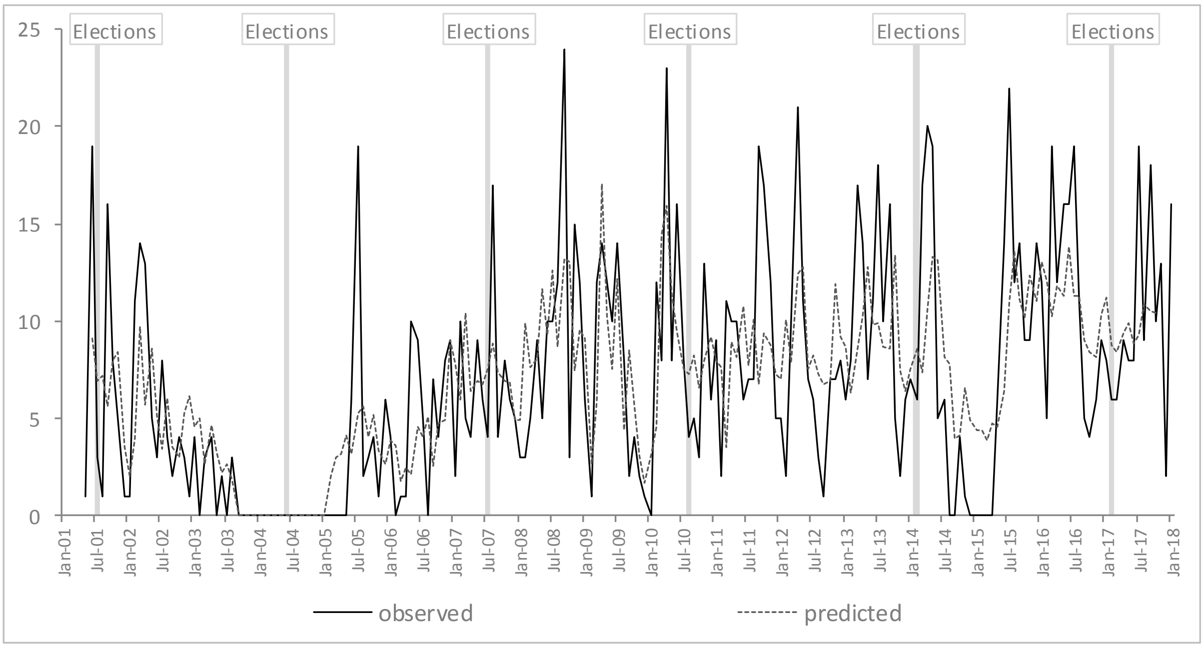
Table 1. The Mining Industry in Kosovo and Albania, 2013

	Kosovo	Albania
Mining industry GVA, % GDP	2.2	0.9
Mining industry GVA, % total industrial GVA	16.8	7.6
Mineral exports, % total exports	8.3	9.3

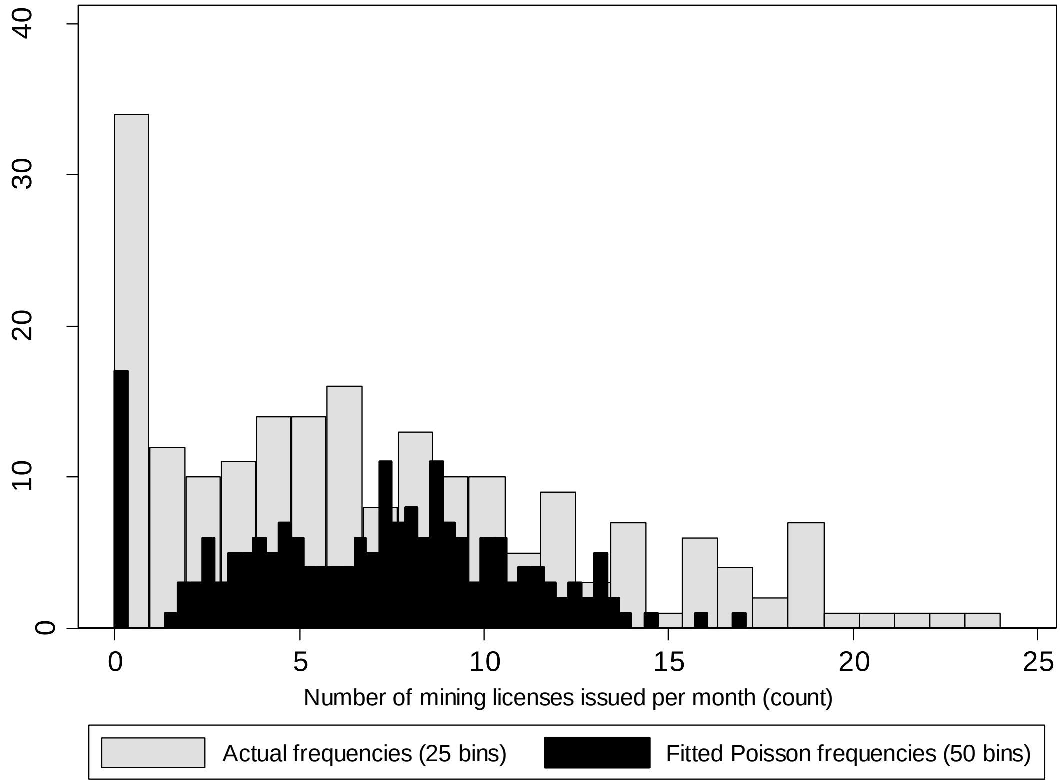
Notes: authors' calculations based on data from the Statistical Agency of Kosovo and INSTAT (Albania). GVA = gross value-added. The figures do not include the oil and gas industry. By ‘mining’ and ‘minerals’, we understand ores and concentrates.

Figure 1. Number of mining permits issued per month: time-series plot



Notes: the predictions are based on model (5) in Table 2.

Figure 2. Number of mining permits issued per month: actual and fitted frequency distributions



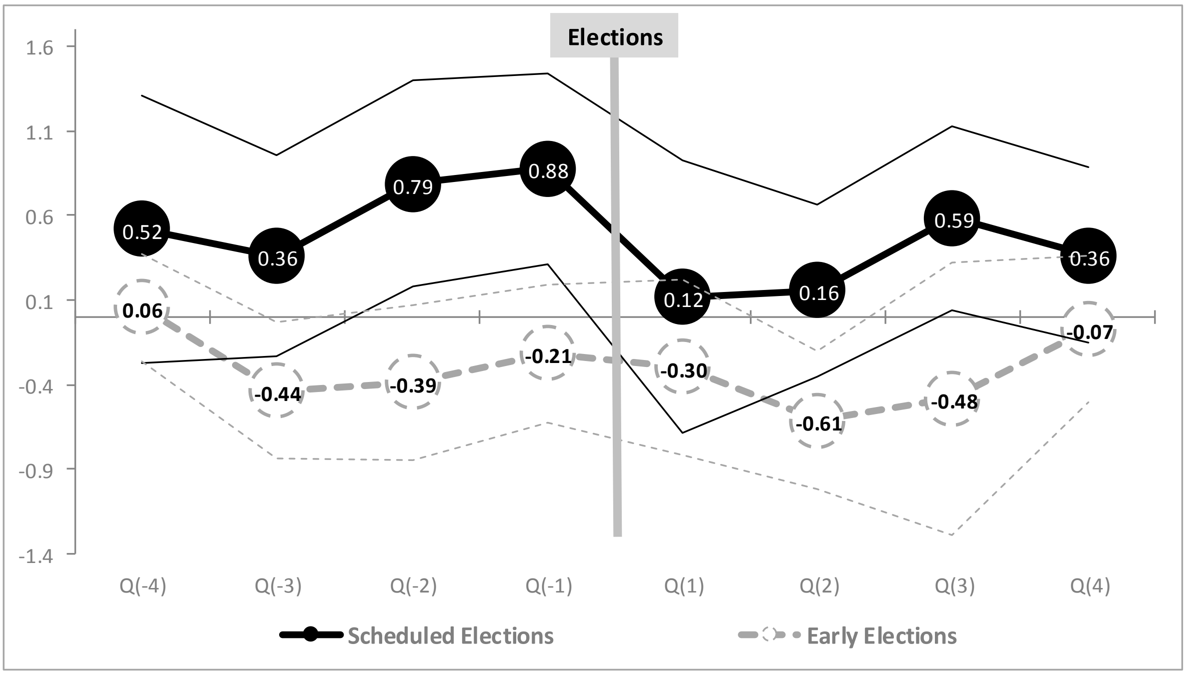
Notes: the fitted frequencies are based on model (5) in Table 2.

Table 2. Determinants of mining licensing: annual election dummies, Poisson GLM
Dependent variable: the number of mining licenses issued per month

	(1)	(2)	(3)	(4)	(5)
ln (y_{t-1})	0.263***	0.261***	0.244***		0.245***
	(0.064)	(0.061)	(0.067)		(0.062)
ln (y_{t-1}) (for $y_{t-1} \neq 0$)				0.175**	
				(0.075)	
Dummy (for $y_{t-1} = 0$)				-0.753**	
				(0.368)	
All Elections:					
A_{-1}	0.157				
	(0.128)				
Scheduled Elections:					
A_{-1}		0.549**	0.443*	0.536**	0.721***
		(0.261)	(0.275)	(0.255)	(0.279)
A_{+1}					0.299
					(0.231)
Early Elections:					
A_{-1}		-0.026	0.014	-0.055	-0.166
		(0.115)	(0.118)	(0.145)	(0.174)
A_{+1}					-0.254
					(0.195)
Control Variables:					
Independence	0.407*	0.677**	0.583**	0.704**	0.792***
	(0.222)	(0.290)	(0.255)	(0.282)	(0.294)
Law 2005	-0.078	-0.096	-0.085	-0.143	-0.287
	(0.447)	(0.464)	(0.258)	(0.470)	(0.456)
Law 2010	-0.067	-0.174	-0.069	-0.357	-0.346
	(0.754)	(0.778)	(0.291)	(0.794)	(0.757)
Suspension Period	-16.2***	-16.4***	16.4***	-16.1***	-16.6***
	(0.389)	(0.432)	(0.419)	(0.468)	(0.439)
Demand side-variables:	Yes	Yes	No	Yes	Yes
c (st. error, if estimated)	0.5	0.5	0.5	0.013 (0.045)	0.5
Ljung-Box statistic [*p*-value]	5.72	6.26	5.50	8.65	6.96
	[0.838]	[0.793]	[0.856]	[0.565]	[0.729]
Pseudo R^2	0.38	0.39	0.32	0.41	0.41
Observations	197	197	200	197	197

Notes: *** p<0.01, ** p<0.05, * p<0.1; robust standard errors in parentheses. The lags (from -4 to -15) of the demand-side variables (Metal price index and LIBOR), and the constant term, are not reported to save space. c is the constant used to rescale the zero values of the dependent variable before logging. The *Ljung-Box statistic* is based on the first 10 autocorrelations of the Pearson residuals. The *Pseudo R^2* is the squared correlation between observed and fitted values.

Figure 3. Determinants of mining licensing: quarterly election dummies (coefficient plot), Poisson GLM



Notes: the coefficients and 90% confidence intervals are based on a fully specified models that includes the demand-side variables. The Ljung-Box statistic is 6.34 (*p*-value = 0.786) and the Pseudo-R^2 is 0.44. Full results are available upon request.

Table 3. License sub-sets, Poisson GLM
Dependent variable: the number of mining licenses issued per month

	Type of mineral		Type of license	
	High-value	Low-value	Exploration	Extraction
	(1)	(2)	(3)	(4)
ln (y_{t-1})	0.043	0.229***	0.231***	0.117*
	(0.032)	(0.067)	(0.079)	(0.070)
ln (y_{t-2})	0.101***		0.138*	
	(0.032)		(0.072)	
Scheduled Elections:				
A_{-1}	0.564	0.737**	0.598	0.818***
	(0.533)	(0.296)	(0.463)	(0.315)
A_{+1}	0.180	0.284	0.153	0.482
	(0.391)	(0.246)	(0.295)	(0.295)
Early Elections:				
A_{-1}	-0.140	-0.176	-0.247	-0.109
	(0.254)	(0.205)	(0.199)	(0.224)
A_{+1}	-0.350	-0.326	-0.222	-0.515*
	(0.275)	(0.215)	(0.235)	(0.263)
Control Variables:				
Independence	1.078**	0.720**	1.441***	0.068
	(0.490)	(0.306)	(0.415)	(0.329)
Law 2005	2.603**	-0.580	1.412**	-0.901
	(1.120)	(0.486)	(0.626)	(0.604)
Law 2010	2.730**	-0.752	1.291*	-0.768
	(1.373)	(0.821)	(0.835)	(1.089)
Suspension Period	-11.9***	-17.9***	-15.5***	-17.4***
	(1.123)	(0.460)	(0.603)	(0.487)
Demand-side Variables	Yes	Yes	Yes	Yes
Ljung-Box statistic [*p*-value]	12.38	7.76	11.97	14.13
	[0.260]	[0.652]	[0.287]	[0.167]
Pseudo R^2	0.26	0.37	0.35	0.14
Observations	196	197	196	197
N. of licenses issued (2001-18)	347	1048	754	641
[fraction of total n. of licenses]	[0.25]	[0.75]	[0.54]	[0.46]

Notes: *** p<0.01, ** p<0.05, * p<0.1; Robust standard errors in parentheses. The demand-side variables (metal price index and LIBOR) and the constant term, are not reported to save space. Low-value minerals refer to construction materials (sand, gravel, etc.); high-value minerals include metallic and industrial minerals (e.g. lead, zinc, nickel).

Table 4. Falsification test, Poisson GLM
Dependent variable: the number of mining licenses issued per month.

	(1)	(2)	(3)
	From Table 2, Col.5		
ln (y_{t-1})	0.245***	0.264***	0.239***
	(0.062)	(0.063)	(0.061)
Scheduled elections:			
A_{-1}	0.721***	-0.101	-0.409
	(0.279)	(0.290)	(0.315)
A_{+1}	0.299	0.237	0.305
	(0.231)	(0.235)	(0.324)
Early Elections:			
A_{-1}	-0.166	0.114	0.202
	(0.174)	(0.165)	(0.188)
A_{+1}	-0.254	-0.187	0.034
	(0.195)	(0.154)	(0.153)
Control Variables:	Yes	Yes	Yes
Demand-side Variables	Yes	Yes	Yes
Pseudo R^2	0.41	0.39	0.39
Observations	197	197	197

Notes: *** p<0.01, ** p<0.05, * p<0.1; Robust standard errors in parentheses. The other control variables, the demand-side variables (Metal Price Index and LIBOR) and the constant term, are not reported to save space.

Table 5. Alternative econometric specifications
Dependent variable: number of mining licenses issued per month (the log of this number in column 3)

Conditional mean	Zeqer-Qaqish	Zeqer-Qaqish	Log-lin	GLARMA	Brannas INAR
Distribution	Poisson	NB2	Normal	Poisson	
Estimator	GLM	ML	OLS	GLM	NLS
	(1)	(2)	(3)	(4)	(5)
ln (y_{t-1})	0.231***	0.234***	0.369***		
	(0.068)	(0.069)	(0.081)		
e_{t-1}				0.086***	
				(0.025)	
y_{t-1}					0.262***
					(0.077)
Scheduled Elections:					
A_{-1}	0.553**	0.642**	0.592**	0.731***	0.759**
	(0.279)	(0.272)	(0.269)	(0.255)	(0.379)
A_{+1}	0.190	0.176	0.237	0.236	0.227
	(0.184)	(0.199)	(0.219)	(0.191)	(0.279)
Early Elections:					
A_{-1}	-0.012	0.004	0.003	-0.025	-0.048
	(0.128)	(0.134)	(0.163)	(0.128)	(0.174)
A_{+1}	-0.133	-0.162	-0.404*	-0.257	-0.262
	(0.159)	(0.164)	(0.226)	(0.164)	(0.220)
Control Variables:					
Independence	0.630**	0.638***	0.599**	0.848***	0.883**
	(0.255)	(0.241)	(0.272)	(0.238)	(0.348)
Law 2005	-0.064	-0.064	-0.004	-0.064	-0.146
	(0.259)	(0.251)	(0.274)	(0.248)	(0.376)
Law 2010	0.060	0.030	0.227	0.128	0.054
	(0.306)	(0.308)	(0.348)	(0.301)	(0.442)
Suspension period	-16.6***	-17.0***	-1.439***	-16.9***	-
	(0.423)	(0.387)	(0.318)	(0.351)	-
Constant	1.097***	1.097***	0.504*	1.301***	1.062***
	(0.236)	(0.228)	(0.274)	(0.200)	(0.332)
Demand-side variables	No	No	No	No	No
Pseudo R^2	0.323	0.316	0.301	0.304	0.313
Observations	200	200	200	200	200

Notes: *** p<0.01, ** p<0.05, * p<0.1; robust standard errors in parentheses. In model 5, the parameter for the suspension period dummy is taken as the constant term in the model.