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The political economy of local government in Croatia: winning coalitions, corruption, and taxes

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Abstract

This paper represents the first comprehensive effort to provide a theoretical and empirical explanation of systemic corruption in Croatian local government. It follows the logic of the selectorate theory, according to which staying in power for long periods of time depends on creating a small group of loyal but powerful supporters (the winning coalition). Mayors that exist within such environments not only maximize their chances of staying in power; they also engage in greater corruption and set higher taxes. Its citizens are stuck in a negative spiral of corruption, high taxes, and a politician that regardless of this keeps winning elections. The paper makes two main contributions to the current literature. First it provides a theoretical extension of the selectorate theory to Croatian local government by explicitly modeling the link between corruption and winning coalitions, and second, it empirically verifies the theoretical findings using a novel matching approach called entropy balancing.

Keywords: political economy, winning coalition, selectorate theory, corruption, taxes

1 INTRODUCTION

Corruption in Croatia is systemic, embedded deeply in the core of both national and local politics, and consequently of the domestic economy. Politicians holding power have developed and successfully maintained a crony system that helps them win office and reward their loyal cronies with exclusive deals and political protection. The result is a system where competitive advantage on the market is achieved mainly because of political connections instead of competence. Existing within such a crony microcosm enables a politician freely to extract rents, not having to worry about being punished by the voters.

The best way to illustrate this is with a few examples. Over the past two electoral cycles a number of mayors from major Croatian cities (Zagreb, Dubrovnik, Zadar, Varaždin, Vukovar, and the Sisak-Moslavina County) originating from both center-left and center-right political parties were arrested or indicted on corruption charges for abuse of power, bribery, and making favorable deals with selected private firms. They were either convicted or are still in trial. The Dubrovnik mayor was found guilty, served his sentence of six months of probation and then came back to win a local by-election in 2015 (Tomičić, 2015). The coalition government passed a special law in the months prior to the regular 2017 local election with the sole purpose of preventing him from running and most likely winning again (Romić, 2016). The mayor of Zagreb, in power for 16 years, after being arrested was immediately bailed out after paying the largest ever bail of 15 million kuna, arrested again for breaking his probation by intimidating witnesses, bailed out again after a few months thanks to a very quick Constitutional Court decision in his favor, and is still due to face trial on various charges (Petrušić, 2015). None of this prevented him from running in the parliamentary elections in 2015 (regular) and 2016 (repeated) with his new party, both times being elected as a member of

parliament, or from winning a new four-year term as mayor in the 2017 local elections. For this entire period, even when he was in jail for six months, he was still the official mayor of Zagreb. These are just two of the most egregious cases that suggest there is something very wrong with how Croatian voters perceive corrupt politicians.

Why don't voters in Croatia punish their politicians for corruption? Many Croatian politicians have held their local offices for more than 20 years (the average tenure is over 10 years!). Have these politicians managed to design a system of key supporters so powerful and so efficient that even after being arrested or even convicted, they can still remain in power? Can political economy theory in any way explain these outcomes in which politicians in Croatia can stay in power for long periods of time with virtually no accountability to their voters? Who are they accountable to if not the voters?

This paper aims to provide both a theoretical and empirical explanation of systemic corruption in Croatian local government. The underlying theoretical logic of the paper is based on the *selectorate theory* from Bueno de Mesquita et al. (2002, 2005). The selectorate theory suggests that politicians can stay in power for long periods of time and simultaneously extract rents and shun any accountability if they create a powerful enough group of key supporters. This relationship provides mutual benefits to all players: the politicians stay in power while the supporters get direct benefits, either in favorable legislation, exclusive procurement contracts, jobs in the public sector, etc. (Shleifer and Vishny, 1994; Verdier, 1995; Root and Nellis, 2000; Faccio, 2006). Accordingly politicians are free to extract rents without hurting their election chances if they operate in *small winning coalition environments* within large electorates. In other words, the smaller the group of key supporters that delivers votes with respect to the entire electorate, the easier it is for a politician to bind the networks of key supporters to himself and hence maximize his chances of staying in power. Another unique prediction of the selectorate theory is that in small winning coalition environments a politician is free to impose higher taxes in order to fund the allocation of public resources towards private interests. Theoretical implications similar to those of the selectorate theory were given by North, Wallis and Weingeist (2009). They define such an environment as a typical *limited access order* in which institutions are personalized so that all political and economic outcomes depend on close interpersonal relationships between the ruling elites and various powerful groups that support them.

The empirical part of the paper proves the given theoretical predictions. It shows that Croatian systemic corruption can indeed best be explained by the selectorate theory, where longevity in office is dependent upon having small winning coalition environments and hence greater opportunities for corruption. Furthermore, mayors in small winning coalition cities or municipalities set higher local taxes. The greater the corruption of a mayor, the higher the tax rate he or she sets. Therefore, a mayor operating within a small winning coalition environment is likely to

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stay in power longer, be more corrupt, manipulate corruption to keep himself in office, and keep tax rates high. Lack of accountability in such environments will enable the mayor to get away with it, thus sending the entire community down a negative spiral of poor governance and loyal supporters of bad politicians.

The paper delivers two main contributions. First, it provides a theoretical extension of the Bueno de Mesquita et al. (2005) theory by establishing a direct link between the winning coalition and corruption. The selectorate theory accounts for this link indirectly, by explaining the motivation for higher corruption in small winning coalition systems. Leaders within small winning coalitions use corruption to reward their crony supporters, they use it to finance their own kleptocratic aspirations, and they discourage any attempt to reduce corruption as this strengthens the rule of law, thus weakening their hold on power. The theory therefore does predict a greater motivation for corruption, but it does not model it explicitly nor does it test empirically. This is the second contribution of the paper – its empirical verification of the theory and its main findings in the case of Croatia. It uses a direct proxy for corruption that confirms the motivation provided by the original selectorate theory.

In the following sections the paper first presents the underlying logic of the selectorate theory in section 2, after which it creates a new model in section 3 to explicitly define the relationship between winning coalition size and corruption. These two sections offer direct testable hypotheses for the empirical part of the paper. Section 4 lists the sources of data and describes the process of coding the main variables. Section 5 presents the empirical strategy and the subsequent results for each of the main theoretical implications. Section 6 concludes.

2 THEORETICAL FRAMEWORK: SELECTORATE THEORY

According to the selectorate theory, defined and introduced by Bueno de Mesquita et al. (2002, 2005), politicians win office and stay in power for long periods of time if they are integrated enough into the system, or in other words, if they create a powerful enough group of supporters/voters they can buy with various concessions, and which in return offer them persistent political support – the so called *winning coalition* (W).

A winning coalition represents a group of voters essential for the politician to keep hold of power. This may include various influential organizations and interest groups such as unions, religious groups, civil society groups, entrepreneurs, or even the police and the army in autocratic regimes. A winning coalition is a subset of a larger group called the *selectorate* (S) representing all voters eligible to vote in a society. The smaller the winning coalition with respect to the size of the selectorate (the so-called *W/S ratio*), the greater the chances of political survival. This is due to the fact that members of small winning coalitions can easily be replaced by members outside of the coalition if they fail to remain loyal. The costs of defection of members within the winning coalition are too large (they lose their privi-

leges and benefits). This testifies to a mutually dependent relationship between political elites and the groups they choose to include in their winning coalitions.

The main emphasis of the selectorate theory is to explain how the existence of a small coalition in a large electorate induces a lack of political accountability to the voters, which further leads to a lack of transparency in public decision making, the personalization of institutions, and the creation of clientelistic relationships that secure re-election. In environments with really small winning coalitions and large selectorates, like autocracies, politicians in power use private goods to satisfy their coalition members. In larger coalition environments, like democracies, politicians use public goods to satisfy a wide enough number of citizens for electoral support, which often implies an interconnection with corporate interests in the form of campaign donations and lobbying, and returning the favor via favorable legislation (McChesney, 1997; Grossman and Helpman, 2002; Faccio, 2006). The logic of political survival is the same in both systems; create a group of essential people powerful enough to secure reelection, and reward them with either direct monetary benefits, or indirect concessions.

Politicians in democracies can also use the support of so-called *key members* of the group within their winning coalitions. These are individuals who influence a large number of people and thus carry a greater weight in the electoral process. For example, politicians in democracies use the support of labor union leaders, which secures large amounts of votes by utilizing union membership. Politicians return the favor by rewarding the union leader, which effectively reduces the cost of maintaining the winning coalition. Despite the perception of a large coalition of voters who support them, they actually have a very small coalition made up of only a few key members. Other examples are military generals in autocracies, religious leaders, or even representatives of various ethnic groups, which in certain countries (or areas) are "worth" a significant number of votes.

Business interests can also influence the creation of small winning coalitions, where politicians reward entrepreneurs within the coalition by giving them exclusive government contracts via fraudulent public procurement procedures (Bandiera, Prat and Valletti, 2009; Ferraz and Finan, 2011). In return, the quasi-entrepreneurs, as key members of their own networks, use their influence to secure votes for the politicians. It is in the best interest of such corruptly-intralinked quasi-entrepreneurs for their politician to remain in power as long as possible (Verdier, 1995; Root and Nellis, 2000).

One of the main drawbacks of the selectorate theory is that its assumptions rest upon rational choice theory which assumes complete information and perfectly rational agents. This implies that not all of its theoretical findings will be supported in the data. In line with that critique Clarke and Stone (2008) find that its empirical results do suffer from an omitted variable bias, while Gallagher and Hanson (2015) find that the application of the theory to case studies is selective

and that it fails to fully explain outcomes in autocratic regimes. Nevertheless, even though the theory is not a magic bullet for understanding democracies and dictatorships, it does offer an excellent framework for a testing ground.

An approach similar to the selectorate theory was developed by North, Wallis and Weingast (2009) through their theory of the institutional development of a society with respect to its response to violence. The authors define an institutional framework in which well-organized ruling elites can manipulate the economy by generating privileges based on the personalization of governing institutions. They call it a limited access order. In such systems the base of organization in a society is intrapersonal relationships between the powerful and the political elites, which successfully discourage the development of civil society and ensure long-term persistence of the constellation of political relationships. When the institutions of a system are depersonalized it is much harder to create clientelistic relationships. This is the case of an open access order, in which the foundation of intrapersonal interactions is a well-defined, depersonalized legal framework, and not politically generated privileges. Ivanković and Šonje (2011) use the framework set out in North et al. (2009) to define Croatia as a system of "undemocratic capitalism" in which institutions are personalized and there is a delicate balance of power between politics and politically-generated winners.

Croatia is therefore a good candidate for analysis under the selectorate theory framework, particularly its local government, in which mayors are allowed to stay in power indefinitely. In addition, the new electoral system imposed in 2009 redefined electoral rules to allow mayors to be elected directly (prior to 2009, voters voted for parties, where the mayor was usually the leader of the local party branch). This created some institutional limits for mayors, but it arguably gave them even more direct power than before.

Furthermore, Croatia is a country prone to corruption and budgetary misappropriation. It ranks among the five most corrupt EU member states (along with Greece, Italy, Bulgaria, and Romania) (European Commission, 2014), while Transparency International (2016) gives it an average score of 49 out of 100, making it a mostly corrupt country. There is ample research done in Croatia that confirms this. Bićanić (1997), Franičević (1997), Madžarević-Šujster and Mikulić (2002), Bejaković (2002), and Ott (2002) have all tried to measure and explain the extent of Croatia's grey economy and its systemic corruption. Štulhofer (2004) analyzes its impact on Croatia's depleting social capital and Budak and Rajh (2012) on the trust in the country's institutions. Bađun (2011) has shown a positive correlation between public investment spending and corruption, Ateljević and Budak (2010) successfully uncover corruption practices in domestic public procurement tenders, while Mačkić (2014) finds evidence of local political business cycles in Croatian cities that indicate opportunistic and rent-seeking behavior of politicians. Each of these research efforts painted an individual piece of the puzzle of why Croatia is so vulnerable to political distortions, poor institutions, corrup-

tion, and lack of transparency. This paper aims to connect all the dots and provide an explanation rooted in the selectorate theory that can explain Croatia's persistently bad outcomes when it comes to dealing with bad politicians.

3 MODEL

One thing that is missing from the theoretical predictions of the selectorate theory is an explicit link between small winning coalitions and corruption. It confirms empirically that winning coalition size matters for taxation and political longevity in office but apart from justifying the motivation for corruption in kleptocratic regimes, it does not include a precise prediction on how winning coalition size is suppose to impact the potential corruption of politicians in power in democratic systems. This section combines the local government political alignment model of Brollo and Nannicini (2012) with the simple probabilistic voting model of Persson and Tabellini (2000) in order to make a theoretical extension to the selectorate theory.

The model starts by defining the incumbent mayor objective function. Politicians holding local power engage in corruption for two reasons: (1) to maximize their welfare (via obtaining rents), and (2) to maximize their probability of winning by buying the electoral support of key members of the winning coalition (Besley, 2006; Ferejohn, 1986; Brennan and Buchanan, 1980). The objective function of the mayor in power can then be defined as:

$$W^A\left(r_i\right) = r_i + \Pr\left[V_t^I \geq \frac{1}{2}\right] + ty\left(r_i\right) \tag{1}$$

Where r_i represents total rents received by a politician i during the entire term in office. The second term denotes the probability of winning for the politician (vote share V_t^I for the incumbent I in election year t needs to be higher than 1/2 to win). This is a probabilistic function that is further defined below. The final term represents a budget constraint, where t denotes the local tax rate[1], and $y(r_i)$ is total income, which is by definition a negative function of rents. This means that greater rent extraction will lower incomes of the population, encouraging a negative reaction from voters which lowers political utility. The budget constraint therefore represents the opportunity cost of rents.

Before performing the maximization of this objective function with respect to rents, we must first express the probability of winning for the local incumbent party as a function of vote share and rent-extraction. The vote share of the incumbent party can be expressed by:

$$V_t^I = V_{t-1}^I - \frac{W}{S}r_i^2 + \mu \tag{2}$$

The first term V_{t-1}^I is the vote share of the incumbent party in the previous election depicting the simple incumbency advantage (a typical implication of the political

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[1] In Croatia this is the income tax surcharge rate (*prirez*) that is used to finance the local government.

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science literature, e.g. Gelman and King, 1991; Ansolabahere and Snyder, 2002; or Lee, 2008). The final term is a random stochastic shock ε distributed normally on $\mu \sim \left[-\dfrac{1}{2\sigma}, \dfrac{1}{2\sigma} \right]$, and it represents all the possible factors that could influence electoral chances apart from rents (corruption) and the incumbency advantage (such as ideological preferences of the voters, specific politician idiosyncrasies, informational asymmetry, better campaign, more experience, etc.). The middle term represents a unique prediction of the selectorate theory (Bueno de Mesquita et al., 2005): negative rents squared (following the logic of concave rent-extraction, according to which too much in rents can reduce vote share at the election) multiplied by the ratio of the winning coalition (W) over the entire selectorate (S). The smaller the winning coalition with respect to the selectorate (i.e. for low levels of W/S), the greater the scope for corruption, and the higher the probability of winning. Small coalition environments are characterized by low levels of transparency and political accountability, and hence greater scope for corruption. On the other hand, in such environments it is much easier for an incumbent politician to gather a small winning coalition of key supporters, enough to keep him in power, whilst distributing budgetary funds both to them and himself. Therefore in equation (2) a small W/S ratio will imply a lower punishment for corruption, whereas a large W/S ratio implies greater punishment, reducing the incentives for corruption. We should observe higher corruption in small winning coalition environments, and lower corruption in large winning coalition environments.

The probability of winning for the mayor is the probability that more than half of the electorate[2] votes for him; $V_t^I \geq \dfrac{1}{2}$.

$$P_w^I = \Pr\left[V_t^I \geq \frac{1}{2} \right] \tag{3}$$

$$= \Pr\left[V_{t-1}^I - \frac{W}{S} r_i^2 + \mu \geq \frac{1}{2} \right] \tag{4}$$

Given that this is a probabilistic voting function based on Persson and Tabellini (2000), the winning probability becomes a particular point on the cumulative distribution function of μ:

$$= \Pr\left[\mu \geq \frac{1}{2} - V_{t-1}^I + \frac{W}{S} r_i^2 \right] \tag{5}$$

$$= 1 - \frac{\left[\dfrac{1}{2} - V_{t-1}^I + \dfrac{W}{S} r_i^2 \right] + \dfrac{1}{2\sigma}}{\dfrac{1}{\sigma}} \tag{6}$$

_____

[2] In reality, with multiple party races it takes much less than 50% of the electoral vote to win. A random parameter could have been used instead of the ½ threshold, however this would not have changed the theoretical prediction.

Which yields the following probabilistic function continuous in r_i:

$$\Pr\left[V_t^I \geq \frac{1}{2}\right] = \frac{1}{2} - \sigma\left(\frac{1}{2} - V_{t-1}^I + \frac{W}{S}r_i^2\right) \tag{7}$$

We can now insert this into the government objective function from equation (1), and formulate the following maximization problem of the incumbent:

$$\max_{r_i}\left[r_i + \frac{1}{2} - \sigma\left(\frac{1}{2} - V_{t-1}^I + \frac{W}{S}r_i^2\right) + ty\left(r_i\right)\right] \tag{8}$$

The first order condition is:

$$1 - 2\sigma\frac{W}{S}r_i + ty'\left(r_i\right) = 0 \tag{9}$$

For which we get:

$$r_i = \frac{ty'\left(r_i\right) + 1}{2\sigma\dfrac{W}{S}} \tag{10}$$

which holds if $\sigma \neq 0, \forall \sigma \in <0,1>$, and if $W, S > 0$. Total rents can therefore be expressed as a function of the opportunity cost of rents, the W/S ratio, and the distribution of the asymmetric shock, σ. According to equation (10), rents increase when the winning coalition is small $W \lim 0$, adhering to the logic of the selectorate theory, and when asymmetric shocks are large $\sigma \lim 0$. The distribution of $\varepsilon \sim u\left[-\dfrac{1}{2\sigma}, \dfrac{1}{2\sigma}\right]$ implies that a σ value approaching 0 entails a wide shock, which could mean a greater difference in political abilities of candidates, a high level of informational asymmetry (lack of transparency), or high levels of ideological uniformity which also induces lack of transparency and hence encourages rent-extraction. The implication is that, for example, with large differences in ideology between the incumbent and the challenger in an ideologically uniform city or municipality, mayors can get away with corrupt activities more easily. The empirical section approximates σ as political polarization which measures ideological uniformity of a city or municipality.

The model allows us to draw several important empirical implications from this extension to the selectorate theory. The first is that a small W/S ratio is good for rents, meaning that we should observe *greater corruption in small winning coalition environments*. In addition, higher levels of political polarization, higher local taxes, and greater total income should all increase the scope for corruption. This is explicitly tested and proven in the empirical section. Second, the selectorate theory itself implies that *small winning coalition environments are supposed to result in higher taxes and longer years in power* of an incumbent politician. This too, although not modeled, is explicitly tested for and proven in the empirical part of the paper.

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4 DATA AND VARIABLES

The dataset comprises several main variables: the W/S ratio, proxy for corruption, local income tax surcharge rates, political re-election, and total years in power. This section describes the data sources as well as the operationalization of each variable (a table of all variables used, summarizing their definitions and sources is available in the appendix).

The main independent variable in testing each theoretical hypothesis is the *W/S ratio*: the winning coalition divided by the entire selectorate. While measuring the size of the selectorate (S) is rather straightforward (total number of eligible voters), measuring the winning coalition requires some approximation. One way of doing so would be to look at the absolute number of voters who voted for the incumbent mayor. Defining a winning coalition (W) this way does not necessarily include only politically-connected voters, but it does show how many voters the incumbent managed to attract, some of whom were certainly politically connected. In other words the absolute number of voters who voted for the incumbent includes a subset of politically-connected voters, but not every vote for the mayor came from his loyal cronies. Bueno de Mesquita et al. (2005) approximate the winning coalition in a similar way, particularly in cases of democratic countries. The *W/S ratio* therefore captures the ratio of votes for the incumbent candidate over the total number of eligible voters in a local unit. It is a continuous variable with a mean of 0.27 and standard deviation of 0.1. The mean value of 0.27 is a good representation of what the W/S ratio actually stands for, as it implies that an incumbent needs only 27% of those eligible to vote to cast their vote for him or her to win a local unit. This means that, on average, politicians have a rather low voter base they need to satisfy in order to win. For example, the minimum value of 3.4% is a particularly worrisome number, given that politicians in this city or municipality can easily manipulate a small subset of voters into their winning coalitions. The point of the ratio is therefore to show how likely it is for a crucial group of politically-connected voters to generate more clout in local low-turnout environments.

In addition to taking the W/S ratio at face value it makes sense to separate it based on its median value (0.26) in order to distinguish between large and small W/S environments, and thus code them as an indicator as well (*W/S_binary*). The median is taken so as to generate a binary indicator with exactly one half of all cities and municipalities on either side of the cutoff. This will be used as a robustness check of the validity of the indicator.

Corruption in itself is very difficult to measure, primarily since anyone engaged in corrupt activities has every incentive to hide them. The proxy for corruption was taken from Vuković (2017) at the level of 556 cities and municipalities in Croatia. The original source of this data was the Official Gazette of the Republic of Croatia, and its registry of public procurement. It contained more than 35,000 contracts signed during a single term in local office, from 2009 to 2013. The total value of

all contracts was 28,765 billion kuna. Each corruption proxy was created as the ratio of suspicious procurement contracts to total procurement contracts in a given city or municipality for the entire 4-year period.

The definition of a suspicious procurement entails several different categories. The main three characteristics of a corrupt procurement include:
  – cases in which firms with zero employees won multiple public tenders (fake firms),
  – cases in which the value of the procurement contracts exceeds the annual average revenues of a firm (several categories are used), and
  – cases in which the value of the contract is being given to firms with losses, where the contract covers the entire loss and enables the firm to operate at a profit in the following year.

In this paper only the total index of corruption for each city or municipality was used, which is the weighted average of all previous categories of corruption (and their extrapolations), designed carefully so as not to include the same value more than once (labeled *Corruption Index*). The total index thus represents the finite level of corruption in each local unit, accounting for all potential frauds.

It should be noted however that this index does not measure corruption directly (given that this is impossible without data on criminal investigations and court rulings), but it only offers an approximation of corruption, where any case of potentially fraudulent procurement offers a reasonable doubt as to the validity of the procedure of assigning the contract. Further justification of the usage of this index can be found in Vuković (2017), who cites interviews with local govern-ment and law enforcement officials, public sector bureaucrats, and entrepreneurs in legitimizing the usage of the aforementioned characteristics in describing potential corruption.

The next set of data was local electoral results. These were taken from the State Electoral Commission of the Republic of Croatia (DIP, 2014) for the following elections for all 556 Croatian cities and municipalities: local elections in 2009 and 2013, and parliamentary elections in 2011. The variables for political use were taken from Glaurdić and Vuković (2016, 2017). These include a number of politi-cal and mayoral characteristics: turnout in the 2013 local election (*Turnout*), vote share of parties and candidates received in the 2011 national election (*HDZ and SDP vote share*), political experience (*Years in power*), political polarization (*EPI*), size of governing coalition in 2009 (*Size of gov coalition*), mayor majority in 2009 (*Mayor majority*), size of local assembly (*Assembly size*), an indicator of whether the mayor was from the same party that held national power (*Mayor gov*), and an indicator variable of re-election (*Re-election*). A few of these require further clarification. The indicator of re-election assigns the value of 1 if a local mayor was re-elected or, in cases where the mayor retired or left for higher office, when his or her party remained in power. A value of 0 is assigned in all other

cases. Political polarization was measured as an index of vote share between two major political blocs[3]; the left and the right, within a given local unit. The larger the index the greater the difference between the two electoral blocs (the more ideologically uniform a local unit), whereas a small difference depicts ideologically divided municipalities with strong support for either the left or the right.

TABLE 1
Summary statistics

Variable	Observations	Mean	Std. dev.	Min.	Max.
W/S ratio	546	0.270	0.102	0.034	0.681
W/S_binary	546	0.469	0.499	0	1
Corruption Index	546	0.083	0.127	0	0.75
Tax rate (prirez)	546	0.042	0.043	0	0.18
Re-election	546	0.731	0.444	0	1
Years in power	546	8.668	3.532	2	12
Turnout	546	0.478	0.100	0.231	0.886
HDZ vote share	546	0.206	0.115	0.003	0.675
SDP vote share	546	0.244	0.098	0.019	0.523
Mayor gov	546	0.676	0.468	0	1
Size of gov coalition	546	2.170	1.136	1	7
Mayor majority	545	0.585	0.112	0.231	1
EPI	546	0.818	0.177	0	0.996
Unemployment	546	0.193	0.093	0.04	0.567
Income per capita	546	1,696.8	438.5	519.9	2,973.6
Grants per capita	546	507.3	998.5	0	11,575.9
Croats	546	0.889	0.172	0.018	1
Average age	546	42.642	3.330	33.1	63.3
Years of education	546	9.842	0.875	5.925	12.132
War disabled per 1,000	546	8.539	8.082	0	64.98
Settlement size	546	3.006	0.545	1.65	5.78
Assembly size	546	13.385	3.897	7	51

Note: The data was collected for all 556 municipalities but 10 of them had no procurements in the given years so they were excluded from the dataset. Including them as a zero value does not influence the findings.

Sources: See appendix.

The size of governing coalition includes all parties which were a part of the local governing coalition (varying from 1 where only 1 party was holding local office to 7 where it took the coalition of 7 parties to get a legislative majority). This is not to be confused with the winning coalition, which is a more complex variable.

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[3] The index is calculated using the following formula, adapted from Glaurdić and Vuković (2017) and originally proposed by Reynal-Querol's (2002) ethnic polarization index: $EPI = 1 - \sum_{i=1}^{N} \left(\frac{\frac{1}{2} - \pi_i}{\frac{1}{2}} \right)^2 \pi_i$, where π_i represents the proportion of votes given to three political blocs: the conservative bloc led by the HDZ containing all parties that classify as right-wing, the social-democrat bloc led by SDP containing all parties that classify as left-wing, and a few unaffiliated parties and independent candidates. This variable measures to which extent party competition within a city or municipality was polarized between the two main left and right blocs.

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The size of the local assembly is measured by the number of local councilmen in each local legislature (varying based on population size from 7 to 51 for the capital Zagreb). The mayor majority measures the proportion of deputies in the local council that support the mayor. Each of these variables is supposed to capture how difficult it was for the mayor to govern. Even though after 2009 the change in the law gave local mayors more authority as they were being elected directly, having a majority in the legislative assembly was still important for a mayor to pass the local budget.

Economic and demographic variables were taken from the Croatian Bureau of Statistics (CBS, 2011; 2013), the Croatian Unemployment Bureau (HZZ, 2014), and the Ministry of Finance (2014). These include data on local income tax surcharge rates (*Tax rate – Prirez*), total unemployment (*Unemployment*), total intergovernmental grants received by the local government, which are used to alleviate fiscal inequalities in municipalities (*Grants p/c*), incomes per capita (*Income p/c*), average number of Croats (*Croats*), average age (*Average age*), total years of education (*Years of education*), relative population size of a city or municipality (*Settlement size*) and total number of war-disabled as a proxy for casualties of the 1991-1995 Croatian Homeland War (*War disabled*). All the data are taken on the unit of cities and municipalities. Table 1 shows the summary statistics of each variable used.

5 EMPIRICAL ANALYSIS
5.1 EMPIRICAL STRATEGY
As stated in the theoretical section and derived from the model, there are three predicted effects of small winning coalition environments: (i) greater chances of political survival, (ii) higher levels of corruption, and (iii) higher tax rates. This section directly tests the given implications using several methodological approaches. It is important to note that the availability of data has limited the usage of time series-based approaches given that corruption data were only available during a single term of office, from 2009 to 2013. This means that the dataset at hand is a cross-section where levels of corruption are observed as a single time unit over the duration of the local mandate. This has forced the usage of all other independent variables as a single time point, in most cases taking 4-year averages, as well as looking at the value of political indicators in the 2009 elections.

In order to prove the theoretical predictions, the first step of the empirical analysis is to examine the impact of small winning coalition environments on political tenure and chances of political survival. The following equation takes political longevity in office (*Years in power, T_i*) as the dependent variable and the *W/S ratio* as the main independent variable:

$$T_i = \rho WS_i + X_i \gamma + \mu_i \qquad (11)$$

Where X_i is a vector of all political, economic, and municipality-specific controls for city or municipality i. Political controls include the level of political polarization in a local unit (*EPI*), the relative size of the mayor's majority in terms of deputy support (*Mayor majority*), the size of the mayor's governing coalition (*Size of gov coalition*), an indicator variable of whether the mayor was from the party that held national power (*Mayor gov*), and the corresponding vote shares of HDZ and SDP in the national election in 2011 to control for change in relative party positions on a national level (*HDZ and SDP vote share*). Economic controls include average *Unemployment* during the term, average *Income per capita*, and average *Grants per capita* received from the central government. Finally the unit-specific controls include for each city and municipality the total percentage of *Croats*, *Average age*, total *Years of education*, number of *War disabled* per local unit, local *Assembly size* and the average size of the city or municipality (*Settlement size*).

The second step is to test the relationship between winning coalitions and corruption. A similar equation is estimated, where the *W/S ratio* is again the main independent variable, while the dependent variable is the total index of corruption per city or municipality i (*Corruption index*, k_i), featuring the same control variables as before:

$$k_i = \rho WS_i + X_i \gamma + \mu_i \tag{12}$$

Additionally, following the conclusions from Vuković (2017) which shows that corruption entails a concave effect on chances of re-election, meaning that voters reward corrupt behavior up until a certain level after which they do in fact punish corruption, the same relationship is tested here, but with respect to varying levels of the W/S ratio. It takes corruption and the squared value of corruption as the main independent variables regressing them against the indicator variable of re-election:

$$P\left(R_i = 1|k_i\right) = \beta_1 k_i + \beta_2 k_i^2 + X_i \gamma + \mu_i \tag{13}$$

The same controls are used as before, however the sample is split into small and large winning coalition environments based on *W/S_binary* and the equation is estimated for each environment. This way it was possible to draw conclusions from two sets of environments: how corruption affects re-election in local units with small winning coalitions, and how corruption affects re-election in local units with large winning coalitions. The prediction from the theory is that there should be a positive and concave effect of corruption on re-election only in small winning coalition environments. Both a probit and a linear probability model (LPM) are used in estimating equation (13).

Finally, the third step is to test how the W/S ratio affects local tax surcharge rates, where the average tax surcharge rate in the year of the election 2013, t_i is taken as the main dependent variable (*Tax rate*), featuring the all the same right-hand side variables as in the first two equations:

$$t_i = \rho W S_i + \boldsymbol{X_i}\boldsymbol{\gamma} + \mu_i \qquad (14)$$

All of the aforementioned equations are estimated using the simple ordinary least squares (OLS) estimation (except in equation 13 where a probit is used as well). However each of them could still be biased due to endogeneity issues. For example, there could be an unobserved factor like a specific local mentality that is simultaneously affecting both a small W/S environment and a higher tendency for corruption or even taxation. In order words corruption and taxes could be higher not because of a small W/S environment, but because of something intrinsic to the community that we cannot observe. If this is the case then each estimated $\hat{\rho}$ coefficient is likely to be biased.

In order to address these concerns the empirical strategy will employ several matching techniques seeking to solve potential endogeneity problems and get an unbiased estimate of the effect a small W/S environment has on corruption, taxation, and tenure in office. Matching is a conditioning-on-observables approach for the estimation of causal effects, meaning that its application depends on selection of comparable units by the researcher. This selection depends on finding a balance between a set of observable covariates in order to make the comparison between very similar treated and controlled units. Its primary goal is to establish counterfactual inference when the data are not generated by a random process or by an experiment.

The key to a successful matching strategy is to achieve a balance based on the distribution of covariates for the treated and control units. In other words, we must make the treatment group as similar as possible to the control group in order to get a good estimate of a potential effect of the treatment on the outcome (in our case the impact of the W/S environment on corruption, taxation, and tenure). Therefore the first step before even beginning the estimation of the treatment effect is to generate comparable datasets of treated and control units, based on the set of observable covariates from the sample at hand. This paper will use several matching techniques (propensity score matching, nearest neighbor matching, radius matching, and kernel matching) and compare the results between them as a robustness check. After this it will apply a much more robust method of entropy balancing (Hainmueller, 2012; Hainmueller and Xu, 2013) to generate an even better balanced sample, and compare the results. After performing the balancing of the dataset, making sure that we have comparable treatment and control units, equations (11), (12) and (14) will be estimated again in order to get the unbiased average treatment effect.

In the subsequent sections we first present the initial OLS estimates for each effect. Section 5.5 then applies the matching strategies and estimates the same equations once again using a balanced dataset to get a more precise inference on the causal effect of the W/S ratio on corruption, taxes, and tenure.

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5.2 THE W/S RATIO AND POLITICAL SURVIVAL

Table 2 shows the results of equation (11) on how the W/S ratio impacts political survival. Overall the results imply that a higher W/S ratio decreases total tenure in office: a switch from a small winning coalition to a large winning coalition reduces time in office between one and one and a half years. This effect is very large given that the estimates are made using the period of a single political mandate. Across three terms in office this implies that a politician operating within a small winning coalition is likely to have one full term in office more than a politician operating within a large winning coalition. This result remains robust to the inclusion of a number of political, economic, and municipal-specific characteristics.

TABLE 2

Impact of winning coalition size on tenure in power

Dependent variable: Years in power (T_i)	(1)	(2)	(3)	(4)
W/S ratio	-4.135 (1.55)***	-3.779 (1.56)**	-4.898 (1.61)***	-3.913 (1.71)**
Corruption index		2.581 (1.2)**	2.446 (1.24)**	2.36 (1.18)**
Size of gov coalition	-0.613 (0.15)***	-0.641 (0.145)***	-0.661 (0.143)***	-0.681 (0.143)***
Mayor majority	8.889 (1.13)***	8.851 (1.14)***	9.078 (1.11)***	9.332 (1.107)***
EPI	-1.281 (0.97)	-1.454 (0.98)	-1.725 (0.97)	-1.172 (1.07)
Mayor gov	0.259 (0.41)	0.25 (0.40)	0.343 (0.40)	0.275 (0.41)
HDZ vote share	2.294 (1.57)	2.196 (1.59)	2.843 (1.54)*	2.818 (1.93)
SDP vote share	6.424 (2.15)***	5.915 (2.18)***	3.339 (2.42)	3.738 (2.61)
Unemployment			-5.77 (2.28)**	-7.29 (2.55)***
Income p/c			-0.0007 (0.0004)	-0.0007 (0.0006)
Grants p/c			0.0002 (0.0001)	0.0003 (0.0001)**
Controls	No	No	No	Yes
N	546	546	546	546
R²	0.1869	0.1951	0.2102	0.2194
F-test (p-value)	27.53 (0)	24.29 (0)	21.42 (0)	15.70 (0)

*Note: OLS regressions performed throughout. Control variables include the municipality and city specific controls: Croats, Average age, Years of education, War disabled, Settlement size, and Assembly size. Standard errors are robust to heteroskedasticity. *** denotes significance at 1%, ** at 5%, and * at 10%.*

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Column (1) shows the results while including only the initial set of political variables as covariates. Moving from a local unit with a W/S ratio of 0.1 (small winning coalition environment) to a local unit with a W/S ratio of 0.4 (large winning coalition environment) decreases tenure in office by 1 year and 3 months. Column (2) adds corruption to the equation, under the hypothesis that longer tenure is likely to be conditioned by higher corruption. This is indeed the case as greater levels of corruption – moving from a zero corruption local unit to a unit where corruption is moderate, at an index value of 0.25 – increase tenure by an average of 8 months. However a switch from a zero corruption to a high corruption environment (levels of corruption at 0.5) increases tenure by 1 year and 4 months on average. A corrupt politician therefore stays in power longer. This also aligns well with the overall hypothesis that political survival is dependent to some extent on corruption. Interestingly the inclusion of the corruption variable decreases the magnitude of the W/S ratio effect, but it does not lose its statistical significance. Columns (3) and (4) add other economic and municipal-specific controls. In both cases the magnitude of the corruption effect is still roughly the same, while the effect of the W/S ratio varies from 1.1 to 1.5 fewer years in office for a switch from a small winning coalition local unit to a large winning coalition local unit.

The control variables all exhibit an expected effect on tenure in office. In all cases, each additional party that is a member of the governing coalition lowers tenure by about 8 months. This makes sense as more parties in the governing coalition make that coalition unstable which affects a mayor's probability of staying in power. Furthermore, a higher proportion of deputies in the local council that support the mayor (an increase of one standard deviation of 0.145) increases tenure between 1.2 and 1.3 years. This is also an expected effect as more deputies from the same party as the mayor, the more likely that a mayor has a stable governing coalition in the local assembly. Essentially these two variables measure the same thing with an opposite sign. Interestingly a higher vote share for SDP in a local unit (a one standard deviation increase of 0.097) increases tenure by roughly 6 months, however this effect loses statistical significance once economic covariates are included. From the economic covariates, unemployment exhibits an expected negative effect on tenure, where higher unemployment (a one standard deviation increase of 0.092) in a local unit decreases time in office between 6 and 8 months on average. All of these estimated effects may seem small in absolute terms, however keep in mind that each is estimated during a single political term in office, so the effects are actually quite large.

5.3 THE W/S RATIO AND CORRUPTION

The results from section 5.2 have confirmed a positive impact of corruption on tenure in office, and have established that both corruption and a small winning coalition environment entail a positive impact on political tenure. This section empirically tests the second prediction of the theoretical section (and the unique prediction of the model) according to which small winning coalition environments should exhibit higher levels of corruption. Table 3 shows the results. As

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expected and hypothesized in each case a higher W/S ratio decreases total levels of corruption. Corruption is therefore more likely to occur in small winning coalition environments. The magnitude of the effect is roughly the same across all regression equations. A one standard deviation increase (decrease) of the W/S ratio decreases (increases) corruption by 1.1 to 1.4 percentage points.

TABLE 3

Impact of winning coalition size on corruption

Dependent variable: Corruption index (k_i)	(1)	(2)	(3)	(4)
W/S ratio	-0.138 (0.06)**	-0.122 (0.06)**	-0.121 (0.06)*	-0.113 (0.06)*
Tax rate (prirez)				0.267 (0.129)**
Size of gov coalition	0.01 (0.005)*	0.006 (0.005)	0.006 (0.005)	0.005 (0.143)
Mayor majority	0.014 (0.039)	0.02 (0.038)	0.031 (0.039)	0.033 (0.04)
EPI	0.067 (0.032)**	0.083 (0.031)**	0.078 (0.032)**	0.074 (0.032)**
Mayor gov	0.003 (0.01)	0.001 (0.01)	0.005 (0.01)	0.006 (0.01)
HDZ vote share	0.038 (0.068)	0.008 (0.069)	0.018 (0.088)	0.014 (0.089)
SDP vote share	0.197 (0.10)*	0.05 (0.10)	0.058 (0.12)	0.063 (0.12)
Unemployment		0.117 (0.09)	0.092 (0.101)	0.079 (0.10)
Income p/c		8.38×10^{-5} (1.48×10^{-5})***	4.59×10^{-5} (2.43×10^{-5})*	4.06×10^{-5} (2.4×10^{-5})*
Grants p/c		-8.14×10^{-6} (3.17×10^{-6})**	-9.42×10^{-6} (4.2×10^{-6})**	-9.44×10^{-6} (4.1×10^{-6})**
Controls	No	No	Yes	Yes
N	546	546	546	546
R^2	0.0498	0.0989	0.1088	0.1158
F-test (p-value)	4.56 (0)	8.54 (0)	6.90 (0)	6.99 (0)

*Note: OLS regressions performed throughout. Control variables include the municipality and city specific controls: Croats, Average age, Years of education, War disabled, Settlement size, and Assembly size. Standard errors are robust to heteroskedasticity. *** denotes significance at 1%, ** at 5%, and * at 10%.*

Graph 1 below plots the impact of the W/S ratio on corruption. Predicted levels of corruption (predicted from the estimates of column 3) are plotted on the y-axis, and the W/S ratio is plotted on the x-axis. As anticipated by the underlying theory the effect is clearly negative and confirms the results from table 3.

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GRAPH 1

Smaller winning coalition size environments on average exhibit greater levels of corruption
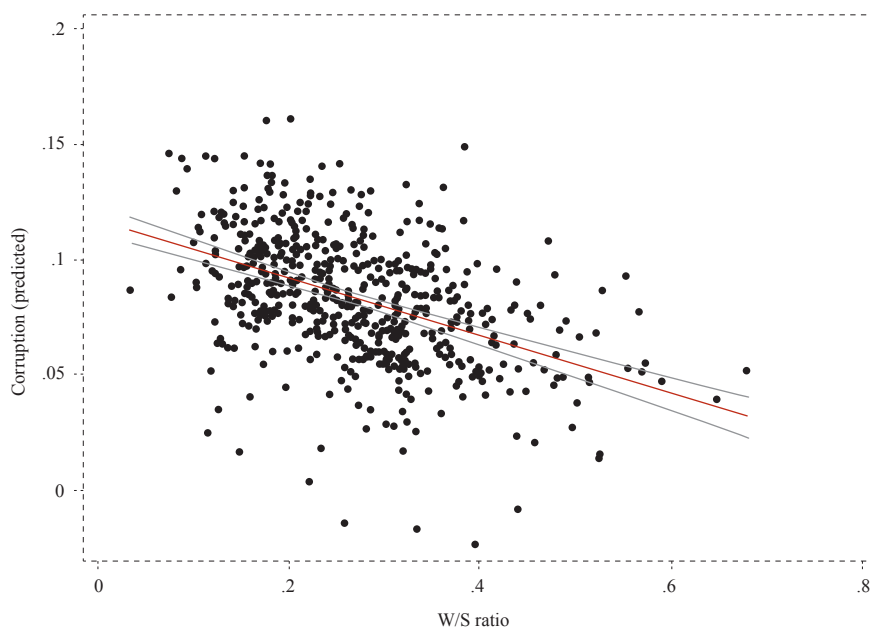
Note: Corruption is predicted from equation (12).
Source: Author's own calculations.

The theoretical prediction from the model in equation (10) was also that corruption should be higher in areas with greater political polarization ($\psi \underline{\lim} 0$), i.e. in areas of higher ideological uniformity, and in areas with higher income and higher local taxes. This too has been shown in table 3. Greater political polarization, i.e. greater ideological uniformity of a city or municipality, results in higher corruption. This makes sense given that ideologically uniform local environments are likely to accept higher corruption from their local mayor if he or she comes from the "correct" political spectrum. Also, greater income per capita in a local unit implies higher corruption, whereas more grants received by a local unit (i.e. the poorer the city or municipality) imply lower corruption.

The final column includes the local tax surcharge rate as an additional covariate. As predicted by the model in equation (10) cities or municipalities whose mayors set up higher tax rates tend to have greater levels of corruption. The effect is robust even after controlling for total tax revenues (not shown, available upon request). The reason for this could be simple: higher taxes imply higher revenues and hence more opportunities to engage in corruption. This is not necessarily a causal effect, but it testifies of an interesting relationship that will be examined in further sections.

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5.3.1 Corruption and re-election in large and small winning coalitions

The previous two sections confirmed a relationship between small winning coalitions and political longevity as well as potential corruption. This could mean that the relationship between corruption and re-election is bounded and encouraged only within a small winning coalition environment. In order to test whether this is true, this section looks at the relationship between corruption and re-election in two different sets of winning coalition environments. It estimates equation (13) and follows the assumption of a quadratic relationship between corruption and re-election according to Vuković (2017).

The expectation is that there will be a positive (yet non-linear) relationship between corruption and re-election only in small winning coalition local units (those with lower levels of transparency and political accountability), and that there will be no such relationship in local units with a larger winning coalition (which have higher levels of transparency and political accountability).

TABLE 4

Impact of corruption on re-election in small and large winning coalitions

Dependent variable: Re-election (R_i)	Small winning coalition (W/S < 0.26)		Large winning coalition (W/S > 0.26)	
	(1) Probit	**(2) LPM**	**(3) Probit**	**(4) LPM**
Corruption k_i	0.948 (0.508)*	0.954 (0.524)*	0.229 (0.353)	0.318 (0.447)
Corruption squared k_i^2	-1.868 (0.948)**	-1.891 (0.927)**	-1.182 (0.795)	-1.576 (1.401)
Controls	Yes	Yes	Yes	Yes
N	295	295	250	250
Pseudo R^2	0.1276	0.1522	0.2717	0.1821
Wald/F-test (p-value)	43.67 (0)	4.36 (0)	40.93 (0)	1.26 (0.19)

Note: Columns 1 and 3 report the results from the probit estimations (and they have the Wald test and a Pseudo R^2 reported), while columns 2 and 4 report the results from the OLS estimations (which have the F-test and the regular R^2 reported). The full list of control variables is the same as in column 4 of tables 2 and 3. Standard errors are robust to heteroskedasticity. *** denotes significance at 1%, ** at 5%, and * at 10%.

Table 4 presents the results and confirms the intuition where the effect of corruption on reelection is both higher and statistically significant in municipalities with a smaller W/S ratio (smaller winning coalition). This effect should be calculated as $k_i^* = \left| \dfrac{\hat{\beta}_1}{2\hat{\beta}_2} \right|$, where k_i^* is the marginal cutoff value of corruption while $\hat{\beta}_1$ and $\hat{\beta}_2$ are estimated regression coefficients from equation (13). The cut-off value for corruption in small winning coalition environments is 25.2%. This means that politicians maximize their chances of re-election when the share of suspicious procurements[4] in total procurements is about one fourth. This is a relatively high level of

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[4] This is the official definition of the Total index corruption proxy, according to Vuković (2017).

corruption persistence given that voters usually lack the necessary information to punish such behavior. In municipalities with a high W/S ratio (larger winning coalition) the effect is much smaller – the cut-off value of corruption is 10.1% – however it is statistically indistinguishable from zero. This suggests that corruption does not aid in re-election chances in large winning coalition environments, but only in small ones. In other words, in places where it is easier for politicians to form long-term connections with quasi-entrepreneurs, it is more likely that by engaging in such activities they will increase their chances of re-election. No such effect exists in areas where it is more difficult to successfully satisfy a group of key supporters.

5.4 TAXES, CORRUPTION AND THE W/S RATIO

The final theoretical prediction is that small winning coalition environments should have higher tax rates. This section will present the results from equation (14) and will in addition test the reverse impact of corruption on tax rates. The main dependent variable represents the tax surcharge rate applied on top of the main income tax rate; however it will be referred to simply as *local tax rate*, given that this rate is under full jurisdiction of the local mayor and his or her administration. There is an upper boundary for how high each city or municipality can set this rate, which is based on population size, but any decision on tax rates within this boundary is carried out by the local mayor. It therefore makes sense to use this variable as yet another measure of a local mayor's power structure.

GRAPH 2

A small winning coalition environment has on average higher local tax rates
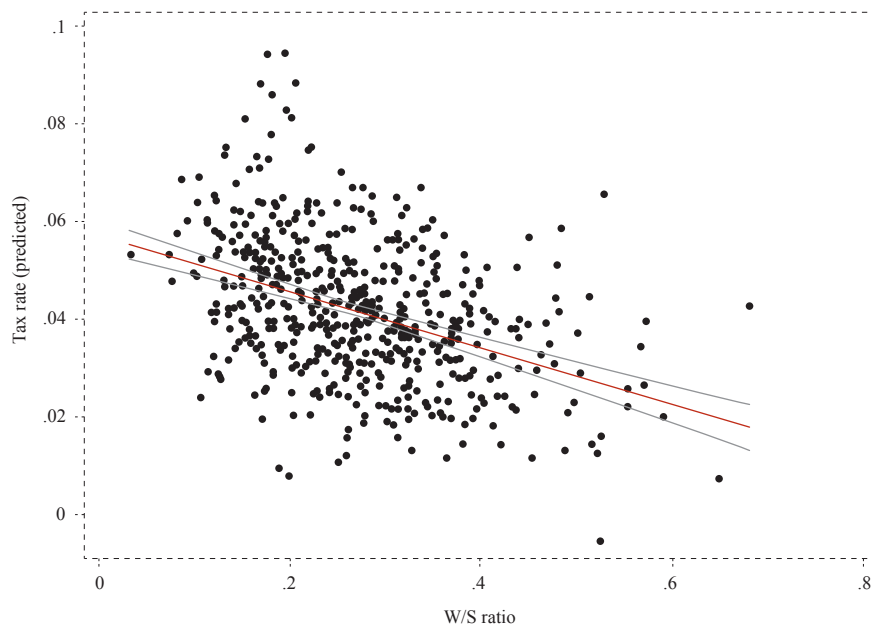
Note: Tax rates are predicted from equation (14).
Source: Author's own calculations.

Table 5 shows the results. Both the W/S ratio and the corruption index entail an effect on local taxes of roughly similar magnitude, but in different directions. For example, moving from a small winning coalition environment (W/S = 0.1) to a large winning coalition environment (W/S = 0.4) the local tax rate is estimated to be between 1.2 and 1.5 percentage points lower. This is a very large effect given that the mean local tax rate is 4.15%. In other words a small winning coalition environment is likely to be characterized by much higher tax rates (around 30% higher taxes), even when controlling for population size and all other important factors. Graph 2 shows this relationship graphically estimating the level of tax rates using the regression equation from column (4). Similarly, a move from a zero corruption to a high corruption environment (from 0 to 0.5 value of the corruption index) will increase local tax rates by between 1.5 and 2.6 percentage points. This effect, just like the previous one, is very large and statistically significant.

TABLE 5

Impact of winning coalition size on local tax rates

Dependent variable: Tax rate (t_i)	(1)	(2)	(3)	(4)
W/S ratio	-0.051 (0.016)***	-0.039 (0.016)**	-0.041 (0.018)**	-0.023 (0.019)
Corruption index	0.053 (0.016)***	0.034 (0.015)**	0.030 (0.015)**	0.029 (0.014)**
Size of gov coalition			0.005 (0.002)***	0.004 (0.002)**
Mayor majority			-0.012 (0.015)	-0.007 (0.016)
EPI			0.003 (0.012)	0.011 (0.012)
Mayor gov			-0.004 (0.005)	-0.005 (0.005)
HDZ vote share			0.022 (0.023)	0.013 (0.025)
SDP vote share			-0.017 (0.03)	-0.021 (0.03)
Unemployment		0.084 (0.024)***	0.063 (0.027)**	0.046 (0.031)
Income p/c		3.05×10^{-5} (5.87×10^{-6})***	3.11×10^{-5} (6.13×10^{-6})***	1.88×10^{-5} (8.2×10^{-6})**
Grants p/c		-2.37×10^{-6} (1.7×10^{-6})	-2.59×10^{-6} (1.7×10^{-6})	3.72×10^{-7} (1.8×10^{-6})
Controls	No	No	No	Yes
N	546	546	546	546
R^2	0.0442	0.0958	0.1208	0.1513
F-test (p-value)	11.62 (0)	10.37 (0)	6.19 (0)	5.75 (0)

Note: OLS regressions performed throughout. Control variables include the municipality and city specific controls: Croats, Average age, Years of education, War disabled, Settlement size, and Assembly size. Standard errors are robust to heteroskedasticity. *** denotes significance at 1%, ** at 5%, and * at 10%.

Together these two effects confirm the intuition set out in the theoretical section, according to which tax rates are likely to be higher in small winning coalition environments where corruption is larger. In other words, cities and municipalities which are unlucky to be existing within a small winning coalition environment, where politicians can easily exploit the logic of political power to keep themselves in office for long periods of time, will have a mayor who is more likely to be corrupt and keep tax rates high.

5.5 MATCHING

In order to get a better estimation of the potential causal effect of winning coalition environments on corruption, taxes, and tenure, a matching procedure is done in order to balance the treated and control units based on a set of used covariates. In our case the non-random treatment is the previously defined binary indicator of a small and large winning coalition environment, *W/S_binary*, where a small winning coalition is given the value of 0, and a large winning coalition is given a value of 1. The covariates, with their summary statistics given in table 1, are the same control variables used throughout tables 2, 3 and 4. The balancing of the dataset is done with respect to these covariates in order to make the small and large winning coalition environments as similar as possible to draw a causal inference.

In order to balance the dataset several matching techniques will be applied. This is done primarily as a robustness check so that our results do not depend on method specification. The first matching method is the default propensity score (PS) matching. In order to evaluate how good it balanced the dataset we need to perform visualizations of balance checks. Graph 3 reports how well each of the four used algorithms matched the covariates by showing the standardized biases (PS matching is shown in the first upper left panel). It reports the graphical results of a t-test on whether the mean value of each covariate is the same in the treatment and control groups. The closer the matched values (denoted by X in each of the graphs) are to the horizontal line, the lower the standardized bias across covariates, and the better matched the sample gets. By observing the pre-matched means of the covariates (denoted by black dots in all charts of graph 3) it is easy to see that there were clear differences in the distributions of covariates before performing the matching. For example, small winning coalition environments were already more likely to have smaller governing coalitions, greater mayor majorities, and lower levels of political polarization which could have affected both total corruption and tenure in office. Running a regression without taking this into account could have overestimated the effect of the W/S ratio on both corruption and tenure. This is why it is important to perform a balancing of treatment and control groups with respect to the given covariates.

GRAPH 3

A comparison of sample balancing using four different matching algorithms
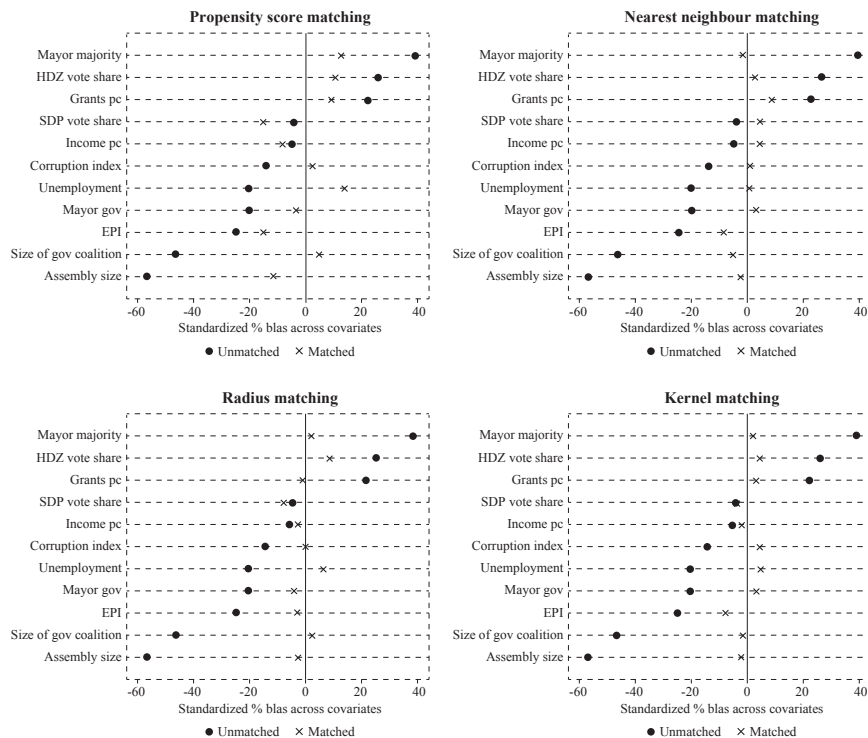
Note: *Each x-axis shows the standardized bias across covariates (the lower the bias, i.e. the closer the matched values to 0, the better the balancing), while each y-axis shows the list of covariates used in the matching procedure.*

Source: *Author's own calculations.*

After matching, it is obvious that a better balance is achieved across all covariates (shown by X in each graph). The percentage differences in means in treated and control units across each covariate for all the algorithms are very low and statistically insignificant (the lowest p-value across all balance tests is 0.23), meaning that we can safely conclude that each matching algorithm produced a well-balanced dataset.

There are obvious differences however. Different matching techniques are usually applied until one of them delivers the optimally balanced dataset. Even though the default propensity score procedure did produce a reasonably well-balanced dataset (shown in the upper left panel of graph 3), other algorithms, in particular kernel, were matching even better. It should be mentioned that several different variations of each algorithm were tested until reaching the one with the lowest bias. Several values of the nearest neighbour algorithm were used (2, 3, 5, 10, and 12 nearest controls per one treatment). The nearest neighbour matching matches the best controls for each unit in the treatment group based on the relative distance of the

control to the treatment unit (hence neighbour). The upper right panel in graph 3 shows the matched dataset using 10 controls per one treatment, and it is the one that had the lowest standardized bias. For radius matching, which establishes a maximum propensity score radius, and matches units based on an established calliper, the following calliper values were used: 0.2, 0.1, 0.08, 0.05, 0.02., and 0.01. The final one, with a calliper of 0.01 was the best matched and is shown in the lower left panel of graph 3. Finally, the lower right panel shows kernel matching, where a normal and a tricube kernel types were used. Kernel matching uses a kernel-weighted average of the outcome for all control units, where the weight indicates a proportion of distance between control and treatment units. The tricube kernel type was the one that offered the lowest bias. The kernel algorithm balanced the sample better than the other three chosen algorithms (it had the lowest overall standardized bias), with nearest neighbour matching coming in second.

5.5.1 Estimating the effect after matching

After having successfully achieved a balance between treatments and controls across all the covariates, the next step is to engage in estimation. Table 6 shows the estimates for a linear OLS regression of the treatment effect of a small winning coalition environment on tenure, corruption, and local taxes, across each of the balanced datasets for different matching algorithms (each column represents results for a different matching algorithm). The coefficients do not differ substantially from each other. They all suggest a highly significant and negative effect of the W/S ratio on tenure, corruption, and taxes, as we have shown in sections 5.2 to 5.4. In terms of effect size, a city or municipality with a small winning coalition increases tenure by an average of 6 to 12 months, increases corruption by an average 4.6 to 5.2 percentage points, and increases tax rates by an average of 0.8 to 1 percentage points when compared to a similar city or municipality with a large winning coalition. These average treatment effects are not too different from the estimated coefficients reported in tables 2, 3 and 5, except in the case of the impact on corruption, where the effect is now several times larger. The reason for this is a different unit of comparison – in section 5.3 the estimated effect was looking at a one standard deviation increase in the W/S ratio (a change of the ratio by 0.1), whereas now we are interpreting an indicator variable change (the equivalent of which would be an increase of the W/S ratio by 0.3, which would according to table 3 yield an effect on corruption of 3.6 percentage points, which is still lower but it is relatively close).

TABLE 6

Comparing the results for different matching algorithms

	(1) PS matching	(2) NN matching	(3) Radius matching	(4) Kernel matching
Impact of W/S on tenure (T_i)	-0.987 (0.440)**	-0.543 (0.32)*	-0.644 (0.338)*	-0.669 (0.302)**
R^2	0.3059	0.2221	0.2144	0.2430
F-test (p-value)	12.9 (0)	13.09 (0)	10.29 (0)	13.89 (0)
Impact of W/S on corruption (k_i)	-0.0558 (0.038)	-0.0526 (0.026)**	-0.0367 (0.028)	-0.046 (0.026)*
R^2	0.0846	0.088	0.0713	0.0912
F-test (p-value)	1.85 (0.04)	2.56 (0)	2.59 (0)	2.67 (0)
Impact of W/S on tax rates (t_i)	-0.0102 (0.005)*	-0.0086 (0.004)**	-0.0097 (0.003)**	-0.0099 (0.003)***
R^2	0.1855	0.1290	0.0962	0.1168
F-test (p-value)	4.85 (0)	3.81 (0)	4.25 (0)	4.29 (0)
Controls	Yes	Yes	Yes	Yes
N	257	484	546	537

Note: The four matching algorithms are: (1) propensity score matching, (2) nearest neighbor matching, (3) radius matching, and (4) kernel matching. After each matching procedure an OLS regression was performed to estimate the impact of the W/S ratio on tenure, corruption and tax rates. The full list of control variables is the same as in column 4 of tables 2 and 3. Standard errors are robust to heteroskedasticity. *** denotes significance at 1%, ** at 5%, and * at 10%.

### 5.5.2 Entropy balancing and comparison of average treatment effects

The different matching algorithms all performed a decent balancing of the dataset and have all suggested a positive effect of small winning coalition environments on tenure, taxes, and corruption. In addition to the used algorithms there is another matching method that has proven to be even better and superior to all of the previously used ones. Entropy balancing (Hainmueller, 2012) is a multivariate reweighting method that depends on specifying the moment conditions of covariates in the treatment and control groups in order to match them. Its algorithm prevents the loss of information by predefining the balance conditions and as such "always [...] improves on the covariate balance achieved by conventional preprocessing methods" (Hainmueller and Xu, 2013). This method also enables balancing across joint distributions of covariates, which is achieved by including interaction terms. Its main advantage over all other methods is its accuracy in balancing the sample, and consequentially its accuracy in establishing the treatment effect.

When applying entropy balancing, the results are clear: it provides a much better match across the covariates than all four matching methods used in the previous section. Graph 4 compares on a single graph the differences in means for all the covariates (again looking at the standardized bias, as in graph 3). It compares the entropy balanced sample with the unmatched sample and the default propensity score matching procedure (replicated for the same interaction variables used in entropy balancing). A simple visual inspection suggests that the dataset was

extremely well-balanced using the entropy balancing approach. All of the mean values of for the treatment and the control covariates are almost exactly the same. This clearly suggests a superior matching method and it implies that we are now able to perform a quasi-experimental estimation of the average treatment effect on exactly the same treatment and control units that differ in one single factor: having a small or large winning coalition environment.

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GRAPH 4

A balanced sample using entropy balancing (triangles), compared with an unbalanced sample (diamonds), and a sample matched using propensity score matching (circles)
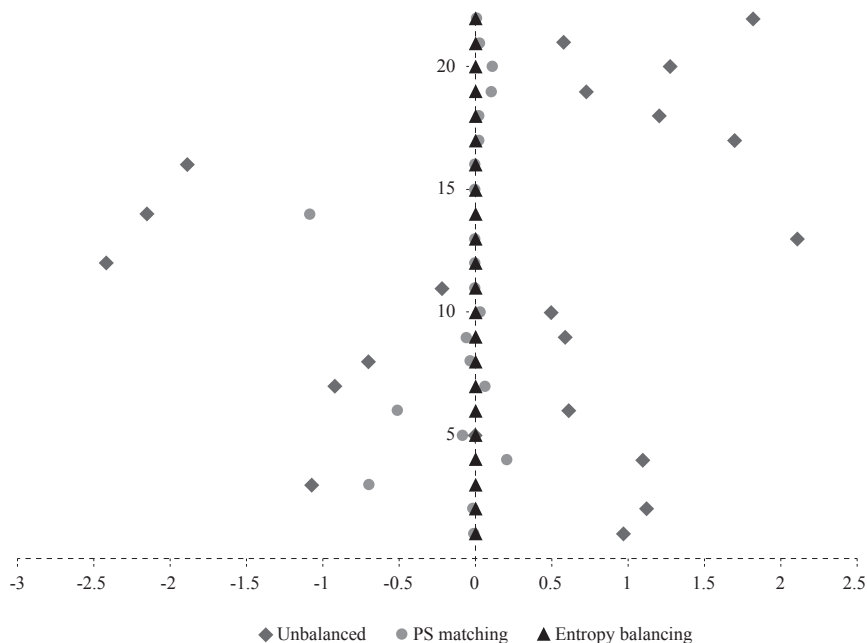
Note: Entropy balancing performs an almost perfectly balanced sample against a set of 22 covariates in total on the y-axis (these include all the covariates used in the initial estimations plus a few interaction terms created by the ebalancing algorithm). The x-axis measures the standardized bias, the same as in graph 3.

Source: Author's own calculations.

We now use the entropy balanced dataset to again generate the treatment effect of a small winning coalition on tenure, corruption, and taxes. The results are reported in table 7 comparing the estimates after entropy balancing, propensity score matching, and an OLS for the unmatched sample (these results differ from the initial ones estimated in tables 2, 3 and 5 given that the main independent variable is a binary indicator rather than a continuous variable). The coefficients for entropy balancing are slightly higher than the propensity score matching estimates, and carry greater statistical power. When compared to an unmatched dataset, the entropy balancing estimated coefficients are again larger and in the same expected direction. According to the entropy balancing estimates, a city or municipality

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with a small winning coalition increases tenure by an average of 10 months, increases corruption by an average of 6 percentage points, and increases local tax rates by 1.2 percentage points when compared to a similar city or municipality with a large winning coalition. Each of these effects are not too different from the unmatched sample estimates, however they do present us with more robust evidence of the actual treatment effect in each of the hypothesized scenarios.

TABLE 7
Entropy balancing

	(1) Entropy balancing	(2) Propensity score matching	(3) Unmatched sample
Impact of W/S on tenure (T_i)	-0.857 (0.275)***	-0.987 (0.440)**	-0.62 (0.30)**
R^2	0.2519	0.3059	0.2018
F-test (p-value)	23.88 (0)	12.9 (0)	20.49 (0)
Impact of W/S on corruption (k_i)	-0.0602 (0.030)**	-0.0558 (0.038)	-0.054 (0.026)**
R^2	0.0856	0.0846	0.0752
F-test (p-value)	3.09 (0)	1.85 (0.04)	5.21 (0)
Impact of W/S on tax rates (t_i)	-0.012 (0.004)***	-0.0102 (0.005)*	-0.0097 (0.037)**
R^2	0.184	0.1855	0.1245
F-test (p-value)	3.51 (0)	4.85 (0)	6.45 (0)
Controls	Yes	Yes	Yes
N	546	257	546

Note: The first column reports the estimates of the three given relationships after entropy balancing, the second replicates the same results for propensity score matching from column 1 in table 6, and the final are OLS estimates from an unmatched sample reported in column 4 of tables 2, 3 and 5. The full list of control variables is the same as in column 4 of tables 2 and 3. Standard errors are robust to heteroskedasticity. *** denotes significance at 1%, ** at 5%, and * at 10%.

Based on the entire set of results from all matching procedures applied, the conclusion is that there is a positive causal effect of a small winning coalition environment on tenure, corruption and taxes (adhering to the assumptions of the selection on observables approach). This means that in Croatia the systemic corruption is indeed supported by an environment in which institutional and legal frameworks are personalized and hence discourage development. A city or municipality that has these characteristics is more likely to be subject to usurpation of budgetary resources for private gains and higher taxes by a politician that is likely to stay in power for a long period of time.

6 CONCLUSION

This paper has made an important contribution to the existing literature on corruption and institutional deficits in Croatia. It has defined a specific environment in which corruption, lack of transparency, lack of accountability, and high taxes coexist. It is the same environment in which, most likely because of low account-

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ability and low transparency, local mayors can stay in power for long periods of time. Such mayors, because their tenure in office depends on satisfying only a selected group of loyal voters, are free to engage in corrupt activities as well as set high tax rates. The paper's two main contributions have been a successful theoretical extension of the selectorate theory to include a direct link between small winning coalition environments and corruption, and a successful empirical verification of the selectorate theory in Croatia that confirms the adverse effect of small winning coalitions on corruption, taxes, and tenure. We now have a systematic explanation of why Croatian voters do not punish corrupt politicians.

This explanation is arguably only one part of a much wider issue that includes mentality, culture, historical heritage, and a range of other factors. In addition, there are several potential drawbacks of the research. The dataset is limited to only a cross-section of corruption during a single political mandate. It would have been more intuitive to observe a time series of corruption, particularly to look at how things have changed before and after the 2009 elections when the new law granting more power to mayors had been enacted. Finally, the very definition of corruption could easily underestimate total corruption in a municipality given that it does not include an arguably even bigger source of local corruption – changes in urban planning laws. Future research should attempt to address these drawbacks in order to make an even better indicator of corruption, and test it across several political terms.

Even with these limitations the main contributions of the paper redefine Croatia within a new theoretical framework for analysis. The findings confirm that Croatia is a worrying phenomenon; however the logic of political survival of Croatian politicians is no different than in any other country. The difference is only in institutions that either prevent or encourage corrupt behavior. Countries with clear and enforceable rules manage to prevent power-hungry individuals from usurping social and market outcomes, while countries that lack such rules or lack enforceability of existing rules are sentenced to vicious cycles of bad equilibria. A change of such a social order must come from an institutional push, encouraging both greater transparency and accountability to the voters, and credibly punishing politicians when they break the rules.

Disclosure statement
I state that I have no conflict of interest that might influence the results or interpretation of my manuscript.

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APPENDIX

TABLE A1

Variables used in analysis, their description and sources

Variable	Description	Source
W/S ratio	Winning coalition divided by the entire selectorate (votes for incumbent divided by all eligible voters)	State Electoral Commission of the RC: http://www.izbori.hr/ izbori/dip_ws.nsf/public/ index?open&id=BF0A&
Corruption index	Weighted average of all the given proxies of corruption defined through various cases of potentially fraudulent public procurements	Vukovic (2017), based on interviews with key stakeholders in the procurement process. Source of procurements: Official Gazette of the Republic of Croatia: https://eojn.nn.hr/ Oglasnik/
Tax rate (prirez)	Local tax surcharge rate on income tax	Ministry of Finance, Tax Administration: https://www.porezna-uprava.hr/ HR_porezni_sustav/Stranice/ Popisi/Stope.aspx
Re-election	=1 when incumbent mayor gets re-elected, =0 otherwise	State Electoral Commission of the RC
Years in power	Total years in power until 2013	State Electoral Commission of the RC
Turnout	Municipality-level turnout in the 2013 local election	State Electoral Commission of the RC
HDZ vote share	Proportion of municipal vote for the HDZ in the 2011 national election	State Electoral Commission of the RC
SDP vote share	Proportion of municipal vote for the SDP in the 2011 national election	State Electoral Commission of the RC
Mayor for government	=1 when mayor from the same party as the national government; =0 otherwise	Glaurdić and Vuković (2017); source of original data: State Electoral Commission of the RC
EPI	Electoral polarization index, based on municipality-level results in the directly preceding national election	Glaurdić and Vuković (2017); source of original data: State Electoral Commission of the RC
Size of governing coalition	Number of seats the governing coalition has in the local assembly	Glaurdić and Vuković (2017); source of original data: State Electoral Commission of the RC
Mayor majority	Proportion of deputies in the local council supporting the mayor	Glaurdić and Vuković (2017); source of original data: State Electoral Commission of the RC
Unemployment	Average monthly unemployment rate	Croatian Unemployment Bureau: http://www.hzz.hr/ default.aspx?id=10052
Income	Income per capita (deflated to 2010 levels), natural log transformed	Croatian Bureau of Statistics: https://www.dzs.hr/Hrv_Eng/ publication/2014/SI-1526.pdf

Variable	Description	Source
Grants	Per capita grants (deflated to 2010 levels), natural log transformed	Ministry of Finance: http://www.mfin.hr/hr/lokalni-proracun-arhiva
Croats	Proportion of population ethnically Croatian	Croatian Bureau of Statistics
Average age	Average age for entire population in municipality	Croatian Bureau of Statistics
Years of education	Average years of education for population older than 15 years of age	Croatian Bureau of Statistics
War disabled per 1,000	Number of disabled persons whose disability was caused by war	Glaurdić and Vuković (2016); source of original data: Croatian Bureau of Statistics
Settlement size	Weighted average of settlement size in a municipality as a measure of urban-rural cleavage	Glaurdić and Vuković (2016); source of original data: Croatian Bureau of Statistics
Assembly size	Total number of representatives in the local assembly	State Electoral Commission of the RC

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The effects of intercompany lending on the current account balances of selected economies in the Western Balkans

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Abstract

This paper quantifies the effect of intercompany lending on the current account balances in the economies that previously made up the Socialist Federative Republic of Yugoslavia. This kind of transaction is a controversial part of the foreign direct investments, often criticised for its indebting nature and its involvement in tax evasion strategies. However, the data shows that in the post-crisis period it was the driving force of foreign capital inflow and investments in the region. For this study, a novel model averaging approach was employed, as it allows cross-country and country-specific analysis, and provides a sound basis for future policymaking. Additionally, as a way to overcome the limited data availability problem, and provide an additional robustness check, panel regression fixed effects for 17 CESEE economies was done. The results of both models are significant and indicate the stabilising effect of the observed transaction as it provided a steady inflow of funds.

Keywords: intercompany lending, foreign direct investment, Jackknife model averaging

1 INTRODUCTION

Intercompany lending, is a constituent part of foreign direct investments (FDI) (IMF, 2009) that represents a controversial element, frequently underestimated because of its indebting nature and involvement in tax evasion strategies. However, in this paper, the goal is to observe it from another perspective and determine the beneficial effect it had on the current account balances (CABs) of Western Balkan economies when the global financial crisis struck.

The effects of FDI have been the topic of many papers (Lim, 2001; Li and Liu, 2005; Hermes and Lensink, 2003; Borensztein, De Gregorio and Lee, 1998). Some find that, especially for transition economies, the presence of foreign investors has created positive externalities, and helped them to integrate into the world market (DiMauro, 2000), and some indicate that it has contributed to the restructuring of formerly state-owned enterprises (Estrin et al., 2009). Positive findings further spurred a vast literature about the FDI determinants in transition economies (Bevan and Estrin, 2004; Resmini, 2000). However, all papers observe FDI as a total and disregard the different nature of its components: equity investment (EI), reinvested earnings and intercompany lending (ICL).

The existing literature on ICL is mostly focused on corporate finance and the possibilities for transfer pricing. However, in this paper, we will shift focus onto the FDI nature of cross-border ICL and show that since the crisis hit the observed region, ICL became an essential element of total FDI inflow. Furthermore, it will explain that the magnitude of the effect differs and depends on the country's characteristics.

The contribution of this paper is that until now, at least to the best of the author's knowledge, no paper has analysed ICL in the region from this perspective. More-

over, it quantifies the effect with the use of two models. The novel technique of model averaging is employed as the primary model, while the panel regression fixed effects model, serves as a robustness check. The former focuses on individual countries, provides separate estimates for each of them and enables cross-country and country-specific analysis. Therefore, it can serve as a sound basis for policymaking.

The paper is organised as follows: the next section will give an overview of ICL. The occurrences in the post-crisis period that inspired this analysis are given in section 3. The econometric analysis will be explained in detail in section 4. Finally, the last part will contain the conclusion, potential policy recommendations and indicate the limitations of the study. Additionally, the appendix provides data that can be used for further analysis.

2 INTERCOMPANY LENDING

The power of multinational companies (MNC) depends on their ability to make the most out of their global presence. Therefore, when deciding where to establish a new unit, they look for cost-cutting opportunities, such as tax reduction schemes (Devereux and Griffith, 1998; Barrios et al., 2009). ICL is a common part of these plans as it allows exploitation of the tax rate differentials to shift profit (Buettner and Wamser, 2007; Stewart, 1977) leaving both source and host countries with lower tax bases and tax revenues.

However, ICL should not be observed only as a part of a tax evasion strategy but also as a part of a risk-management strategy that MNC employ to optimise resources and protect previously invested funds. Since the outbreak of the financial crisis, risk aversion among investors has risen, leading to an increase in ICL (ECB, 2012). There are at least two reasons for this. First, affiliated companies located in emerging markets were faced with a tightening of credit conditions and were unable to obtain the necessary funds under conditions acceptable to them. Second, as EI represents pricier and riskier forms of capital, direct investors needed something that would allow them to finance their subsidiary while preserving flexibility. The solution was ICL since it provides a flow of funds for taffiliates, and at the same time, creates enough pressure to make them step up and work better in the crisis period.

The latter view fits better with the observed FDI nature of ICL. This dimension became noticeable only after the crisis when many posed the question regarding the sustainability of the benefits of FDI to emerging economies (Starnawska, 2015). Hebous and Weichenrieder (2010) noted that ICL amplified the stabilising role FDI had during the crisis. Moreover, data show that in Central, Eastern and South-eastern Europe (CESEE) in the aftermath of the crisis, the increase of ICL sustained total FDI, and stabilised capital flows (Gardo and Martin, 2010). That occurrence is precisely what this paper analyses, with the focus on the Western Balkan region.

3 THE CRISIS PERIOD

Before the crisis, CESEE countries, experienced positive growth rates. They were resilient until the first half of 2008, since the region was not as exposed to sub-prime markets as developed countries (Gardo and Martin, 2010). However, in September 2008, the crisis spread to the whole financial sector, and growth rates started to decline.

Fast growth and return on investment that existed before 2008 created keen interest from foreign banks in the whole CESEE region. It is noteworthy that at the end of 2008, almost 80% of the banking sector in this area was held by foreigners (Gardo and Martin, 2010). Although this type of vulnerability was not in focus before the crisis, it created significant liquidity problems after. The problem occurred when parent banks decided to withdraw funds from these markets so they could consolidate at home (Herrmann and Mihaljek, 2010). This spillover had several implications. First, parent banks were not able to extend the same amount of loans as earlier to their CESEE subsidiaries. Secondly, the interbank market was disrupted (ECB, 2012). Subsequently, this reduced the available funds of domestic banks, which reacted by tightening credit conditions. Finally, problems were transferred to the private sector, which was faced with liquidity issues.

Conversely, the share of the ICL flows in the total FDI inflow increased right after the start of the global financial crisis. Additionally, the same situation repeated itself when the Euro debt crisis began. Figure 1 highlights this in the region and allows comparison with similar movements in the advanced economies.

FIGURE 1
The share of ICL in the total FDI, inflow in bn USD
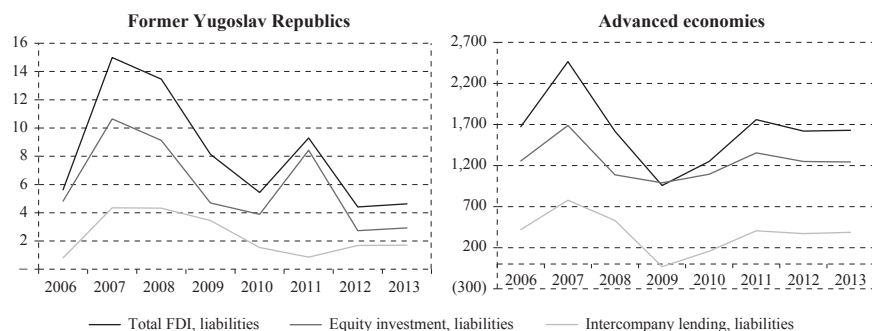
Source of data: IMF.

This post-crisis reversal in capital flows is crucial for this paper. Therefore, it is necessary to compare the inflow of loans, classified as *Other investments* in the Balance of Payments (BoP) and inflow of ICL in the region (figure A1 in appendix). The assumption is that as a solution of mentioned liquidity issues, foreign affiliates sent additional funds to their subsidiaries in the form of increased ICL.

During the observed period countries had a very volatile inflow of foreign loans. The majority of them had a significant decline, and some even had a negative inflow, indicating deleveraging by the banks. In addition, it is visible that although countries differ in their levels and dynamics of ICL inflow, there were no precipitous declines or reversals. This is where we can see the FDI nature of ICL and the concern of foreign affiliates for their subsidiaries. Some would argue that additional funds were not provided in sufficient amounts and that they are essentially a loan. However, it must be taken into account that the whole world was faced with the crisis and that MNC had to manage risk on the global level.

The final comparison is the structure of FDI inflow (figure A2 in appendix). First, Croatia and Serbia, in 2008, experienced, at the same time, a plunge in EI and a growth in ICL. Furthermore, the opposing movement of EI and ICL in 2011 in Croatia represents debt-for-equity swap, which changed the nature of the initial transaction and reduced the Croatian external debt by €0.7bn (CNB, 2012). Second, in Macedonia and Bosnia and Herzegovina, the decline of EI was so severe that ICL, although also on a declining trend, had a stabilising role since the decrease was not as steep. Next is Montenegro, where EI peaked because part of their energy sector being privatised in 2009. Finally, in Slovenia, ICL had an adverse effect on FDI inflow in 2009. However, this changed in 2010 and 2011 when it surged. For Kosovo, ICL apparently is not a crucial source of funds, since EI dominates in the overall inflow of FDI.

Based on these movements, we can see that countries differ when it comes to the structure of the total FDI inflow. However, it is noticeable that even in comparison to EI, ICL shows a certain robustness, and that it is a constant source of funds, available when other, more "important", sources of foreign capital come to a halt.

4 ECONOMETRIC ANALYSIS

Data availability for this particular topic and the region observed posed several constraints for the econometric analysis. First, the methodology for the compilation of the BoP has changed. Several shifts within its main components made old and newly compiled data incomparable and disenabled their simple combination and creation of longer time series. An additional impediment is that the majority of the observed economies are not EU members and do not have the same obligation to revise past data. For them, relevant BoP data can be found only as of 2007. Second, variables included in the models are not methodologically consistent and available at the same frequency. Therefore, these issues led to a decision to use available annual data from relevant databases, such as IMF, World Bank, Eurostat and WIIW and preserve comparability across countries.

Table 1 lists the starting year from which the data are available for each country. The last year included in the analysis is 2016. It is important to note that for those countries where final data were unavailable, preliminary data from the statistical offices and national banks were used.

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Table 1
Starting year of the available dataset

Country	Year	Country	Year	Country	Year
Albania[a]	2007	Hungary	2002	Poland[a]	2004
Bosnia and Herzegovina[a]	2007	Kosovo[a]	2006	Romania	2002
Bulgaria[a]	2007	Latvia	2002	Serbia[a]	2007
Croatia	2002	Lithuania[a]	2004	Slovakia[a]	2004
Czech R.	2002	Macedonia	2002	Slovenia	2002
Estonia	2002	Montenegro[a]	2010		

[a] *For some of the variables one year of observation is N/A.*

The goal of this analysis was to get country-specific estimates for the observed economies, which is why the model averaging is chosen as the primary model. This approach rests on the averaging of the results of several candidate models and gives more robust results (Hansen, 2007). The major benefit is that it does not impose a limitation in the form of one preferred model with predefined variables. Also, it allows the use of the larger set of variables out of which the model creates various combinations and discrete sub-models. In the end, the final estimate is obtained by averaging results across all sub-models, where each receives a proper weight based on the suitable criterion.

The panel data fixed effects model serves as a robustness check for the results obtained by the primary model. For that purpose, we broadened the dataset and populated it with all countries listed in table 1, as they had similar movements in their cross-border financial transactions during the crisis.

The dependent variable is one of the most important macroeconomic indicators, the CAB to GDP. It represents one of the core indicators the United Nations Commission on Sustainable Development uses to observe a sustainable development of the country (UN, 2007). In addition, the importance of this indicator for the observed economies lies in the fact that they are all small and open, and highly dependent on the movement of foreign capital and international trade.

4.1 DETERMINANTS OF THE MODELS
The focus of this paper is the effect of ICL on the CAB. Therefore, other determinants will be briefly explained here, while the expected sign and source of data can be found in table A1 in appendix.

All selected variables can be divided into two parts. The choice of the first set was influenced by relevant publications (Aristovnik, 2006; Caivano and Coniglio, 2016; Loayza, Chong and Calderon, 1999).

Lagged values of the CAB to GDP ratio (*CAB_lag*) – persistence of the CAB deficit exists, and countries need time to overcome shocks.

Oil balance to GDP ratio (*Oil*) – oil prices affect the observed economies significantly since all of them are net oil importers.

Trade Openness (*Open*) – represents the share of a country's foreign trade in GDP.

Macroeconomic uncertainty (*Vix*) – a proxy for instability in the global markets.

The second part of the variables were chosen based on the intertemporal approach, which seeks to explain movements in the CAB as a result of the changes in savings and investment (UN, 2008; IMF, 2009). The following identity is crucial for this part, and it breaks savings and investment down into private and government parts:

$$S - I = Sp + Sg - Ip - Ig \rightarrow CAB = (Sp - Ip) + (Sg - Ig)$$

General Government Budget to GDP ratio (*F.bal*) – budgetary balance of the Government can have a significant impact on the CAB (IMF, 2009).

GDP growth (*GDP_gr*) – a standard measure of the health of the economy.

Unemployment rate (*Unem*) – its impact on expectation and living standard affects both savings and the investment in the economy.

Foreign direct investments – one of the most important sources for financing CAB deficits in the transitional economies. The effect of the FDI depends on several factors, and the sign of the coefficient can be both positive and negative. A positive sign on the coefficient can mean that investments created positive effects and increased country's exports or that there is not enough foreign capital inflow to allow a higher CAB deficit. On the other hand, a negative sign can indicate that FDI produced a higher outflow of capital in the form of income or interest payments, or that the import of equipment increased, which again is beneficial in the long run, since it is assumed that it will create future growth. Therefore, the interpretation, which applies to all of its components, must be country specific.

As noted, for this analysis FDI is broken into EI and ICL. Both variables are expressed in terms of GDP and only those that represent the liabilities of the country were taken into account.
  – **Intercompany lending** (*ICL*) – The effect of this is the topic of the paper. As a part of FDI, it provides additional funds that countries use to generate future growth. However, as debt position, it can also create an outflow in the form of interest and of principal repayment.
  – **Equity investment** (*EI*) – Commonly perceived as the only form of FDI, this type of flow brings growth and future improvement of the CAB. However, it can also generate outflows in the form of profit repatriation.

4.2 MODEL AVERAGING

The idea for the model came from the article by Urosevic, Nedeljkovic and Zildzovic (2012), which observed determinants of the CAB for five CESEE economies. Although the model is similar, the goal of the analysis is different. First, the focus is only on ICL. Second, FDI is broken down into EI and ICL, which makes their effects comparable. Third, the model aims to make a comparative analysis of the economies and show the heterogeneity in the ICL effects among them. Finally, the data used are comparable across countries, which allows a robustness check with a panel regression.

Previously it was mentioned that model averaging, unlike model selection, enables that all relevant information is taken into account by averaging the results of the candidate sub-models. Its resulting estimates consider both the uncertainty and the bias that exist in each of the sub-models, which makes them more robust. However, the crucial element necessary for the optimal results is the criterion that assigns the weights to each of the sub-models. Two methods are proposed in the literature, Bayesian and frequentist. While the first relies on the subjective determination of probabilities and weights, the second one uses well-known criteria. Although many criteria can be found in the literature, many of them exclude heteroscedasticity and non-nested setup, which makes them unsuitable for the analysis of the CAB (Urosevic, Nedeljkovic and Zildzovic, 2012). Therefore, as proposed by Hansen and Racine (2012) for these conditions the best results are given by jackknife model averaging (JMA), which selects the weights by minimising a leave-one-out cross-validation criterion and provides an estimator that is an asymptotic equivalent to the lowest expected squared error. Liu (2012) further adjusted this model and made it applicable for the time series analysis, and this is the approach that will be used in this paper.

The regression model used can be described as follows[1]:

$$y_n = X_n \beta + u_n \tag{1}$$

$$E(u_n | X_n) = 0 \tag{2}$$

$$E(u_n^2 | X_n) = \sigma^2(X_n) \tag{3}$$

where y_n is the CAB to GDP ratio, X_n is the vector of independent variables, and β is the ordinary least square estimator. Also, u_n is the error term that does not preclude heteroscedasticity. If there is an assumption about M number of models, m = 1, 2 … M, where each model is a sub-model and unique combination of independent variables X_n, then for each model there is a set of linear estimators that can be written as:

$$\tilde{\beta}_m = \left(X'_m X_m \right)^{-1} X'_m y \tag{4}$$

[1] Description of the model is taken from Urosevic, Nedeljkovic and Zildzovic (2012).

The final estimate of the model is a weighted average of all sub-model estimations:

$$\tilde{\beta}_m = \sum_{m=1}^{M} \omega_m \tilde{\beta}_m \tag{5}$$

where ω_m is a set of non-negative weights, that sum up to one. As previously mentioned JMA selects the weights by minimizing the leave-one-out cross-validation criterion (CV), defined as:

$$CV_n(\omega) = \frac{1}{N} \omega' \tilde{u}'_{-i} \tilde{u}_{-i} \omega \tag{6}$$

where $\tilde{u}_{-i} = (\tilde{u}_{-i,1}, ..., \tilde{u}_{-i,M})$ is a *NxM* matrix of leave-one-out residuals, where $\tilde{u}_{-i,m}$ are the residuals from the mth model estimated by least squares, excluding the ith observation. Finally, the JMA chooses ω_m which minimizes the $CV_n(m)$.

4.2.1 Model averaging results

Before running the JMA, the stationarity was checked. For each variable within each country, a KPSS test was done, as it shows better performance on small samples (Bart, Franses and Ooms, 1998). Its null hypothesis states that the observed time series is stationary. Results can be found in table A2 in appendix, and for all, except for Montenegrin GDP growth, we failed to reject the null. For this country the problem is the limited sample availability since the complete data set is available only as of 2010. Also, the JMA gave insignificant coefficients for Montenegro, which is why this country is excluded from the subsequent analysis.

Previously it was noted that only the effects of ICL would be in focus and heterogeneity across the economies (table 2). Results obtained are significant, and for all, except for Kosovo, the effect is negative. Remaining variables and their p-values can be found in table A3 in appendix and can serve for future analysis. Furthermore, in order to understand better how important ICL was in comparison to other variables within the specific economy, the standardised coefficients are also included in the analysis. Those were calculated as the product of the estimated coefficient and the ratio between the independent's and dependent's variable standard deviations.

TABLE 2
Jackknife model averaging, ICL coefficients, by country

Country	ICL	p-value
Bosnia and Herzegovina	-1.868	0.022**
Croatia	-0.802	0.002***
Kosovo	3.837	0.003***
Macedonia	-2.818	0.000***
Serbia	-3.488	0.008***
Slovenia	-0.650	0.004***

*** Significance at 5% level, *** significance at 1% level.*

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The negative values of the estimates imply that for those economies that have persistent current account deficits, such as Bosnia and Herzegovina, Macedonia and Serbia, an increase in the share of ICL in GDP is followed by the growth of the CAB deficit in the GDP. Inversely, a similar movement can be spotted for Slovenia and Croatia, which in recent years have had a surplus in the CAB, followed by the negative inflow of total FDI. The only positive effect of ICL on CAB is derived for Kosovo, and the possible explanation is that since the inflow of ICL was at a very low level, it did not provide additional funds that would finance a higher share of CAB deficit in GDP.

As for the magnitude, the strongest negative effect is for Serbia, then Macedonia and Bosnia and Herzegovina, while Slovenia and Croatia demonstrate a somewhat weaker effect. The magnitude of the results serves as an indicator of the differences that exist among the observed countries in their level of dependence on foreign investment inflows.

Furthermore, we include standardised coefficients into the analysis to find the relative importance each variable has within the country (table 3). Again, we can note that for all observed economies, ICL had a high contribution to the CAB. However, it is the largest for Kosovo and Serbia. As previously noted, the positive sign for Kosovo signals an insufficient inflow of ICL. Here, the stock of EI is at a low level, and foreign investors do not have an incentive to send additional ICL. Furthermore, the negative value of the EI coefficient proves that this is the FDI component that serves as a source for CAB financing. For Serbia, it is interesting to see that, when observed separately, ICL has a negative effect, while EI has a positive impact on the CAB, which is a result of successful privatisations that induced export. Therefore, here we can argue that in Serbia ICL is sent as a constant fuel that allows further growth of the initial investment and finances the CAB deficit, caused by the import of intermediary goods and equipment.

TABLE 3

Jackknife model averaging, standardised coefficients, by country

Variable	B&H	Croatia	Kosovo	Macedonia	Serbia	Slovenia
ICL	-0.482	-0.270	**1.236**	-0.890	**-0.688**	-0.154
CAB_lag	0.174	0.301	-0.081	0.519		**0.367**
EI	**-0.879**	-0.019	-0.800	-0.375	0.202	-0.105
Oil	0.597	0.231	-0.203	-0.151	0.282	0.259
Open		0.099	-0.033	0.327		0.257
GDP_gr	0.088	**-0.542**	-0.664	0.717	-0.187	-0.152
Unem	0.369	0.289	-0.031	-0.039	0.495	0.281
F.bal	-0.151	0.119	0.331		0.451	0.002
Vix	-0.233	0.127	-0.405	**-0.916**	-0.485	

Bolded figures indicate the strongest effect.

For Macedonia and Bosnia and Herzegovina, the contribution of ICL is also among the highest ones. For Bosnia and Herzegovina, which struggles to attract

FDI (U.S. Department of State, 2015), the highest contribution is derived for EI. That result signals that this country still needs to attract the critical amount of this type of FDI and that those already obtained had an important effect on the financing of CAB deficit. After that happens, the relative significance of ICL might increase. For Macedonia the highest contributions are for VIX and ICL, which can suggest that this is a country highly dependent on foreign capital and highly sensitive to the movements of the world economy. For Slovenia and Croatia, the ICL contribution is not as important as for the mentioned three countries. Additionally, the importance of other variables is somewhat balanced. That may indicate that their CAB is not overly dependent on one particular variable.

The analysis can be further broadened when we observe how the contribution of ICL to the CAB changed over the years (figure A3 in appendix). We can see that for Serbia and Macedonia, the strength of the effect fluctuates but remains high, which is in line with the previously stated high dependency on foreign investment inflow. However, for Bosnia and Herzegovina, this contribution declines and can be explained by the very slow increase in the EI inflow, and subsequent drop in the ICL inflow. As previously noted, its primary concern should be attracting EI. In Slovenia and Croatia, the contribution was modest across years and became positive when countries started achieving a surplus in the CAB, and this is also in line with the previous analysis of the results.

4.3 PANEL REGRESSION FIXED EFFECTS

The second model serves as a robustness check of the results obtained the JMA. This model uses the same variables as JMA, however, here it was used on an unbalanced panel with 17 economies (table 1). The estimated model can be written as:

$$y_{in} = \alpha + X_{in}\beta + \gamma t_i + u_{in} \tag{7}$$

where y_{in} is the CAB to GDP ratio, X_{in} is the vector of independent variables, t_i is the trend dummy, and β and γ are coefficient estimates. Also, u_{in} is the error term.

The unit root tests were done for the whole dataset. Since it failed to reject its existence for dependent and several control variables, the trend dummy variable was added to control for the trend. Results obtained before and after adding the trend dummy (table 4) did not differ substantially, which is why we decided to accept the results of the panel regression and continue with the analysis.

Results in table 4 show that the ICL effect is significant, while the same is not true for that of EI. One way to explain this is that once initial investment happens in the form of EI, foreign investors opt to send additional funds to the CESEE region in the form of ICL. This supports the view that since the crisis started, the FDI inflow has been sustained by the ICL in the observed countries. The negative sign is also expected, since FDI in the transitional economies represents funds used for financing the CAB deficit.

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TABLE 4

Panel regression results, with and without trend dummy

Coefficients	Without trend dummy		With trend dummy	
	Estimate	**p-value**	**Estimate**	**p-value**
ICL	-0.372	0.0291**	-0.368	0.0306**
EI	-0.012	0.8576	-0.011	0.8787
Oil	0.665	0.0044***	0.634	0.0047***
Open	0.186	0.0000***	0.176	0.0000***
GDP_gr	-0.487	0.0000***	-0.469	0.0000***
Unem	0.249	0.0005***	0.253	0.0007***
F.bal	0.130	0.3883	0.132	0.3764
VIX	-0.008	0.3651	-0.008	0.3483
Dummy trend			0.0004	0.7613

** Significance at 5% level, *** significance at 1% level.

5 CONCLUSION

5.1 SUMMARY OF FINDINGS

The assumption from which this analysis was derived is that in the crisis periods, countries that are dependent on foreign capital can face many problems if a sudden stop occurs. For the observed economies, September 2008 marked the beginning of the crisis and the sudden decline in the inflow of foreign funds.

As noted, the financial sector in the CESEE is mostly foreign owned, and pre-crisis growth rates were greatly fuelled by the loans provided by the major banking groups. Additionally, the crisis affected the private sector. However, those that were foreign owned got help in the form of ICL. Although it shares some characteristics with commercial banking loans, ICL is an FDI. The core element of FDI is the long-term commitment that an MNC makes when it invests initial equity. Afterwards, the subsidiary is expected to function on its own and generate profits. Additional funds can be sent because there are either growth opportunities or problems. After the crisis, growth opportunities became rare, so we assume that ICL was sent as a support.

Reasons why the ICL and not some other form of capital was sent can be observed from the risk management point of view. As explained before, for an MNC they provide flexibility, an additional channel of control and lower costs. The worldwide crisis and perception of risk in the region made ICL a reasonable option and influenced the amount sent. An additional element that proves this point is the debt-for-equity swaps. If we look again at the structure of the FDI for the observed economies (figure A2 in appendix), movements in the FDI inflows show the existence of several debt-for-equity swaps, which imply risk management strategy. If the investors perceive the economy to be risky, and they want to keep flexibility in their capital flows, they can initially send funds in the form of the loan, and then, depending on the situation transform it into equity or pull it back. In these economies, all plunges in ICL were followed by the surges in EI[2].

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[2] In Croatia in 2015, debt-for-equity swaps prevented even larger decline of total FDI since EI plunged due to value adjustments and negative reinvestment of earnings (EBL News, 2016).

Furthermore, if we add that the majority of investments came into the region as the brownfield investments and that this source of EI has almost dried up, then we can note that ICL is an FDI component that sustains the total inflow. Therefore, if we envision ICL as funds that fuel previously established direct investment and are an "introductory" form of current EI, it is easy to comprehend the magnitude of the effect derived in the model and how the benefits of this FDI component are created.

Finally, the opponents of ICL can state that the outflow of capital through interest payments is a drain on the economy and that it increases external debt. The former is not true for the observed economies since they have not experienced a substantial outflow of funds through ICL-based interest payments. As for the latter, debt exposures toward affiliated companies are considered to be more sustainable because of the FDI relationship. Furthermore, the frequently recorded debt-for-equity swaps have the same effect on external debt level as debt forgiveness. Overall, this debt is more stable.

5.2 POLICY IMPLICATIONS AND RECOMMENDATIONS

The goal of the paper was to observe whether or not the ICL inflow had a significant effect on the observed economies. Two models that were employed indicated that it did. For the majority of countries, that effect is negative, which is not surprising since in many of the observed economies domestic savings are not sufficient to finance investments, and additional funds must be imported. Therefore, when deciding which policy to pursue, an increase in the CAB deficit because of the surge in investment should not be considered problematic, since, in the long run, it will generate growth.

The first set of results obtained by JMA allows cross-country comparison and indicates which countries are more dependent on foreign financing, while the second set, with the standardised coefficients, allows country-specific analysis and shows areas where policy actions are most important.

Essentially, when it comes to the policy implications for ICL, one crucial element must be taken into account. ICL are conditioned with the existence of the direct investment relationship, and its inflow is dependent on the policies that promote EI. Therefore, countries should concentrate on the implementation of policies that attract it, and consequently, inclusion in the multinational networks will bring higher performance (Alfaro and Chen, 2010).

Therefore, the set of recommendations should be foremost concerned with the attraction of new EI. In the past, foreign investors were mostly interested in privatisation and brownfield investments. Since the potential for those is decreasing, it would be beneficial to promote greenfield investments. Furthermore, as noted by Estrin and Uvalic (2016), FDI in the region was mostly focused on the financial sector, and not enough in the manufacturing, which may be a reason why these countries still are not able to integrate better into the global market and benefit from the higher export potential. In addition to that, for the majority of these economies,

full integration into the European Union would bring significant benefits. One of the biggest ones would be an improvement in the institutional quality that these economies lack. The severe effect of this is the increased perception of risk investors have, that ultimately undermines the higher inflow of foreign capital (Estrin and Uvalic, 2013).

Finally, when giving a policy recommendation, one must take into account the other side of the ICL. Even though in the paper the focus was on the identification of the benefits from ICL, it must not be forgotten that it can induce transfer pricing. That is why the implementation of policies and participation in the initiatives such as the BEPS are of utmost importance. Ultimately, this can lead to a situation in which ICL is no longer perceived as a rent-seeking transaction and starts to be perceived as a significant advantage to those companies that are a part of the FDI circle.

5.3 LIMITATIONS OF THE STUDY AND FUTURE RESEARCH POSSIBILITIES

Certain limitations of the study exist, and they are mostly concerned with the availability of the data. The issue of the short time series may have impeded an econometric analysis that would have yielded more convincing results. The methodological changes and difficulties in finding relevant quarterly data further complicated the analysis. Additionally, it restrained the full potential of the model averaging technique, since the number of variables had to align with the number of observations available.

When we observe the region of interest, only two countries are members of the EU and have more stringent rules when it comes to the availability of data sets. Others choose whether to comply with them, as it is not compulsory. Since in this paper the emphasis was on the comparability across economies, the same number of variables was used for all models. Therefore, for those countries where longer series are available, the data limitation can be overstepped, and a more detailed analysis with additional variables can be made. Additionally, this can also provide a solid ground for other economies when the longer time series become available. Finally, the analysis does not have to stop at the effect of ICL, and the data in the appendix can also be a starting point for further research.

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TABLE A1
Variable description

Variable	Expected sign	Source
Lagged values of the CAB to GDP ratio *(CAB_lag)*	+	Vienna Institute for International Economic Studies (WIIW)
Oil balance to GDP ratio *(Oil)*	+	Eurostat, International Trade in Mineral fuels, lubricants and related materials (SITC Rev. 4) served as a proxy
Trade Openness (*Open*)	Ambiguous	World Bank
Macroeconomic uncertainty (*Vix*)	Ambiguous	Federal Reserve Bank of St. Louis, CBOE Volatility Index, % change from a year ago, annual, not seasonally adjusted
General Government Budget to GDP ratio (*F.bal*)	+	WIIW
GDP growth (*GDP_gr*)	–	WIIW
Unemployment rate (*Unem*)	Ambiguous	WIIW
Foreign direct investments: Intercompany lending (*ICL*) Equity investment (*EI*)	Ambiguous	International Monetary Fund

TABLE A2
Jackknife model averaging, unit root tests results

Variables	B&H	Croatia	Kosovo	Macedonia	Montenegro	Serbia	Slovenia
				KPSS test results			
CAB	0.392*	0.440*	0.107	0.306	0.326	0.416	0.729**
ICL	0.390*	0.230	0.302	0.185	0.144	0.347*	0.131
EI	0.251	0.282	0.332	0.223		0.205	0.348*
Oil	0.211	0.137	0.314	0.175	0.287	0.407	0.148
Open	0.314	0.363*	0.231	0.389*	0.089	0.698**	0.391*
GDP_gr	0.153	0.260	0.306	0.144	1.235***	0.186	0.239
Unem	0.135	0.292	0.261	0.560**	0.514**	0.196	0.382*
F.bal	0.144	0.124	0.349*	0.226	0.253	0.165	0.267
Vix	0.085	0.050	0.097	0.050		0.085	0.050

The null hypothesis for KPSS is existence of stationarity, and it is rejected if obtained value is above the critical. Critical values are 0.739 (*** significant at 1%), 0.463 (** significant at 5%), 0.347 (* significant at 10%).
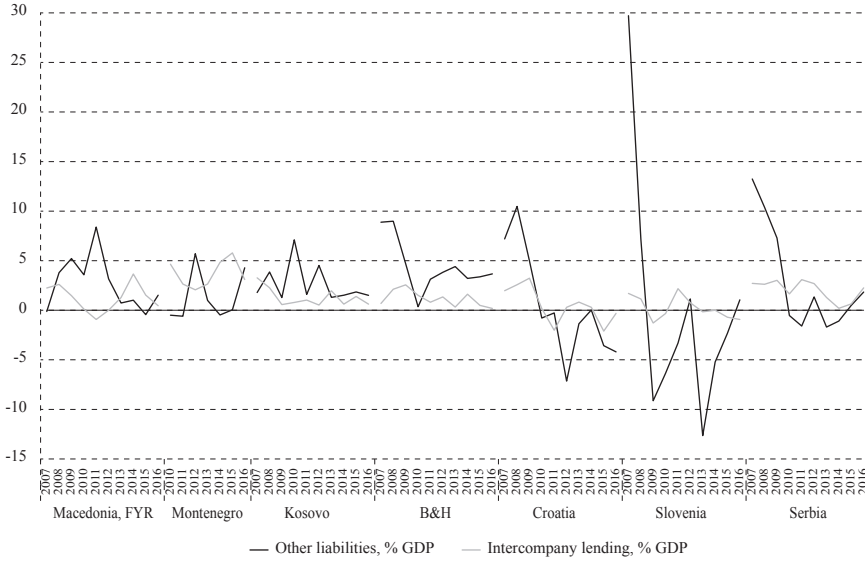
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TABLE A3

Coefficient estimates obtained by the Jackknife model averaging approach

Variable	B&H		Croatia		Kosovo		Macedonia		Montenegro		Serbia		Slovenia	
	Coeff.	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**	**Coeff.**	**p-value**
							Model averaging coefficients							
Intercept	-0.147	0.020	-0.062	0.000	-0.056	0.003	-0.087	0.000			-0.075	0.008	-0.077	0.000
ICL	-1.868	0.022	-0.802	0.002	3.837	0.003	-2.818	0.000	0.421	0.124*	-3.488	0.008	-0.650	0.004
CAB_lag	0.134	0.011	0.322	0.000	-0.082	0.013	0.517	0.000	-0.394	0.016			0.412	0.000
EI	-0.759	0.010	-0.009	0.000	-0.995	0.004	-0.688	0.000			0.566	0.011	-0.292	0.000
Oil	0.619	0.014	1.036	0.000	-0.378	0.015	-0.277	0.001	0.576	0.024	1.101	0.006	0.665	0.000
Open			0.067	0.000	-0.021	0.003	0.075	0.000	-0.179	0.024			0.060	0.000
GDP_gr	0.092	0.060	-0.624	0.000	-0.962	0.010	1.284	0.001	-0.191	0.072	-0.380	0.011	-0.156	0.001
Unem	0.561	0.020	0.428	0.000	-0.007	0.014	-0.034	0.000	-0.183	0.075	0.867	0.008	0.617	0.000
F.bal	-0.266	0.011	0.284	0.025	0.351	0.003	0.000		-0.305	0.079	1.443	0.014	0.002	0.002
Vix	-0.015	0.021	0.015	0.026	-0.029	0.007	-0.092	0.001			-0.067	0.004		

* Significance at 10% level, ** significance at 5% level, *** significance at 1% level.

FIGURE A1

Other investments and intercompany lending, liabilities, % of GDP

Source of data: WIIW.

FIGURE A2

The structure of FDI, inflow, in bn USD

Source of data: IMF.

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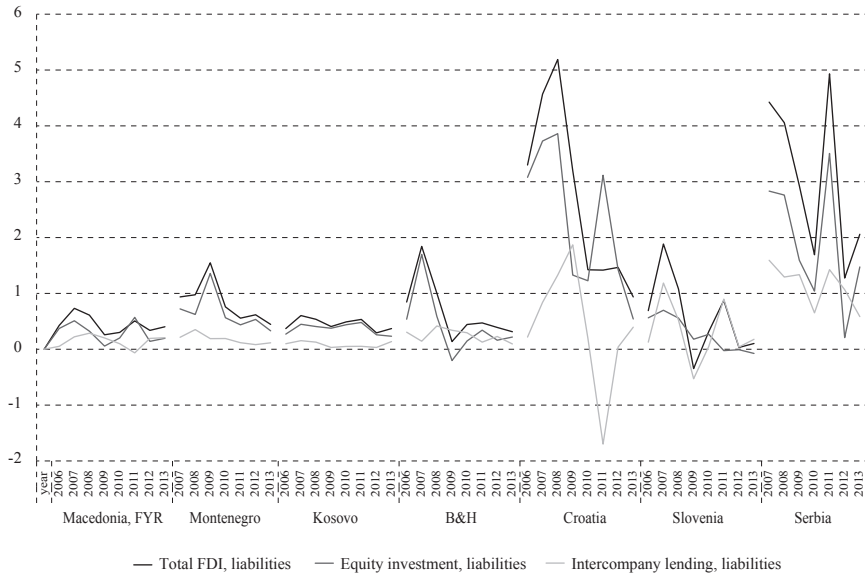
FIGURE A3

The contribution of ICL to the CAB, two and three-year averages, % of GDP
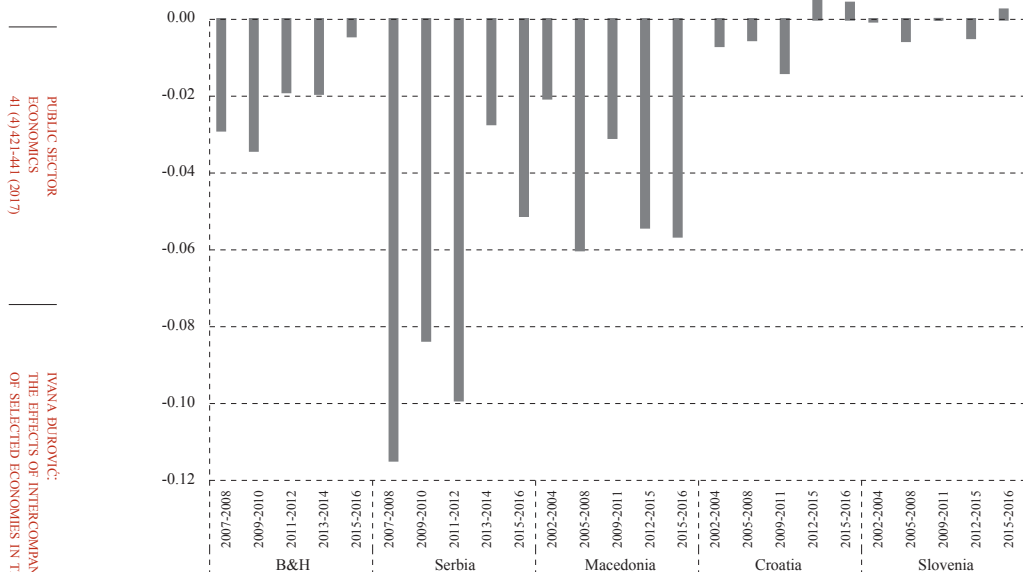
Source: Results of the author.

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Economic uncertainty and its impact on the Croatian economy

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Abstract

The aim of this paper is to quantify institutional (political and fiscal) and non-institutional uncertainty (economic policy uncertainty, Economists' recession index, natural disasters-related uncertainty, and several disagreement measures). The stated indicators are based on articles from highly popular Croatian news portals, the repository of law amendments (Narodne novine), and Business and Consumer Surveys. We also introduce a composite uncertainty indicator, obtained by the principal components method. The analysis of a structural VAR model of the Croatian economy (both with fixed and time-varying parameters) has showed that a vast part of the analysed indicators are significant predictors of economic activity. It is demonstrated that their impact on industrial production is the strongest in the onset of a crisis. On the other hand, the influence of fiscal uncertainty exhibits just the opposite tendencies. It strengthens with the intensification of economic activity, which partially exculpates the possible utilization of fiscal expansion as a counter-crisis tool.

Keywords: economic uncertainty, Economic Policy Uncertainty Index, VAR model with time-varying parameters

"The only relevant thing is uncertainty – the extent of our own knowledge and ignorance. The actual fact of whether or not the events considered are in some sense determined, or known by other people, and so on, is of no consequence."

De Finetti, 1975:vi

1 INTRODUCTION

The collapse of Lehman Brothers and the downturn of the US economy in 2008 triggered a domino effect of spillovers on almost all of the world's national economies. The crisis affected some countries only marginally (e.g. Poland and Slovakia), and some European countries recovered from the global shock in relatively short periods (Germany, Austria, Benelux countries), while the crisis in Croatia proved to be extremely resilient and long lasting. Two years passed from the evident onset of the crisis in Q3 2008 (Krznar, 2011) to the first transition of GDP growth to the positive domain. However, the positive shift was only short-lived. A double dip recession became evident in Q4 2011, followed by as many as 12 consecutive quarters of declining economic activity. This negative streak was one of the longest recessions in the whole of Europe (Buturac, 2017:23) and has triggered an entire series of domestic studies.

Croatian economic practitioners and academics mostly agree that the crisis was the multi-layered consequence of numerous factors. Many authors (from several scientific disciplines) have referred to the demographic issues, the meta-determinants of weak competitiveness such as the egalitarian syndrome (Burić and Štulhofer, 2016), the (too) high levels of public debt (Buturac, 2017), etc. How-

ever, the international literature also offers relevant evidence on the considerable influence of uncertainty on the decreased economic activity during the recent crisis (Bloom, 2009; Bachmann, Elstner and Sims, 2013; Baker, Bloom and Davis, 2016, etc.). These studies are (with rare exceptions such as Arčabić, 2015) not followed by corresponding analysis for the Croatian economy. Therefore, a question is raised: to what extent can the longevity and intensity of the recent Croatian economic crisis be attributed to uncertainty itself? That is the main research question of this paper, and its basic contribution to the literature.

This paper introduces an entire set of uncertainty indicators to the Croatian practice of economic research: political uncertainty, recession-related uncertainty (The Economist, 2002), fiscal uncertainty resulting from changes of fiscal legislation, uncertainty stemming from natural disasters and a Croatian version of the Baker, Bloom and Davis (2016) economic policy uncertainty index. We also consider indicators of prognostic disagreement (following Bachmann, Elstner and Sims, 2013) as proxy variables for economic uncertainty. Additionally, we employ the principle components method to generate a composite indicator of aggregate uncertainty as a linear combination of the previously stated variables.

The second specificity of this paper is the interdisciplinary approach used to elucidate the macroeconomic phenomenon of recession by utilizing the psychological concept of uncertainty (while controlling for other relevant micro- and macroeconomic predictors of economic activity). The third contribution to the literature is the analysis of the potentially time-varying influence of uncertainty on economic activity, for previous papers have found that uncertainty is counter-cyclical (in the sense of correlation; Bloom, 2014), but do not analyse whether the uncertainty effect is different in various phases of the economic cycle. This paper employs a VAR model with time-variable parameters (Primiceri, 2005; del Negro and Primiceri, 2015) to test the hypothesis that economic activity is more responsive to uncertainty (in terms of the impulse response function) in a recession than in the expansionary phase of the economic cycle.

The econometric analysis has shown that the observed uncertainty measures on average have a significant, but only temporary effect on economic activity. Such an inference regarding the negative short-run relationship between uncertainty and economic activity is quite robust to several alternative model specifications (structural VAR model with fixed parameters, the inclusion of foreign variables due to the fact that Croatia is a small open economy, and alternation of the control variables in the VAR model with time-varying parameters). Nevertheless, we show that the negative impact of most of the uncertainty measures employed becomes more pronounced in the contractionary phase of the cycle.

The remainder of the paper is structured as follows. Section 2 presents the main theoretical considerations of the relationship between uncertainty and economic activity and an overview of the empirical approaches to uncertainty quantification.

Section 3 deals with the data and methodological issues. Section 4 lays out the obtained econometric results, while the final section synthesizes the paper and sums up the main policy implications.

2 THEORETICAL CONCEPTS OF THE INTERRELATIONSHIP BETWEEN UNCERTAINTY AND THE REAL ECONOMY

The term *uncertainty* was introduced to the literature through a famous book by Frank Knight (1921), resulting in the concept nowadays often being referred to as Knightian uncertainty. The concept implies a situation in which economic agents are not able to anticipate future events or estimate the likelihood of their occurrence. Although uncertainty per se is a latent variable, the channels of its influence on economic activity are well elaborated in the literature.

The most meticulous overview to date of theoretical frameworks describing the uncertainty-economic activity link is presented by Bloom (2014). Arčabić (2015) is a rare example of the genre among the domestic Croatian authors. The two stated authors emphasize two negative, but also two positive transmission channels of uncertainty.

Bernanke (1983) introduces the first negative channel, and Bloom (2014) later on refers to it as the *real options* channel. Firms faced with uncertainty postpone their irreversible decisions in order to enrich their information set and reduce the uncertainty related to making such decisions. Therefore, the *wait and see* mechanism (Bachmann and Bayer, 2013) is activated because activities such as investments, employment, and buying durable goods generate certain sunk costs. Consumers can also activate the exact same mechanisms in the sense of postponing the consumption of durable goods or investing in financial assets.

It should also be highlighted that there are four conditions under which the real options channel is efficient (Bloom, 2014:163-164): (i) the existence of considerable adjustment costs and irreversible decisions, (ii) the possibility of postponing decisions (partially or completely eliminated by perfect competition), (iii) firms operating in imperfectly competitive markets with decreasing returns to scale, and (iv) uncertainty should be variable (not constant).

The second channel of influence is *risk aversion*. In a corporative sense, growth in uncertainty increases the risk premium and increases the cost of borrowing. It also discourages the company's management from taking part in risky endeavours (especially if the management has shares in the company), and it ultimately decreases the aggregate investments and the overall economic activity. In the consumer sector, risk aversion is effectuated through precautionary savings, which again decreases personal consumption and consequently lowers GDP (Carroll, 1992).

The two positive transmission channels are the *growth options* and the *Oi-Hartman-Abel* effect. The first one is the antipode to the real options channel, relating

to the fact that uncertainty can also be an incentive for risky investments in situations of high potential gain (e.g. the global economic growth due to the Internet revolution, the so-called *dot-com boom* in the second half of the 1990s). The Oi-Hartman-Abel effect is based on hedging against negative business results, or situations in which the potential gain from a positive outcome considerably surpasses the potential losses from bad business results. In such scenarios, firms often tend to be risk-seeking and to take part in new investments.

The existing empirical findings offer far more evidence in favour of the negative influences of uncertainty (see Bloom (2014) for an excellent literature review), so it remains to be seen if the same will be corroborated for the Croatian economy.

The latent character of uncertainty has conditioned the fact that it was empirically treated only descriptively until the global financial crisis in 2008. The prevailing turbulent economic conditions in 2008 triggered several methodological strands in the quantification of economic uncertainty. Their basic principles will be presented in the following subsections.

2.1 BAKER BLOOM DAVIS INDEX OF ECONOMIC POLICY UNCERTAINTY

Prompted by the hypothesis that uncertainty is one of the major causes of such an atypically long and strong recession of the US economy after the Lehman Brothers bankruptcy, Baker, Bloom and Davis (2012) published their index of economic policy uncertainty (EPU index). The index is based on three pillars. The first is the archive of media news reports on economic uncertainty from 10 leading US newspapers. Searching through that archive using specific keyword combinations (e.g. *economic + uncertainty/uncertain + congress/deficit/legislation*, etc.) enabled Baker, Bloom and Davis (2012) to publish a scaled measure of monthly frequencies of articles dealing with economic uncertainty in the US.

As the second pillar of their index, the authors establish the number of federal tax code provisions expiring in the following 10 years.

The last EPU pillar is a disagreement measure based on the responses from the Survey of Professional Forecasters (SPF). To be specific, the authors use disagreement in forecasts of the consumer price index and the purchases of goods and services by the federal/state/local governments. The final estimate of EPU index is obtained each month by attaching a specific weight to each EPU component. A 1/2 weight is attached to the media index of economic uncertainty, and a 1/6 is attached to the expiring tax code provisions, price level disagreement, and government purchases.

The US public has mostly perceived the EPU index as a republican criticism of the Obama administration. However, it has also drawn the attention of researchers from various scientific fields. Baker, Bloom and Davis (2016) published a revised version of their research on uncertainty, showing that it affects the micro level

through higher stock price volatility on the US market. This effect (analysed through panel regressions) is shown to be particularly strong in companies closely tied to the government (defence sector, health care, and the construction sector).

The authors also use VAR modelling to show that the macro effects are very similar. Uncertainty reduces investments, employment levels in US companies, and US industrial production. The exposed conclusions are corroborated using various robustness checks and adding different control variables, and are also confirmed by a panel VAR model comprising 12 world economies that have had their own versions of the EPU index at the time.

It is also worth mentioning Karnizova and Li (2014), who confirm the predictive characteristics of EPU in forecasting the US recessionary episodes; while Ajmi et al. (2015) point to the existence of (bi-directional) Granger causality between EPU and the US stock market yields.

Considering the strong evidence of the significant influence of EPU index on the US economic activity, one of the goals of this paper is to quantify the Croatian version of a media-based uncertainty index, and scrutinize if it is characterized by the same effects as that for the USA.

2.2 THE ECONOMIST'S RECESSION INDEX

The Economist has conceptualized its R-word (recession) index as early as the beginning of the 2000s (The Economist, 2002). This indicator is based on the quarterly frequency of recession-related articles in the Washington Post and New York Times. The index witnessed its full affirmation in the scientific sense only with the onset of the global financial crisis, when it started to be calculated and published in other countries as well as the USA. For example, Grossarth-Maticek and Mayr (2008) point to the significance of the R-word index in dating the German business cycle by a probit regression model. Iselin and Siliverstovs (2013) employ the autoregressive-distributed lag (ARDL) models of Swiss GDP, using the R-word index as the independent variable in the model. By doing so, they increase the predictive accuracy of simple univariate benchmark models. Iselin and Siliverstovs (2016) provide very similar results for Switzerland, while the same models for the German economy show to be less precise.

Considering the well-known fact that aggregate economic uncertainty is considerably higher in recessions (in comparison to the stable growth periods) (Bloom, 2009; 2014), the R-word index can also be viewed as a particular type of an uncertainty indicator. Since Croatia has not yet seen its own version of the R-word index, it will be interesting to quantify one and analyse its influence on the Croatian economic activity, and this will be carried out in the remainder of the paper.

2.3 FORECASTING DISAGREEMENT AS A PROXY FOR UNCERTAINTY

Bachmann, Elstner and Sims (2013) provide a follow-up on the third element of the EPU index, and postulate that uncertainty can be proxied by the degree of disagreement among economic forecasters. Their basic premise is that, in times of high uncertainty, the respondents in SPF would give remarkably heterogeneous answers to the questions focused on relevant variables from their economic surroundings. However, a high level of response dispersion does not necessarily need to imply high uncertainty. That is the reason why some authors insist on a clear distinction between the terms *disagreement* and *uncertainty* in the classical Knightian sense (Jurado, Ludvigson and Ng, 2015; Krüger and Nolte, 2016).

In the context of this paper, four separate indicators of disagreement are considered as proxy variables for economic uncertainty. All of them are based on the data from regular monthly Business and Consumer Surveys (BCS), published by the European Commission. The first indicator relates to consumers' disagreements on future tendencies of the general economic situation in Croatia. The second and third indicators reflect the uncertainty of consumer expectations of future unemployment and the timeliness of buying permanent goods. The fourth indicator is a disagreement measure stemming from all forward-looking questions in BCS. All four stated indicators will be explained in more detail in section 3.1.

Bachmann, Elstner and Sims (2013) use the described methodology to show that (*ex ante*) prognostic disagreement, just as the *ex post* prognostic errors of economic agents, considerably reduce economic activity in the short run (measured by industrial production, working hours, and employment level). Their results for the US turn out to be more significant than those for the German economy.

2.4 ALTERNATIVE EMPIRICAL APPROACHES

Apart from the widely recognized uncertainty measures (which might be valid for any analysed economy), one should also take into account the specificities of post-transition economies such as Croatia. Bloom (2014:162) emphasizes three specific uncertainty sources in economic systems of that type. First of all, they are often exposed to remarkable political instabilities and international conflicts. Due to their lower level of socio-economic development and insufficient infrastructure, they are also more prone to natural disasters such as floods or epidemics. The third group of relevant uncertainty drivers stems from the inefficient channels of fiscal and monetary policy.

In the last two decades, Croatia has not witnessed a coup d'état or a war, but the Croatian form of crony capitalism has revealed the link between the state and economic activity as a very important one. The elections play a particularly important role in generating political uncertainty (Mačkić, 2014) (especially with regards to the repeated parliamentary elections in September 2016 after only 10 months of the Orešković' government). Other sources of political uncertainty include the prevailing perception of public companies as the spoils of politics as well as an entire series of corruption scandals related to prominent political officials.

Natural disasters also seem to be potentially relevant for the Croatian disasters, especially considering the dependency of agricultural production on the weather conditions, such as droughts and hailstorms, as well as floods, like those in Eastern Slavonia in May 2014.

Considering the continuity of Croatian monetary policy in employing the kuna/ euro exchange rate as the nominal anchor for maintaining price stability, we postulate that monetary uncertainty (if there is any at all) is not relevant in Croatia, and will not be examined here.[1] On the other hand, fiscal uncertainty in Croatia is a problem that was recognized a long time ago (Bejaković, 2009).

The three stated uncertainty sources (political instability, natural disasters and fiscal instability) will be scrutinized in the following sections, along with their corresponding quantitative uncertainty indicators.

3 DATA AND METHODOLOGY

This section presents a brief overview of the analysed uncertainty indicators (as well as the other examined macroeconomic variables). The methodological basis of the utilized structural VAR models (with fixed and time-varying parameters) is presented afterwards.

3.1 DATA

The main methodological innovation of this paper is the introduction of several uncertainty indicators that have so far not been considered for the Croatian economy. To be specific, four monthly media-based indicators are formed: Baker Bloom Davis index (*BBD*)[2], political uncertainty index (*polit*), the R-word index (*R_index*), and an indicator of uncertainty related to natural disasters (*nature*). With the goal of quantifying the named indicators, we formed a database of media articles from the Internet archives of the most read Croatian news portals: Jutarnji list, Večernji list, 24sata, index.hr, dnevnik.hr, and Poslovni dnevnik. Besides the fact that these are the most popular Croatian daily newspapers, their web portals are regularly among the most visited of all Croatian web sites (Gemius Audience, 2017). Further on, these are (to the best of the authors' knowledge), the only news portals to have a coherent article archive without the breaks constituted by methodological changes. It is, therefore, reasonable to conclude that the described database will serve its purpose as the foundation for quantifying media-based economic uncertainty in Croatia. The database comprises articles dating from November 2002 to December 2016, but each of the six examined web portals has its own starting date. Finally, the database comprises 1,030,768 articles for

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[1] The readers may consult Governor Vujčić's frequent public presentations (e.g. Vujčić, 2016a; b) as indications of the strict focus on the stability of the HRK/EUR exchange rate.

[2] The Croatian version of the *BBD* index, as suggested in this paper, is founded solely on the media reports on uncertainty. The uncertainty stemming from tax changes will be evaluated separately, just as the influence of forecasting disagreement (quantified through equation 1) on economic activity, will be treated separately. Therefore, a separate label is introduced for the Baker Bloom Davis index (*BBD*), different from the US EPU index, comprising all three original components.

the observed period, out of which 10,491 are employed for the calculation of EPU index, 225,691 are used for the political uncertainty index, 36,810 for the natural disasters index, and 77,843 for the R-word index.

Using *Structured Query Language* (SQL), we search the database by pre-formulated combinations of keywords, different for each uncertainty indicator. A detailed list of keywords is given in appendix 1. For example, an article enters the quantification of *BBD* index if it contains at least one of the keywords in part (a) of appendix 1 (e.g. economy, economic, etc.), at the same time at least one of the words in column "Logical conjunction I" (e.g. minister, prime minister), and some of the words from column "Logical conjunction II" (e.g. uncertainty, uncertain, etc.). The same logic (with the complete list of keywords given in appendix 1) is also applied for the *polit*[3], *R_index* and *nature* indicators. After separating the articles that satisfy the stated criteria, we formed the monthly relative frequencies (number of extracted articles divided by the number of total articles). The obtained results for each individual web portal are then standardized (in accordance with Baker, Bloom and Davis (2016) and Girardi and Reuter (2016) and averaged by months.

The fifth analysed uncertainty indicator refers to changes in fiscal legislation. To quantify it, we again use SQL to filter the legal acts from the web repository of law amendments (Narodne novine). The keywords utilized for this purpose are also given in appendix 1. From the total set of law acts containing the term "tax" in their title, we exclude those acts concerning the appointment or dismissal of officials in the Government or the Tax Administration, as well as specific changes in the tax rulebooks valid at the municipality/city/county level. Out of the total 49,749 legal acts, we extracted 298 "true" tax changes. In this specific case, we did not form the uncertainty indicator out of the obtained monthly frequencies of tax changes since a considerable number of months did not see any law changes at all (satisfying the adopted criteria). The intensity of the influence of a particular law change (in the sense of (un)certainty) is the strongest immediately before and after its passing and publication in Narodne novine. To account for that, we form a 6-months cumulative (3 months before and 3 months after the passing of a law) of the number of tax changes (*PI*6), and the econometric analysis is continued with that indicator.[4]

The basic five indicators analysed in this paper are presented in graphs 1-5. The BBD index shows a below average level of uncertainty right until the onset of the 2008 crisis, followed by an obvious, considerable and long lasting increase of uncertainty in the economic system.

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[3] It should be noted that the list of keywords for the *polit* indicator does not comprise the terms "country prefect", "mayor", etc. The authors' intention was to extract only those political shocks that might significantly influence the Croatian macroeconomic tendencies. Following that approach, our analysis did not include (inter alia) the corruption scandals related to local potentates like the mayors Vlahović (Dubrovnik), Sabo (Vukovar) or Bandić (Zagreb), and the local prefect Lovrić Merzel (Sisačko-Moslavačka County).
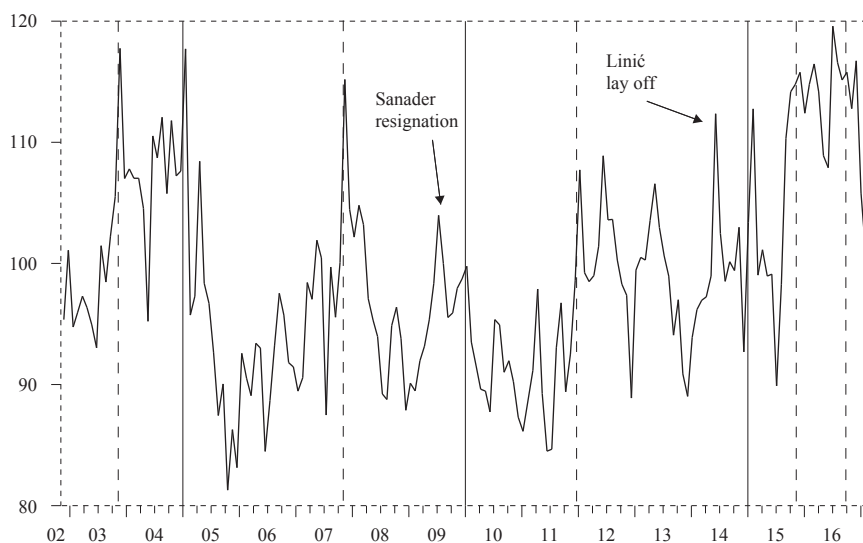
[4] The results are very similar if a 4-month or 8-month cumulative is chosen.

GRAPH 1

Baker Bloom Davis index
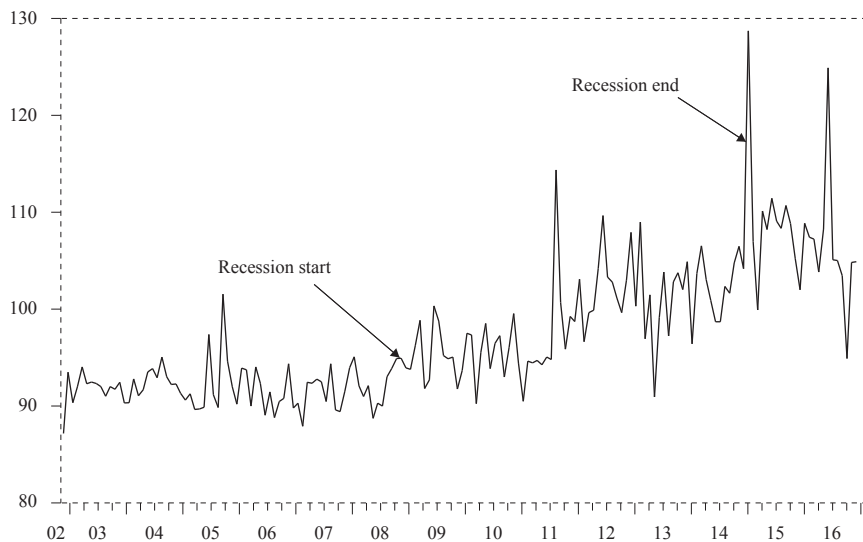
Source: Authors' calculations.

GRAPH 2

Political uncertainty index

Note: Dotted lines point to the dates of parliamentary elections, while full lines point to presidential elections.
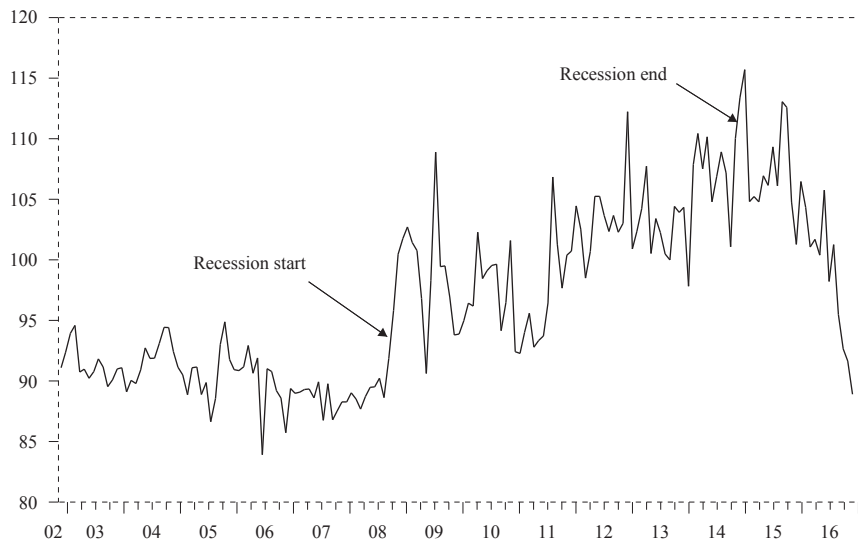
Source: Authors' calculations.

Graph 2 vividly shows that the local maxima of political uncertainty correspond to the pre-election campaigns (related to parliamentary or presidential elections). Therefore, political uncertainty cycles mostly coincide with the electoral cycles. Rare exceptions from that regularity are the political shocks like the resignation of

former prime-minister Ivo Sanader in July 2009 and the reconstruction of Milanović's government (Slavko Linić lay off in May 2014 and the changes of ministers Jovanović-Mornar and Ostojić-Varga one month later).

GRAPH 3

The R-word index

Source: Authors' calculations.

The R-word index graphically looks very similar to the *BBD* index, presenting a cyclical match between economic activity and uncertainty.

GRAPH 4

Natural disaster uncertainty index
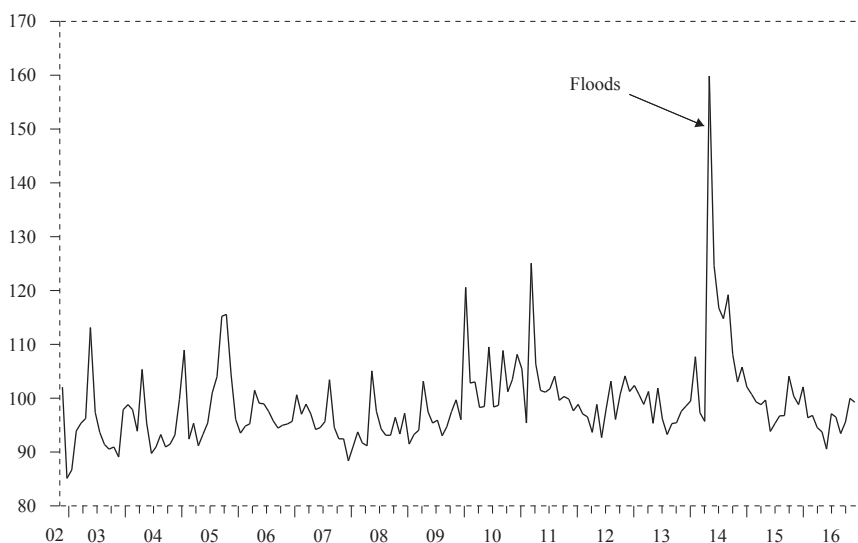
Source: Authors' calculations.

The indicator of natural disaster-related uncertainty is characterized by a remarkable low-level stability. Practically one single shock is recorded, corresponding to the floods in Eastern Slavonia in May 2014.

GRAPH 5

Fiscal uncertainty indicator
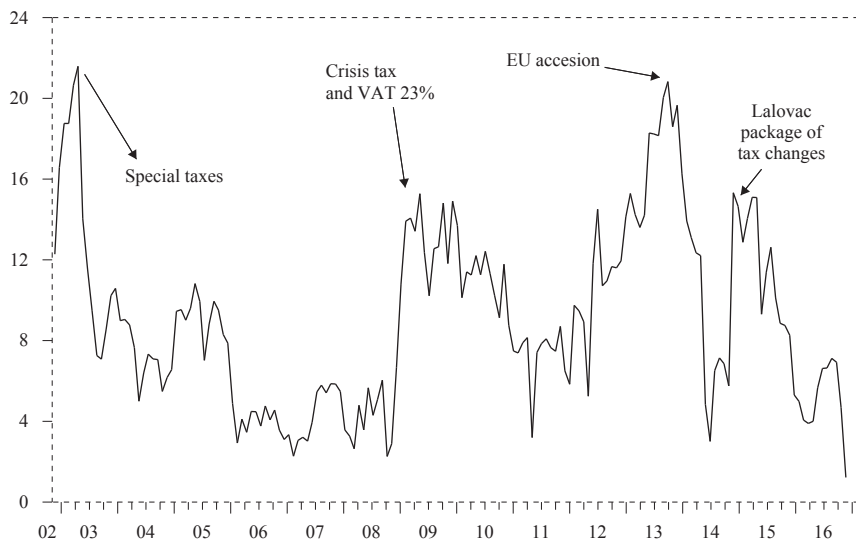
Source: Authors' calculations.

Graph 5 identifies four extreme increases of fiscal uncertainty. The first one occurred in late 2002, when a series of special tax (on tobacco products, personal vehicles, oil derivates, non-alcoholic beverages, luxury products, etc.) laws was passed. The other one refers to the introduction of the "crisis tax" and the increase of the VAT tax rate from 22% to 23%. The third and the most intensive uncertainty shock happened in 2013 when, apart from the harmonisation of domestic fiscal legislation with EU regulations, a series of VAT, income tax and profit tax changes was also passed. The fourth episode corresponds to the package of law changes of the then finance minister Lalovac in January 2015 (law and rulebook changes related to VAT, income tax, lottery taxation, real estate transfer tax and consumption tax).

Besides the stated five media-based uncertainty indicators, we also analyse four measures of disagreement. The first one is calculated from the proportions of responses to BCS question 4 (European Commission, 2014):

How do you expect the general economic situation in this country to develop over the next 12 months? It will: a) get a lot better, b) get a little better, c) stay the same, d) get a little worse, e) get a lot worse, and f) don't know.

Based on the methodology of Bachmann, Elstner and Sims (2013), a disagreement indicator (*dis*_4) is calculated as a standard deviation of (positive and negative) response shares of question 4:

$$dis\_4_t = \sqrt{frac_t^+ + frac_t^- - \left(frac_t^+ - frac_t^-\right)^2},\qquad(1)$$

where $frac_t^+$ is the share of positive answers ($a + b$), and $frac_t^-$ is the share of negative answers ($d + e$).[5] In the next step, *dis*_4 is standardized, and finally scaled to have an expected value of 100 and a standard deviation of 10 (for easier interpretation).

Following the exact same methodology, we also calculate the disagreement stemming from questions on future unemployment rate (Q7, indicator *dis*_7) and the assessments of appropriate time for major purchases (Q8, indicator *dis*_8).

In line with the methodology of Girardi and Reuter (2016), we additionally consider a combination of all *forward-looking* questions from the BCS. That is a total of 18 questions (17 questions just as in the online appendix of Girardi and Reuter (2016)[6], plus question 11 from the Consumer Survey, focusing on the likelihood of saving any money in the household).[7] In calculating this aggregate disagreement measure (*dis*_agr), all 18 individual standard deviations of responses (relation 1) are standardized, then averaged and finally scaled for easier interpretation. That way, an aggregate indicator of forecasting disagreement is obtained, covering five different business sectors: manufacturing industry, construction, services, retail trade and the consumer sector. The results of all analysed disagreement measures are presented in graph 6. Respecting the fact that *dis*_agr is available only from May 2008, very similar tendencies are noticed in *dis*_4, *dis*_7, and *dis*_agr. The disagreement of economic agents is the highest during the crisis onset, followed by a continuous and long-term decline. On the other hand, *dis*_8 exhibits a somewhat longer-lasting growth after the outbreak of the crisis. This shows that the consumers did not considerably revise their uncertainty assessments as related to major purchases, even until the end of 2016.

That way we obtain a total of 10 individual uncertainty indicators. Besides them, we also utilize the following macroeconomic variables in VAR modelling: 3-month money market interest rate (*int*), real average wage index (*rwage*), HICP index (*HICP*), and the industrial production index (*ind*).[8] Considering the fact that Croatia is a small open economy, we also utilize the following control variables: STOXX 600 stock index (*STOXX* 600), oil prices (*oil*), euro area industrial

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[5] Answers "don't know" are excluded from further consideration.

[6] Girardi and Reuter (2016) obtain rather standard results in comparison to the related literature. All the observed disagreement measures have a negative and strictly short-run effect on the euro area GDP.

[7] For the exact wording of all BCS questions, see European Commission (2014).

[8] The industrial production index is analysed as a proxy for GDP to increase the data frequency and ensure an adequate sample size. The same procedure is also used in similar studies: Bloom (2009), Bachmann, Elstner and Sims (2013), Jurado, Ludvigson and Ng (2015), and Baker, Bloom and Davis (2016).

production (*ind**), and the 3-month euro area money market interest rate (*int**). All analysed variables are seasonally adjusted using the ARIMA X12 method. A description of individual variables, together with data sources and the available data spans, are given in appendix 2.

GRAPH 6
Disagreement measures
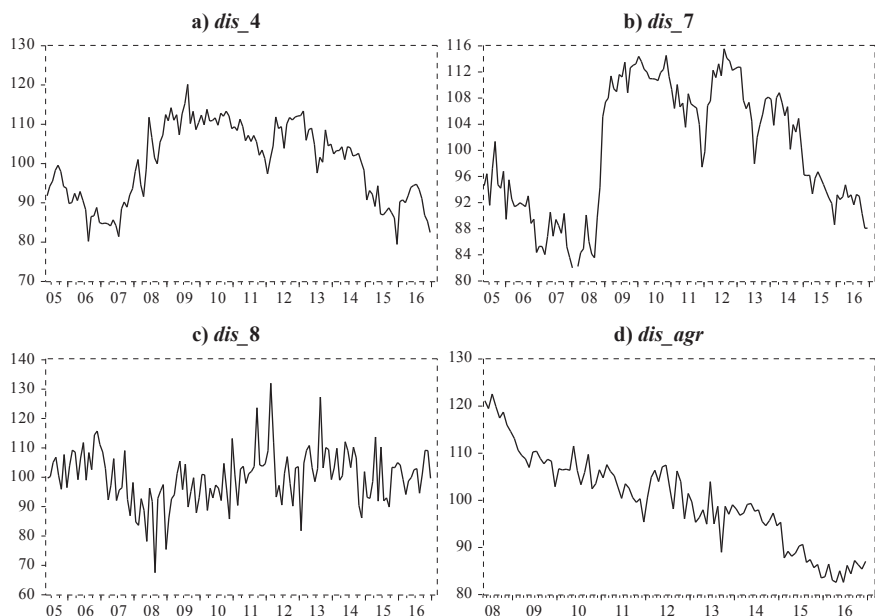
Source: Authors' calculations.

3.2 METHODOLOGICAL GROUNDS
The influence of economic uncertainty on the Croatian economy will be tested using two classes of econometric models. The first step entails a customary VAR model with time-fixed parameters, while the second step includes a more complex VAR model with time-varying parameters (Primiceri, 2015; and del Negro and Primiceri, 2015).

In the literature it is common to analyse the interrelationship between uncertainty and economic activity using VAR models (Bloom, 2009; Bachmann, Elstner and Sims, 2013; Jurado, Ludvigson and Ng, 2015; Girardi and Reuter, 2016). The same methodological principle is also the foundation of this paper. Its starting point is the estimation of a structural VAR model in the vein of Blanchard and Quah (1989), while the analysed dataset and their ordering in the innovation analysis is mostly taken from the above studies. Therefore, the starting VAR model comprises the following variables (in the exact order): *unc* (uncertainty indicator), *int*, *rwage*, *HICP*, and *ind*. Complying with the mentioned papers (Bloom, 2009; Bachmann, Elstner and Sims, 2013; Jurado, Ludvigson and Ng, 2015; Girardi and Reuter, 2016), all variables except *int* and *unc* are in logs.

We estimated ten different versions of the initial VAR model. In each of them *unc* is represented by a different uncertainty indicator: five media-based, four disagreement measures, and finally a composite uncertainty indicator (*score*), obtained by principal component analysis from the five media-based measures.[9] Its basic principle is that each uncertainty type (the media, recession, tax changes, politics, and natural disasters) affects economic activity in some (not necessarily the same) way. Here we present an effort to aggregate the stated sorts of uncertainty to a macro level, and analyse its impact on the economic activity.

The next step of the analysis is to estimate the following structural VAR model:

$$\sum_{i=0}^{p} A_i Y_t = \varepsilon_t, \tag{2}$$

where $Y_t = \begin{bmatrix} STOXX\,600_t & oil_t & ind_t^* & int_t^* & unc_t & int_t & rwage_t & HICP_t & ind_t \end{bmatrix}'$ is a vector of system variables, A_i are matrices of structural parameters, and ε_t is a vector of i.i.d. normally distributed (structural) error terms.[10]

Since previous studies of economic uncertainty (Bloom, 2009; Bachmann, Elstner and Sims, 2013; Jurado, Ludvigson and Ng, 2015; Girardi and Reuter, 2016) deal with developed western economies like the USA, the euro zone or Germany, it comes as no surprise that their VAR models are restricted solely to a set of *domestic* variables. Considering the fact that Croatia is a small open economy, we augment the initial VAR model by three separate foreign variables: *STOXX*600, *oil*, *ind**, and *int**. Taking these into account, the vector of system variables can be disaggregated into two distinct blocks:

$$Y_t = \begin{bmatrix} Y_{1,t}, & Y_{2,t} \end{bmatrix}', \tag{3}$$

where $Y_{1,t}$ is the foreign, and $Y_{2,t}$ is the domestic block of variables. To be specific, the stated blocks are given as $Y_{1,t} = \begin{bmatrix} STOXX\,600_t & oil_t & ind_t^* & int_t^* \end{bmatrix}'$ and $Y_{2,t} = \begin{bmatrix} unc_t & int_t & rwage_t & HICP_t & ind_t \end{bmatrix}'$.

Further on, the matrices of structural parameters A_i (quantifying the interrelationships among the system variables up to period *p*) can be presented as follows:

$$A_i = \begin{bmatrix} A_{11}^i & A_{12}^i \\ A_{21}^i & A_{22}^i \end{bmatrix}, \; i = 0, \ldots, p \tag{4}$$

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[9] *score* is obtained as the first principal component of the five analysed variables (the obtained eigenvalues is equal to 2.1645). The corresponding weights (loadings) are: 0.61 for *BBD*, 0.22 for *nature*, 0.32 for *PI*6, 0.28 for *polit* and 0.63 for *R_index*. It can therefore be concluded that the lion's share of aggregate uncertainty can be attributed to media reports (*BBD*) and recession (*R_index*). They are followed by tax changes and political instability, while the natural disasters have the weakest influence. The proportion of the total variance explained by the first component is 0.4329. It should be noted that Baker, Bloom and Davis (2016) apply a similar strategy of aggregating different types of uncertainty, but they apply arbitrarily chosen weights.

[10] It is important to notice that the constant term is not included in equation (2) for simplicity. However, it has been included in the empirical estimation of the model in this paper.

The assumption that external shocks influence the domestic economy (and not vice versa) effectuates in the block exogeneity restriction $A_{12}^i = 0$.

Let A_0 be the matrix of parameters quantifying contemporaneous interrelationships among system variables. Then, the reduced-form VAR model can be obtained by multiplying equation (2) by matrix A_0^{-1}:

$$Y_t = \sum_{i=1}^{p} B_i Y_{t-i} + u_t,\qquad(5)$$

where B_1, B_2, ..., B_p are parameter matrices of the reduced-form VAR model, and u_t is a vector of i.i.d. error terms. Lütkepohl (2005) shows that matrices B_i keep the block exogeneity restriction:

$$B_i = \begin{bmatrix} B_{11}^i & 0 \\ B_{21}^i & B_{22}^i \end{bmatrix},\ i = 1, ..., p\qquad(6)$$

To identify the structural shocks $\varepsilon_t = A_0 u_t$, it is necessary to introduce additional restrictions.

Considering the fact that in all $9 \cdot 8/2 = 36$ restrictions are needed for the analysed system of 9 variables, Cholesky decomposition is employed. Since A_0 is a lower triangular matrix, this enables the exact identification of the system. The variable ordering in Cholesky decomposition is given as follows: $STOXX\,600_t \quad oil_t \quad ind_t^* \quad int_t^* \quad unc_t \quad int_t \quad rwage_t \quad HICP_t \quad ind_t$.

The set of analysed variables is mostly taken from similar international studies (Bloom, 2009; Bachmann, Elstner and Sims, 2013; Jurado, Ludvigson and Ng, 2015; Girardi and Reuter, 2016). It should be noted that two of the thereby examined variables (working hours and employment level) that are utilized by Jurado, Ludvigson and Ng (2015), and Girardi and Reuter (2016) are not recorded in Croatia at all, or the series at hand is too short.

As the last relevant remark regarding the structural VAR model, it should be noted that all analysed variables are examined in levels; despite the fact the ADF test results indicate nonstationarity.[11] The first reason is that structural VAR models tolerate nonstationarity (Lütkepohl and Krätzig, 2004), and the other one is that the exact same procedure was also followed by the authors of similar studies (Bloom, 2009; Bachmann, Elstner and Sims, 2013; Jurado, Ludvigson and Ng, 2015; Girardi and Reuter, 2016).

In the next step of the analysis, we estimate a structural VAR model with time-varying parameters and a time-varying covariance matrix of residuals (Primiceri, 2005; del Negro and Primiceri, 2015). The main motivation for the utilization of this model is the questioning of the hypothesis that uncertainty has a different

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[11] The results are available upon request.

effect on economic activity in different phases of the economic/political cycles. We start from the following reduced-form VAR model:

$$Y_t = b_t + B_{1,t}Y_{t-1} + \ldots + B_{k,t}Y_{t-k} + u_t, \qquad (7)$$

where Y_t is a $n \times 1$ vector of endogenous variables, b_t is a vector of time-varying intercepts, $B_{i,t}$ ($i = 1, ..., k$) are the $n \times n$ matrices of time-varying parameters, and u_t is a $n \times 1$ vector of heteroscedastic shocks with a time-varying covariance matrix Ω_t. The stated model can be reformulated into a structural VAR specification of the following form:

$$Y_t' = X_t'B_t + A_t^{-1}\Sigma_t\varepsilon_t,$$
$$X_t' = I_n \otimes [1, Y_{t-1}', ..., Y_{t-k}'], \qquad (8)$$

where A_t is a lower-triangular matrix of the following form:

$$A_t = \begin{bmatrix} 1 & 0 & \cdots & 0 \\ \alpha_{21,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ \alpha_{n1,t} & \cdots & \alpha_{nn-1,t} & 1 \end{bmatrix}$$

It is essential to notice that matrix A_t is time-varying, implying that a shock in a certain variable has a time-varying effect on all other variables in the system. Primiceri (2005:823) emphasizes that this kind of specification has two key advantages in comparison to structural break models. The first one is that a model including *forward-looking* variables (like uncertainty in this paper) favours continuous and smooth changes of model parameters. Additionally, any kind of a learning process adopted by the consumers, firms or policymakers implies continuous (not discrete) parameter shifts.

Further on, Σ_t is a diagonal matrix of the following type:

$$\Sigma_t = \begin{bmatrix} \sigma_{1,t} & 0 & \cdots & 0 \\ 0 & \sigma_{2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \cdots & 0 & \sigma_{n,t} \end{bmatrix}$$

Time-variability of standard deviations of shocks in Σ_t implies multivariate stochastic volatility, included in the model to account for potentially heteroskedastic shocks and potential nonlinearities in the variable relationships (Primiceri, 2005:823).

Time-varying parameters from matrices A_t and B_t are modelled as random walk processes, and the standard deviations σ_t are modelled as a geometric random walk process.

$$B_t = B_{t-1} + v_t \qquad (9)$$

$$\alpha_t = \alpha_{t-1} + \zeta_t \qquad (10)$$

$$log\,\sigma_t = log\,\sigma_{t-1} + \eta_t \tag{11}$$

It is also assumed that all error terms (ε_t, v_t, ζ_t and η_t) are jointly normally distributed, and their covariance matrix has the following form:

$$\mathrm{V} = Var\left(\begin{bmatrix} \varepsilon_t \\ v_t \\ \zeta_t \\ \eta_t \end{bmatrix}\right) = \begin{bmatrix} \mathrm{I}_n & 0 & 0 & 0 \\ 0 & \mathrm{Q} & 0 & 0 \\ 0 & 0 & \mathrm{S} & 0 \\ 0 & 0 & 0 & \mathrm{W} \end{bmatrix}, \tag{12}$$

where I_n is a n-dimensional identity matrix, while Q, S and W are positive definite matrices. S is a block-diagonal matrix, with blocks corresponding to the rows of matrix A_t.

Having in mind that the system of equations (8-12) has an extremely large number of parameters, a Bayesian approach would considerably facilitate its estimation. In accordance with the pioneer of this sort of a time-varying VAR model (Primiceri, 2005), we also estimate equation (8) by OLS on the starting 40 observations (2002 M11 – 2006 M02) to calibrate the prior distributions of system parameters. The estimates of parameters from $\hat{\mathrm{B}}_{OLS}$ are obtained by estimating individual equations of the VAR model, while estimates of $\hat{\mathrm{A}}_{OLS}$ and $\hat{\sigma}_{OLS}$ are obtained through a Cholesky factorization of the covariance matrix Ω_t of residuals from the reduced VAR system ($\Omega_t = \mathrm{A}_t^{-1} \Sigma_t \Sigma_t' (\mathrm{A}_t')^{-1}$). Covariance matrices $\mathrm{V}(\hat{\mathrm{B}}_{OLS})$ and $\mathrm{V}(\hat{\mathrm{A}}_{OLS})$ are also obtained by the standard OLS method. Further on, in accordance with Primiceri (2005), and Cogley and Sargent (2005), the prior distribution of time-varying parameters is taken to be normal, while the prior distribution of hyper-parameters from matrices Q, W, and those from the diagonal blocks of matrix S (S_1 and S_2), is the inverse Wishart (*IW*) distribution. Relations (13-19) give the formalization of prior distributions of the utilized Bayesian approach.

$$\mathrm{B}_0 \sim N(\hat{\mathrm{B}}_{OLS}, 4 \cdot \mathrm{V}(\hat{\mathrm{B}}_{OLS})), \tag{13}$$

$$\mathrm{A}_0 \sim N(\hat{\mathrm{A}}_{OLS}, 4 \cdot \mathrm{V}(\hat{\mathrm{A}}_{OLS})), \tag{14}$$

$$log\,\sigma_0 \sim N(log\,\hat{\sigma}_{OLS}, \mathrm{I}_n), \tag{15}$$

$$\mathrm{Q} \sim IW(k_Q^2 \cdot 40 \cdot \mathrm{V}(\hat{\beta}_{OLS}), 40), \tag{16}$$

$$\mathrm{W} \sim IW(k_W^2 \cdot 4 \cdot \mathrm{I}_n, 4), \tag{17}$$

$$\mathrm{S}_1 \sim IW(k_S^2 \cdot 2 \cdot V(\hat{\mathrm{A}}_{1,OLS}), 2), \tag{18}$$

$$\mathrm{S}_2 \sim IW(k_S^2 \cdot 3 \cdot V(\hat{\mathrm{A}}_{2,OLS}), 3), \tag{19}$$

where the covariance matrix from equations (13-14) is up to a scalar equal to the covariance matrix obtained by the OLS method. Similarly, the scaling matrix (first

parameter of the inverse Wishart distribution) for relations (16, 18, 19) is also up to a scalar equal to the covariance matrix obtained by OLS. The parameter values $k_Q = k_W = 0.01$, and $k_S = 0.1$ are taken from Primiceri (2005). The degrees of freedom of *IW* distributions (equations 16-19) are equal to 40, 4, 2 and 3 (respectively).

Starting from the stated prior distributions, del Negro and Primiceri (2015) introduce a *Markov Chain Monte Carlo* (MCMC) algorithm for estimating the posterior distributions of the model parameters (B^T, A^T, Σ^T and the hyperparameters from matrix V). To be concrete, we apply a specific form of the Gibbs sampler (named "algorithm 3" by del Negro and Primiceri, 2015:1343) on the remaining part of data after the initial 40 observations (2006 M03 – 2016 M12). The empirical results (shown in section 4) are obtained through 10 000 MCMC draws from the posterior distributions of the model parameters.

4 EMPIRICAL RESULTS

Prior to concrete econometric analysis of the observed time series, we first present their correlation matrix (table 1).

The issue of special interest here is the correlation between all the suggested uncertainty measures, as well as the correlation between each individual uncertainty indicator and economic activity. A glance at the last row of table 1 reveals that six out of ten suggested indicators are negatively (and mostly significantly) correlated to the domestic industrial production, as well as to real wages and CROBEX. These inferences are almost entirely in line with the stylized fact of uncertainty being a countercyclical variable (Bloom, 2014). There are only two uncertainty measures that deviate from this conclusion (*dis_4* and *dis_agr*). These two variables are because of that not shown in table 1 to save space, and are not examined in further econometric analysis.[12]

In the following steps, two (time-fixed) structural VAR models are estimated: the initial, reduced-form one (including only domestic variables) and the augmented one (comprising also the foreign variables). Each of the two models is estimated in eight different specifications, i.e. each of them includes a different uncertainty indicator.[13] The lag order in each model is chosen by the Schwarz information criterion.[14] In their final versions, all eight models are estimated with three lags. The forecasting error variance decomposition of *ind* in the augmented model (for the forecasting horizon of 24 months) is presented in table 2.

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[12] We also tried to quantify uncertainty through a GARCH estimation of conditional variance of the industrial production index (in line with e.g. Fountas et al., 2006), but that variable was negligibly positively correlated to economic activity. The same conclusion is also drawn for Google trends data (frequency of Web searches by the terms "economic crisis" and "recession") for Croatia. Therefore, all of these alternatives are, just as *dis_4* and *dis_agr*, excluded from further analysis.

[13] The results of the initial and augmented model (impulse response functions and forecasting error variance decompositions) are qualitatively very similar. To save space, only the augmented model results are shown here.

[14] In cases when the chosen lag order was not sufficient to eliminate autocorrelation from the model, additional lags were successively added to the model up to the non-rejection of the null hypothesis of a Lagrange Multiplier autocorrelation test of 12th order.

TABLE 1
Correlation matrix of the analysed variables

	BBD	polit	PI6	R_index	nature	score	dis_7	dis_8	oil	ind*	int*	STOXX600	int	rwage	HICP	ind
BBD	1															
polit	0.57	1														
PI6	0.08	-0.17	1													
R_index	0.71	0.38	0.25	1												
nature	-0.07	0.04	-0.06	0.14	1											
score	0.88	0.60	0.32	0.89	0.20	1										
dis_7	0.19	-0.20	0.62	0.48	0.30	0.06	1									
dis_8	0.05	-0.01	0.07	0.06	0.19	0.09	0.06	1								
oil	-0.30	-0.46	0.12	-0.15	0.24	-0.23	0.35	0.07	1							
ind*	0.25	0.28	-0.49	-0.03	0.02	0.05	-0.82	-0.03	0.11	1						
int*	-0.54	-0.48	-0.35	-0.52	-0.13	-0.67	-0.60	-0.23	0.25	0.13	1					
STOXX600	0.19	0.33	-0.25	-0.04	0.01	0.07	-0.63	0.19	-0.23	0.69	0.05	1				
int	-0.48	-0.35	-0.04	-0.35	-0.25	-0.50	-0.15	-0.24	-0.08	-0.52	0.64	-0.46	1			
rwage	-0.28	-0.17	-0.35	-0.42	-0.26	-0.46	-0.04	-0.30	-0.46	-0.11	0.45	-0.29	0.55	1		
HICP	0.65	0.54	0.21	0.63	0.14	0.74	0.43	0.06	0.00	0.34	-0.74	0.03	-0.78	-0.73	1	
ind	-0.37	-0.08	-0.53	-0.57	-0.20	-0.58	-0.64	-0.26	-0.26	0.17	0.69	0.11	0.52	0.71	-0.67	1

Note: the dark-, medium- and light-grey cells denote 1, 5 and 10% significance level (respectively). The uncoloured table cells denote non-significant correlation coefficients.

Source: Authors' calculations.

TABLE 2
Forecasting error variance decomposition (augmented model)

Uncertainty indicator unc	STOXX 600	oil	ind*	int*	unc	int	rwage	HICP	ind
BBD	0.25	0.16	0.06	0.04	0.01	0.01	0.09	0.02	0.36
polit	0.24	0.13	0.06	0.04	0.04	0.01	0.08	0.02	0.37
PI6	0.26	0.16	0.05	0.01	0.01	0.01	0.09	0.03	0.34
R_index	0.23	0.15	0.06	0.04	0.04	0.01	0.09	0.02	0.34
nature	0.25	0.15	0.06	0.05	0.01	0.01	0.09	0.02	0.36
score	0.25	0.16	0.06	0.04	0.01	0.01	0.09	0.02	0.36
dis_7	0.12	0.21	0.12	0.10	0.00	0.01	0.08	0.03	0.33
dis_8	0.13	0.19	0.11	0.09	0.03	0.01	0.08	0.02	0.34

Source: Authors' calculations.

It is shown that the lion's share of the variability of Croatian economic activity can individually be attributed (except the *ind* variable itself) to the *STOXX* 600 index and oil prices on the world market. This inference is in line with the results of Krznar and Kunovac (2010), which show that foreign variables explain approximately 50% of Croatian GDP variability. Further on, it is interesting that *STOXX* 600 maintains a considerably high variance share in all eight augmented models. This can also, in a wider sense, be linked to the overall importance of the financial sector for the Croatian economic activity (Dumičić and Krznar, 2013). Additionally, it seems that it is reasonable to postulate that energy prices are the main determinant of economic activity (Croatian National Bank, 2015).

Finally, we show that uncertainty does not explain a particularly large proportion of industrial production variability. *R_index* and *polit* are shown to be the "relative winners" with the share of 4%, while the other indicators have smaller or even trivial effects on economic activity. Besides the forecasting error variance decomposition, it is also important to examine the impulse response functions (IRFs) to gain an insight into the dynamic interdependence of the analysed variables. Graph 7 gives the estimated IRFs of industrial production in response to a shock in uncertainty.[15]

Depicting the 68% confidence interval on graph 6 (in line with similar studies: Bloom, 2009; Primiceri, 2005; del Negro and Primiceri, 2015; Jurado, Ludvigson and Ng, 2015; Belongia and Ireland, 2016) enables indirect conclusions about the statistical significance of IRFs (depending on the confidence interval including zero or not). Most IRFs from graph 7 are rather similar (with the exception of *polit*) and reveal a significant short-run uncertainty effect. Nevertheless, the duration of the effect is extremely short, and it fades away after only a few months. The impact of political uncertainty appears to be relevant only in the medium-term. Economic activity does not seem to respond to political turmoil in the short run.
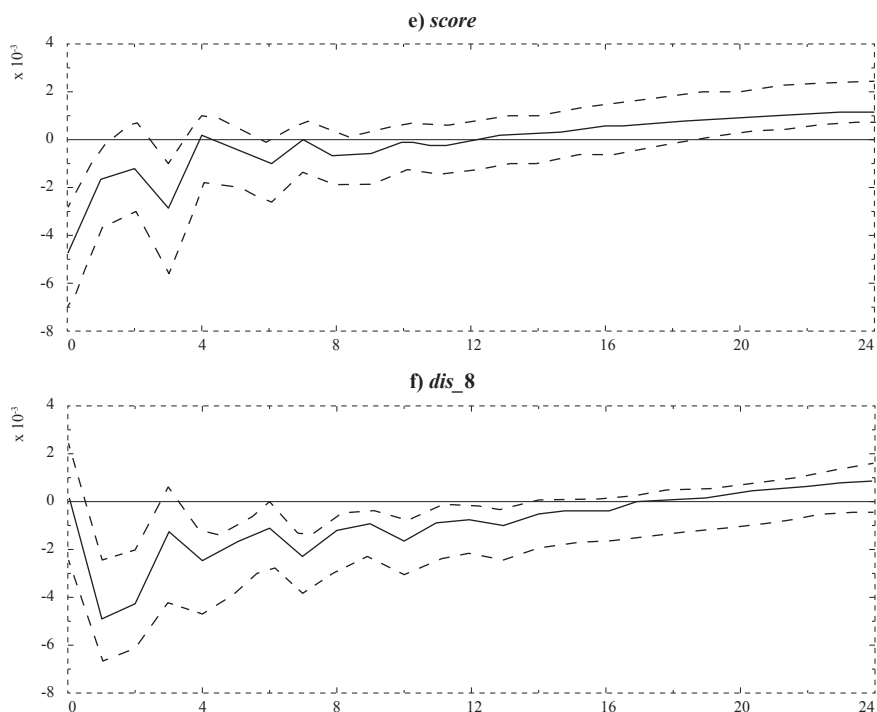
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[15] We suppress the IRFs resulting from shocks in variables *BBD* and *nature*, since these were not shown to be significantly different from zero. They are available upon request.

GRAPH 7
Industrial production IRFs (shock in economic uncertainty)

a) *BBD*

b) *polit*

c) *PI6*

d) *R_index*

e) *score*

f) *dis_8*
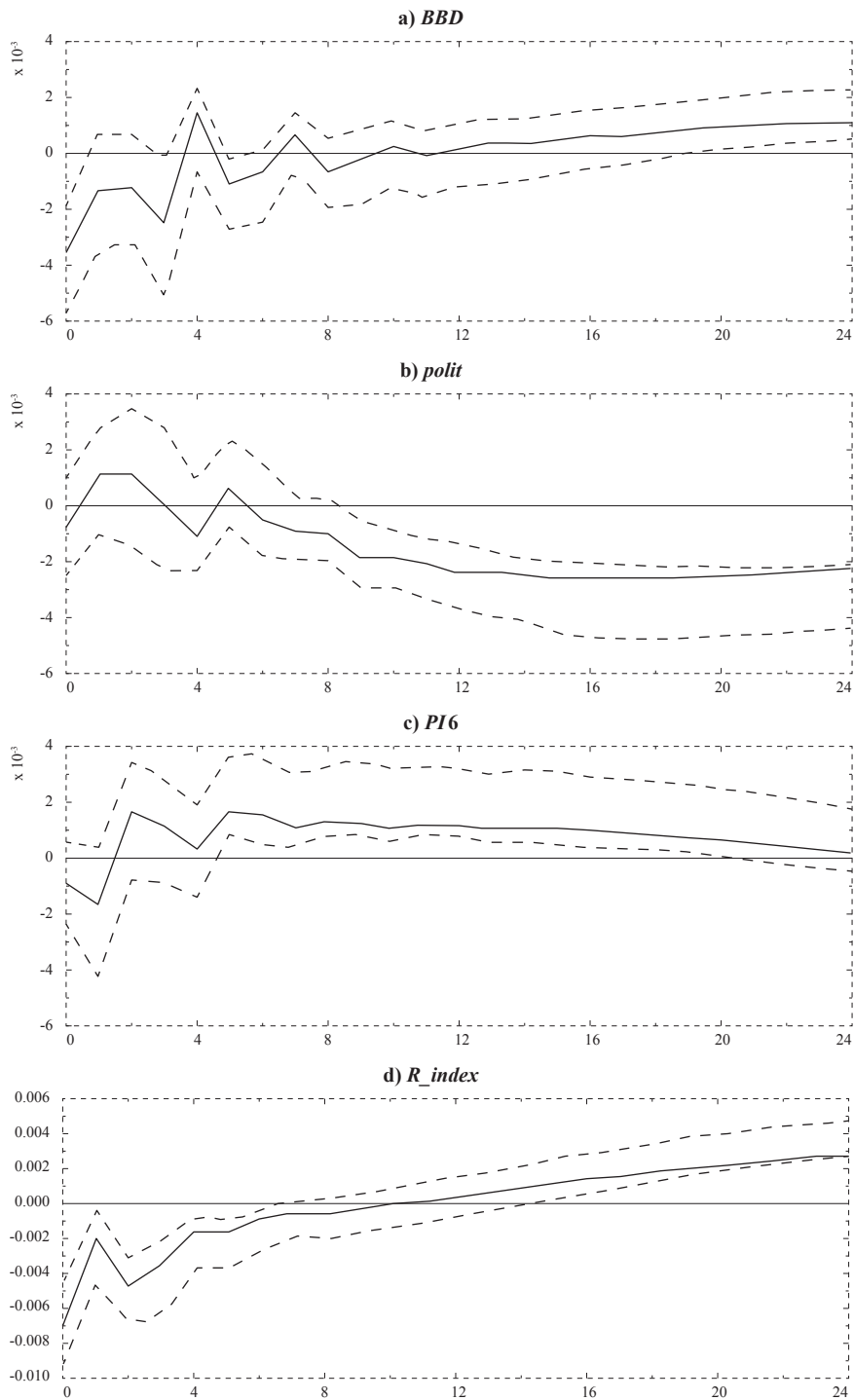
Note: Time on the x-axis is measured in months.
Source: Authors' calculations.

The estimated IRFs are of rather similar intensity. A one standard deviation shock in the variables *BBD*, *PI*6, *R_index*, *score*, and *dis_*8 has the strongest effect in the initial period or in the first month after the shock. The stated influence corresponds to a reduction of economic activity by 0.45%, 0.3%, 0.66%, 0.47%, and 0.57% (respectively). For instance, the standard deviation of residuals from the VAR equation with *PI*6 as the dependent variable is equal to 1.84. Considering that *ind* is in logs, this result implies that each additional change of fiscal legislation in a 6 month period causes a decrease of industrial production by 0.16% (conditional on all other variables in the system).

It should be noted that Bloom (2009) applies a similar analysis for the USA, detecting an abrupt short-run activity reduction as a response to an uncertainty shock. Such a downturn is followed by a recovery, ending with an overshooting effect. In the case of the analysis presented here, the IRFs are characterized by rather similar effects. As the end of the 24th month horizon comes near, the uncertainty effects even cross over to the positive domain. An obvious exception is again *polit*, whose negative effect is apparent only in the medium run.

The following empirical step is the estimation of a time-varying VAR model (del Negro and Primiceri, 2015). Having in mind that the analysed model includes a large number of parameters that are numerically challenging to estimate (equa-

tions 8-12), empirical application of this model regularly include three variables at most (Primiceri, 2005; del Negro and Primiceri, 2015; Belongia and Ireland, 2016). Since it was not possible to examine the widest possible model specification, we estimated several trivariate models. Each of them comprises a single uncertainty indicator and the industrial production index. The third, alternating position is held by one of the variables with the highest *ind* variance decomposition share in the structural VAR with time-fixed parameters. Variables *STOXX* 600, *unc* and *ind* (in this exact Cholesky ordering) are considered for the first time-varying parameters VAR. Namely, those are the variables with the highest individual share in the forecasting error variance of *ind*: 21.63%, 1.88%, and 35.00% (respectively, averaged by rows of table 2, i.e. by uncertainty indicators), summing up to a respectable 58.5%.[16]

Graph 8 shows the time-varying IRFs of industrial production to a shock in six analysed uncertainty indicators. The date points for which the IRFs are calculated are conceptually similar to those of Primiceri (2005). Namely, Primiceri performs a comparable VAR analysis for the US business cycle turning points. A similar principle is also applied here, so the targeted date points are: 2008 M06 (peak of the cycle, followed by a drastic fall of economic activity), 2010 M04 (start of the economic recovery programme by the government of Jadranka Kosor), 2014 M12 (end of the recession), and 2016 M09 (the highest value of the industrial production index in the post-crisis period). The identified date points are analysed for the indicators *BBD*, *PI*6, *R_index*, and *score*. For the *nature* variable, 2010 M04 is modified to 2014 M05 to account for the Eastern Slavonia floods at that time. For the *polit* indicator, the time points are determined by the dates of parliamentary elections (2007 M11, 2011 M12, 2015 M11, and 2016 M09), and by the middle points of electoral cycles as representatives of politically "calmer" times (2009 M11 and 2013 M11).[17] The first conclusion obtained from graph 8 is that most of the analysed uncertainty measures exhibit a time-varying effect on economic activity. This is particularly pronounced for variables *BBD*, *PI*6, *R_index*, and *score*, while the influence of political turmoil and natural disasters is relatively constant throughout the entire analysed period. The basic principle of the detected time-variability is that the uncertainty effect is the strongest at the very onset of the crisis (2008 M06), and it gradually fades away with time. It is the weakest in the last analysed time point (2016 M09), when it practically converges to zero. The only exception from the observed pattern is *PI*6. Fiscal instability has a "less negative" short-run effect at the crisis epicentre, while the intensity of the stated negative influence increases as the economy heats.

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[16] Alternatively, *STOXX* 600 was replaced by *oil* (with the average share in the forecast error variance of industrial production equal to 16.38%). The results obtained that way showed to be qualitatively very similar.
[17] For each of the stated indicators, somewhat different date points are considered, but the basic conclusions have remained the same.

GRAPH 8

Comparison of time-varying IRFs of industrial production (shock in economic uncertainty)
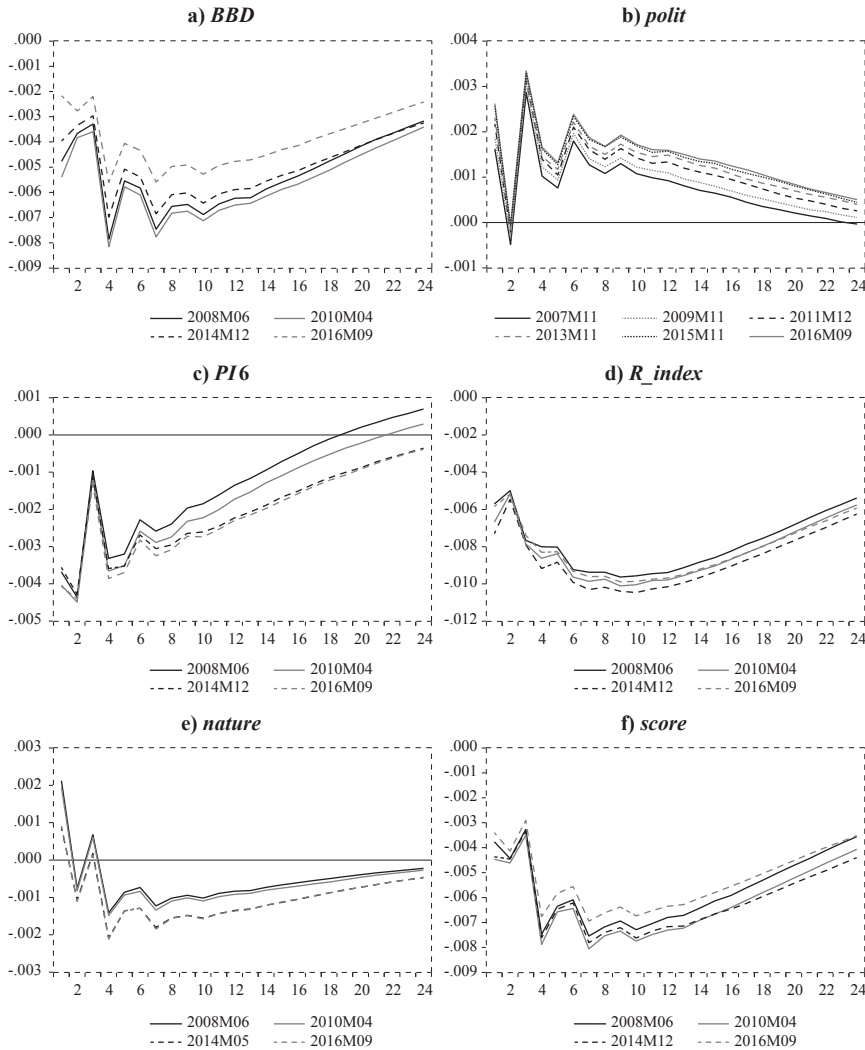
Note: Time on the x-axis is measured in months.
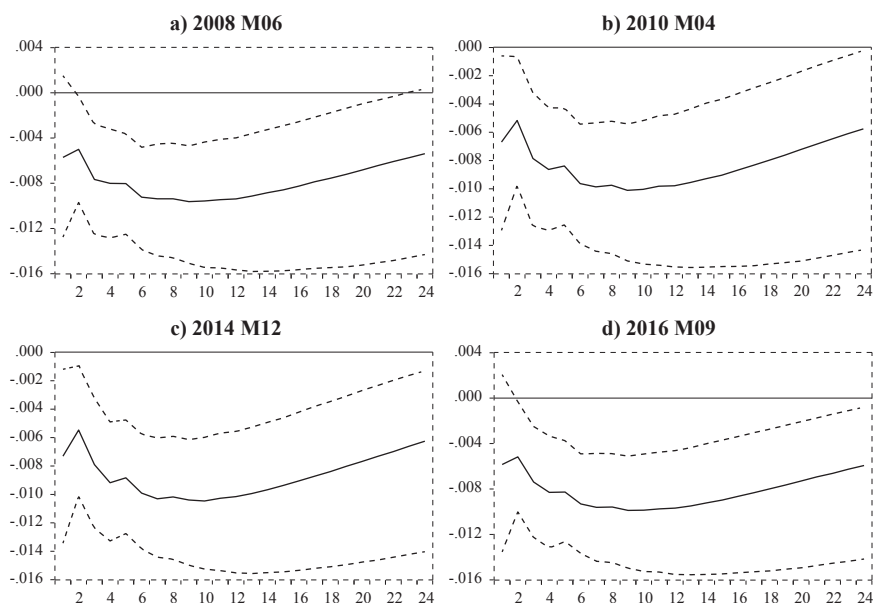Source: Authors' calculations.

In the context of policymaking, a question arises as to whether the introduction of large and frequent fiscal changes in the system is justifiable. Considering that *PI6* has the weakest effect in the crisis, this partially exculpates fiscal expansion as a counter-recession measure. It should be highlighted, of course, that the negative effect of fiscal instability noticed here is of a merely short-run nature.

Graphs 9-11 present very similar information. Except from the sole IRFs (posterior distribution means), they also contain its 16th and 84th percentile. This is in line

with similar reports (Bloom, 2009; Primiceri, 2005; del Negro and Primiceri, 2015; Jurado, Ludvigson and Ng, 2015; Belongia and Ireland, 2016). That way we obtain the middle 68% of the posterior distribution, which comes down to the interval of ±1 standard error in the normal distribution case.[18] A short overview of graphs 9-11 reveals that the effects of shocks in all three uncertainty indicators are depicted by J-curves. The function shapes are somewhat different from those in graph 7, and their maximum values (in absolute terms) are somewhat smaller. However, all the main conclusions remain intact: the short-run uncertainty effect is significant, negative, and quickly diminishing. The IRFs estimated at the very start of the crisis are quite persistent, but with the heating of the economy (as 2016 M09 becomes closer), they get closer to zero.[19]

GRAPH 9

Time-varying IRFs of industrial production (shock in R_index)

Note: Time on the x-axis is measured in months.

Source: Authors' calculations.

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[18] IRFs for *BBD*, *polit*, and *nature* are left out because they were not statistically significant.

[19] As a robustness check, we also attempted to redo the analysis with quarterly data for the Croatian economy (comprising GDP as the dependent variable and quarterly versions of other variables, obtained as averages of the corresponding monthly observations). However, due to there being too few data points and to the complexity of the assumed relationships between the observed variables, the utilized numerical methods used to estimate the model parameters were not able to converge to stable estimates.

GRAPH 10

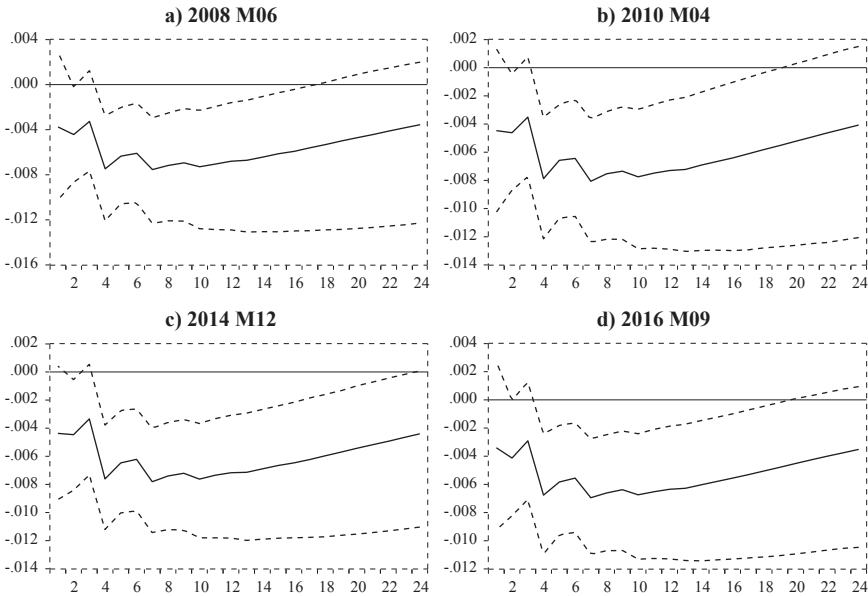*Time-varying IRFs of industrial production (shock in PI6)*
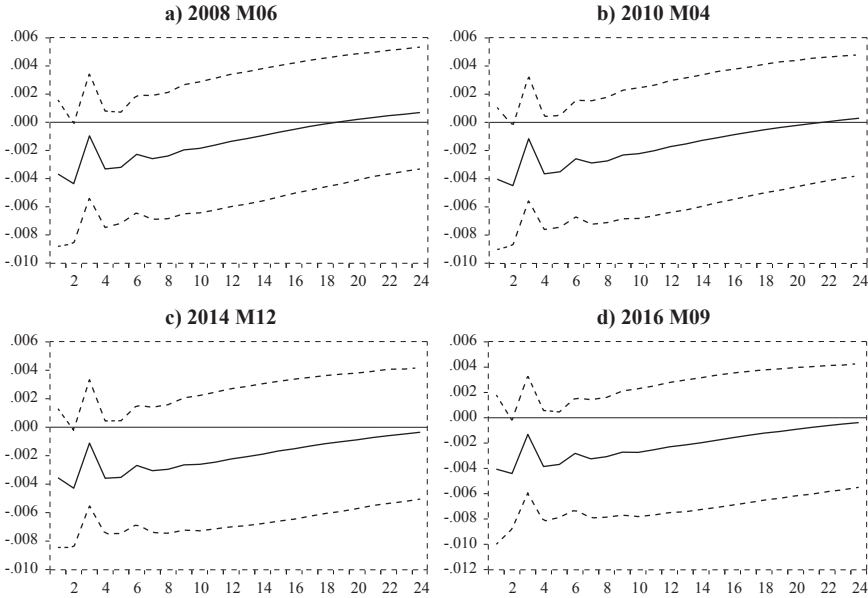
Note: Time on the x-axis is measured in months.

Source: Authors' calculations.

GRAPH 11

Time-varying IRFs of industrial production (shock in score)

Note: Time on the x-axis is measured in months.

Source: Authors' calculations.

5 DISCUSSION AND CONCLUSION

This paper offers an explanation of the intensity and persistence of the recent crisis by introducing the psychological concept of uncertainty in macroeconomic analysis. With that goal, we quantified and analysed eight different uncertainty indicators by utilizing a web database of articles from the most popular Croatian news portals, the Narodne novine repository of legal amendments, and BCS. Out of the considered uncertainty types, the following media based indicators seem to be especially relevant: the R-word index, fiscal legislation uncertainty, and the composite uncertainty indicator obtained by principal component analysis of individual uncertainty measures.

The overall conclusion is, in short, that the uncertainty effect is statistically significant, but short-lived and rather weak. The results of the VAR model with time-varying parameters are especially relevant; confirming that the uncertainty effect (quantified through the R-word index and the composite uncertainty indicator) changes throughout the phases of the business cycle. Its negative influence is the strongest at the very epicentre of the crisis in mid-2008, and it gradually disappears as the economy starts to heat again. These results can be put in relation to the long since observed phenomenon that consumer confidence significantly affects consumption expenditures only upon abrupt and unexpected downturns of economic activity (Garner, 1991). Although confidence (first moment, expected value) is a concept inherently different from uncertainty (second moment, distribution variance), sudden falls of economic activity obviously activate the same mechanisms (wait and see, risk aversion) in economic agents.

The implications of these results for policymakers primarily refer to the importance of adequate crisis communication, but also the general informing of all involved parties about the planned measures of fiscal, monetary, or any other type of economic policy. Recession for sure cannot be avoided by mere timely communication of the economic policy measures to the public (consumers, firms, banks), but the effects of a crisis can probably be held under control to a certain extent. The necessity of sound communication of the planned policy moves was recognized as a *conditio sine qua non* of efficient economic policy a long time ago (Kramer et al., 2008). On the other hand, recent domestic trends such as the then government's refusal to face the evident crisis in November 2008 (Jutarnji list, 2015) give an example of the negative effect of uncertainty due to inadequate communication.

A special comment should also be devoted to the performance of the fiscal uncertainty indicator, especially with regards to the current implementation of the tax reform in Croatia. To avoid the obviously existing negative short-run effects caused by the large number of fiscal changes, it is vital to present a clear and coherent workflow of tax reform activities in the 4-year period. In addition, without going into any normative issues of equity and wealth distribution, if lowering the tax burden is the proclaimed direction of the tax reform, then all further reform

activities should keep the same direction (independent of the current economic and political circumstances). That would diminish fiscal instability, which is recognized as a source of confusion and as an obstacle to entrepreneurship and the attraction of FDI even by tax reform working group members (Zrinušić and Vuraić Kudeljan, 2016:31). This premise is also in line with the fourth Bloom (2014:164) condition for the efficiency of the real options channel: uncertainty has a negative effect on the economy (investments, entrepreneurial activities, consumption) only when it is variable (volatile). In situations of permanent uncertainty, economic agents adapt to it. Following the same logic, they will be capable of adjusting their business plans to the coordinated legislation changes in the direction of tax disburdening.

Similar conclusions should also be drawn for the possible future implementation of a macro-reform of revoking various administrative barriers to business (Čučković and Bartlett, 2007) and the cancelling a part of the enormous number of para-fiscal charges that block the development of small and medium entrepreneurship. Impinging on the existing system of business-related legal, administrative, and fiscal regulations would surely generate a certain level of economic uncertainty, but that is by no means a reason to walk-away from reform in the segment of business facilitating. On the contrary, that segment should also be operationalised taking into account the necessity of adequate communication to the interested parties, and by insisting on keeping the long-run course of deregulating business conduction.

This paper presents one of the initial steps needed to elucidate the phenomenon of uncertainty and its influence on macro and micro tendencies. We see a sectoral analysis in the vein of Arčabić (2015) as a potentially fruitful direction for future research, where the savings and borrowing of consumers, companies, and the state would be scrutinized separately. It remains to be seen if the measures of prognostic disagreement (quantified from BCS as the central source of data on psychological concepts in economics) become significant predictors of economic activity, once the number of data points at hand gets larger.

Disclosure statement
No potential conflict of interest was reported by the authors.

APPENDIX 1

KEYWORDS USED FOR THE QUANTIFICATION OF MEDIA-BASED UNCERTAINTY INDICATORS

TABLE **A1**
BBD

Keywords	Logical conjunction I	Logical conjunction II
economy	prime minister	uncertainty
economic	minister	uncertain
	parliament	not certain
	Croatian National Bank	risk
	CNB	risky
	European Central Bank	non-reliable
	ECB	not reliable
	International Monetary Fund	
	IMF	
	European Commission	
	EC	

TABLE **A2**
R_index

Keywords	Logical conjunction
economy	recession
economic	crisis

TABLE **A3**
polit

Keywords	Logical conjunction
elections	prime minister
electoral	president
dismissal	minister
dismissed	parliament
migrant	Croatian National Bank
asylum	CNB
Schengen	European Central Bank
trial	ECB
accusation	International Monetary Fund
accused	IMF
investigation	European Commission
investigated	EC
corruption	
corruptive	
corrupt	
affair	
prison	
custody	

TABLE A4

nature

Keywords	Logical conjunction
flood	Croatia
hail	
drought	
earthquake	
epidemic	

TABLE A5

PI6

Keywords	Logical negation
tax	dismissal
	dismissed
	appointment
	appointed
	municipality
	city
	county

APPENDIX 2

TABLE A6

Dataset description

Variable	Description	Time span	Source
int	3-month money market interest rate (ZIBOR)		Eurostat
rwage	Real net wage; 2010=100; deflated by the Harmonized Index of Consumer Prices		Eurostat
HICP	2010=100; all categories of goods included		Eurostat
ind	Industrial production index; 2010=100		Eurostat
*STOXX*600	Stock index of 600 companies from 17 developed European economies	2002 M11 – 2016 M12	Thomson Reuters database
oil	Brent Europe; in USD		U.S. Energy Information Administration
ind*	Industrial production index; 2010=100		Eurostat
int*	3-month money market interest rate (EURIBOR)		Eurostat
BBD, R_index, PI6, nature and *polit*	Uncertainty indicators	2002 M11 – 2016 M12	Authors' calculation
dis_4, dis_7 and *dis_8*	Disagreement indicators	2005 M05 – 2016 M12	European Commission
dis_agr		2008 M05 – 2016 M12	European Commission

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Simulation of an application of the Hartz-IV reform in Austria

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Abstract

*This paper examines the application of the German Hartz-IV model in Austria. If the Hartz-IV reform were to be transferred to Austria, this would imply that instead of unemployment assistance (*Notstandshilfe*), the social-assistance-type minimum income benefit (*Bedarfsorientierte Mindestsicherung*) would be follow-up assistance after unemployment benefit expires. The analysis is carried out using the tax-benefit microsimulation models EUROMOD and SORESI based on the latest EU-SILC 2015 data for Austria. We simulate a baseline scenario according to the minimum income benefit regulations of the nine federal states for the year 2017 and a scenario including a proxy for an asset check of capital income. In addition, following current political discussions and developments, we simulate a ceiling scenario, in which the sum of minimum standards per household is capped at EUR 1,500 per month. The direct (monetary) effects of the potential reform are analysed on three levels: fiscal implications; number of receiving households including socio-demographic characteristics; income distribution and risk of poverty.*

Keywords: social assistance, public expenditure, household income, Austria, microsimulation

1 INTRODUCTION

The report by the commission Modern Services on the Labour Market (chaired by Peter Hartz) from 2002 contains recommendations for reforms in labour market policy and placement service in Germany. At the same time, measures were developed to increase the incentives for job search for the (long-term) unemployed. The resulting legal regulations were implemented in 2003 (Hartz I and II), 2004 (Hartz III) and 2005 (Hartz IV).

Basically, Hartz IV combines the hitherto existing unemployment assistance (*Arbeitslosenhilfe*) and social assistance for persons capable of gainful employment into unemployment benefit II (*Arbeitslosengeld II*), a new basic security provision for jobseekers. Compared to the former unemployment assistance, a change from standard of living to basic security took place: the benefit amount is now based on a lump-sum assessment of needs and is independent of income from previous employment. Unemployment benefit II (like the previous unemployment assistance) is paid once the entitlement to the insurance-based unemployment benefit (I) has expired. The latter was left more or less unchanged after the reform.

Since its introduction in Germany, a transfer of the Hartz-IV reform has been a recurring topic of public and political debate in Austria. If the reform were to be transferred, this would imply that after the expiration of unemployment benefit (*Arbeitslosengeld*), instead of unemployment assistance (*Notstandshilfe*), a minimum income benefit (*Bedarfsorientierte Mindestsicherung*) would be the follow-up benefit. As was the case in Germany, unemployment assistance represents a means-tested insurance benefit, whereas minimum income benefit is a social-assistance-type benefit.

Some representatives from Austrian institutions like the Federal Ministry of Finances or the Chamber of Commerce argue that a regulation similar to Hartz IV would reduce social expenditure and contribute to unemployment reduction by increasing incentives for taking up employment, especially for the long-term unemployed (cf. e.g. Schmidt, 2016:2), while some other representatives from institutions like the Federal Ministry for Social Affairs and the Chamber of Labour point to the negative outcomes of the reform in terms of increased risk of poverty and the expansion of the low-pay sector as has happened in Germany (AK Wien, 2016).

Given the controversial political debate, the European Centre Vienna was contracted by the Federal Austrian Ministry of Finance to analyse the direct monetary effects of a corresponding application of the German model to Austria by using tax-/benefit microsimulation models (EUROMOD and SORESI) based on the latest EU-SILC 2015 micro-data from *Statistics Austria*. The present paper describes the main findings from this analysis.

The direct (monetary) effects of the replacement of unemployment assistance by minimum income benefit are analysed on three levels:
– fiscal implications;
– number of beneficiaries/receiving households, including socio-demographic characteristics;
– income distribution and risk of poverty.

Current political discussions and developments in Austria are tending towards capping the sum of minimum standards within minimum income benefit with a maximum amount per household regardless of its size. This results in the following simulation scenarios:
– Scenario 1: basic scenario according to the minimum income benefit regulations of the nine federal states for the year 2017;
– Scenario 2: basic scenario including a proxy for an asset check of capital income, which is specified in the respective laws;
– Scenario 3: basic scenario with a cap on the sum of minimum standards at EUR 1,500 per month and household including a proxy for an asset check of capital income.

2 SYSTEMATIC FRAMEWORK
Given that the qualifying period is fulfilled, the insurance-based unemployment benefit I (*Arbeitslosengeld I*) in Germany provides a replacement rate of 60 to 67% of the previously earned income from work. The potential receiving period is three to 24 months (usually twelve months, older people 15-24 months). After expiration of unemployment benefit I, the social-assistance-type unemployment benefit II (*Arbeitslosengeld II*) is granted. It is composed of flat-rate amounts for subsistence plus housing support (where appropriate) (AK Wien, 2016; Bock-Schappelwein et al., 2014; Bräuninger, Michaelis and Sode, 2013).

In the Austrian system, unemployment benefit is also granted if the unemployed person has fulfilled the qualifying period. Again, the amount depends on the previously earned income (replacement rate of 55% with possible supplementary amounts). The potential receiving period is dependent on age and insurance history and ranges between 20 and 52 weeks (basically: 20-30 weeks; older people: 39-52 weeks; after vocational rehabilitation up to 78 weeks).

Unemployment assistance is a means-tested insurance benefit, since the income of the spouse is taken into account. The basis for the calculation is also the previous income from work (92 or 95% of unemployment benefit). For the means-test of the income of the spouse, there are amounts of exemption, which are doubled or tripled for persons that become unemployed from the age of 50. Unemployment assistance is generally granted for a period of 52 weeks but may be annually reapplied for.

As in the case of unemployment benefit II in Germany, the previous income from work is not relevant for the calculation of the tax-financed minimum income benefit. There are uniform needs-based standard rates (in 2017 usually EUR 844.46 per month for single persons and single parents including a 25% basic amount for housing needs). In contrast to unemployment assistance, for eligibility for minimum income benefit, a household's own assets that exceed a certain amount of exemption are also relevant.

When unemployment assistance is received there is the possibility of marginal earnings (2017 up to EUR 425.70 per month). For the minimum income benefit, there are certain amounts of exemptions for employment income: in most of the federal states, an amount of 15% of the monthly net income is granted for a period of 18 months after a period of at least six months of minimum income benefit claim, with a lower limit of 7% and a ceiling of 17% of the standard rate for single persons. In Lower Austria and Upper Austria an employment entry bonus of up to one third of the monthly net income has recently been introduced (AK, 2017).

Since the minimum income benefit may also be obtained as a supplement to (low) other incomes, households with unemployment assistance beneficiaries with correspondingly low previous household incomes have already been able to apply for it. In 2015 (providing the latest available figures), the average number of people receiving unemployment assistance per month amounted to 163,040 or 1.9% of the Austrian population. Among these there were, on average, 30,218 recipients (18.5%) whose household was additionally receiving minimum income benefit. The latter group would be little affected by a potential reform (AK Wien, 2016).

3 DATA AND METHODS
The analysis is carried out using the tax-/benefit microsimulation models EUROMOD and SORESI for the policy year 2017 based on the latest EU-SILC 2015 data (with incomes for 2014) for Austria with additional disaggregated

income variables provided by *Statistics Austria*. The data was uprated to 2017 with the use of empirical indicators.

EUROMOD is a tax-/benefit microsimulation model for the European Union, based on household micro-data from representative sources (mostly EU-SILC) and calculates disposable income for each household in the dataset (Sutherland and Figari, 2013). This calculation is made up of elements of income taken from the survey data (e.g. employee earnings) combined with components that are simulated by the model (taxes and benefits).

SORESI is an online tax-/benefit microsimulation tool for Austria based on EUROMOD (Fuchs and Gasior, 2014). It is further adapted to the effective outcomes of the Austrian tax-/benefit system by additional income information being taken directly from the data for policies that are difficult to simulate due to missing information related to individual social insurance histories (e.g. in the case of unemployment benefits). However, as there have been significant policy changes since 2014, for this paper, apart from minimum income benefit, we also simulate social (insurance) contributions, income tax and family allowances.

The basic output from EUROMOD and SORESI is the micro-level change in household disposable income as a result of simulated reforms (in the concrete case: policy system with unemployment assistance vs. policy system with minimum income benefit instead). The data output of EUROMOD and SORESI was analysed using the statistics programs Stata and SPSS.

The sample for the simulation of minimum income benefit contains all households in EU-SILC with at least one household member that is an unemployment assistance beneficiary, who has received the benefit for at least one month.[1] Unemployment assistance is set to zero. Subsequently, minimum income benefit is simulated instead, according to the specific regulations in the nine federal states and the composition and socio-demographic characteristics of the affected households. It is assumed that the regulations for unemployment benefit (for example, receiving period, qualifying period) remain unchanged from the EU-SILC data collection 2015.

Within the simulation of minimum income benefit both means of subsistence and potential housing needs are simulated. For the latter, the basic amount for housing needs (included in the standard rates and usually 25% thereof) and potentially received housing allowances outside minimum income benefit are also taken into account. Incomes to be included in the means-test (implemented according to the nine federal state regulations) as well as actual housing costs are also taken from

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[1] Unweighted, the persons/households with unemployment assistance relevant for the analysis are as follows:
  – households with unemployment assistance recipients: 361;
  – unemployment assistance recipients: 381;
  – members in unemployment-assistance households: 895.

the EU-SILC data and are offset with potential benefit amounts (see appendix for more details).

First, a basic scenario (S1) corresponding to the regulations for minimum income benefit in the nine federal states for the year 2017 is simulated.[2] Own assets are also relevant for minimum income benefit. As there is no corresponding information in the EU-SILC data, a proxy is applied in the basic scenario with an asset test (S2): test for capital income according to EU-SILC data, which, assuming a net interest rate of 1%, exceeds the amounts of exemption on assets (in 2017 in most federal states EUR 4,222.30 per household).[3]

Current political discussions and developments related to minimum income benefit aim to cap minimum standards with a maximum amount per household regardless of its size. The upper limit also includes any supplementary housing need. In the federal state Lower Austria the sum of the minimum standards within a household is limited to EUR 1,500 per month. Even if other income is earned and this is topped up by minimum income benefit, the latter is only granted as long as the total net household income does not exceed EUR 1,500.[4] Therefore, in Scenario 3 (S3) the sum of minimum standards (=total net household income) is capped at EUR 1,500 per month/household[5] (plus including a proxy for an asset check of capital income as in S2).

We simulate only the first-round effect (without accounting for behavioural changes) and the direct outcomes related to fiscal costs and disposable household income of the reform. Indirect consequences such as, e.g. future lower pension entitlements, are not considered. Likewise, additional costs for society (e.g. regarding health, social participation) arising from an increased risk of poverty (cf. e.g. Lamei et al., 2017) are not taken into account.

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[2] However, in all federal states without ceiling for minimum standards, in order to make the differences with respect to Scenario 3 more visible.

[3] The difference in estimated values for gross household incomes between national accounts and EU-SILC decreases from 10.8 to 3.6% if incomes from assets are not taken into account. This fact points to the under-recording of incomes from assets in EU-SILC (Statistik Austria, 2016:51). Thus, the share of households passing the means-test for minimum income benefit in terms of capital income might be over-estimated, although we assume a relatively high interest rate for capital income. However, exclusion from entitlement to minimum income benefit in the case of home ownership (not reported here) showed similar results to the test for capital income.

[4] In the federal state of Burgenland for example, also in which a ceiling was also introduced, the amount of the minimum income benefit (and not the total household income) is capped at EUR 1,500 per month. As this regulation shows only a minor additional impact compared to the basic scenario, it is not reported here.

[5] The legal explanations justify the monthly ceiling of EUR 1,500 by arguing that the benefit income should not lead to a higher household income than one that would be received from an average employment income. For a household with minimum income benefit as single income source located in Vienna, the ceiling of EUR 1,500 per month would roughly correspond to the minimum standards (incl. basic amount for housing, excl. additional housing allowance) of a couple plus one child.

4 RESULTS

4.1 STATUS QUO UNEMPLOYMENT ASSISTANCE:
EXPENDITURE AND BENEFICIARIES (-HOUSEHOLDS)

According to EU-SILC data (uprated to 2017), the following statistics are obtained for persons/households who have received unemployment assistance for at least one month in the reporting year:

- 277,000 households (7.3% of all households in EU-SILC) with 296,000 beneficiaries and 736,000 household members (3.5%/8.7% of all persons in EU-SILC);
- the average receiving duration (within the reporting year) is 6.6 months;
- expenditure for unemployment assistance amounts to 1,454 million EUR (0.42% of GDP 2016 and 1.45% of total social expenditure 2015);
- out of the 277,000 households, 56,000 households (20.2%) receive additional benefits to prevent social exclusion (minimum income benefit and/or social assistance), the expenditure amounts to 338 million EUR (0.1% of GDP 2016 and 0.34% of total social expenditure 2015);
- in total, the 277,000 households receive 1,792 million EUR of unemployment assistance and benefitsaimedat preventing social exclusion (0.51% of GDP 2016 and 1.79% of total social expenditure 2015).

Men account for 59% of unemployment assistance recipients whereas they receive 65% of expenditure, having on average higher average amounts than women. In unemployment-assistance households, 40% of the members are men, 37% are women and 23% are children (see table 1).[6]

TABLE 1

Unemployment assistance: recipients, members in households and total expenditure by women and men (and children)

Gender/ children	Unemployment assistance recipients		Members in unemployment-assistance households		Expenditures (related to unemployment assistance recipients)	
	in 1,000	in %	in 1,000	in %	in million €	in %
Women	122	41.0	270	36.6	503	34.6
Men	175	59.0	296	40.2	951	65.4
Children	–	–	171	23.2	–	–
Total	296	100.0	736	100.0	1,454	100.0

Source: Own analysis EU-SILC 2015 (uprated to 2017).

Related to household type, by far the largest share of unemployment assistance both according to number of households and according to expenditure, of about one third, is accounted for by single persons, followed by couples without children (18 to 19%; see table 2).

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[6] Children are minors who live with at least one adult person in a household. Adult "children" are counted as women or men.

TABLE 2

Unemployment assistance: receiving households and total expenditure by household type

Household type	Households		Expenditure	
	in 1,000	in %	in million €	in %
Single persons	93	33.5	518	35.6
Couple without children	49	17.6	275	18.9
Other households w/o children*	31	11.2	131	9.0
Single parents	19	6.9	94	6.5
Couple, 1-2 children	39	14.0	179	12.3
Couple, 3+ children	16	5.9	102	7.0
Further households w children**	30	10.9	155	10.6
Total	277	100.0	1,454	100.0

* Households with more than two adults; ** households with more than two adults and at least one child.

Source: Own analysis EU-SILC 2015 (uprated to 2017).

In the unemployment-assistance households, approximately three quarter of the households receive (additional) income from work in the reporting year. The same is true with respect to (other) unemployment benefits. These are followed by benefits for sickness and care (56%), family benefits (46%) and capital income (42%). Benefits to prevent social exclusion, pensions (incl. for work accidents) as well as housing allowances (outside benefits to prevent social exclusion) are drawn in about one fifth of the households each. 15% of households also receive maintenance benefits as income. In turn, maintenance payments are provided by 12% of the affected households (see table 3).

TABLE 3

Unemployment-assistance households: (additional) income types and paid maintenance

Income types/ paid maintenance	Households	
	absolute in 1,000	in % of all 277,000
Income from work	212	76.4
(Other) unemployment benefits	207	74.7
Sickness, care benefits	155	56.0
Family benefits	128	46.0
Capital income	117	42.2
Benefits to prevent social exclusion	56	20.2
Pensions (incl. for work accidents)	56	20.2
Housing benefits	53	19.1
Maintenance benefits	40	14.6
Income of children < 16 years	12	4.3
Education benefits	10	3.8
Maintenance payments	33	11.8

Source: Own analysis EU-SILC 2015 (uprated to 2017).

4.2 MINIMUM INCOME BENEFIT-SCENARIOS: EXPENDITURE AND RECIPIENTS (-HOUSEHOLDS)

The results for all households, for which we alternatively simulate minimum income benefit for at least one month instead of unemployment assistance, are presented below. Since the results for the three scenarios go in a similar direction as regards the duration of benefit receipt and socio-demographic characteristics, corresponding results are only shown for the basic scenario without asset check (S1).[7]

In the basic scenario without asset check (S1), minimum income benefit is simulated for 158,000 households (57% of original 277,000) with 348,000 members (47% of the original 736,000). The average duration of the simulated minimum income benefit receipt during the reporting year increases by 0.6 months to 7.2 months.[8] The simulated expenditure for minimum income benefit amounts to EUR 781 million (54% of the originally EUR 1,454 million for unemployment assistance).

The proxy for the asset check (S2) further reduces the number of simulated recipient households to 131,000 (47%) and the sum of the simulated expenditures to EUR 670 million (46%).

Significant additional reductions in expenditure result from the capping scenario (S3),[9] in which case reductions also affect the benefits preventing social exclusion from the data (for more details see appendix): only 93,000 households (34%) receive simulated minimum income benefit, simulated expenditure amounts to EUR 450 million (31%). Benefits preventing social exclusion from the EU-SILC data would still be received by 42,000 households (75% of the originally 56,000), with an expenditure of EUR 240 million (71% of the originally EUR 338 million). Together, the simulated minimum income benefits and/or the benefits preventing social exclusion are drawn by 108,000 former unemployment-assistance households (39%); the total expenditure amounts to EUR 690 million (39% of the original EUR 1,792 million for unemployment assistance and benefits against social exclusion) in the reporting year (see table 4).

However, in all simulated scenarios, because of the abolishment of unemployment assistance there is a reduction in income tax revenue of about EUR 30 million.[10]

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[7] Results for other scenarios are available upon request.

[8] At the level of the individual household, the number of simulated minimum income benefit months corresponds to the number of original unemployment assistance months, since minimum income benefit can only be received alternatively for these months.

[9] In 54% of those 153,000 unemployment-assistance households, which in the basic scenario with asset check (S2) receive simulated minimum income benefit and/or benefits against social exclusion from the SILC data, the average monthly total net household income is above the ceiling of EUR 1,500.

[10] Unemployment assistance represents a non-taxable income but is taken into account for the determination of the average tax rate for taxable income.

TABLE 4

Status quo (unemployment assistance) and simulated minimum income benefit-scenarios: number of households and expenditure

Scenario	Households in 1,000	Expenditure in million EUR
Status quo		
Unemployment assistance (UA)	277	1,454
Benefits against social exclusion/data	56	338
Total	277	1,792
Scenario 1: Base without asset check		
Minimum income benefit (MIB) simulated	158	781
Benefits against social exclusion/data	56	338
MIB simulated plus social exclusion/data (=MIB total)	179	1,119
Difference unemployment assistance minus MIB simulated	-119 (-43%)	-673 (-46%)
Difference UA + social exclusion/data minus MIB total	-98 (-35%)	-673 (-38%)
Scenario 2: Base with asset check capital income		
Minimum income benefit simulated	131	670
Benefits preventing social exclusion/data	56	338
MIB simulated plus social exclusion/data (=MIB total)	153	1,008
Difference unemployment assistance minus MIB simulated	-146 (-53%)	-784 (-54%)
Difference UA + social exclusion/data minus MIB total	-124 (-45%)	-784 (-44%)
Scenario 3: Ceiling 1,500 with asset check capital income		
Minimum income benefit simulated	93	450
Benefits preventing social exclusion/data	42	240
MIB simulated plus social exclusion/data (=MIB total)	108	690
Difference unemployment assistance minus MIB simulated	-184 (-66%)	-1,004 (-69%)
Difference UA + social exclusion/data minus MIB total	-169 (-61%)	-1,102 (-61%)

Source: Own analysis with EUROMOD and SORESI.

For the analysis of sociodemographic characteristics, it shows that related to the original main recipients of unemployment assistance,[11] there is a higher proportion of men among recipients and that expenditure on them is proportionally higher (plus 5.5/3.0 percentage points) for the simulated minimum income benefit entitlement.

The proportion of children has risen by 5.5 percentage points compared to unemployment assistance when looking at household members (see tables 1 and 5).

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[11] If there is more than one unemployment-assistance recipient per household, the person with the higher number of receiving months was defined as the main beneficiary.

TABLE 5

Simulated minimum income benefit (S1): recipients, household members and expenditure by women, men (and children)

Gender, children	Original unemployment assistance main recipients		Members in simulated minimum income benefit-households		Expenditure (related to original unemployment assistance main recipients)	
	in 1,000	in %	in 1,000	in %	in million €	in %
Women	56	35.5	114	32.6	247	31.6
Men	102	64.5	134	38.6	534	68.4
Children	–	–	100	28.7	–	–
Total	158	100.0	348	100.0	781	100.0

Source: Own analysis with EUROMOD and SORESI.

According to household type, simulated minimum income benefit compared to unemployment assistance for both number of households and expenditure shows a significantly higher proportion for single persons (plus 16.4/9.0 percentage points) and single parents (plus 2.7/3.7 percentage points). By trend these are household types with a smaller number of household members, in which there are no or only few possibilities of income provision by other persons than by the unemployment assistance recipient him/herself (see tables 2 and 6).

TABLE 6

Simulated minimum income benefit (S1): households (hh) and expenditure by household type

Household type	Households		Expenditure	
	in 1,000	in %	in million €	in %
Single persons	79	49.9	348	44.6
Couple without children	21	13.6	139	17.8
Other hh w/o children	2	1.4	20	2.6
Single parents	15	9.6	80	10.2
Couple, 1-2 children	21	13.1	104	13.3
Couple, 3+ children	10	6.3	43	5.5
Other hh w children	9	6.0	48	6.1
Total	158	100.0	781	100.0

Source: Own analysis with EUROMOD and SORESI.

4.3 MINIMUM INCOME BENEFIT-SCENARIOS: IMPACT ON INCOME DISTRIBUTION AND RISK OF POVERTY

Depending on the scenario, the replacement of unemployment assistance by minimum income benefit leads to a reduction in household income for 81 to 95% of all (736,000) household members in unemployment-assistance households. In the case of the 5% to 19% of household members where the household income is increased by the reform, a previous non-take-up of topping-up supplementary minimum income benefit should be the main reason.[12]

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[12] In the reform scenarios, 100% take-up of entitled minimum income benefit is assumed (see also appendix). Numerous research studies (for example, Hernanz et al., 2004; Matsaganis et al., 2013; for Austria: Fuchs, 2009) prove that this is a simplifying assumption.

The average loss of equivalised annual total household income per member in an unemployment-assistance household is 1,344 EUR in the basic scenario (S1) and increases in the capping scenario with asset check for capital income (S3) up to 2,292 EUR. Compared to the mean equivalised income of members in unemployment-assistance households of 18,264 EUR before the reform, this would imply an average income reduction of 7.4% in the basic scenario (S1) and an average income reduction of 12.5% in the capping scenario (S3).[13]

In the basic scenario (S1) the number of people at risk of poverty[14] is increased by 86,000 persons, with asset check for capital income (S2) by a further 6,000, and in the capping scenario (S3) by a further 63,000, i.e. a total increase of 155,000 persons at risk of poverty in S3. Correspondingly, the risk of poverty rate increases from 13% in the status quo to 14% in the basic scenario (S1) and to 15% in the other two scenarios (S2 and S3).

The poverty gap increases in the scenario with proxy for capital income (S2) as well as in the capping scenario (S3) from 19% to 20%, while in the basic scenario (S1) the indicator remains unchanged at 19%.

The Gini coefficient increases from 0.26 to 0.27 with the exception of the basic scenario (S1) (see table 7).

TABLE 7

Status quo (unemployment assistance) and simulated minimum income benefit-scenarios: impact on income (inc.), income distribution and risk of poverty

Scenario	hh members w reduced inc. in % all	Loss per capita-inc./year	People at risk of poverty in 1,000	At-risk-of-poverty rate in %	Poverty gap in %	Gini
Status quo	–	–	1,137	13	19	0.26
Basic without asset check (S1)	80.8	-1,344	1,223 (+86)	14	19	0.26
Basic with asset check capital income (S2)	82.2	-1,536	1,229 (+92)	15	20	0.27
Ceiling 1,500 with asset check capital income (S3)	95.3	-2,292	1,292 (+155)	15	20	0.27

Source: Own analysis with EUROMOD/SORESI.

Results by socio-demographic characteristics are again presented only for the basic scenario without asset check (S1) since the results for the other two scenarios tend to go in a similar direction.

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[13] Compared to the mean equivalised income of all persons in EU-SILC (uprated to 2017 with policies 2017) before the reform of EUR 26,892, this would imply an average income reduction of 5.0%/8.5%.

[14] 60% of median income is used as threshold for risk of poverty.

The average loss of equivalized annual household income per member in an unemployment-assistance household (all: 1,344 EUR) is the highest for men (1,630 EUR), followed by women (1,283 EUR) and children (1,008 EUR). The household types with highest losses because of the reform are single persons (1,824 EUR) and households with at least two adults without children (1,752 EUR).

Of the 86,000 persons that are additionally at risk of poverty because of the reform, 34,000 are men (40%), 27,000 women (31%) and 26,000 children (30%). Household types with the highest increase in absolute numbers are those with at least two adults without children (+19,000), households with at least two adults with three and more children as well as single persons (+18,000 each).

A reform would increase the at-risk-of-poverty rate of the total population from 13 to 14%. It would increase most for children (from 16% to 18%). For women it would increase from 13 to 14% and for men from 12 to 13%. The household types facing the highest increase in the at-risk-of-poverty rate would be single parents (from 24 to 27%) and households with at least two adults with three and more children (from 23 to 26%) followed by single persons (from 23 to 25%).

5 CONCLUDING REMARKS

If the Hartz-IV reform were to be transferred to Austria, this would imply that after expiration of unemployment benefit, instead of unemployment assistance as a means-tested insurance benefit, the social-assistance-type minimum income benefit would be the follow-up benefit. The analysis of the direct (monetary) effects of a corresponding application of the German model is carried out using the tax-/benefit microsimulation models EUROMOD and SORESI for the policy year 2017 based on the latest EU-SILC 2015 micro-data from *Statistics Austria*.

First, a basic scenario (S1) corresponding to the regulations for minimum income benefit in the nine federal states for the year 2017 is simulated without an asset check. However, for the entitlement to minimum income benefit, own assets are also relevant. As there is no relevant information in the EU-SILC data, a proxy is applied in the second scenario (S2): exclusion from entitlement to minimum income benefit where there is capital income, which, assuming a net interest rate of 1%, exceeds the stipulated amounts of exemption. Following the current political discussions and developments, a third scenario (S3) with a capping of the monthly household income of households receiving minimum income benefit at EUR 1,500 is simulated for all federal states (including the proxy for asset check).

According to the EU-SILC data there are 277,000 households (7.3% of all households in EU-SILC) receiving unemployment assistance (for at least one month in the reporting year) with 736,000 household members. The total expenditure (uprated to 2017) for unemployment assistance amounts to 1,454 million EUR (0.42% of GDP 2016 and 1.45% of total social expenditure 2015). Out of the unemployment-assistance households, 56,000 households receive in addition

benefits aimed at preventing social exclusion of EUR 338 million. In sum, both benefits amount to 1,792 million EUR (0.51% of GDP 2016 and 1.79% of total social expenditure 2015).

The basic simulation scenario for minimum income benefit (S1) with still 158,000 receiving households leads to a lower expenditure of 673 million EUR (-0.19% of GDP 2016 and -0.63% of total social expenditure 2015) in comparison to the status quo. The reduced expenditure increases by a further 111 million EUR (total -784 million EUR or -0.22% of GDP 2016 and -0.78% of total social expenditure 2015) with the use of the proxy for capital income (S2; still 131,000 receiving households). Significant additional reductions result from the capping scenario (S3). In this case the benefits aimed at preventing social exclusion from the SILC-data are also affected by cuts: compared to the basic scenario with asset check, there is additional reduced expenditure of 318 million EUR (total -1,102 million EUR or -0.32% of GDP 2016 and -1.10% of total social expenditure 2015), with 93,000 households receiving simulated minimum income benefit and 42,000 households receiving benefits preventing social exclusion from the SILC-data. However, in all scenarios, the revenue from income tax is reduced by about 30 million EUR (0.01% of GDP 2016) due to the abolishment of unemployment assistance.

Concerning the distribution of income and the risk of poverty, the replacement of unemployment assistance by minimum income benefit, depending on the scenario, leads to a reduction in household income for 81 to 95% of all household members in unemployment-assistance households. The average loss of equivalized annual total household income per household member in the basic scenario (S1) is around 1,300 EUR and increases with the capping scenario with asset check (S3) to around 2,300 EUR. Compared to the mean equivalised annual income of members in unemployment-assistance households of 18,300 EUR before the reform, this would imply an average income reduction of 7.4% in the basic scenario (S1) and an average income reduction of 12.5% in the capping scenario (S3).[15]

The number of people at risk of poverty increases in the basic scenario (S1) by 86,000 persons, with asset check (S2) by a further 6,000 and in the capping scenario with asset check (S3) by a further 63,000, i.e. a total increase of 155,000 persons at risk of poverty in S3. Correspondingly, the at-risk-of-poverty rate increases from 13% in the status quo to 14% in the basic scenario (S1) and to 15% in the other two scenarios. The Gini coefficient increases from 0.26 to 0.27 with the exception of the basic scenario (S1).

In sum, it can be concluded that the replacement of unemployment assistance by minimum income benefit on the one hand would lead to considerable reductions

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[15] Compared to the mean equivalised income of all persons in EU-SILC (uprated to 2017 with policies 2017) before the reform of EUR 26,900, this would imply an average income reduction of 5.0%/8.5%.

in social expenditure. As there is more a strategic than a concrete interest in this hypothetical reform in Austria, it is not known how the potential resulting spending reduction would be allocated by the Federal Government and/or the governments of the federal states. On the other hand the reform would cause significant changes in income distribution and increases in risk of poverty. However, this purely monetary analysis fails to estimate the potential additional social costs resulting from an increase in poverty. Such an investigation is outside the scope of EUROMOD and SORESI.

Based on experiences in Germany (cf. e.g. Bäcker, Bosch and Weinkopf, 2011; Baethge-Kinsky, Bartelheimer and Wagner, 2010; Bräuninger, Michaelis and Sode, 2013; Brussig and Knuth, 2011; Dörre et al., 2013; Fitzenberger, 2009; Hassel and Schiller, 2010; Jacobi and Kluve, 2006; Klinger, Rothe and Weber, 2013; Launov and Wälde, 2013; Schneider, 2006) it is difficult to judge the potential behavioural and labour-market-related outcomes of such a reform. Some argue that in Germany it increased the efficiency of the welfare system by shortening unemployment durations and that it contributes to the reduction of unemployment by increasing incentives for taking up employment especially for the long-term unemployed.

However, others are of the opinion that measures for labour market support, personal services and job placement (Hartz I-III) were more important for the labour market integration of clients than the concrete benefit design (Hartz IV). Furthermore it is said that the "positive" employment effects of Hartz IV result more from a deterrent than from a supportive effect. Particularly, the job-seeking activities tends to intensify in particular just before the insurance-related unemployment benefit I expires.

In addition, the change to the new benefit (unemployment benefit II) was accompanied by a primary orientation towards a quick, but often not lasting, reduction of need for assistance. Also, the limits of activation became obvious: transition into unsubsidised employment is rare due to labour market shortage, especially for unemployed persons with multiple restrictions who are difficult to place. Furthermore, unemployment benefit II turned into an instrument of a means-tested wage top-up, which subsidises low wages.

In sum, experiences with the German Hartz-IV reform are mixed but also need further research for a final evaluation.

Disclosure statement
No potential conflict of interest was reported by the authors.

APPENDIX

MINIMUM INCOME BENEFIT REGULATIONS OF THE FEDERAL STATES (OVERVIEW)
A person whose means of subsistence plus housing needs are not (sufficiently) secured by his/her own resources (income, assets) or prior social benefits (e.g. unemployment benefit, maintenance payments) is legally entitled to minimum income benefit. The right to permanent residence in Austria is another fundamental eligibility requirement.[16]

Minimum income benefit is provided by means of flat-rate cash benefits to secure subsistence costs and housing needs. The initial value for single persons and single parents is EUR 844.46 in 2017. The agreement between the federal government and the federal states stipulates that the minimum standards for additional persons are certain percentages of this base:
- 75% for adult persons living with other adult persons in a common household;
- 50% for the third and subsequent entitled adult persons;
- 18% for the first three minor children;
- 15% for the fourth child and subsequent children.

The minimum standards are basically granted twelve times a year and include a basic amount of 25% for housing needs (in 2017 EUR 211.12 for single persons and single-parent households). If the appropriate housing needs cannnot be fully covered with these basic amounts, the federal states may provide additional benefits.

Based on the agreement, all federal states have passed minimum income decrees. The implementation shows a number of federal state-specific features.

Higher minimum standards
In Upper Austria, higher minimum standards apply than those laid down in the agreement between the federal government and the federal states. The included basic amount for housing is 18% (instead of the usual 25%).

Special payments
In Vienna, persons who have reached the regular retirement age or who are classified as incapacitated also receive higher benefits by special payments. There are also special payments in Tyrol and – limited to minors – in Salzburg and Styria. In these three federal states, the special payments depend on the length of the minimum income receiving period (entitlement from three months receiving duration onwards).[17]

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[16] It is assumed that there is no reason for exclusion from minimum income benefit with regard to the right to permanent residence in case of entitlement to the insurance benefit unemployment assistance (no corresponding information is available in EU-SILC data).

[17] For these federal states, given that other entitlement conditions are provided, special payments were simulated (aliquot) for those households, which show an unemployment assistance receiving period of three or more months in SILC. However, in the SILC data only the unemployment assistance months from the reporting year are shown but not any receiving months already dating from the previous year(s).

Children's standard rates
With the exception of Carinthia, all federal states grant higher minimum standards for minor children than is provided for in the federal government/federal states agreement:
- – Burgenland for all children 19.2%;
- – Lower Austria for all children 23%;
- – Upper Austria (based on the higher initial value) for the first three children 23%, for all other 20%;
- – Salzburg for all children 21%;
- – Styria for the first four children 19% and for all other 23%;
- – Tyrol and Vorarlberg (in each case based on subsistence costs without housing costs) for all children 33% and 29%;
- – Vienna for all children 27%.

Housing need
Tyrol and Vorarlberg assume 75% of the initial value for means of subsistence and provide a more generous regulation for the housing need than the usual 25% share of basic housing need since the actual housing costs are covered to certain maximum limits. In Vienna and Styria, there is a legal right to additional benefits for housing. Salzburg also provides for additional benefits, without legal entitlement, and takes into account different regional housing costs, just like Styria. In Burgenland, in Carinthia, as well as in Lower- and Upper Austria, additional benefits for the purpose of covering housing needs (beyond the basic amount for housing) are on principle not granted. Differences in the minimum income benefit regulations also concern the extent to which general housing allowances (outside minimum income benefit) are taken into account in the means-test (only included in the means-test for housing need or also in the means-test for means of subsistence, etc.) and whether the basic amount for housing is reduced if housing costs are lower (cf. Pratscher, 2016).

Overview on differences between federal states
Differences in the detailed regulations between the federal states (S: minimum income benefit laws and decrees; Armutskonferenz, 2012; Mundt/Amann, 2015) can be discerned mainly regarding the following issues:
- – types and amount of standard rates; with/without basic amount for housing; with/without special payments;
- – potential reduction of the basic amount for housing costs where there are no or lower housing costs;
- – additional benefits for housing needs provided by the federal states;
- – general housing allowances (outside minimum income benefit) only included in the means-test for housing needs or also in the means-test for means of subsistence, etc.;
- – (non-)including of other incomes in the means-test;
- – amounts of exemptions for income from work;
- – payment obligations (e.g. maintenance payments) to be taken into account;
- – accessibility of heating cost allowance for minimum income benefit recipients.

As far as enabled by the information available in the EU-SILC data, the relevant regulations of the individual federal states were modeled for the simulation of minimum income benefit 1:1.

(ADDITIONAL) CORRECTIONS MADE IN THE SILC DATA FOR THE SIMULATION OF MINIMUM INCOME BENEFIT

In addition to corrections of income data not or only partly counting in the means-test for minimum income benefit, etc. (see above), for the unemployment assistance months the following corrections were made in the SILC-data:

– unemployment assistance: set to 0 because minimum income benefit is simulated instead;

– unemployment benefit of unemployment assistance recipients: for unemployment assistance months set to 0 since simultaneous receipt with unemployment assistance (apart from overlaps) can be excluded;

– further unemployment allowances (transitional allowance, education allowances, etc.), expense allowance (e.g. for public employment service course attendance) of unemployment assistance recipients: set to 0 since as a rule, further unemployment benefits should be drawn up before or after unemployment assistance. Expense allowances are also received in addition to unemployment assistance. However, if a corresponding expense allowance is also paid out to minimum income benefit recipients (and does not count as income in the means-test), it is a zero-sum game;

– benefits preventing social exclusion: remain basically unchanged within the simulation of minimum income benefit for unemployment assistance months and are taken into account in the means test 1:1; these benefits are not affected by the proxy for the asset check since it is assumed that in reality a basic asset check has already been completed and that it continues to exist unchanged; as regards simulation Scenario 3 (capping), a correction of the simulated minimum income benefit with the EUR 1,500 exceeding amount of the total household income is first carried out, in case the simulated minimum income benefit is reduced to 0 a corresponding correction of the benefit against social exclusion from the SILC-data is also made since the capping was newly introduced;

– employment and self-employed income of unemployment assistance recipients within the unemployment assistance receiving period (entitlement to unemployment assistance is possible up to the marginal earnings threshold): the respective amounts of exemption for earnings within the minimum income benefit regulations were applied where there is unemployment assistance receipt for at least six months;

A basic problem related to the income variables in the SILC-data is that income that is not received in all twelve calendar months is difficult to allocate to the individual calendar months. Apart from alternatively simulated minimum income benefit (on the basis of the number of unemployment assistance months) as well

as other specifications for employment and self-employment income,[18] basically all income from the data is included in the calculations as a year-twelfth.

ANALYSIS OF COSTS AND RECIPIENTS (-HOUSEHOLDS)
As far as possible, the adjusted income data for the simulation of minimum income benefit apply for one (unemployment assistance) month each. Simulated minimum income benefit (per month) is multiplied by the number of months of unemployment assistance receipt. A comparison of original unemployment assistance and alternatively simulated minimum income benefit with regard to expenditure and recipients(-households) is carried out.

The evaluations are carried out exclusively on the basis of unemployment-assistance households for the receiving months according to EU-SILC. Unweighted, the persons/households with unemployment assistance relevant for the analysis are as follows:

– households with unemployment assistance recipients: 361;
– unemployment assistance recipients: 381;
– members in unemployment-assistance households: 895.

ANALYSIS OF IMPACT ON INCOME DISTRIBUTION AND RISK OF POVERTY
Apart from unemployment assistance, simulated minimum income benefit and benefits against social exclusion (see above), all incomes are also included in their original form in the reform scenarios. That is, the further income corrections for the simulation of minimum income benefit (income not or only partly to be considered in the means-test, amounts of exemption for earnings, etc.) are reversed (the corresponding EU-SILC original data is used) as these incomes actually were accrued by the respective households.

The analyses are based on all persons in the EU-SILC-data on a yearly basis. For the calculations of the risk of poverty, the at-risk-of-poverty threshold from the original data is left unchanged (= fixed poverty line).

Unweighted, the total number of persons/households in EU-SILC 2015 is as follows:

– households: 6,045;
– persons: 13,213.

POTENTIAL NON-TAKE-UP OF MINIMUM INCOME BENEFIT
The simulation results are influenced to a certain extent by current (before the potential reform) and future (following the potential reform) non-take-up of minimum income benefit:

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[18] Estimation of the receiving period for income from work on the basis of the following variables: employment status January-December, annual amount, number of receiving months for income from work, number of receiving months for unemployment assistance in the SILC data.

– Current non-take-up of minimum income benefit: persons with unemployment assistance below the standard rates of minimum income benefit that would be entitled to a supplement: with the relatively low minimum income benefit amounts they are entitled to, it can be assumed that some unemployment-assistance households will abstain from application for supplementary minimum income benefit. However, with the abolition of unemployment assistance, the gap in income will be greater and thus the application for minimum income benefit is more likely. Since there are also households with income gains in all simulated minimum income benefit scenarios, it can be assumed that of the 277,000 unemployment-assistance households there are up to 20% of households that are entitled to supplementary minimum income benefit but do not apply for it.

– Potential future non-take-up of minimum income benefit: especially in the case of households (or persons) that are only entitled to a relatively low supplementary minimum income benefit after the abolition of the unemployment assistance (due to other household income), it can be assumed that, in reality, there will be some non-take-up of minimum income benefit.

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Poverty and Shared Prosperity 2016: Taking on Inequality

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Global trends in inequality and their economic effects have become the subject of vigorous debate since the global financial crisis. The United Nations and the World Bank, which have focused on poverty alleviation since long before the crisis, have significantly strengthened their analytical and policy work on income and wealth inequality. One result of these efforts is the World Bank's inaugural flagship report *Poverty and Shared Prosperity 2016: Taking on Inequality*, which showcases the Bank's latest thinking and policy work on these issues. It emphasises the role of inequality-reduction in ending poverty, gives examples of successful approaches to reducing inequality and fighting poverty in different countries, and offers some guidelines on the design of policies to reduce inequality.

The report does not break new ground as did, for instance, Piketty's *Capital in the 21st century,* in terms of paradigms or data, but offers a very useful summary of research, policy arguments and data sources on poverty and inequality issues. It is perhaps most useful for researchers and policymakers who are new to this field, and want to benefit from country experiences and research highlights. But it will also be welcomed by the interested public and civil society organisations, who will gain valuable knowledge on the subject due to the simple explanations, useful definitions and a good overview of the evolution of research and policy work on these issues. In sum, a reader will get a coherent picture of poverty and inequality around the world and a sense of how they might evolve under different scenarios in the future.

Poverty and Shared Prosperity 2016 consists of six chapters. The first sets the stage by explaining the goals of the World Bank and how their achievement is being assessed. It also dedicates a section to data availability, which varies greatly around the world.

The second chapter, "Global Poverty"*,* presents the latest data and trends on global and regional poverty. The extreme poor are defined as those living on a daily income below USD 1.90 (in 2011 purchasing power parity). It gives a profile of the poor: most of them are young, living in rural areas in large households with numerous children, are uneducated, and work in agriculture. Sub-Saharan Africa has over half of the world's poor.

The third chapter deals with shared prosperity, a new concept of the World Bank measured as the income growth of the bottom 40% of the population. The goal of boosting shared prosperity can be achieved with larger average income growth rates for the bottom 40%, which causes the most disadvantaged sectors to make quicker progress. The key message of the chapter is that achieving shared prosperity and at the same time reducing poverty is not easy to achieve in an environment of low and declining growth rates. A simulation of poverty rates in 2030 suggests that a significantly greater prosperity-sharing will be needed than seen so far in order to reduce the poverty rate to 3% of the world population from 10.7% in 2013.

The fourth chapter, *"Inequality"*, deals with inequality in incomes and consumption expenditure and its role in poverty reduction over long time horizons. The authors decompose changes in poverty into contributions of economic growth and those of changes in inequality. The conclusion of an extensive analysis is that *global* income inequality – i.e. the inequality among all citizens worldwide, as if the world's population lived in a single state entity – increased from the 1820s through the 1980s. Since the early 1990s, and especially since 2008, there has been an exceptional period of falling global inequality, which can be explained almost entirely by the rapid growth of middle-class incomes in China and India relative to the stagnating incomes of the middle classes in advanced economies. Regionally, Latin America and the Caribbean, and more than half the African countries have very high levels of income inequality, with Gini indices exceeding 40. It is a pity that the analysis in this chapter does not extend to wealth inequality, to which reference is not made, even though data for a number of countries, such as the World Wealth and Income Database[1] of the Paris School of Economics, are readily available and adduced in the report.

Particularly interesting for policymakers are five case studies analysed in the fifth chapter, "Reductions in Inequality: A Country Perspective". Brazil, Cambodia, Mali, Peru and Tanzania have successfully reduced poverty and inequality in recent years. The common feature of their experiences is that economic growth is the most important driver of reductions in inequality and poverty, and hence of improvements in shared prosperity. Almost half of the reduction in income inequality and poverty comes from improved employment opportunities, alignment of macroeconomic and fiscal policies with sound sector-specific interventions, and public investment. As the authors repeatedly state, there is no room for complacency as various internal (fiscal discipline, political conflicts) and external challenges (climate change) remain.

The final chapter, "Reductions in Inequality: A Policy Perspective", describes some general policy interventions that may usefully guide policymakers in their efforts to reduce inequality. These policies include early childhood development (breast-feeding and nutrition, parenting skills, preschool attendance), universal healthcare, education for all, conditional cash transfers, investment in rural roads, electrification, and redistribution through taxation. The benefits of good healthcare, early childhood development, and quality schooling – not just increased enrolment – are long-term. They consist of improved well-being of the poorest, which enhances human capital accumulation in the society and raises expected lifetime earnings of the poor, thereby lowering income gaps. Conditional cash transfers and redistribution through taxation can have a strong short-term impact on inequality, but should be well targeted, constantly monitored, evaluated and adjusted.

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[1] World Wealth and Income Database: <http://wid.world/>.

Compared with some other World Bank studies dealing with the same topic, this report is more accessible and hence intended for the general public and policy-makers rather than, for instance, the two-volume collection of academic papers *Inequality and Growth: Patterns and Policy*. It also compares favourably with the World Bank's *World Development Indicators 2017*, which is more comprehensive in terms of the data but offers much less information on inequality and poverty issues. The report has an accompanying website[2] with access to data sources and a short version of the report.

Taking on inequality is an evidence-based, easy-to-read report that fulfils expectations. While it focuses more on developing countries, not only in chapter 5 but also when describing approaches to dealing with income inequality and poverty, this is understandable in view of the scale of the challenge of poverty in regions such as Africa and Latin America. Researchers interested in these issues in higher-income countries will benefit slightly less, but should nevertheless find the report useful.

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[2] Poverty and Shared Prosperity publication website: <http://www.worldbank.org/en/publication/poverty-and-shared-prosperity>.

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