Central bank independence and inflation - new insights from a meta-regression analysis

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Abstract

The question of whether central bank independence (CBI) has a negative effect on inflation has long been analyzed in the economic literature, both theoretically and empirically. This study is among the first to conduct a quantitative analysis of empirical studies. We collected 894 CBI coefficients from 85 primary studies. Using meta-regression analysis (MRA) techniques, we analyze the effects that the differences between the primary studies with regard to scope, independence measures, estimation methods and control variables have on the CBI coefficients. We find that some characteristics of the estimation approach and the geographical and time scope of the studies significantly affect their results. In addition, contrary to previous findings, we provide evidence that the choice of the independence measure matters for CBI coefficients. We do not find any robust effects of the control variables in the primary estimations.

Keywords: Central banks, Central bank independence, Inflation, Meta-analysis, Meta-regression

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1 Introduction

The question of whether the independence of the central bank from government, especially from the executive branch, has any beneficial effects has interested economists for decades.

Because of the lack of strong and immediate democratic control that accompanies independence by its very nature, it is a reasonable and important question whether the benefits of an autonomous monetary authority are worth the perceived weakness of its democratic legitimacy. Seemingly, as illustrated by recent political discussions regarding the role and policies of the European Central Bank, the likelihood for contemporary politicians to answer 'no' to this question is a monotonically increasing function of the degree of their dislike for specific policies being pursued by the central bank.

Based on the influential theoretical framework of Barro and Gordon (1983) and subsequent work that employs and extends this framework, economists have hypothesized that a higher level of central bank independence (CBI) would ceteris paribus decrease inflation. Early studies generally confirmed the empirical validity of this hypothesis but later theoretical and empirical studies incorporated a slightly more skeptical perspective and questioned the validity of these early confirmations.

The primary argument against the early empirical findings is their lack of reasonable *control variables*. Later studies resolved this problem by including different sets of controls, depending on their specific focus.

Already in the early studies (see Cukierman, Webb and Neyapti (1992)), it was recognized that the *scope* of countries that were included in the estimations may have affected the results.

The control variables and scope considerations have evolved over time. The same is true for the *estimation techniques*. The authors of the early studies primarily employed simple cross-section or pooled time-series OLS regressions while later studies employed time-series techniques, estimated heteroscedasticity and autocorrelation robust standard errors and considered the possibility of CBI's endogeneity by employing instrumentation.

Although the general ideas behind the *CBI indicators* have remained comparably stable since the beginning, the indicators as such have evolved and become more sophisticated (a good example are Vuletin and Zhu (2011)'s enhancements of the measurement of the central bank governor turnover rate).

Similar to any field of research that has produced a vast number of empirical results, it is challenging to consolidate the results and develop conclusions from the overall body of literature. Several qualitative reviews of empirical studies regarding the effects of central bank independence have been conducted (for example, Eijffinger and de Haan (1996) or more recently, Petrevski, Bogoev and Sergi (2012)); however, to our knowledge, with the notable exception of Klomp and de Haan (2010*b*), there has not been an attempt to apply meta-regression (MRA) techniques to systematically aggregate the results of prior empirical studies. For quite some time, MRA has been used in other disciplines, such as medicine, that face the same challenge of aggregating empirical results. Recently, MRA has begun to be used more broadly in the context of economic questions (see, for example, Feld and Heckemeyer (2011)).

This study conducts a meta-regression analysis of the question whether central bank independence leads to lower inflation. We collected 894 CBI coefficients from 85 primary studies and analyze if the choice of sample scope, independence indicators, estimation methodology and control variables affect the results of these primary studies.

The paper is organized as follows: Section 2 provides an overview of the data set. Section 3 provides an in-depth discussion of the choice of our meta-regressors. After the meta-regression estimation methodology has been introduced in section 4, section 5 presents our empirical results. Section 6 concludes.

2 Study sample

We collected 894 primary coefficient estimates from 85 studies that analyze the effect of central bank independence on inflation. The studies were published as journal articles, working papers, books and book contributions. They were identified by searching the Research Papers in Economics (RePEc) database using the IDEAS interface for 'central bank independence' and 'inflation', by searching on Google and Google Scholar and by reviewing selected references from the studies that were found during the search.[1] Single-country case studies were excluded from our sample. Generally, all the coefficient estimates of the identified studies were collected. The only exceptions to this rule are estimations where inflation was regressed on a broad set of highly atomic independence indicators. The estimations considered in our analysis include up to two different independence measures (generally a combination of one de jure and one de facto measure).

The studies and the number of coefficients extracted from these studies are provided in table A1 in the appendix. Our sample ensures a broad coverage of extant literature regarding this topic and is significantly larger than the sample employed in Klomp and de Haan (2010*b*).

For all estimations, we surveyed four sets of variables that we use to explain how central bank independence affects estimates: Scope (reflects geographic and time coverage of the primary estimations), methodology (reflects differences in estimation and testing techniques), independence measures (reflect differences in how independence is measured) and controls (reflect non-independence-related variables included in the estimations). These four sets of variables are discussed in the next section.

[1]Clearly, this search strategy favors studies that *focus* on central bank independence; therefore, it is not surprising that 56 of the 85 studies include a reference to central bank independence in their title. In section 3.4, we introduce a control variable that reflects if a study is 'focused' in this sense or not.

3 Meta-regression variables

3.1 Analysis scope

Early studies recognize that the effects of central bank independence on inflation depend on the combination of the type of independence measure and the countries that are included in the sample. For example, Cukierman et al. (1992) reported that a significantly negative relationship exists between independence and inflation when using a legal measure of central bank independence and restricting the sample to industrial countries. However, for developing countries, Cukierman et al. (1992) estimate a significant effect only when using the turnover rate of central bank governors as a measure for central bank independence. For this reason, we analyzed the sample composition of the studies, which could be determined from the primary sources in many cases. For these studies we introduced a variable, *SC_OECD*, which represents the percentage of OECD countries in the overall set of sample countries. Alternatively, we rated countries according to the World Bank's income classification, which distinguishes high and upper-middle income countries (percentage in sample: *SC_HIGHINCSHARE*) and lower-middle and low income countries.[2]

To reflect changes in the importance of central bank independence vis-à-vis other factors, we also consider decade dummy variables (*TC_50* to *TC_20*), which indicate the time covered by the sample (from the 1950s to the first decade of the new century).

3.2 Methodology

Many of the early studies regarding the effects of central bank independence on inflation employ bivariate regression models (Petrevski et al. (2012)). Therefore, in addition to specific meta-regressors that account for different controls that studies consider (or do not consider, as in the case of bivariate regressions), we introduce a variable *MT_BIVARIATE* which indicates the estimation of a purely bivariate model. By introducing this variable, we also implicitly include control variables that are not specifically considered by our meta-

[2]These variables have been established based on data for January 2015. This may not correctly reflect the situation for some of the studies, particularly for studies that use time scopes that reach far back into the past, because the economic situation has significantly improved over time in some countries. These countries would actually need to be categorized into a lower income range for the time to which the sample relates.

regressors. In the presence of other factors that drive inflation, bivariate CBI coefficients are likely biased in either direction.

In regards to the study design, we analyze if time series models (dummy meta-regressor *MT_TIMESERIES*) yield different results than cross-section estimations. Time series models generally employ more observations which provides the degrees of freedom necessary to include a broader set of controls. This potentially alters effects that can be attributed to central bank independence. Additionally, a large number of observations increases the precision of the estimations, making significant effects of independence on inflation ceteris paribus more likely, if they occur. The time series character also ameliorates the problem of idiosyncratic shocks that temporarily affect the relationship between central bank independence and inflation and distort the results of cross-sectional estimations. Finally, unlike cross-section estimations, time series models make it possible to consider country-specific effects. To analyze these effects in more detail, we introduce a dummy variable *MT_FE* which indicates whether a model includes fixed effects (FE). The country-specific effects may implicitly include (if they are not included as an explicit control) drivers of de facto independence, including respect for the rule of law, the general acceptance/prevalence of corruption, the tendency of the public sector to engage in partisan behavior or the inflation preferences of society. Not including fixed effects may lead to (wrongly) attributing these effects to central bank independence. To estimate a model with fixed effects, a time-varying independence measure is required. It is not surprising that among the FE studies, governor turnover rate (TOR) measures are more frequently used (47.3 % of coefficients) than in the entire sample (26.1 %) and the converse is true for legal measures that are more difficult to update for different points in time.

Furthermore, we consider whether the dependent variable, inflation, is averaged over longer periods of time (dummy meta-regressor *MT_AVG*). Averaging reduces the risk of developing inaccurate conclusions (particularly in cross-section analyses) due to purely random effects on inflation caused by factors that are not explicitly considered in the model specification.

Statistical significance is an important characteristic of estimation results and influences the odds that the results will be published in leading journals (Stanley (2005)). However, in the presence of heteroscedasticity (and/or autocorrelation), standard errors are likely to be either too large or too small, which could result in an incorrectly reported significance. Therefore, we introduce the dummy meta-regressor *MT_ROBUSTSIMPLE* to account for any mitigation measure against the effects of heteroscedasticity (and/or autocorrelation), whether the mitigation is estimating weighted least squares models or employing corrected standard errors. The assumption in determining the value of *MT_ROBUSTSIMPLE* is that estimations are only protected against the effects of heteroscedasticity (and/or autocorrelation) if it is explicitly stated in the study.[3]

When analyzing the relationship between central bank independence and inflation, the direction of causality is not always clear. In fact, causality may well run from inflation to central bank independence. Societies that have suffered from periods of extreme inflation may attempt to secure their hard-fought inflation containment by strengthening central bank independence assuming this would support their course (even if in fact, it does not). This will lead to a negative relationship between central bank independence and inflation. However, if independence-increasing reforms are pursued as a means to achieve disinflation during a high-inflation period, then econometric analyses may determine that a positive relationship exists between the variables, with causality running from inflation to independence. This is particularly likely in cross-section estimations. In addition to these concerns, there may be variables that drive both independence and (low) inflation. For example, a society's sensitivity to inflation may lead to stronger central bank independence and the appointment of monetarily conservative central bankers in the Rogoff (1985) sense. Even if central bank independence was completely irrelevant for inflation outcomes, studies may still find an effect that in reality is only attributable to the policies of conservative central bankers. To account for the endogeneity of central bank independence and omitted variables that drive both independence and inflation, we introduce a dummy meta-regressor *MT_INSTRUMENT* that is set to one if and only if central bank independence has been instrumented in the original estimations.

Certain studies (e.g., Brumm (2000), Brumm (2002) and Brumm (2006)) explicitly con-

[3]Clearly, authors may have found negative heteroscedasticity/autocorrelation test results and concluded that mitigation measures were not necessary. If they did not report this conclusion in their study, the value of *MT_ROBUSTSIMPLE* would nevertheless be zero.

sider that all indicators of central bank independence are necessarily imperfect measures for this complex, multi-faceted construct. Covariance structure analysis is applied to estimate the effect of true independence for which the available indicators are only imperfect proxies. This approach is notable because it makes it possible to concurrently consider the effect of de jure and de facto independence measures by using only one indicator. We account for these latent variable approaches by including the meta-regressor *MT_LATENT*.

Finally, by introducing the meta-regressor *MT_CBILAG* we consider whether the indicator for central bank independence enters the estimations with a time lag. It is natural to assume that changes in independence need time to affect inflation. *First*, changes in independence affect monetary policy with a time lag and *second*, the effects of monetary policy on inflation only occur after some time has passed. If central bank independence affects inflation, then time series models with time-varying and lagged central bank independence variables should confirm this effect clearer than those without lagged values.

3.3 Independence measures

Independence measures represent different aspects of independence. Personal independence (reflected in our dummy meta-regressor *IN_PERSONAL*) represents the circumstances of appointment, tenure and dismissal of leading central bank personnel. Financial independence (*IN_FINANCIAL*) reflects how well the central bank controls its own budget. This is particularly important for funding a sufficiently large and skilled staff, without which the central bank may need to rely on government expertise. Finally, instrument independence (*IN_INSTRUMENT*) reflects whether the central bank can make monetary policy decisions without *direct* interference of the government (indirect interference, e.g., through purposeful appointments of leading personnel, is of course still possible).

Legal measures of central bank independence (such as those used in Cukierman et al. (1992) and Grilli, Masciandaro, Tabellini, Malinvaud and Pagano (1991)) usually represent more than one aspect of independence. The popular governor turnover rate (e.g., Cukierman et al. (1992) and Vuletin and Zhu (2011)) and vulnerability indicators (e.g., Cukierman and Webb (1995)) focus solely on personal independence (and particularly on the independence of the central bank *governor*). Survey-based (de facto) measures such as those used in Cukierman et al. (1992) or Ilieva and Gregoriou (2005) generally represent a broader perspective of independence.

Certain indices of central bank independence (e.g., Cukierman et al. (1992)) consider whether price stability is among the central banks primary objectives and if central bank credits to the government are explicitly prohibited. However, the relationship between these indicators and central bank independence is highly indirect. These legal provisions may even prevent the central bank from freely choosing their own objectives and instruments, which might be considered a lack of goal and instrument independence. As de Haan and Kooi (1997) note, the existence of a statutory price stability target and the prohibition of lending to the government are indicators of monetary conservatism in the Rogoff (1985) sense, rather than central bank independence. To consider that numerous measures of legal central bank independence are combined measures of independence and conservatism, we include the meta-regressor *IN_CONS* in our models. This variable is based on the composition of the independence measures; however, we also employ a broader meta-regressor that reflects monetary conservatism in the central bank and society. This meta-regressor is discussed in subsection 3.4.

A common method used to determine a monetary institution's independence is to assess the degree of freedom it enjoys through laws. We refer to these measures of central bank independence as legal or *de jure* measures, for which Cukierman et al. (1992) and Grilli et al. (1991) provide some of the most prominent examples. These measures can be constructed from publicly available sources and are relatively stable over time; however, they do not holistically represent the complex construct of independence. As is well known from a comparative government perspective, there may be significant differences between the text of the law and its actual implementation. This occurs for various reasons, including historically established practices and habits (see Cukierman et al. (1992) for an impressive example regarding the Argentinian central bank), the inability to foresee and regulate every contingency, and a lack of respect for the rule of law. In addition, as Mangano (1998) demonstrates, de jure measures of independence are not as objective as they first appear to be.

Therefore, the measurement of central bank independence has been enhanced by the development of *de facto* measures that are either based on actual changes in leading central bank personnel (generally, the governor) or questionnaires that focus on the central bank's actual behavior. Beginning with simple central bank governor turnover rates, the first category has been refined by considering the coincidence of changes in the central bank governor with changes in the composition of the government, which leads to measures reflecting political vulnerability (pioneered by Cukierman and Webb (1995)). The first category has also been refined by considering the retirement of governors ahead of their terms end and their replacement with political allies of the incumbent government (Vuletin and Zhu (2011)). Examples for the second category (survey-based independence measures) include Cukierman and Webb (1995) and Ilieva and Gregoriou (2005). We reflect the type of the independence measures employed in the primary studies in our dummy meta-regressors *IN_DEJURE* and *IN_DEFACTO*.[4]

The most common de jure independence measures used in the literature are the indices that were developed by Alesina (1989), Cukierman et al. (1992), Grilli et al. (1991) and Eijffinger and Schaling (1993). These indices have been used in various studies, some of which have proposed certain modifications to the original measures. Therefore, we classify de jure independence measures with the dummy meta-regressors *IN_TYPEALESINA*, *IN_TYPECUK*, *IN_TYPGMT*, and *IN_TYPEES*, according to the underlying original measure. The variables *IN_TYPECUKCORE* and *IN_TYPEGMTCORE* indicate whether the two most popular measures that were proposed by Cukierman et al. (1992) and Grilli et al. (1991) were used *without* modifications (but likely included data updates). Our dummy meta-regressor *IN_TYPETOR* identifies all de facto independence measures based on governor turnover, *IN_TYPEVUL* all measures based on vulnerability rates.

3.4 Controls

The experience of high inflation in the past may have a positive impact on the preference for price stability in a society, effectively making it monetarily more conservative. When the majority of individuals back this type of policy, a central bank that pursues a conservative monetary policy in such an environment is less likely to come under attack by the government at any level of de jure independence and more likely to implement a stability-oriented

[4]Since latent variable approaches make it possible to combine de jure and de facto measures into one single independence indicator, there may be studies that employ *mixed* indicators.

policy. For this reason, our meta-regression includes the dummy variable *CT_LAGGED* which indicates whether the primary estimation considers any measure of lagged inflation.[5]

As a meta-regressor, we also consider political instability (dummy moderator *CT_POLSTABIR*), which is included in numerous primary estimations. There are at least two reasons to suspect that political stability influences the effect of central bank independence on inflation: *First*, in countries with unstable political regimes, legal independence is worth less than in more stable environments because of a lack of respect for the rule of law and the permanent possibility of a non-constitutional regime change. In these situations, given any level of (de jure) central bank independence, central bankers fare better by establishing a reputation for avoiding disagreement with those currently in power. *Second*, as Cukierman et al. (1992) argue, politically instable regimes may not want their successors (who may operate at the opposite end of the political spectrum, particularly in highly polarized societies) to inherit a functioning tax system and prefer revenue creation by seignorage over establishing an effective tax administration, which is more of an effort. To achieve this, the central bank must accommodate the requests for seignorage and could be more likely doing so given the prospect of being (unlawfully) replaced by more docile monetary leadership. *Third*, as Heylen and van Poeck (1995) note, governments that expect to be replaced soon are less likely to engage in building credibility, in contrast to governors that know they are part of a repeated game.

Our dummy meta-regressor *CT_OPEN* reflects the openness of the economy as a control variable in the primary estimations. Romer (1993) argues that the real depreciation that is triggered by surprise inflation is particularly harmful to more open economies and subsequently reduces the incentives of policymakers to pursue an expansive monetary policy. This leads to lower inflation in equilibrium. Alternative explanations have been offered for the negative relationship between inflation and openness (see, for example, Lane (1997)). Most argue that in some manner, openness alters the incentives for policymakers when determining monetary policy, which subsequently affects inflation.

[5]One might consider that in addition to including a relevant factor that influences the effect of CBI on inflation, it would also be important to consider lagged inflation values in the meta-regression from a purely technical perspective; including lagged inflation in the primary estimations reduces the magnitude of inflation from the equation and thereby influences the effect size of the CBI. However, as discussed in detail in Section 4, we eliminate the magnitude of inflation from the CBI coefficients by using t-values.

Another control that is included in many of the primary estimations is the fiscal situation (government expenditures, revenues, budget deficits and debt). Deteriorating public finances may tempt the government to increase pressure on the central bank to pursue an expansive monetary policy for several reasons. The government could hope to render a positive effect on output, thereby increasing tax revenues and decreasing expenditures. An expansive monetary policy could appear in the form of a direct purchase of government debt by the central bank, which would alleviate immediate (re-)financing needs. In addition, the government could seek to lower interest rates to improve (re-)financing conditions. Finally, the executive branch could urge the central bank to generate seignorage revenue, which would positively impact the governments budget deficit. For any level of central bank independence that is below 'full independence', an expansive monetary policy demanded by the government would tend to increase inflation. Therefore, we include the variable *CT_FISCSIT* in our meta-regression models, which indicates whether the primary estimations include any reference to the fiscal situation.

Since the influential work of Calmfors and Driffill (1988), the importance of wage bargaining centralization for inflation outcomes has been widely recognized. Calmfors and Driffill (1988) argue that increasing the centralization of wage bargaining results in two counteracting effects: *First*, it gives firms the ability to increase product prices in response to nominal wage increases because the elasticity of product demand is lower as more firms with substitutable products cooperate. This reduces the loss of employment from any given increase in wages, and makes increases in the price level more likely. *Second*, the greater the centralization, the stronger the effect of wage increases in one sector of the economy on the overall price level. In the extreme case of total centralization, bargainers will completely internalize this effect and there will be no incentive to increase wages above the market-clearing level; subsequently, there will be no inflation. For an intermediate level of centralization, when bargaining units are large enough to avoid being price takers (which provides no incentive to increase wages and prices) but not large enough to completely internalize the effects that their decisions have on total prices and employment, the dominant strategy for unions is to increase nominal wage claims and the dominant strategy for firms is to accede to such claims and increase product prices. This leads to the hump-shaped relationship that Calmfors and Driffill (1988) proposed for the degree of wage bargaining centralization on one hand and unemployment, real wages and inflation on the other hand. The effects of monetary policy are neglected in the original Calmfors and Driffill (1988) model; however, later studies explicitly considered central bank activity. For example, Hall and Franzese (1998) argue that in a centralized bargaining setting, the bargainers will react to signals from the central bank regarding the likely course of monetary policy. This can lead to wage restraint and moderate inflation without the central bank ever having to take action. To consider the effect of labor market institutions in the primary estimations, we include the dummy meta-regressor *CT_LABINST*.

In fixed exchange rate regimes, the monetary authority cannot pursue an independent domestic monetary policy; therefore, we include a meta-regressor *CT_FXREGIME* which reflects the inclusion of an FX regime variable in the primary estimations.

Finally, we introduce a dummy meta-regressor *CT_CONSERVAT* which indicates whether the primary estimations include one of several measures of monetary conservatism (in the Rogoff (1985) sense) of the central bank and its environment. Three potential sources of monetary conservatism are considered in our variable. The *first* source is the central bank's aversion to inflation, which at least partially results from the careers of its decision makers (Adolph (2004)). The *second* source is the government's inflation aversion, which is assumed to be stronger as the government is more politically conservative on a left-right scale. The *third* source is the outside support that a monetarily conservative central bank receives from interest groups; this can be proxied by the size of the financial sector as an important opponent to price instability. Franzese (2003) uses this approach, while Posen (1995) constructs a more complex indicator of 'effective financial opposition to inflation'. He does not consider the size of the financial sector, but rather its ability to influence policy making and implementation and the potential sources of conflict with the central bank which might make it more difficult for the financial industry to support the monetary authority.

As an alternative indicator, we combine *CT_CONSERVAT* with our dummy for conservatism that is included in the independence measures, *IN_CONS*, and create a new meta-regressor *CT_CONSERVATBROAD*.

3.5 Characteristics of the publications

In addition to the variables that represent the characteristics of the scope of the analysis, the methodology applied, the independence measures employed and the controls that are included in the primary estimations, we consider certain features of the publications themselves.

First, in alignment with Klomp and de Haan (2010*b*), we include a dummy variable *FOCUS* which indicates whether the primary study focuses on the analysis of the relationship between central bank independence and inflation. To avoid complex criteria all publications that refer to central bank independence in their title are denoted as 'focused'.

Second, we use a common practice i meta-regression analysis by including the standard error of the primary regressions (variable *STDERR*) into our meta-regression models to account for possible publication bias. This approach is discussed in more detail in the following section.

Third, we control for the year of publication (meta-regressor *PUBYEAR*). By including this variable, we consider the effects of the evolution of this type of study over time to the extent that these effects have not yet been covered by the other moderators.

4 Estimation approach

Traditional meta-analysis can be applied to identify the true or genuine effect of central bank independence on inflation by using the primary estimates of the studies that are included in our sample. $\widehat{CBI}_{ij}$ denotes the effects of central bank independence that were found in the ith estimation of study j. Depending on the assumed error structure, meta-analysis can either be performed as fixed effects or as random effects/mixed effects meta-analysis (Feld and Heckemeyer (2011)).

Fixed effects (FE) meta-analysis assumes that the effect estimates vary randomly around the genuine effect, CBI_0:

$$\widehat{CBI}_{ij} = CBI_0 + \varepsilon_{ij} \tag{1}$$

where ε_{ij} is a random error term that fulfils the assumptions $E(\varepsilon_{ij}) = 0$, $V(\varepsilon_{ij}) = \sigma_{ij}^2$. Although the ε_{ij} are likely to be heteroskedastic, estimating equation (1) does not pose a problem because estimates $\widehat{\sigma}_{ij}$ of the random error terms' standard deviations ε_{ij} are available in the primary studies. FE meta-analysis assumes that estimated coefficients only deviate from the genuine effect because of sampling error.

In contrast, *random effects/mixed effects* (RE) meta-analysis considers the specifics of the individual estimations as an additional source of heterogeneity beyond a pure sampling error. This yields the relationship:

$$\widehat{CBI}_{ij} = CBI_0 + \mu_{ij} + \varepsilon_{ij} \tag{2}$$

where μ_{ij} is an estimation-specific error term with $\mu_{ij} \sim iid(0, \sigma_\mu^2)$. After estimating σ_μ^2 in a first step equation (1) can be estimated with weighted least squares (WLS) using the inverse of $(\sigma_\mu^2 + \sigma_{ij}^2)$ as weights.

Although RE meta-analysis considers model-specific heterogeneity, it does not provide much information regarding the actual source of heterogeneity.

Meta-regression analysis (MRA) fills this gap and includes the characteristics of the primary estimations as regressors into models (1) and (2), which lead to fixed effects and random effects meta-regression models. The random effects version of MRA can be written as

$$\widehat{CBI}_{ij} = \alpha + \sum_{k=1}^{K} \beta_k Z_{ik} + \mu_{ij} + \varepsilon_{ij} \tag{3}$$

where the Z_k are the K estimation characteristics (or moderator variables) that are considered relevant and β_k represents their respective coefficients.

Several problems occur when estimating (3). *First*, as noted by Stanley and Jarrell (2005), the primary effect estimates $\widehat{CBI}_{ij}$ can be of different dimensionalities. In our sample, some studies use the percentage inflation rate π as the dependent variable, other use the transformed inflation rate $\pi/(1-\pi)$. Even if all studies had used the percentage inflation rate, it is occasionally not immediately clear if the author uses a decimal fraction (0.01 for one percent) or counts percentages in ones (1.0 for one percent). *Second*, the measures of central bank independence are proxies for a latent variable that cannot be observed. Therefore, quite a few CBI measures exist many of which differ in their magnitude. As in the case of the inflation measure, these differences in magnitude directly impact the size of the CBI coefficients. Therefore, it is necessary to remove the dimensionality of the various CBI estimates from equation (3). This is done by dividing (3) by the standard errors of the primary estimates (Stanley and Jarrell (2005)):

$$\begin{aligned} \frac{\widehat{CBI}_{ij}}{\widehat{\sigma}_{ij}} &= \frac{\alpha}{\widehat{\sigma}_{ij}} + \sum_{k=1}^{K} \frac{\beta_k Z_{ik}}{\widehat{\sigma}_{ij}} + \frac{\mu_{ij}}{\widehat{\sigma}_{ij}} + \frac{\varepsilon_{ij}}{\widehat{\sigma}_{ij}} \\ \widehat{t}_{ij} &= a_{ij} + \sum_{k=1}^{K} \beta_k z_{ik} + u_{ij} + e_{ij} \end{aligned} \tag{4}$$

The resulting model uses t-values as effect sizes. Because $e_{ij} \sim iid(1,0)$, model (4) is homoscedastic and can be estimated with ordinary least squares (OLS).

This approach, which was also used by Klomp and de Haan (2010*b*), appears plausible and arithmetically straightforward; however, it is worthwhile to reconsider the basic concept of the original equation (3). This relationship assumes that the moderators directly explain the *effect sizes* $\widehat{CBI}_{ij}$. However, the effect sizes have very different magnitudes. The size of

α and the β_k depend on the sample composition regarding the CBI and inflation measures employed in the primary estimations. Unlike numerous other meta-regression studies, the dependent variable for this study is scaled very differently across primary studies, which makes it impossible to interpret the α and β_k coefficients from equation (3) in a meaningful manner. For this reason, we estimate an alternative relationship that assumes that the moderators explain *the t-values* of the primary estimates:

$$\begin{aligned} \frac{\widehat{CBI}_{ij}}{\widehat{\sigma}_{ij}} &= \frac{\alpha}{\widehat{\sigma}_{ij}} + \sum_{k=1}^{K} \beta_k Z_{ik} + \frac{\mu_{ij}}{\widehat{\sigma}_{ij}} + \frac{\varepsilon_{ij}}{\widehat{\sigma}_{ij}} \\ \widehat{t}_{ij} &= a_{ij} + \sum_{k=1}^{K} \beta_k Z_{ik} + u_{ij} + e_{ij} \end{aligned} \tag{5}$$

This approach has two primary advantages. *First*, it completely removes the magnitude of $\widehat{CBI}_{ij}$ (which subsequently depends on the magnitude of the central bank independence indicators and inflation measures) from our meta-regression. *Second*, this calculation provides a straight-forward interpretation of the meta-regressor coefficients, specifically (in the case of a dummy meta-regressor, which reflects the inclusion of a control variable) the average ceteris paribus change in the t-value of central bank independence in response to the inclusion of the control variable that is represented by the meta-regressor.

Thus far, our estimation approaches implicitly assume that the residual heterogeneities u_{ij} of the single estimates in our sample are independent from each other. However, it is likely that this is not the case. Most of the studies in our sample contribute more than one CBI coefficient. The u_{ij} of the estimations in a study are correlated because they all share common characteristics that have not been accounted for by the moderator variables, Z_{ik}. This interdependency requires equation (5) to be estimated using hierarchical linear model estimation techniques.[6]

A major challenge for any type of meta-study, whether it is quantitative, such as a meta-analysis and meta-regression, or qualitative, such as a literature review, is the possibility of publication bias. Publication bias is the effect of a selection mechanism that is inher-

[6]To estimate these models, we use the function *rma.mv* from the R package *metafor* with a restricted maximum-likelihood (REML) estimator. A compound symmetric (CS) error structure is assumed, i.e., equal covariances for any two u_{ij} from the same study. See Viechtbauer (2010) for details.

ent in the scientific publication process. This phenomenon was first mentioned by Sterling (1959). Its importance is widely acknowledged for economics as a scientific discipline (Stanley (2005)). The concept of publication bias implies that reviewers may tend to accept studies with statistically significant results and reject others.[7] Authors anticipate this selection and may refrain from submitting (or even writing) their studies if they do not find sufficiently significant results. The danger in this is not only that well-crafted papers do not get published. Sterling (1959) notes that if a proposition is wrong and is considered as the alternative hypothesis in a standard test setup, then in the presence of publication bias, the confirming evidence (which is actually an irregularity) is published while the overwhelming evidence against the proposition is not published.

Generally, a meta-regression analysis can test for the existence of publication bias by including the standard error of the meta-regressions dependent variable as a moderator variable in the meta-regression (Stanley (2005)). In the presence of publication bias, larger standard errors should be associated with higher coefficients because these are needed to keep t-ratios sufficiently high to reach conventional levels of statistical significance (Berlin, Begg and Louis (1989)).

Unfortunately, things are not so easy in our case. Because the CBI coefficients of the primary studies have different levels of magnitude, which are driven by the choice of the CBI and inflation measures, primary studies employing CBI measures that are larger on average will ceteris paribus also produce larger standard errors.[8] Therefore, the results will provide evidence of a positive relationship between the CBI coefficients and their standard errors, even in the *absence* of publication bias.

Klomp and de Haan (2010*b*) include the standard error in equation (3), which yields a constant meta-regressor in equation (4) after dividing by the standard error. They find that this constant has significant effects and interpret their result as a strong indication of publication bias. This interpretation would be possible for a meta-regression setup with

[7]However, as de Long and Lang (1992) point out, editors may find statistically insignificant results interesting if they contradict conventional wisdom. Conversely, Stanley (2005) argues that studies that *align* with conventional wisdom are preferred by publication bias (which increases their chances to be published even if they do not provide significant results).

[8]This becomes immediately clear if one considers two otherwise identical studies where the CBI measure of the second study is a proportionally scaled version of the first study's measure.

an equally scaled dependent variable; however, it is problematic in our case.
In a similar manner, it can argued that α, the coefficient of the inverse of the standard error in equation (4), cannot be interpreted as the true effect of central bank independence.

We include a constant and the inverse of the standard error as meta-regressors in our model; however, we refrain from interpreting them as indicators for publication bias and the true effect of central bank independence, respectively.

In summary, our estimation model is described as follows:

$$\widehat{t}_{ij} = a_0 + a_{ij} + \sum_{k=1}^{K} \beta_k Z_{ik} + u_{ij} + e_{ij} \tag{6}$$

where a_0 is a constant and a_{ij} represents the inverses of the standard errors from the primary estimations.

5 Results

We begin with a descriptive analysis of the CBI coefficients in our sample. Of the 894 coefficients we collected, 77.4 % are negative, with estimations based on de facto independence measures resulting more often in negative CBI coefficients than estimations that employ de jure measures (85.0 % over 71.4 %).[9]

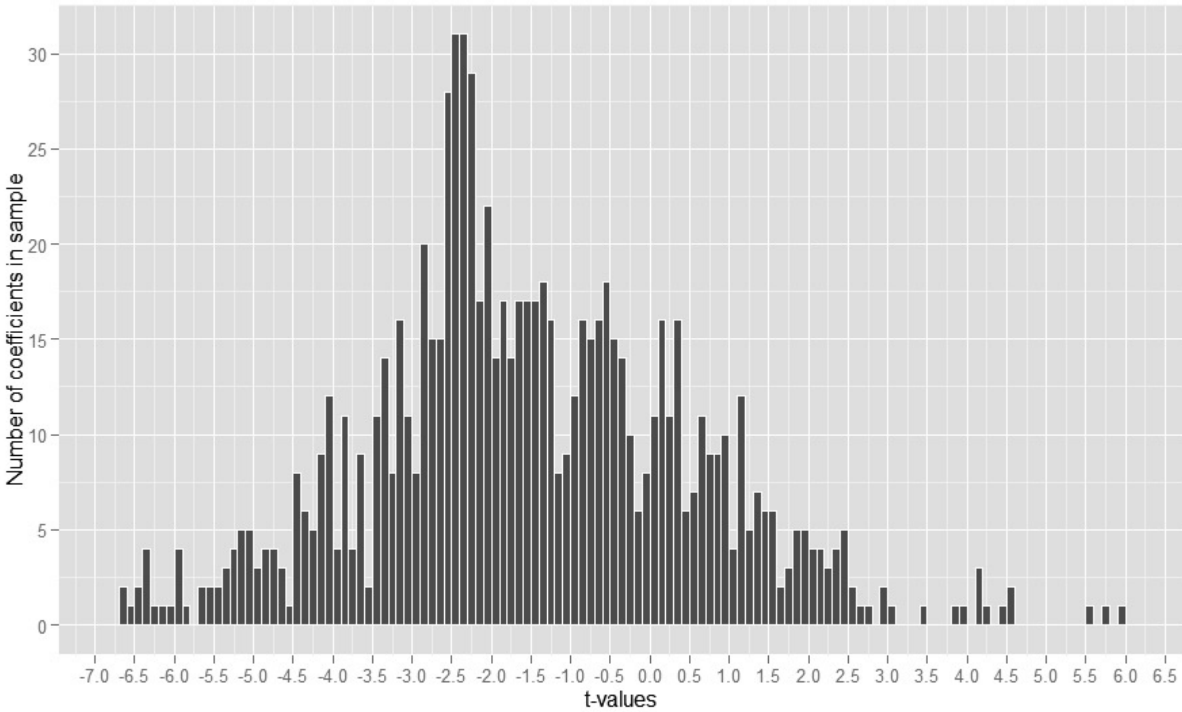

Figure 1: Histogram of t-values in the sample (without outliers)

The majority of estimations that employ de facto measures use the central bank governor turnover rate (TOR) as an independence indicator (ca. 92 %) so that *higher* coefficient values indicate *lower* CBI. Certain authors (e.g., Jenkins (1996)) switch the sign of the TOR to allow for a more intuitive interpretation of their results. However, it can be ruled out that this is a significant phenomenon that results in the large percentage of negative de facto coefficients because this would imply that many of the estimations would have indicated that a positive relationship exists between governor turnover and inflation, which

[9]Half of the studies in our sample report negative *and* positive coefficients, 41 % report only negative coefficients and 9 % report positive coefficients.

generally does not align with the verbal interpretation in the studies.

The median effect of independence on inflation in terms of the coefficients' t-values is -1.78, with de facto measures indicating a more pronounced result with a median of -2.16. Mixed coefficients, i.e., coefficients from independence measures that combine de jure and de facto indicators (such as the latent variable estimations in Brumm (2000), Brumm (2002) and Brumm (2006)), indicate a median effect of -2.38.[10]

More than half (52.0 %) of the coefficients are significantly different from zero at the 5 % level in a two-sided hypothesis test; 47.3 % are smaller or equal to -1.96 and only 4.6 % are equal to or exceed +1.96. In a one-sided test, 52.8 % of the coefficients are significantly negative at the 5 % level.

These descriptive results are reflected in the histogram that is provided in figure 1. The distribution of t-values is strongly right-skewed. Slightly more than a fifth (21.0 %) of all observations are located in the interval [-1.96, -2.50].

The results of our first meta-regression are reported in the first two columns of table 1 and indicate that bivariate estimations of the effect of independence on CBI result in significantly lower t-values of the CBI coefficients when compared to estimations that include controls. The publication year, which is negatively correlated with the estimation of bivariate models, has no significant effects.

Notably, a precise verbal interpretation of the results of our meta-regressions is important. A significantly negative meta-regressor coefficient does not necessarily imply that the author reports "a higher level of significance" (Klomp and de Haan (2010*b*, p. 608)), instead, it indicates significantly lower t-values. If t-values were generally positive and large, a significantly negative meta-regressor coefficient would imply a *lower* level of significance. The interpretation of Klomp and de Haan (2010*b*) is only valid if we restrict the estimations to the subset of negative CBI coefficients (which we routinely do with the second column for each of our meta-regressions in table 1).

[10] These mixed coefficients account for 9.5 % of all the coefficients in our sample, while the majority of coefficients consists of de jure (62.1 %) and de facto measures (28.4 %).

Table 1: Estimation results of meta-regression models

Variable	Model 1		Model 2 Focus: Independence indicators		Model 3 Focus: Independence indicators		Model 4 Focus: Independence indicators	
	Full sample	CBI coeff. < 0	Full sample	CBI coeff. < 0	Full sample	CBI coeff. < 0	Full sample	CBI coeff. < 0
Intercept	-0.783* (0.427)	-2.446*** (0.300)	-1.966*** (0.426)	-3.233*** (0.374)	-0.687*** (0.220)	-1.947*** (0.160)	-1.064*** (0.201)	-2.239*** (0.153)
Inverse of standard error	-0.001 (0.000)	-0.000 (0.000)	-0.001** (0.000)	-0.000 (0.000)	-0.001 (0.000)	-0.000 (0.000)	-0.001* (0.000)	-0.000 (0.000)
Bivariate model	-0.622** (0.258)	-0.049 (0.222)						
Publication year	-0.047 (0.033)	0.000 (0.023)						
Personal independence			0.088 (0.400)	0.319 (0.369)				
Policy independence			1.064*** (0.298)	0.950*** (0.278)				
Financial independence			-1.145*** (0.301)	-0.843*** (0.254)				
Conservatism (indep. measure)			-0.095 (0.275)	0.039 (0.258)				
Type Alesina					-2.121*** (0.468)	-1.446*** (0.384)		
Type Grilli					-1.241*** (0.293)	-0.900*** (0.240)		
Type Eijffinger-Schaling					-0.578 (0.433)	-0.020 (0.376)		
Type Latent					-2.139*** (0.608)	-1.869*** (0.500)		
Type TOR					-1.167*** (0.232)	-1.173*** (0.212)		
Type Vulnerability					-1.307** (0.587)	-0.348 (0.467)		

Variable	Model 1		Model 2 Focus: Independence indicators		Model 3 Focus: Independence indicators		Model 4 Focus: Independence indicators	
	Full sample	CBI coeff. < 0	Full sample	CBI coeff. < 0	Full sample	CBI coeff. < 0	Full sample	CBI coeff. < 0
De facto measure							-0.867*** (0.210)	-0.777*** (0.200)
Num. observations	894	692	861	661	782	692	809	692
Num. studies	85	77	84	76	73	77	79	77
Q-test	4629.884***	2179.102***	4196.015***	1916.108***	3791.009***	1982.098***	4058.237***	2177.523
Intra-class correlation	0.535	0.424	0.547	0.493	0.471	0.443	0.518	0.456

Dependent variable is t-value of CBI coefficients from primary studies. Standard errors in brackets. Meta-regressions estimated with restricted maximum likelihood estimators (REML). Asterisks indicate significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1: Estimation results of meta-regression models (continued)

Variable	Model 5 Focus: Time and country scope		Model 6 Focus: Time and country scope		Model 7 Focus: Methodology		Model 8 Focus: Methodology	
	Full sample	CBI coeff. < 0	Full sample	CBI coeff. < 0	Full sample	CBI coeff. < 0	Full sample	CBI coeff. < 0
Intercept	-2.122***	-3.007***	-0.520	-2.719	1.539***	-0.000	-1.092*	-2.052***
	(0.651)	(0.508)	(0.563)	(0.449)	(0.529)	(0.426)	(0.628)	(0.461)
Inverse of standard error	-0.001**	-0.001*	-0.001***	-0.001	-0.001	-0.000	-0.001*	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Share high-income countries	2.276***	1.637***						
	(0.483)	(0.443)						
60s	0.217	0.243	2.700	1.984				
	(0.401)	(0.325)	(0.450)	(0.440)				
70s	-0.815***	-0.812***	-0.627	-0.663				
	(0.306)	(0.265)	(0.301)	(0.261)				
80s	-0.825**	-0.745**	-0.353	-0.301				
	(0.330)	(0.303)	(0.326)	(0.302)				
90s	-0.100	0.412	-0.483	0.170				
	(0.367)	(0.312)	(0.362)	(0.314)				
00s	0.046	0.184	-0.689	0.004				
	(0.800)	(0.495)	(0.770)	(0.502)				
De jure measure			-0.689	0.004				
			(0.770)	(0.502)				
OECD countries * De jure measure			-3.268	-1.045				
			(0.563)	(0.558)				
Time series					-1.896***	-1.812***	0.107	-0.137
					(0.434)	(0.362)	(0.465)	(0.356)
Bivariate model					-0.576**	-0.092	-0.590**	0.101
					(0.254)	(0.220)	(0.283)	(0.240)
Fixed effects					0.480	0.548*	1.153***	1.233***
					(0.366)	(0.295)	(0.381)	(0.296)
Time dummies					1.059***	0.928***	0.296	0.164
					(0.380)	(0.319)	(0.397)	(0.323)

Variable	Model 5 Focus: Time and country scope Full sample	Model 5 CBI coeff. < 0	Model 6 Focus: Time and country scope Full sample	Model 6 CBI coeff. < 0	Model 7 Focus: Methodology Full sample	Model 7 CBI coeff. < 0	Model 8 Focus: Methodology Full sample	Model 8 CBI coeff. < 0
Inflation averaged					-2.490*** (0.422)	-1.984*** (0.341)	0.416 (0.546)	0.061 (0.391)
Robust standard errors					-0.726** (0.301)	-0.711*** (0.249)	-0.625** (0.297)	-0.529** (0.232)
Independence instrumented							0.320 (0.448)	0.672* (0.374)
Lagged independence							-1.177* (0.675)	-0.061 (0.489)
Number of observations							-0.001*** (0.000)	-0.001*** (0.000)
Num. observations	552	456	552	456	894	692	733	544
Num. studies	56	49	56	49	85	77	78	70
Q-test	2443.737***	1162.133***	2315.793***	1151.633***	4648.085***	2136.691***	3360.427***	1393.941***
Intra-class correlation	0.801	0.637	0.794	0.673	0.667	0.605	0.565	0.447

Dependent variable is t-value of CBI coefficients from primary studies. Standard errors in brackets. Meta-regressions estimated with restricted maximum likelihood estimators (REML). Asterisks indicate significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 1: Estimation results of meta-regression models (continued)

	Model 9	
	Focus: Controls	
Variable	Full sample	CBI coeff. < 0
Intercept	-1.698***	-2.670***
	(0.270)	(0.210)
Inverse of standard error	-0.001*	-0.001
	(0.000)	(0.000)
Lagged inflation	0.057	-0.381
	(0.344)	(0.303)
Openness	-0.335	-0.397*
	(0.237)	(0.204)
Fiscal situation	0.025	0.198
	(0.310)	(0.255)
FX regime	0.507**	0.166
	(0.256)	(0.232)
Labor market institutions	0.019	0.252
	(0.379)	(0.303)
Political (in)stability	-0.507	-0.039
	(0.432)	(0.391)
Political conservatism	0.486***	0.529***
	(0.186)	(0.175)
World inflation	0.642*	0.307
	(0.348)	(0.307)
Num. observations	861	661
Num. studies	84	76
Q-test	4314.254***	1900.760***
Intra-class correlation	0.565	0.484

Dependent variable is t-value of CBI coefficients from primary studies. Standard errors in brackets. Meta-regressions estimated with restricted maximum likelihood estimators (REML). Asterisks indicate significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 2: Results of models with robust / significant meta-regressors (including random-sample estimations)

		Model A						**Model B**					
		Full sample			**CBI coeff. < 0**			**Full sample**			**CBI coeff. < 0**		
Variable	**Expected effect direction**	Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.
Intercept		-0.765 (0.839)	6.10	6.60	-2.614*** (0.791)	0.00	98.60	0.375 (1.005)	61.80	0.00	-0.252 (0.784)	5.00	0.30
Inverse of standard error		-0.001*** (0.000)	0.00	93.20	-0.001* (0.000)	0.40	27.20	-0.001 (0.001)	0.30	23.50	-0.001* (0.000)	0.00	40.30
Bivariate model	negative	-0.598** (0.299)	99.7	45.90	-0.079 (0.271)	51.80	0.50	-0.745*** (0.286)	100.00	84.30	0.077 (0.237)	19.50	0.80
Policy independence	positive	1.530*** (0.530)	100.00	89.30	1.324*** (0.468)	100.00	89.50						
Financial independence	negative	-0.786** (0.354)	99.6	56.60	-0.565 (0.362)	97.30	27.10						
Type TOR	negative							-1.802* (0.951)	98.30	42.00	-2.060*** (0.765)	100.00	87.30
Type Vulnerability	negative							-1.577 (1.092)	96.60	12.00	-1.208 (0.865)	99.80	3.90
Type Grilli	negative							-0.907*** (0.330)	99.40	71.80	-0.548** (0.280)	98.70	46.70
De facto measure	negative	-1.796** (0.764)	99.8	63.40	-0.721 (0.757)	87.70	6.40						
Share of OECD countries	positive	4.026*** (0.467)	100.00	100.00	2.417*** (0.549)	100.00	100.00						
OECD countries * De jure measure	negative	-4.915*** (0.577)	100.00	100.00	-2.269*** (0.688)	100.00	97.00						
Share high-income countries	positive							0.052*** (0.016)	100.00	78.80	0.003 (0.014)	55.20	0.90
Share high-income countries * de jure measure	negative							-0.059*** (0.017)	100.00	95.38	0.006 (0.016)	38.42	3.71

Variable	Expected effect direction	Model A						Model B					
		Full sample			CBI coeff. < 0			Full sample			CBI coeff. < 0		
		Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.
60s	positive	-0.614 (0.420)	1.00	35.30	-0.436 (0.337)	1.70	22.10	0.053 (0.333)	46.40	1.20	0.053 (0.260)	64.70	2.80
70s	negative	-0.324 (0.339)	86.80	8.00	-0.736** (0.306)	100	71.80	-0.037 (0.267)	39.90	1.90	-0.386* (0.216)	98.30	37.80
80s	negative	-0.253 (0.308)	80.80	2.20	-0.251 (0.314)	75.70	5.80	-0.236 (0.270)	73.10	2.20	-0.291 (0.229)	84.10	8.90
Number of observations	negative	-0.002*** (0.000)	99.90	97.40	-0.002*** (0.000)	99.80	92.30	-0.001*** (0.000)	98.70	83.50	-0.001*** (0.000)	98.50	85.40
Robust standard errors	negative	-0.639* (0.353)	99.50	39.50	0.106 (0.288)	22.70	1.60	-0.602* (0.305)	99.30	30.90	-0.413* (0.236)	98.10	31.00
Time dummies	positive	0.467 (0.361)	90.70	17.70	0.296 (0.316)	72.70	6.50	0.263 (0.363)	77.90	2.50	0.270 (0.283)	84.00	6.90
Fixed effects	positive	1.194*** (0.379)	99.60	80.10	1.013*** (0.324)	99.60	82.60	1.183 (0.359)	99.90	81.70	1.203*** (0.281)	100.00	97.00
Lagged independence	negative	-0.703 (0.829)	96.20	19.60	-1.568** (0.694)	100.00	81.60	-1.385* (0.772)	100.00	45.50	-0.781 (0.587)	97.40	24.70
Political conservatism	positive	0.085 (0.420)	53.80	0.10	0.219 (0.464)	82.60	0.20	-0.142 (0.942)	45.90	0.10	-0.946 (0.756)	0.20	0.40
Political (in)stability	negative	-0.079 (1.045)	67.30	0.00	0.101 (0.712)	37.60	0.00	-0.655* (0.394)	99.90	24.30	0.024 (0.335)	49.00	0.00
Number of observations		402			312			627			459		
Number of studies		43			36			67			59		
R^2		0.310			0.295			0.1200			0.1230		
Pseudo R^2		0.282			0.250			0.1104			0.1398		
Q-test		1132.812			569.801			2290.106			929.970		
Intra-class correlation		0.789			0.455			0.588			0.447		

Dependent variable is t-value of CBI coefficients from primary studies. Standard errors in brackets. Meta-regressions estimated with restricted maximum likelihood estimators (REML).

Asterisks indicate significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
'Coeff.' is the coefficient from the full sample estimation.
'Share expected' is the percentage of runs in a 1,000-runs random-sample estimation (each based on randomly chosen two thirds of the full sample) in which the respective coefficient has the sign expected from previous models ('Expected effect direction').
'Share signif.' is the percentage of in runs a 1,000-runs random-sample estimation (each based on randomly chosen two thirds of the full sample) in which the respective coefficient is significant at the 5 % level in a one-sided test.
Pseudo R^2 is the share of residual heterogeneity that is 'explained' by the moderators in the model.

Table 3: Estimation results of general-to-specific models

Variable	Expected effect direction	Based on model A (Full sample)			Based on model A (CBI coeff. < 0)			Based on model B (Full sample)			Based on model B (CBI coeff. < 0)		
		Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.
Intercept					-2.340*** (0.618)	0.00	99.50						
Inverse of standard error		-0.001*** (0.000)	0.00	95.40	-0.001** (0.000)	0.10	59.10				-0.001** (0.000)	0.00	84.80
Bivariate model	negative	-0.598** (0.295)	99.80	48.90				-0.733*** (0.281)	100.00	85.40			
Policy independence	positive	1.136*** (0.378)	100.00	91.80	1.168*** (0.444)	100.00	85.90						
Financial independence	negative	-0.783** (0.350)	99.80	61.30									
Type TOR	negative							-1.654*** (0.252)	100.00	100.00	-2.410*** (0.245)	100.00	100.00
Type Vulnerability	negative							-1.459** (0.576)	100.00	86.80	-1.568*** (0.465)	100.00	99.50
Type Grilli	negative							-0.949*** (0.308)	100.00	91.90			
De facto measure	negative	-2.512*** (0.420)	100.00	100.00	-1.021* (0.605)	97.20	29.80						
Share of OECD countries	positive	3.830*** (0.446)	100.00	100.00	2.273*** (0.533)	100.00	100.00						
OECD countries * De jure measure	negative	-5.284*** (0.519)	100.00	100.00	-2.377*** (0.665)	100.00	99.20						
Share high-income countries	positive							0.050*** (0.016)	100.00	81.70			
Share high-income countries * de jure measure	negative							-0.059*** (0.017)	100.00	97.50			

Variable	Expected effect direction	Based on model A (Full sample)			Based on model A (CBI coeff. < 0)			Based on model B (Full sample)			Based on model B (CBI coeff. < 0)		
		Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.	Coeff.	Share expected	Share signif.
60s	positive	-0.716* (0.399)	0.10	51.90									
70s	negative				-0.820*** (0.274)	100.00	95.60				-0.394** (0.196)	99.70	43.50
Number of observations	negative	-0.002*** (0.000)	100.00	98.70	-0.002*** (0.000)	99.80	94.60	-0.001*** (0.000)	99.00	83.90	-0.001*** (0.000)	99.60	88.70
Robust standard errors	negative	-0.881*** (0.320)	100.00	83.90				-0.538** (0.257)	99.90	36.40	-0.425* (0.222)	99.50	41.50
Fixed effects	positive	1.343*** (0.354)	100.00	87.20	1.122*** (0.267)	100.00	96.10	1.149*** (0.337)	100.00	81.60	1.179*** (0.260)	100.00	98.40
Lagged independence	negative				-1.521** (0.622)	100.00	85.80	-1.316* (0.758)	100.00	44.30			
Political (in)stability	negative							-0.639* (0.382)	99.80	22.10			
Political conservatism	positive										-1.391*** (0.220)	0.00	100.00
Number of observations		402			312			627			459		
Number of studies		43			36			67			59		
R^2		0.2712			0.2603			0.1322			0.1039		
Pseudo R^2		0.3190			0.3815			0.1485			0.1843		
Q-test		1206.183			582.913			2348.392			976.911		
Intra-class correlation		0.777			0.359			0.570			0.403		

Dependent variable is t-value of CBI coefficients from primary studies. Standard errors in brackets. Meta-regressions estimated with restricted maximum likelihood estimators (REML).
Asterisks indicate significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
'Coeff.' is the coefficient from the full sample estimation.
'Share expected' is the percentage of runs in a 1,000-runs random-sample estimation (each based on randomly chosen two thirds of the full sample) in which the respective coefficient has the sign expected from previous models ('Expected effect direction').
'Share signif.' is the percentage of in runs a 1,000-runs random-sample estimation (each based on randomly chosen two thirds of the full sample) in which the respective coefficient is significant at the 10 % level in a two-sided test.

Our next model (columns 3 and 4 in table 1) analyzes the effects of the composition of the CBI measures from the primary studies. We find that CBI indicators that cover policy independence lead to higher t-values of CBI while the coverage of financial independence decreases CBI coefficients. Indicators that reflect personal independence lead to higher t-values, but not significantly so. These results, which can also be found in the subsample for CBI coefficients from de jure independence measures, show that although the literature typically refers to central bank independence as though it were a monolithic construct, it is critical to look at the different facets of independence which may not all have the same effect. However, a warning is in order: We claim that an independence measure refers to a specific facet of independence. But this implies only that the respective measure includes *some indicators* that reflect *certain aspects* of the facet of independence. Clearly, this does not mean that the indicator fully captures the concept of that facet or even significant components of the facet. It is the very nature of conceptual constructs that different approaches for measurement are possible and may be judged on plausibility grounds but per definitionem cannot be empirically validated.

After analyzing the components of independence, we consider the CBI measures that are generally employed in the literature. De jure independence measures are generally constructed as composite measures that reflect more than one aspect of independence. Conversely, most of the de facto measures (the majority of which are turnover (TOR) or vulnerability indicators) only reflect personal independence. Because Cukierman et al. (1992) with their two legal indices (same indicators, one time weighted, one time unweighted) are de facto the market leader (329 coefficients in our sample, compared to 233 TOR coefficients and 123 coefficients based on the legal index of Grilli et al. (1991)), we use the Cukierman indices as the base category. The results are provided in columns 5 and 6 of table 1. All other independence measures, including the TOR and vulnerability measures, yield lower t-values than the estimations that employ Cukiermans legal indices. The only non-significant independence indicator is ES, the measure used by Eijffinger and Schaling (1993), which is a modification of Cukierman's index. The strong effect of AL, the measure used by Alesina (1989), may have occurred because more than half of the models that employ Alesina's index are bivariate models where the negative effect of omitted variables correlated with Alesinas measure may be inaccurately attributed to central bank independence.[11] [12] For this meta-regression model, we consider primary estimations that employ the original measures as well as slight modifications (relevant for the Cukierman, Grilli and Eijffinger-Schaling measures that have been modified by other authors). However, the results do not change if we use only estimations based on the original measures.

Our next model, of which the results are provided in columns 7 and 8, abstracts one level from the specific independence measures and considers the broader classes of de facto and de jure measures instead. Primary estimations that employ de facto measures yield lower t-values than de jure measures. This result occurs because the de jure measures are dominated by Cukierman's legal indices (59.3 % of all de jure coefficients in our sample are from estimations that employ Cukierman's original indices or modifications of these indices) which, as we know from model 3, yield more positive independence coefficients than all other independence measures. The dominance of the TOR measures (91.7 %) among the de facto portion of the sample explains why the vulnerability measures, which yield higher CBI coefficients than certain de jure measures in model 3, do not suffice to make de facto measures cause higher CBI coefficients than de jure measures.

The next model is reported in columns 9 and 10 in table 1. In this model we consider the geographic and time scope of the primary studies. The results for the time dummies reveal that compared to the 1950s (the base category in our meta-regression), estimations for the 1960s report higher CBI coefficients and studies that utilize data from the 1970s and 1980s (at least in the full sample, column 9) yield lower CBI t-values. This result, which is quite robust to variations of the specification of our meta-regression, is consistent with the traditional hypothesis that higher central bank independence leads to lower inflation. After the dissolution of the Bretton Woods system, central banks regained monetary policy as a tool to fight inflation and independent monetary authorities may have used this tool more successfully than others. The results of this model also indicate that primary

[11] In comparison, only 16 % of the models that employ the TOR measure are bivariate models, 15 % of the models that employ the index of Grilli et al. (1991) and even only 13 % of the models that use Cukiermans legal indices.

[12] However, the significant proportion of bivariate models is not the only reason. The coefficients of models that employ latent variable measures of independence also significantly lower t-values and only 7 % of these are bivariate models. Although this result should not be overrated because it includes only 29 latent coefficients from 5 studies, a similar case can be made for the TOR and vulnerability measures, which represent a significantly larger proportion of the sample.

study samples that include a larger proportion of high-income countries result in higher, i.e., more positive, CBI coefficients. This result is robust to measuring high-income by the percentage of countries that are categorized by the World Bank as high and upper-middle income countries, rather than the percentage of OECD countries in the sample, as in table 1.[13] However, the results change when we restrict the sample to the estimations that are based on de jure independence measures. In this case, a larger proportion of high-income countries decreases the t-values of the independence coefficients. This result is supported by an alternative specification provided in columns 11 and 12 of table 1, where we include a dummy for de jure independence measures and the interaction of that dummy with the percentage of high-income countries in the sample of primary studies. The de jure dummy is significantly positive, which is consistent with the previous findings; however, the interaction effect is negative indicating that the coefficients of de jure measures are lower in high-income countries. Under the assumption that the classic hypothesis of higher CBI leading to lower inflation indeed holds, this would support Cukierman et al. (1992)'s assertion that de jure independence may play a larger role in high-income countries.[14]

For the models that are reported in columns 13 and 14, we focus on the methodological differences between the primary studies. Including country fixed effects and time dummies causes the independence coefficients to be more positive. These variables reflect the effects of other variables that are correlated with independence but that are not included in the primary estimations. Their effects would otherwise be (wrongly) attributed to central bank independence. Time series models appear to result in lower, i.e., more negative, independence coefficients. Initially, the reason for this effect appears to be a larger number of observations (see columns 15 and 16). If included in the meta-regression, the number of observations has a significantly negative effect on the independence coefficients and the time series variable loses its significance. Is the effect of the number of observations caused by a more precise estimation? No, apparently not, since we control for the size of the standard errors (which, as previously discussed, also reflect the effect of differences in the magnitude of the CBI coefficients that result from different CBI and inflation measures in the primary studies). But the effect of the number of observations may come from the fact

[13]Detailed results can be obtained from the author by request.

[14]Furthermore, the negative effect of turnover-related CBI measures is even more negative in studies that have a large proportion of low-income countries, which also supports this interpretation (detailed results are not reported here).

that time series models with more observations can account for more controls (the number of which we do not consider directly in our meta-regression models). If the effect direction of these controls on inflation and their covariation with CBI have different signs, a negative effect of the number of observations on the CBI coefficients would result.

In addition, we note that the primary models that estimate robust standard errors lead to stronger negative CBI coefficients. This result also applies to models that only include the CBI measure with a time lag and is consistent with the traditional hypothesis that CBI decreases inflation. Because changes in CBI need time to affect inflation due to various time lags in policy decision making and implementation, the CBI effect should be stronger if the time lag is explicitly considered.

The effect of the variable that reflects the use of averaged inflation values instead of yearly values as the dependent variable in the primary estimations is significant here but not in most of the other specifications that we estimated. In addition, this variable has mixed effect directions.

The variable that indicates the use of instrumentation in the primary estimations is generally positive (with the exception of the subsample of cross-section models) and partly significant at conventional levels. Under the assumption of the classical hypothesis that CBI decreases inflation, this would imply that either inflation or other variables that cause CBI have a positive effect on CBI. This would align with the hypothesis that high-inflation countries are more likely to establish independent monetary authorities.

Finally, we analyze the role of controls in the primary estimations. The results of this model are provided in columns 17 and 18 of table 1. The effect of conservatism is positive and significant; models that did not consider conservatism report negatively biased coefficients for independence. This result is consistent with the concept that conservatism is positively correlated with independence and exerts a negative influence on inflation. In this case, part of the negative effect of conservatism would, if left out from the specification, be incorrectly attributed to independence. If, instead of *CONVERVAT_BROAD*, the narrower definition *CONSERVAT* is used (not reported here), the effect keeps its sign but loses its significance.

The meta-regressor for including political (in)stability in the primary estimations is negative and at least for the models that employ the narrow definition of conservatism, it is also significant. This implies a positive bias of the independence coefficient estimates when political (in)stability is not included. This bias would require the effects of political (in)stability on inflation and its correlation with CBI to have the same signs. A negative relationship between political instability (most authors consider instability rather than stability) and independence appears plausible, the negative impact of political instability on the level of inflation is much less plausible. Therefore, the interpretation of this result is not very intuitive.[15] For the other variables, which are partially significant, the interpretation of the effects is complicated because the studies lack a clear hypothesis regarding the relationships between each of these variables and central bank independence. However, this assumption would be necessary to provide a sound assessment of the omitted variable bias in these cases.

To test our results for robustness, we constructed two models based on the meta-regressors that generally proved significant and robust (in terms of effect direction) in the models of table 1 and certain variations of these models. The results are provided in table 2. In addition, we estimated these two models based on 1,000 randomly selected subsamples that are two thirds the size of the entire sample. Table 2 reports these results in terms of the percentage of estimations where the coefficient had the same sign as expected (based on the prior results) and the percentage of runs where the coefficient was significant at the 10 % level.

A set of meta-regressors in the *first* model not only consistently delivers the expected effect directions but is also significant in a large number of repeated estimations that included the use of an independence indicator that reflects policy independence, the percentage of OECD countries and its interaction with de jure independence measures, the number of observations and the estimation with fixed effects. Interestingly, in the model based on the subsample of negative CBI coefficients, the meta-regressor that indicates that hypothesis

[15] Our (in)stability meta-regressor reflects whether primary estimations include a control for the presence of coups, civil wars and other non-regular events. If we consider a broader definition of political (in)stability (including for example, cabinet changes and government crises, as in Krause and Méndez (2008)), then the coefficient of the meta-regressor is positive in both the full sample and subsample of negative CBI coefficients, but it is never significant.

tests in the primary studies were based on robust standard errors is positive in most estimations (though rarely significant). This result suggests that the CBI coefficients from these estimations were significant less frequently. In addition, the inclusion of lagged values of the CBI indicator lowers the CBI coefficients. This effect is highly robust and often significant in the subsample of negative CBI coefficients.

The *second model* resulted in robust and often significant effects, particularly for the meta-regressors for bivariate primary estimations, independence measures of the Grilli et al. (1991) type, the proportion of high-income countries in the primary samples, the number of observations and the existence of fixed effects. If the sample is restricted to negative CBI coefficients, the relevance of the proportion of high-income countries in the sample is less pronounced and the effect of a turnover-type CBI indicator presents more clearly than in the model that is based on the full sample of positive and negative CBI coefficients. Even more interesting, however, is that bivariate primary estimations apparently result ceteris paribus in higher, i.e., less often significant, CBI coefficients.

Finally, similar to Klomp and de Haan (2010*b*), we constructed models by following a general-to-specific approach and eliminated the variable with the highest p-value in each round. This step-wise elimination was performed for each of the models in table 2. The results of these estimations are provided in table 3. This table also reports the results of random-sample estimations following the same approach as for the models that appear in table 2. The results that were previously presented are qualitatively confirmed by our general-to-specific models.

6 Conclusion

Since the influential work of Barro and Gordon (1983), the potential effects of central bank independence (CBI) on inflation outcomes have been widely discussed in the literature. The conventional wisdom that greater central bank independence is conducive to lower inflation is reflected in the institutional design of numerous central banks around the world. This conventional wisdom has been confirmed by various empirical studies, though to a different degree.

This paper is, after Klomp and de Haan (2010*b*), among the first to apply meta-regression analysis to analyze the reasons for differences in the results that are reported by empirical studies. We collected 894 CBI coefficients from 85 empirical papers and estimate the effects of a broad set of meta-regressors that reflect the sample scope, the estimation methodology, the sort of independence measure used, and the controls included in the primary studies.

We find that the number of observations has a strong negative effect on the t-values of CBI coefficients in the primary studies and estimations that use fixed effects models generally report higher, i.e., more positive, t-values for CBI.

Unlike Klomp and de Haan (2010*b*), who use a different estimation approach, we note significant differences in the CBI coefficients caused by the choice of the independence measure: *First*, on the level of broad categories, measures that quantify de facto independence result in lower, i.e., more negative, CBI coefficient t-values. This implies that in the subsample of negative CBI coefficients, significance is found more often with de facto than with de jure independence measures. *Second*, on the level of independence facets covered by the measures, our findings suggest that independence measures reflecting policy independence result in more positive CBI coefficients. *Third*, on the level of single independence measures, the turnover- and vulnerability-related (de facto) measures and CBI measures of the Grilli et al. (1991) type result consistently in lower, i.e., more strongly negative, CBI t-values than do the legal independence indices of Cukierman et al. (1992), which are broadly used in the literature.

In accordance with Klomp and de Haan (2010*b*), we also find that primary studies covering the 1970s usually report stronger negative CBI coefficients. Whereas primary studies with samples including a high share of high-income countries in general, or OCED countries in particular, generally find higher CBI coefficients, we identify a negative interaction effect between a high share of those countries in the primary samples and the use of a de jure independence measure.

We do not find robust effects of the specific control variables that are included in the primary estimations, including political conservatism, political stability, openness or labor market institutions. On a more general level, we note that bivariate models have a positive effect on the t-values of CBI coefficients that are estimated in these models. However, these results are not robust when the sample is restricted to the subset of negative CBI coefficients.

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Appendix

Table A1: Primary studies

Study	Mean CBI coefficients	Variance CBI coefficients	Number of coefficients	...thereof negative (in %)	...thereof significant at 5 % (in %)
Adolph (2004)	-3,319	0,053	4	100,0	100,0
Al-Marhubi and Willett (1995)	-2,311	12,584	15	80,0	93,3
Alpanda and Honig (2014)	-0,536	1,108	12	58,3	25,0
Athanasios (2009)	-2,787	0,256	10	100,0	100,0
Banaian, Burdekin and Willett (1998)	-1,035	1,684	8	75,0	37,5
Banaian and Luksetich (2001)	-0,777	0,738	10	80,0	30,0
Banaian, Burdekin and Willett (1995)	0,438	3,402	9	44,4	55,6
Belke, Freytag, Keil and Schneider (2012)	-0,080	3,073	9	44,4	33,3
Bleaney (1996)	-0,908	2,227	6	50,0	50,0
Bogoev, Petrevski and Sergi (2012)	-3,787	0,437	8	100,0	100,0
Borghijs (2009)	-4,935	0,132	4	100,0	100,0
Bowdler and Nunziata (2005)	-1,190	1,571	4	100,0	25,0
Broz (2002)	1,802	0,320	8	0,0	62,5
Brumm (2000)	-4,068	39,778	6	66,7	66,7
Brumm (2002)	-1,635	4,502	6	66,7	66,7
Brumm (2006)	-1,701	2,389	4	75,0	50,0
Campillo and Miron (1997)	0,312	1,556	29	20,7	20,7
Cargill (1993)	-1,854	2,282	25	88,0	48,0
Carmignani, Colombo and Tirelli (2008)	-4,127	0,134	11	100,0	100,0
Cecchetti and Krause (2002)	1,284	0,811	2	0,0	50,0
Chou (2000)	-2,028	0,851	5	100,0	80,0
Chou (2001)	-2,175	1,498	15	100,0	66,7
Clark (2003)	-1,068		1	100,0	0,0
Crowe (2006)	-0,760	2,455	4	50,0	25,0
Crowe and Meade (2008)	-1,600	2,991	7	71,4	57,1
Cukierman, Miller and Neyapti (2002)	-0,988	2,023	3	66,7	33,3
Cukierman (2010)	-1,660	3,160	10	80,0	80,0
Cukierman, Webb and Neyapti (1992)	-2,420	10,181	9	66,7	66,7
Daniels, Nourzad, VanHoose (2005)	-5,190	2,636	6	100,0	100,0
Daniels and VanHoose (2009)	-4,561	3,557	5	100,0	100,0
de Gregorio (1996)	-1,920	3,967	10	70,0	60,0
de Haan and Sturm (1992)	-2,728	3,145	12	91,7	75,0
de Haan Leertouwer, Meijer and Wansbeek (2003)	-3,134	1,671	25	100,0	80,0
de Haan and Kooi (2000)	-1,757	1,246	11	81,8	63,6
de Haan and Siermann (1996)	-2,237	10,766	18	66,7	44,4
de Jong (2001)	-5,842	1,474	6	100,0	100,0
Destefanis and Rizza (2007)	-1,755	0,613	41	95,1	63,4
Dolmas, Huffman and Wynne (2000)	-1,303	0,198	24	100,0	29,2
Dreher, Sturm and de Haan (2007)	-2,135	2,465	8	100,0	50,0
Eijffinger and Schaling (1993)	-2,669	0,913	12	100,0	83,3
Eijffinger and van Keulen (1995)	-0,518	1,359	8	62,5	25,0
Eijffinger and de Haan (1996)	-0,549	7,572	8	50,0	62,5

Study	Mean CBI coefficients	Variance CBI coefficients	Number of coefficients	...thereof negative (in %)	...thereof significant at 5 % (in %)
Eijffinger, Schaling and Hoeberichts (1998)	-3,838	1,400	24	100,0	100,0
Eijffinger and Stadhouders (2003)	-0,493	0,898	27	51,9	14,8
Farvaque Héricourt and Lagadec (2008)	-2,065	2,973	4	75,0	75,0
Flanagan and Hammermann (2007)	-1,236	2,093	2	100,0	50,0
Franzese (1999)	-2,569		1	100,0	100,0
Franzese and Hall (2000)	-2,033		1	100,0	100,0
Franzese (2003)	3,491	68,984	2	50,0	100,0
Fujiki (1996)	-2,058	0,152	7	100,0	71,4
Grilli, Tabellini, Malinvaud and Pagano (1991)	-2,630	3,767	15	100,0	53,3
Gutiérrez (2004)	-1,087	1,428	5	80,0	40,0
Hayo and Voigt (2005)	1,149	2,266	4	25,0	50,0
Heylen and van Poeck (1995)	3,887	0,685	3	0,0	100,0
Ilieva an Gregoriou (2005)	-2,150		1	100,0	100,0
Iversen (1999)	-1,070	2,551	6	83,3	16,7
Jácome and Vásquez (2008)	-2,342	2,156	22	100,0	63,6
Jenkins (1996)	-0,499	2,550	27	51,9	48,1
Jonsson (1995)	-3,224	4,854	10	100,0	80,0
Keefer and Stasavage (1999)	0,245	2,549	6	50,0	33,3
Keefer and Stasavage (2001)	3,044	0,809	10	0,0	100,0
Stasavage and Keefer (2003)	1,498	0,988	4	0,0	50,0
Kilponen (1999)	-2,346	0,786	3	100,0	100,0
King and Ma (2001)	-0,410	3,852	3	33,3	33,3
Klomp and de Haan (2010)	-3,893	16,924	14	100,0	57,1
Krause/Méndez (2008)	-2,062	0,259	4	100,0	75,0
Lane (1997)	-1,319	10,821	9	66,7	66,7
Lin (2009)	1,315		1	0,0	0,0
Maliszewski (1997)	-2,744	1,190	4	100,0	100,0
Maslowska (2011)	-2,570	10,230	56	83,9	80,4
Mazhar and Méon (2012)	-0,036	0,189	3	66,7	0,0
Neyapti (2012)	0,610	2,116	7	28,6	28,6
Oatley (1999)	-0,033	3,410	12	58,3	41,7
Pehnelt (2007)	-1,813	0,241	4	100,0	50,0
Petrevski, Bogoev and Sergi (2012)	1,272	0,070	4	0,0	0,0
Posen (1995)	0,893	0,553	10	10,0	10,0
Prast (1997)	-1,233	0,787	9	100,0	33,3
Quintyn and Gollwitzer (2010)	-2,510	0,081	5	100,0	100,0
Romer (1993)	5,847	0,031	2	0,0	100,0
Siklos (2008)	-0,250	0,168	2	50,0	0,0
Sturm and de Haan (2000)	-1,647	0,813	26	100,0	50,0
Sturm and de Haan (2001)	-2,030	0,351	6	100,0	66,7
Treisman (2000)	-0,891	1,650	14	78,6	35,7
Vuletin and Zhu (2011)	-1,940	1,772	59	88,1	64,4
Zervoyianni, Anastasiou and Anastasiou (2012)	-1,843	1,042	18	100,0	55,6