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央行獨立性與通貨膨脹

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中 文 摘 要 ： 本文針對追蹤資料分量迴歸模型提出兩階段配適值法，探討央行獨立性和通貨膨脹之關係。本文有以下貢獻：首先，本文提出之模型說明央行獨立性的抗通膨效果在通貨膨脹較高時，效果較大，這可以解釋Franzese (1999) 的理論。其次，本文提出之方法可以解決用央行總裁更替率當央行獨立性變數的內生性問題。因此，本文能完整地分析央行獨立性和通膨之關係。

中 文 關 鍵 詞 ： 央行獨立性；內生性；通貨膨脹；分量迴歸

英 文 摘 要 ： This paper investigates the empirical relationship between central bank independence (CBI) and inflation by proposing a two-stage fitted value approach to a quantile regression for panel data
models. This approach has several advantages. First, Franzese (1999) proposes theoretically that the anti-inflationary effect of CBI is heterogeneous and is stronger when inflation is higher. Our method,
which estimates the anti-inflationary effect of CBI for various rates of inflation, can expose the conditional heterogeneity of inflation. Second, a simple two-stage approach to the quantile regression for panel data models is proposed to solve the endogeneity problem by using the turnover rate of a central bank governor as a measure of CBI. Third, by exploiting an extensive panel data set, our empirical findings show that the anti-inflationary effect of CBI is stronger in higher
inflation episodes, and is weaker in lower inflation episodes. As we explore this method, the CBI--inflation relationship becomes more convincing.

英 文 關 鍵 詞 ： central bank independence; endogeneity; inflation; quantile regression

The Anti-Inflationary Effect of Central Bank Independence: Heterogeneity and Endogeneity

Abstract

This paper investigates the empirical relationship between central bank independence (CBI) and inflation by proposing a two-stage fitted value approach to a quantile regression for panel data models. This approach has several advantages. First, Franzese (1999) proposes theoretically that the anti-inflationary effect of CBI is heterogeneous and is stronger when inflation is higher. Our method, which estimates the anti-inflationary effect of CBI for various rates of inflation, can expose the conditional heterogeneity of inflation. Second, a simple two-stage approach to the quantile regression for panel data models is proposed to solve the endogeneity problem by using the turnover rate of a central bank governor as a measure of CBI. Third, by exploiting an extensive panel data set, our empirical findings show that the anti-inflationary effect of CBI is stronger in higher inflation episodes, and is weaker in lower inflation episodes. As we explore this method, the CBI–inflation relationship becomes more convincing.

JEL classification: E31; E52; E59

Keywords: Central Bank Independence; Endogeneity; Inflation; Quantile Regression

1 Introduction

Central bank independence (CBI) refers to the ability of central banks to make decisions that are independent of the government. The degree of CBI affects the rates at which the money supply and credit expand, and also has an impact on important aspects of macroeconomic performance. Rogoff (1985) and Lohmann (1992) argue that conservative central bankers attach substantial weight to inflation-rate stabilization, and to reducing the inflationary bias resulting from the time-inconsistency problem in their monetary policies. Their theories imply that the effective conservativeness or independence of a central bank reduces inflation. Cukierman (1992) also points out that the higher the degree of CBI, the more the monetary authority becomes committed to fighting inflation. Thus, independent central banks have been associated with lower inflation rates.

Franzese (1999) constructs a political–economic model where the observed inflation is a weighted average of "commitment" inflation if the conservative central bank autonomously controls the monetary policy, and "discretionary" inflation if instead the current government controls monetary policy, with the degree of CBI weighting the former. From this model, the anti-inflationary effect of CBI is stronger with higher discretionary inflation relative to commitment inflation. For example, in a high-inflation environment, the political economy imposes substantial inflationary pressure on the government, and the discretionary inflation is higher than the commitment inflation. Thus, the anti-inflationary effect of CBI is stronger in situations of higher inflation. On the other hand, if the political economy puts minimal inflationary pressure on a government when inflation is low, then the anti-inflationary effect of CBI is weaker. Therefore, the anti-inflationary effect of CBI is heterogeneous at different inflation levels.

The empirical research on the relationship between CBI and inflation is consistent with the theoretical arguments. Alesina (1988, 1989), Grilli, Masciandaro, and Tabellini (1991), Alesina and Summers (1993), and Franzese (1999) find evidence of a negative relationship in developed and industrial countries. Cukierman, Webb, and Neyapti (1992) observe that the legal indicator of CBI is inversely related to inflation in industrial countries, but not in developing countries. They propose that the turnover rate (TOR) of central bank governors is a more appropriate measure of CBI. Because there may be reverse causality running from inflation to the TOR index, Cukierman, Webb, and Neyapti (1992) introduce instrumental variables for TOR to solve the endogeneity problem and find that the effect of CBI on inflation is negative. Recently, Jácome and Vázquez (2008) and Brumm (2011) have also considered the likely endogeneity of TOR, and affirm the anti-inflationary effects of CBI. Dreher, Sturm, and de Haan (2008) use some political and economic factors as instruments of TOR but find no significant CBI–inflation relationship. In addition, several studies note that the anti-inflationary effect of CBI is influenced by a few influential observations. Temple (1998), de Haan and Kooi (2000), Sturm and de Haan (2001), Bouwman, Jong-A-Pin, and de Haan (2005), Dreher, Sturm, and de Haan (2008), Lin (2010), and Vuletin and Zhu (2011) find that the CBI–inflation relationship is not constant and tends to weaken if high-inflation observations are excluded from the sample. Klomp and de Haan (2010) also confirm empirically that a heterogeneous model is appropriate for estimating the CBI–inflation relationship.

Both the theoretical and empirical literature point out that the anti-inflationary effect of CBI is heterogeneous at different inflation levels. Accordingly, our first contribution is to employ a quantile regression for the panel data model to estimate the anti-inflationary effect of CBI at various inflation rates in order to expose the conditional heterogeneity of inflation. The quantile regression ideally uncovers the relationship between CBI and inflation. We include high-inflation observations and explore the implications of both the theory and the data. In addition, the use of TOR leads to endogeneity. Thus, our second contribution is to propose a fitted value approach for the quantile regression for the panel data model to solve for the endogeneity problem that occurs when using TOR as a measure of CBI. The proposed approach is a two-stage estimation procedure. We first obtain the fitted value of the endogenous variable and we then use it to replace the endogenous variable of the model.

Finally, we provide an empirical justification for the relationship between CBI and inflation. In our empirical study, we consider panel data spanning 93 countries over the period 1974–2010. We also include country fixed effects in all specifications to remove the impact on inflation resulting from fixed country characteristics that are potentially correlated with CBI. Our results show that there exists a significantly positive relationship between TOR and inflation at all inflation levels, i.e., a negative CBI–inflation relationship for the whole distribution of inflation. In particular, the anti-inflationary effect of CBI is stronger in high- and middle -inflation episodes than in low-inflation episodes. The empirical results reveal that the CBI–inflation relationship is nonlinear and heterogeneous across different inflation levels, which is consistent with the theoretical argument of Franzese (1999) and several empirical studies (Temple, 1998; de Haan and Kooi, 2000; Sturm and de Haan, 2001;

Bouwman, Jong-A-Pin, and de Haan, 2005; Klomp and de Haan, 2010; Vuletin and Zhu, 2011). The results are robust to the inclusion of other inflation-related variables such as openness, the exchange rate regime, political instability, and GDP per capita. We further explore the robustness of the results by using an alternative measure of TOR as well as different classifications of fixed exchange rate regimes. We also consider different unit periods and different sample periods. These robustness checks provide similar trends for the anti-inflationary effect of CBI. Moreover, the effect of the heterogeneity of CBI on inflation is robust to various methods adopted to tackle the problem of endogeneity in relation to TOR.

The remainder of this paper is organized as follows. Section 2 provides the literature review. The econometric methodology and definitions of the data are presented in Section 3. In Section 4, we discuss the results and present the robustness check. In Section 5, we conclude the paper. A list of countries used in this paper is included in the Appendix.

2 Literature Review

As noted by Kydland and Prescott (1977) and Barro and Gordon (1983), the time inconsistency problem of monetary policy leads to inflationary bias. Rogoff (1985) shows that inflationary bias can be reduced by delegating monetary policy to an independent central bank that attaches greater emphasis to inflation rate stabilization than employment stabilization. Rogoff (1985) argues that CBI increases the monetary authority's commitment to fighting inflation, wherein private sector agents reduce their wage increases, thus lowering inflation. Lohmann (1992) proposes that the central banker will implement a nonlinear policy rule and reduce the inflationary bias associated with the time-inconsistency problem of monetary policy. Cukierman (1992) argues that the independence of central banks from their respective political authorities can influence the distribution of inflation. He specifies that a higher degree of CBI denotes a stronger commitment on the part of the monetary authority to fight inflation. Therefore, more independent central banks have been associated with lower inflation rates.

Franzese (1999) proposes a political–economic model in which monetary policy is controlled in part by the central bank and in part by the current government; see also Cukierman (2008). The observed inflation is a weighted average of the "commitment" inflation if the conservative central bank autonomously controls monetary policy and "discretionary" inflation if instead the current government controls monetary policy, with the degree of CBI weighting the former. From this model, the anti-inflationary effect of CBI is stronger the higher that the discretionary inflation would have been relative to what the commitment inflation would have been. Franzese (1999) argues that the anti-inflationary effect of CBI is not constant and depends on the characteristics of the broader political–economic environment in which the central bank operates. In a high-inflation environment, the political economy exerts great inflationary pressure on the government, and the discretionary inflation is higher than the commitment inflation. Therefore, the anti-inflationary effect of CBI is stronger in cases of higher inflation. On the other hand, the political economy puts little inflationary pressure on the government during times of low inflation, and thus the anti-inflationary effect of CBI is weaker. Franzese (1999) provides theoretical support to demonstrate that the effect of CBI is heterogeneous and that it varies across levels of inflation.

Empirical work applied to developed and industrial countries supports the negative relationship between CBI and inflation. In considering 16 OECD countries during the period 1973–1985, Alesina (1988) finds that the countries with the most independent central banks have the lowest inflation, whereas those with the most dependent central banks have some of the highest inflation rates. Alesina (1989) obtains similar results for 17 OECD countries during the period 1973–1986. Grilli, Masciandaro, and Tabellini (1991) compare the monetary regimes of 18 OECD countries in the period 1950–1989, and show that lower inflation is associated with higher CBI. In addition, by plotting the cross-country inflation rates for 16 OECD countries against the CBI measure, Alesina and Summers (1993) verify a nearly perfect negative correlation between CBI and inflation during 1955–1988. The empirical results of Franzese (1999) for 18 developed OECD countries in 1972–1990 confirm that the anti-inflationary effect of any given degree of CBI is greater whenever the government has a stronger incentive to pursue inflationary policies.

Cukierman, Webb, and Neyapti (1992) find that the legal indicator of CBI is inversely related to inflation in industrial countries, but not in developing countries, because legal measures of CBI may not reflect the true relationship between CBI and inflation. Cukierman, Webb, and Neyapti (1992) argue that the actual average term in office of a central bank governor may be a better proxy for CBI, and propose using the TOR of central bank governors as an alternative measure of CBI. Of particular note, when using the TOR index as the measure of CBI, it is difficult to determine whether inflation is high because of political interference that leads to the rapid turnover of central bank officials or because central bank officials are tossed out when they cannot keep inflation low. Cukierman, Webb, and Neyapti (1992) recognize the possibility of endogeneity and introduce instrumental variables for TOR. They find a negative effect of CBI on inflation for 72 countries during the 1950–1989 period. Dreher, Sturm, and de Haan (2008) use the conditional logit model for the likelihood that a central bank governor will be replaced in order to take the endogeneity of TOR into account. They find that the relationship between CBI and inflation is insignificant for 137 countries covering the period 1970–2004. Jácome and Vázquez (2008) explore the effect of CBI on inflation in a sample of 24 Latin American and Caribbean countries during the period 1985–2002. After considering the likely endogeneity of CBI, they find that CBI has a negative effect on inflation. Brumm (2011) addresses the endogeneity of the CBI–inflation relationship by using analysis of covariance structures and finds evidence of a negative relationship for 42 countries during the period from the early 1970s to the mid-1990s.

As noted by several studies, the empirical results of the CBI–inflation relationship are affected by a few influential observations (Dreher, Sturm, and de Haan, 2008). Temple (1998) finds that the relationship between CBI and inflation for 18 countries over the 1974–1994 period is extremely sensitive to influential observations, and shows that there exists a negative effect of CBI on inflation in high-income economies. de Haan and Kooi (2000) explore the effect of CBI on inflation in 82 developing countries over the 1980–1989 period, but fail to find any negative effects of CBI. They find that CBI is related to inflation only if high-inflation countries are included in the sample. Sturm and de Haan (2001) extend the data of de Haan and Kooi to include the years 1980–1998 and obtain similar results. Bouwman, Jong-A-Pin, and de Haan (2005) use the quantile regression method to investigate the CBI–inflation relationship in 57 developing countries for the period 1975–1998 and find evidence of a significant relationship only in the higher quantiles of inflation. Lin (2010) revisits the CBI–inflation relationship and shows that the relationship can be positive or negative for different levels of inflation for 44 countries during 1948–1972. Vuletin and Zhu (2011) calculate TOR using a rolling average over the four years that precede a central bank governor change to purge the sample of reverse causality concerns. Their empirical results indicate that the CBI–inflation relationship tends to weaken if the 10% of observations with the highest inflation rates are excluded for 42 countries during the 1972–2006 period.

Most of the existing literature on the CBI–inflation relationship is based on the pooled estimation of panel data, which could provide inconsistent and misleading estimates for the coefficients of interest. Klomp and de Haan (2010) use a random coefficient specification of a panel data model to examine to what extent heterogeneity influences the relationship. While they do not find a significant relationship between CBI and inflation for more than 100 countries during the years 1980–2005, they do find evidence of a significant relationship in some developing countries. Klomp and de Haan (2010) thus suggest that a heterogeneous model is the appropriate model for estimating the relationship between CBI and inflation.

3 Econometric Methodology and Data

Our analysis of the effect of CBI on inflation involves an estimation procedure based on the features of quantile regression and panel data models.

3.1 Endogeneity in Panel Data Quantile Regressions

Consider a location-scale shift panel data model, $\forall i = 1, \cdots, N, t = 1, \cdots, T$,

$$y_{it} = \alpha_1 d_{it} + x'_{it}\beta_1 + \eta_i + (\alpha_2 d_{it} + x'_{it}\beta_2)u_{it}, \tag{1}$$

where y_{it} is a real-valued dependent variable; d_{it} is an endogenous variable; x_{it} is a vector of real-valued, continuously distributed, exogenous explanatory variables; η_i is the parameter that represents the individual fixed effects; α_1, α_2, β_1 and β_2 are unknown parameters; and u_{it} is the error term. The fixed effects η_i in (1) capture some sources of the variability, or "unobserved heterogeneity," that is not adequately controlled by other regressors in the model. The fixed effect η_i is a pure location shift effect. By construction, d_{it} is an endogenous variable and is correlated with the error term. The aim of this section is to propose a fitted value approach to deal with the endogeneity problem in model (1).

First, consider the following regression model where the endogenous explanatory variable is regressed on the instrumental variable z_{it}:

$$d_{it} = z'_{it}\gamma + v_{it}, \tag{2}$$

where z_{it} is a vector of the instrumental variable, γ is a $(d_Z \times 1)$ vector of unknown parameters, and v_{it} is a real-valued unobserved random variable. Here, z_{it} is allowed to contain the explanatory variable x_{it}. For the identification of the model, it should be specified that there is at least one component of z_{it} that is not included in x_{it}. Replacing d_{it} in (1) by the regression model (2) yields:

$$y_{it} = \alpha_1 z_{it}'\gamma + x_{it}'\beta_1 + \eta_i + \alpha_1 v_{it} + \alpha_2 z_{it}'\gamma u_{it} + \alpha_2 v_{it} u_{it} + x_{it}'\beta_2 u_{it}.$$

Note that the error terms v_{it} and u_{it} in the above model are independent of the exogenous explanatory variable x_{it}, and by construction, v_{it} and z_{it} in (2) are independent. Assuming that u_{it} is independent of z_{it}, almost surely, we can then obtain the τ-th conditional quantile function for the above panel data model as follows:

$$Q_{y_{it}}(\tau|x_{it}, z_{it}) = \alpha(\tau) z_{it}'\gamma + x_{it}'\beta(\tau) + \eta_i + c(\tau),$$

where $\alpha(\tau) = \alpha_1 + \alpha_2 Q_{u_{it}}(\tau)$, $\beta(\tau) = \beta_1 + \beta_2 Q_{u_{it}}(\tau)$, and $c(\tau) = \alpha_1 Q_{v_{it}}(\tau) + \alpha_2 Q_{v_{it}u_{it}}(\tau)$. Note that the fixed effect in our specification does not depend on the quantile τ, which is more realistic in studying the CBI–inflation relationship.

Second, to identify the estimation procedure used, we need to assume that $\hat{\gamma}$ is any consistent M-estimator for γ in regression (2). The τ-th conditional quantile function for the panel data model is:

$$Q_{y_{it}}(\tau|x_{it}, z_{it}) = \alpha(\tau) z_{it}'\hat{\gamma} + x_{it}'\beta(\tau) + \eta_i + c(\tau). \tag{3}$$

The penalized quantile regression approach of Koenker (2004) can then be used to obtain consistent estimators of $\alpha(\tau)$, $\beta(\tau)$, and $c(\tau)$ in (3), where $c(\tau)$ is viewed as the coefficient of the constant term. This suggests that the parameters of the quantile regression for the panel data model can be estimated by a two-stage procedure. The first stage is to construct a regression of d_{it} on z_{it} and obtain the fitted value $z_{it}'\hat{\gamma}$. In the second step, the fitted value $z_{it}'\hat{\gamma}$ is substituted in place of the endogenous variable d_{it}, and the penalized quantile regression approach for panel data models is used for (3). Therefore, the two-stage estimation corrects for the endogeneity of the quantile regression for the panel data model by replacing d_{it} with $z_{it}'\hat{\gamma}$ and can be viewed as a variant of the fitted value approach.

Several studies propose solving similar endogeneity problems using the quantile regression for panel data models. For example, Arias, Hallock, and Sosa-Escudero (2001), following the control function approach, suggest a two-stage estimation. Harding and Lamarche (2009), Galvao and Montes-Rojas (2010), and Galvao (2011) introduce an instrumental variable quantile regression method for panel data models. However, the former two papers deal with models where the fixed effect depends on the quantile, while the latter two papers consider a dynamic panel data model without endogenous variables. One of the main contributions of this paper is that it employs a simple two-stage estimation for quantile regressions using a panel data model in which the fixed effect does not depend on the quantile.

3.2 Data

The main data set consists of a panel of 93 countries covering the period 1974–2010 using annual data. A list of constituent countries is provided in the Appendix. Following Dreher, Sturm, and de Haan (2008), the dependent variable used in this paper is transformed inflation, and the explanatory variables include TOR, the interaction of TOR and the OECD countries' dummy, trade openness, a fixed exchange rate regime dummy, political instability, and GDP per capita. Transformed inflation is defined as $\Pi_{it} := (\pi_{it}/100)/(1 + \pi_{it}/100)$, where π_{it} is measured by the annual change in the consumer price index. Trade openness is measured by the ratio of annual imports plus exports to GDP. The data for inflation, GDP per capita and trade openness are obtained from the World Bank's World Development Indicators. TOR is measured by the frequency of turnover for central bank governors taken from Dreher, Sturm and de Haan (2008). The fixed exchange rate regime dummy equals one if the exchange rate is classified as fixed according to the *de facto* classification of exchange rate regimes in Levy-Yeyati and Sturzenegger (2005), and equals zero otherwise.[1] The degree of political instability is measured by the first principal component of the number of assassinations, strikes, guerrilla warfare attacks, major crises, riots, and revolutions taken from the Databanks International Cross-National Time-Series Data Archive (2012).

The instrumental variables that we take into account for TOR come from Dreher, Sturm, and de Haan (2008); they include elections, lagged inflation, the number of coups, the percentage of veto players who drop, and the average share of the legal term in office that has elapsed. The number of coups includes both successful and unsuccessful attempts to overthrow the government (see Powell and Thyne, 2011). The variable percentage of veto players who drop counts the percentage of veto players who drop from the government in a given year and is taken from data

[1]The original data of Levy-Yeyati and Sturzenegger (2005) are updated to 2004. To increase the data availability, we follow the methodology provided by Levy-Yeyati and Sturzenegger in order to construct three classification variables, namely, exchange rate volatility, the volatility of exchange rate changes, and the volatility of reserves, and then identify the *de facto* exchange rate regime by a K-means cluster analysis.

provided by Beck et al. (2001). The election variable measures the post-election period by the part of the year which is within 12 months after a national election. The average share of the legal term in office that has elapsed is the ratio between the actual and legal duration of a governor's term in office taken from Dreher, Sturm and de Haan (2008). If the governor's legal term in office is indefinite or unknown, the term is specified as eight years. All data are updated to 2010.

Table 1 provides basic summary statistics of the data. In panel (A) of Table 1, the median and the third quartile are 7.99% and 14.99%, respectively. However, the mean of inflation is 38.85%, which is much larger than the third quartile of inflation. Clearly, the distribution of inflation is right-skewed and the mean is sensitive to extremely large values. While the estimation result of the mean regression for the panel data model is sensitive to the extremely large values of inflation, the quantile regression estimation results are robust to extreme values. For the explanatory variables, the average TOR is 0.21, which implies a change of central bank governor every four years and nine months on average. Openness has a symmetric distribution, because its mean and median are close. The average fixed exchange rate regime dummy is 0.43, which means that about 43% of the sample observations are subject to a fixed exchange rate regime. The distributions of political instability and GDP per capita are both right-skewed. Furthermore, all three quantiles of the number of coups, and the percentage of veto players who drop are equal to 0, which is smaller than their averages (0.04, and 0.13, respectively). This shows that at least 75% of their values are at the same level (0) and the averages are sensitive to extremely large values.

The variables in panels (B) and (C) have similar statistical properties to those in panel (A). In particular, inflation, openness, and GDP per capita during 1980–2010 are more volatile than during 1990–2010. The TOR, political instability, and coup variables tend to be lower in value and more stable after 1990, which demonstrates that the political environment is more stable. As the higher TOR of central bank governors is indicative of a lower level of CBI, we can see that the independence of central banks increases after 1990. The characteristics of the remaining variables result in no clear distinction between the periods 1980–2010 and 1990–2010.

4 Empirical Results

4.1 Benchmark Results

Following the theoretical argument of Franzese (1999), the effect of CBI on inflation is heterogeneous at different inflation levels. To account for the varying anti-inflationary impact of CBI at different inflation levels, we investigate the relationship between inflation and CBI in a panel quantile model. As discussed in the Introduction and Section 3, the model captures the CBI–inflation relationship of interest, controls for unobserved individual heterogeneity, and reveals the heterogeneous effects of regressors on the dependent variable.

To fully investigate the relationship between CBI and inflation, we consider several inflation-related variables. First, because TOR may not be a good indicator of CBI in industrial countries (Cukierman, Webb, and Neyapti, 1992; Dreher, Sturm and de Haan, 2008), we add to the model not only TOR but also its interaction with an OECD dummy, which is one if a country is developed and zero otherwise. Second, Romer (1993) argues that trade openness is negatively related to inflation because the time inconsistency problem of a given monetary policy is less critical in more open countries. Third, Edwards and Losada (1994), Ghosh et al. (1997) and Calvo and Végh (1999) all point out that an announced policy of a fixed exchange rate regime may serve as a commitment technology preventing the government from subsequent temptations to follow expansionary macroeconomic policies and thus lowers inflation. Moreover, Cukierman, Edwards, and Tabellini (1992) propose that countries with more unstable and polarized political systems will have less efficient tax structures, and will thus collect a larger fraction of their revenues through seigniorage. Thus, political instability is considered in the analysis. Finally, economic growth is also an important inflation-related factor, and the growth rate of real GDP per capita is considered as an additional control variable. We use the logarithmic form of the growth rate of real GDP per capita.

When we use TOR as a proxy for CBI to measure its impact on inflation, there may be a reverse causality that runs from inflation to turnover (Cukierman, Webb, and Neyapti, 1992). Such causality leads to endogenous bias in the estimation. Therefore, we take into account the endogeneity problem when examining the CBI–inflation relationship. In this paper, to deal with the endogeneity and heterogeneity, we employ a fitted value approach for the quantile regression for the panel data model using a two-stage estimation procedure. For comparison purposes, we considered two-stage least squares estimation as well as within estimation for the traditional mean regression using the panel data with endogeneity. Table 2 reports the estimation results of the mean and quantile regressions for the panel data.[2] Table 2 shows that the mean and quantile regression estimates of TOR on inflation are all positive. As a higher TOR represents a lower CBI, the empirical results indicate that CBI is anti-inflationary. The results are in line with most empirical studies as well as the theoretical studies of Rogoff (1985), Lohmann (1992), and Cukierman (1992).

Moreover, the coefficients of TOR on inflation are plotted in Figure 1. In the figure, the horizontal and vertical axes correspondingly denote the quantile and the coefficients of TOR on inflation. The black solid line depicts the quantile regression estimates, and the gray dotted lines represent their 95% confidence intervals. The black dashed line represents the mean regression estimates. Figure 1 shows that the impact of CBI on inflation has a clear trend. The quantile regression estimates of TOR increase monotonically, along with the quantiles, in magnitude and significance. That is, the anti-inflationary effects of CBI on inflation are stronger at high and middle quantiles of inflation, whereas they are smaller at low quantiles of inflation. As we review the empirical findings, we see that the CBI–inflation relationship is heterogeneous at different inflation levels. Our results coincide with those of previous empirical studies, and ideally justify the theoretical argument of Franzese (1999), which states that the CBI–inflation relationship tends to weaken if high-inflation observations are excluded. See Temple (1998), de Haan and Kooi (2000), Sturm and de Haan (2001), Bouwman, Jong-A-Pin, and de Haan (2005) and Vuletin and Zhu (2011).

When we examine the interaction of TOR and the OECD dummy, both the mean and quantile regression estimates are negative for all quantiles except for the 0.1 and 0.2 quantiles. These results show that the impact of CBI on inflation is weaker in OECD countries than in non-OECD countries, which indicates that OECD countries have ways of overcoming the dynamic inconsistency problem. One interesting finding

[2]The adjusted t-statistics of the Levin, Lin and Chu (2002) (LLC) test for inflation, lagged inflation, openness, and GDP per capita are -9.14, -9.04, -1.37, and -1.8, respectively. The adjusted t-statistics of the Im, Pesaran, and Shin (2003) (IPS) test are -11.88, -13.07, -1.27, and 7.13, respectively. As we cannot reject the hypothesis that openness and GDP per capita have a unit root, we thus first-difference these variables and use them as our regressors. Both tests reject the hypothesis that the first difference of openness has a unit root, which is the same as the result for GDP per capita.

is that the estimates of TOR on inflation are homogeneous across quantiles in OECD countries and are heterogeneous across quantiles in non-OECD countries. In terms of other inflation-related variables, first, the estimates of openness are insignificantly positive, which is not consistent with Romer (1993). Second, the exchange rate is an important factor related to inflation. Many countries have used a fixed exchange rate regime as a nominal anchor for lowering inflation. The quantile regression estimates of the fixed exchange rate regime dummy are negative along with the quantiles both in values and in significance. Here we see that the adoption of a fixed exchange rate regime is much more effective against inflation in higher inflation episodes, and this finding is in line with the findings of Edwards and Losada (1994), Ghosh et al. (1997) and Calvo and Végh (1999). Third, the mean and quantile regression estimates of political instability are positive, which shows that political instability is an inflationary factor; see also Cukierman, Edwards, and Tabellini (1992), and Dreher, Sturm and de Haan (2008). Finally, the estimates of the growth of real GDP per capita are all negative and they decrease monotonically along with the quantiles. The negative relationship is also supported by Dreher, Sturm and de Haan (2008), but not by Sturm and de Haan (2001).

4.2 Robustness Check

Following Dreher, Sturm and de Haan (2008), we define a new measure for TOR, which equals one if the central bank governor was replaced in a particular year and country, and zero otherwise. Panel (I) in Table 3 reports TOR estimates with the new TOR dummy variable. The quantile regression estimates of the effect of TOR on inflation are positive and increase along with the quantiles. In particular, the TOR estimates are insignificant at the 0.1–0.3 quantiles, significant at the 5% level at the 0.4 quantile, and significant at the 1% level at the 0.5–0.9 quantiles. The empirical results show that the anti-inflationary effect of CBI is heterogeneous at different inflation levels, and is robust with respect to different measures of TOR. However, the results of the interaction of TOR and the OECD dummy are mixed; this may be because TOR is a good proxy for CBI in developing countries, but not in developed countries (Cukierman, Webb, and Neyapti, 1992).

As an additional test, following Vuletin and Zhu (2011), we use the fixed exchange rate regime based on the classification of Reinhart and Rogoff (2004) (henceforth the RR classification) in the model.[3] Reinhart and Rogoff (2004) classify exchange

[3]The data are available from http://www.carmenreinhart.com/research/publications-by-

rate regimes into 14 categories and we define the fixed exchange rate regime dummy as one if the RR classification lies between 1 and 4, and zero otherwise. Panel (II) in Table 3 shows the estimates of TOR with the RR classification for the fixed exchange rate regime dummy used. According to the results, the quantile regression estimates of TOR are also positive and increase along with the quantiles. Thus, the CBI–inflation relationship is robust to different classifications of exchange rate regimes.

Dreher, Sturm and de Haan (2008) and Arnone et al. (2007) use data based on five-year averages, and Cukierman, Webb, and Neyapti (1992) use data based on 10-year averages. This paper uses an annual panel model instead of a panel with five-year averages. To provide a better comparison, we transform the annual data into five-year and 10-year averages, and check the robustness of our results. As shown in panels (III) and (IV) in Table 3, the quantile regression estimates of TOR increase monotonically along with the quantiles, are statistically significant at the 0.3–0.9 quantiles for the data transformed into five-year averages, and are statistically significant at the 0.6–0.9 quantiles for the data transformed into 10-year averages. The anti-inflationary effect of CBI on inflation is heterogeneous at different levels of inflation. By using one-year, five-year, and 10-year averages, we allow for actual independence and changes in institutional characteristics (Vuletin and Zhu, 2011). Therefore, our results are robust with respect to different unit periods and moderate institutional change.

Finally, we consider different sample periods of the data. Figure 2 plots the quantile regression estimates of TOR on inflation over the periods 1980–2010 and 1990–2010. The solid line represents estimates of the data during 1980–2010, and the dashed line represents estimates of the data during 1990–2010. We find that the anti-inflationary effect of CBI is stronger in middle- and high-inflation episodes and weaker in low-inflation episodes, which confirms the robustness of our results. Furthermore, Franzese (1999) finds that several anti-inflationary factors, such as trade openness and the strength of the financial sector, become stronger and more stable with time, and CBI plays a less important role in restraining inflation. Figure 2 shows that the anti-inflationary effects of CBI during 1990–2010 are lower than those during 1980–2010, which is consistent with the findings of Franzese (1999).

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4.3 The Endogeneity of TOR

We use instrumental variables for the endogeneity problem of TOR. As described before, the average share of the legal term in office that has elapsed is the ratio between the actual and legal duration of a governor's term in office. If the legal term in office of a governor is indefinite or unknown, then the term is specified as the maximum, which is eight years. Following Dreher, Sturm and de Haan (2008), two alternatives of the legal term are (1) the average term in office in the whole sample, and (2) the average legal term in those countries where central bank law specifies a governor's term in office. The former is 3.7 years and the latter is five years. Panels (I) and (II) in Table 4 present TOR estimates with the two alternatives being used. Both panels show that the regression estimates of TOR on inflation are significant in middle- and high-inflation episodes and insignificant in low-inflation episodes. CBI remains more anti-inflationary as inflation becomes higher.

Instead of using instrumental variables to solve the endogeneity problem, Klomp and de Haan (2010) and Vuletin and Zhu (2011) employ a rolling average of TOR over the preceding years to replace the current TOR. Using this method, we do not include current or future turnovers of central bank governors in the calculation of the current value of TOR, so that we avoid reverse causality concerns. We follow Klomp and de Haan (2010) and Vuletin and Zhu (2011) and set the length of the windows equal to four years. Without using the instrumental variable method, we use Koenker's (2004) ordinary quantile regression for the panel data model and within estimation. Panel (III) in Table 4 reports the estimates of TOR and shows that the quantile regression estimates of TOR are positive and increase along with the quantiles. The results are similar to the benchmark results. Thus, the CBI–inflation relationship is robust to various methods used to deal with the endogeneity problem.

5 Conclusions

This paper proposed a fitted value approach to quantile regressions for panel data models to examine the CBI–inflation relationship. With this we attempted to solve the possible endogeneity problem by using the TOR index as a measure of CBI. The econometric method proposed in this paper demonstrates a fruitfully exploitable alternative compromise. Moreover, both the theoretical and empirical literature point out that the anti-inflationary effect of CBI is heterogeneous at different inflation levels. By exploiting an extensive panel data set, the findings in this paper imply that CBI is more anti-inflationary in cases of higher inflation rates. This paper provided an empirical justification for the heterogeneous anti-inflationary characteristics of CBI. The relationship between CBI and inflation becomes more convincing when the panel quantile model with a fitted value approach is used.

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Appendix: List of Countries

Albania†	Algeria* † §	Argentina* † §	Armenia†
Australia* † §	Austria* † §	Bahamas* † §	Bahrain* † §
Barbados* † §	Belarus	Belgium* † §	Belize* † §
Bhutan* † §	Bolivia* † §	Botswana* † §	Brazil* † §
Bulgaria	Burundi* † §	Canada* † §	Central African Rep.* † §
China	Colombia* † §	Congo, Dem. Rep.* † §	Costa Rica* † §
Croatia†	Cyprus* † §	Czech Rep.†	Denmark* † §
Dominican Rep.* † §	Ecuador* † §	Egypt* † §	El Salvador* † §
Equatorial Guinea* † §	Estonia†	Ethiopia* † §	Finland* † §
France* † §	Gambia* † §	Georgia	Germany* †
Ghana* † §	Greece* † §	Guatemala* † §	Guinea-Bissau* † §
Guyana†	Haiti* † §	Honduras* † §	Hungary
Iceland* † §	India* † §	Indonesia* † §	Iran* † §
Ireland	Israel* † §	Italy* † §	Jamaica* †§
Japan* † §	Jordan* † §	Kazakhstan	Kenya* † §
South Korea* † §	Latvia	Lesotho* † §	Libya
Lithuania†	Luxembourg	Madagascar* † §	Malawi* † §
Malaysia* † §	Malta* † §	Mauritius* † §	Mexico* † §
Mongolia†	Morocco* † §	Mozambique* †	Nepal* † §
Netherlands* † §	New Zealand* † §	Nicaragua	Nigeria* † §
Norway* † §	Pakistan* † §	Papua New Guinea* † §	Paraguay* † §
Peru* † §	Philippines* † §	Poland* † §	Portugal* † §
Qatar	Romania* † §	Russian Federation†	Rwanda* †
Saudi Arabia* † §	Singapore* † §	Slovakia†	Slovenia†
Solomon Islands* †	South Africa* † §	Spain* † §	Sri Lanka* † §
Sudan* † §	Suriname* † §	Swaziland* † §	Sweden* † §
Switzerland* † §	Syria* † §	Tanzania* † §	Thailand* † §
Trinidad and Tobago* †	Tunisia* † §	Turkey* † §	Uganda* † §
Ukraine	United States* † §	Uruguay* † §	Yemen, Rep.
Zambia* † §	Zimbabwe* † §		

Note 1: All 118 countries are used in the analysis over the period 1990–2010.

Note 2: * and † indicate the countries used over the period 1974–2010 and 1980–2010, respectively.

Note 3: § indicates the countries used when using the RR classification for the fixed exchange rate regime.

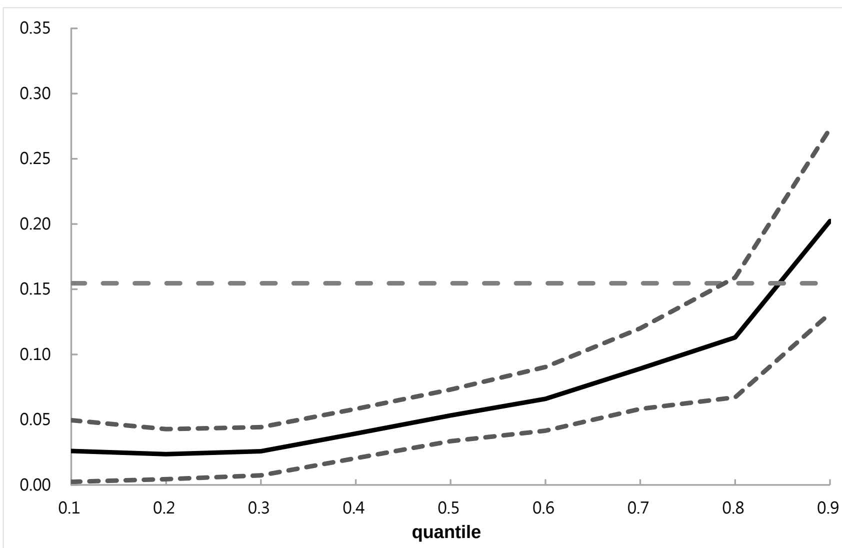

Figure 1: The Impact of TOR on Inflation

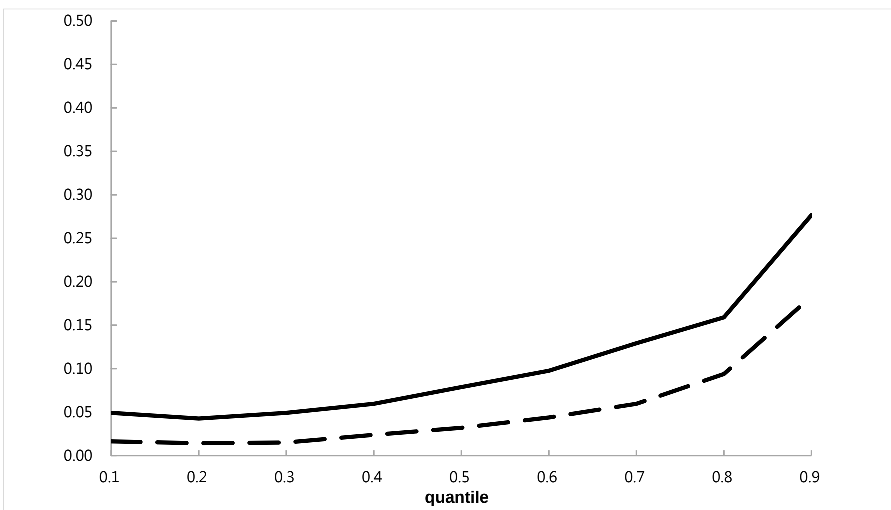

Figure 2: The Impacts of TOR on Inflation for 1980–2010 and 1990–2010

Table 1: Summary Statistics

	Mean	Q1	Median	Q3	S.E.	Min	Max
(A) 93 countries (1974–2010)							
Inflation rate (%)	38.85	3.40	7.99	14.99	504.13	-17.64	23,773.13
TOR	0.21	0.00	0.00	0.00	0.45	0.00	4.00
Openness (%)	72.27	43.23	62.09	88.67	47.89	6.32	460.47
Fixed exchange rate regime dummy	0.43	0.00	0.00	1.00	0.50	0.00	1.00
Political instability	0.19	-0.67	-0.67	0.25	1.66	-0.67	15.40
GDP per capita	6,935.25	594.98	2,125.03	10,749.32	9,156.07	82.67	41,904.21
Coup	0.04	0.00	0.00	0.00	0.22	0.00	4.00
Share of veto players who drop	0.13	0.00	0.00	0.00	0.29	0.00	1.00
Election	0.24	0.00	0.00	0.50	0.33	0.00	1.00
Share of term elapsed	0.65	0.13	0.40	0.88	0.72	0.00	5.60
(B) 104 countries (1980–2010)							
Inflation rate (%)	44.02	2.95	7.00	13.79	534.59	-17.64	23,773.13
TOR	0.20	0.00	0.00	0.00	0.44	0.00	4.00
Openness (%)	76.37	45.86	65.58	95.58	48.90	6.32	460.47
Fixed exchange rate regime dummy	0.40	0.00	0.00	1.00	0.49	0.00	1.00
Political instability	0.12	-0.67	-0.67	0.25	1.53	-0.67	15.40
GDP per capita	6,908.80	627.12	2,171.30	9,757.44	9,216.92	82.67	41,904.21
Coup	0.03	0.00	0.00	0.00	0.19	0.00	4.00
Share of veto players who drop	0.13	0.00	0.00	0.00	0.29	0.00	1.00
Election	0.26	0.00	0.00	0.50	0.33	0.00	1.00
Share of term elapsed	0.64	0.14	0.40	0.88	0.71	0.00	5.60
(C) 118 countries (1990–2010)							
Inflation rate (%)	44.05	2.53	5.74	11.62	550.93	-9.80	23,773.13
TOR	0.19	0.00	0.00	0.00	0.42	0.00	3.00
Openness (%)	82.78	50.95	71.75	101.76	50.11	10.83	460.47
Fixed exchange rate regime dummy	0.40	0.00	0.00	1.00	0.49	0.00	1.00
Political instability	0.03	-0.67	-0.67	0.24	1.38	-0.67	14.87
GDP per capita	7,388.39	728.84	2,359.52	9,649.13	10,247.49	82.67	56,388.99
Coup	0.02	0.00	0.00	0.00	0.15	0.00	2.00
Share of veto players who drop	0.13	0.00	0.00	0.00	0.28	0.00	1.00
Election	0.26	0.00	0.00	0.50	0.33	0.00	1.00
Share of term elapsed	0.64	0.17	0.43	0.88	0.67	0.00	4.83

Sources: Dreher, Sturm and de Haan (2008), World Development Indicators, International Financial Statistics, Cross-National Time-Series Data Archive (2012), Beck et al. (2001), and Powell and Thyne (2011).

Table 2: The Benchmark Results

Dependent variable: inflation

		Quantile								
	Mean	0.1	0.2	0.3	0.4	0.5	0.6	0.7	0.8	0.9
Intercept	0.1015***	0.0065	0.0395***	0.0543***	0.0649***	0.0720***	0.0833***	0.0992***	0.1241***	0.1878***
	(0.0037)	(0.0109)	(0.0071)	(0.0073)	(0.0077)	(0.0082)	(0.0091)	(0.0102)	(0.0132)	(0.0216)
TOR	0.1546***	0.0260	0.0236	0.0258	0.0394**	0.0533***	0.0660***	0.0892***	0.1131**	0.2023***
	(0.0117)	(0.0236)	(0.0192)	(0.0185)	(0.0189)	(0.0198)	(0.0244)	(0.0309)	(0.0459)	(0.0708)
TOR × OECD	-0.1129***	0.0596	0.0026	-0.0127	-0.0381	-0.0493*	-0.0567*	-0.0763*	-0.0962*	-0.1985**
	(0.0240)	(0.0388)	(0.0271)	(0.0243)	(0.0239)	(0.0258)	(0.0314)	(0.0412)	(0.0572)	(0.0937)
Openness	1.13×10^{-4}	2.73×10^{-4}	1.78×10^{-4}	2.29×10^{-4}*	1.59×10^{-4}	1.16×10^{-4}	1.46×10^{-4}	1.30×10^{-4}	6.09×10^{-5}	-2.21×10^{-4}
	(1.94×10^{-4})	(2.58×10^{-4})	(1.29×10^{-4})	(1.18×10^{-4})	(1.25×10^{-4})	(1.14×10^{-4})	(1.31×10^{-4})	(1.42×10^{-4})	(2.90×10^{-4})	(5.32×10^{-4})
Fixed exchange rate	-0.0313***	-0.0016	-0.0104*	-0.0129**	-0.0159***	-0.0163***	-0.0207***	-0.0231***	-0.0288***	-0.0484***
	(0.0051)	(0.0110)	(0.0059)	(0.0054)	(0.0052)	(0.0053)	(0.0062)	(0.0069)	(0.0093)	(0.0177)
Political instability	0.0065***	0.0010	0.0040*	0.0051***	0.0057***	0.0072***	0.0077***	0.0092***	0.0122**	0.0118
	(0.0014)	(0.0033)	(0.0024)	(0.0016)	(0.0016)	(0.0018)	(0.0021)	(0.0030)	(0.0049)	(0.0092)
GDP per capita	-0.2109***	-0.0344	-0.0335	-0.0409	-0.0624	-0.0860*	-0.1109*	-0.1782**	-0.2557**	-0.4850**
	(0.0326)	(0.0508)	(0.0450)	(0.0458)	(0.0466)	(0.0495)	(0.0665)	(0.0858)	(0.1276)	(0.1991)

1. Standard errors are in parentheses. *, **, and *** denote the significance levels of 10%, 5%, and 1%, respectively.

2. Instrumental variables are the number of coups, share of veto players who drop, election, lagged inflation, and share of governor's term in office that has elapsed.

3. The standard errors of the quantile regression estimation are obtained by the bootstrap method. The number of bootstrapping replications is 1,000.

Table 3: The Estimates of TOR

Dependent variable: inflation										
		Quantile								
	Mean	0.1	0.2	0.3	0.4	0.5	0.6	0.7	0.8	0.9
(I) Alternative TOR measure										
TOR	0.1310***	0.0091	0.0123	0.0164	0.0327**	0.0463***	0.0572***	0.0712***	0.0943***	0.1854***
	(0.0130)	(0.0229)	(0.0171)	(0.0161)	(0.0163)	(0.0165)	(0.0200)	(0.0265)	(0.0356)	(0.0593)
(II) RR classification										
TOR	0.1590***	0.0216	0.0154	0.0183	0.0338*	0.0423**	0.0594**	0.0805**	0.1169**	0.1994***
	(0.0128)	(0.0273)	(0.0181)	(0.0182)	(0.0191)	(0.0213)	(0.0242)	(0.0320)	(0.0490)	(0.0759)
(III) Different time unit: five-year averages										
TOR	0.2017***	0.0005	0.0363	0.0661*	0.0899**	0.1077**	0.1296***	0.1305**	0.2053**	0.2974*
	(0.0558)	(0.0443)	(0.0398)	(0.0393)	(0.0432)	(0.0457)	(0.0470)	(0.0546)	(0.0821)	(0.1610)
(IV) Different time unit: 10-year averages										
TOR	0.2877***	0.0041	0.0519	0.0668	0.0669	0.0914	0.1139*	0.1126*	0.1827**	0.3989**
	(0.1057)	(0.0714)	(0.0541)	(0.0558)	(0.0577)	(0.0595)	(0.0645)	(0.0679)	(0.0843)	(0.1928)

1. Standard errors are in parentheses. *, **, and *** denote the significance levels of 10%, 5% and 1%, respectively.
2. Instrumental variables are the number of coups, share of veto players who drop, election, lagged inflation and share of governor's term in office that has elapsed.
3. The standard errors of the quantile regression estimation are obtained by the bootstrap method. The number of bootstrapping replications is 1,000.
4. Because of limitations of space, we report the estimates of TOR. For other estimates of the empirical models, please contact the authors for more information.

Table 4: The Estimates of TOR

Dependent variable: inflation										
		Quantile								
	Mean	0.1	0.2	0.3	0.4	0.5	0.6	0.7	0.8	0.9
(I) Different instrumental variables are used: the maximum legal term is 3.7 years										
TOR	0.1705***	0.0350	0.0223	0.0251	0.0397**	0.0479**	0.0632***	0.0851***	0.1116**	0.2203***
	(0.0119)	(0.0254)	(0.0199)	(0.0189)	(0.0185)	(0.0192)	(0.0239)	(0.0298)	(0.0450)	(0.0689)
(II) Different instrumental variables are used: the maximum legal term is five years										
TOR	0.1588***	0.0288	0.0208	0.0225	0.0382**	0.0478***	0.0637***	0.0872***	0.1077***	0.2000***
	(0.0116)	(0.0225)	(0.0173)	(0.0168)	(0.0167)	(0.0174)	(0.0221)	(0.0282)	(0.0414)	(0.0661)
(III) A rolling average of TOR										
TOR	0.1393***	0.0086	0.0210	0.0315*	0.0495***	0.0588***	0.0787***	0.1165***	0.1675***	0.3034***
	(0.0111)	(0.0251)	(0.0138)	(0.0163)	(0.0147)	(0.0156)	(0.0228)	(0.0354)	(0.0575)	(0.0942)

1. Standard errors are in parentheses. *, **, and *** denote the significance levels of 10%, 5%, and 1%, respectively.

2. Instrumental variables are the number of coups, share of veto players who drop, election, lagged inflation, and share of governor's term in office that has elapsed.

3. The standard errors of the quantile regression estimation are obtained by the bootstrap method. The number of bootstrapping replications is 1,000.

4. Because of limitations of space, we report the estimates of TOR. For other estimates of the empirical models, please contact the authors for more information.

行政院科技部補助國內專家學者出席國際學術會議報告

105 年　10 月 31 日

報告人姓名	林馨怡	服務機構及職稱	國立政治大學經濟學系教授
時間 會議地點	2015 年 12 月 2 日至 12 月 4 日 日本、東京	本會核定補助文號	
會議名稱	2015 The International Symposium on Business and Social Sciences		
發表論文題目	Do Quantitative Monetary Targets Matter?		

報告內容應包括下列各項：

一、參加會議經過

二、

12/2 至會場報到及領取會議相關資料。

12/3 參加不同場次的研討會。

12/4 在 Economics (1)場次主持會議，並於該場次報告本次會議發表之文章。

三、與會心得

1. 分組會議的報告後和與會學者的討論對論文的方向和發表獲得建議。
2. 主持會議時，與來自美國，泰國及中國大陸學者互動及討論。
3. 參加大會其他場次的會議及討論，對其他同領域題目及研究方向有新的靈感。
4. 與其他國家的計量及總體經濟學者認識、交流。
5. 此次會議較屬綜合性會議，大會專題演講題目，並非本人領域，所以無法得到較大啟發。未來本人將選擇參加與本人領域相同的專門性會議。

四、考察參觀活動(無是項活動者省略)

五、建議

感謝科技部補助本人參加此次會議。

六、攜回資料名稱及內容

所有會議相關資料皆可由網路上下載，網址為 http://tisss.org/index.asp?id=22

七、其他

科技部補助計畫衍生研發成果推廣資料表

日期:2016/11/03

科技部補助計畫	計畫名稱: 央行獨立性與通貨膨脹	
	計畫主持人: 林馨怡	
	計畫編號: 104-2410-H-004-009-	學門領域: 數理與數量方法

無研發成果推廣資料

104年度專題研究計畫成果彙整表

計畫主持人：林馨怡				計畫編號：104-2410-H-004-009-		
計畫名稱：央行獨立性與通貨膨脹						
成果項目				量化	單位	質化（說明：各成果項目請附佐證資料或細項說明，如期刊名稱、年份、卷期、起訖頁數、證號...等）
國內	學術性論文	期刊論文		0	篇	
國內	學術性論文	研討會論文		1	篇	參加2016年台灣計量年會並發表。
國內	學術性論文	專書		0	本	
國內	學術性論文	專書論文		0	章	
國內	學術性論文	技術報告		1	篇	本計劃之結案報告。
國內	學術性論文	其他		0	篇	
國內	智慧財產權及成果	專利權	發明專利 申請中	0	件	
國內	智慧財產權及成果	專利權	發明專利 已獲得	0	件	
國內	智慧財產權及成果	專利權	新型/設計專利	0	件	
國內	智慧財產權及成果	商標權		0	件	
國內	智慧財產權及成果	營業秘密		0	件	
國內	智慧財產權及成果	積體電路電路布局權		0	件	
國內	智慧財產權及成果	著作權		0	件	
國內	智慧財產權及成果	品種權		0	件	
國內	智慧財產權及成果	其他		0	件	
國內	技術移轉	件數		0	件	
國內	技術移轉	收入		0	千元	
國外	學術性論文	期刊論文		1	篇	已寫成期刊論文投稿中。
國外	學術性論文	研討會論文		1	篇	參加2015TISSS年會並發表。
國外	學術性論文	專書		0	本	
國外	學術性論文	專書論文		0	章	
國外	學術性論文	技術報告		0	篇	
國外	學術性論文	其他		0	篇	
國外	智慧財產權及成果	專利權	發明專利 申請中	0	件	
國外	智慧財產權及成果	專利權	發明專利 已獲得	0	件	
國外	智慧財產權及成果	專利權	新型/設計專利	0	件	
國外	智慧財產權及成果	商標權		0	件	
國外	智慧財產權及成果	營業秘密		0	件	
國外	智慧財產權及成果	積體電路電路布局權		0	件	
國外	智慧財產權及成果	著作權		0	件	
國外	智慧財產權及成果	品種權		0	件	
國外	智慧財產權及成果	其他		0	件	

	技術移轉	件數	0	件	
		收入	0	千元	
參與計畫人力	本國籍	大專生	0	人次	
		碩士生	3		
		博士生	0		
		博士後研究員	0		
		專任助理	1		
	非本國籍	大專生	0		
		碩士生	0		
		博士生	0		
		博士後研究員	0		
		專任助理	0		
其他成果（無法以量化表達之成果如辦理學術活動、獲得獎項、重要國際合作、研究成果國際影響力及其他協助產業技術發展之具體效益事項等，請以文字敘述填列。）					

科技部補助專題研究計畫成果自評表

請就研究內容與原計畫相符程度、達成預期目標情況、研究成果之學術或應用價值（簡要敘述成果所代表之意義、價值、影響或進一步發展之可能性）、是否適合在學術期刊發表或申請專利、主要發現（簡要敘述成果是否具有政策應用參考價值及具影響公共利益之重大發現）或其他有關價值等，作一綜合評估。

1. 請就研究內容與原計畫相符程度、達成預期目標情況作一綜合評估 ■達成目標 □未達成目標（請說明，以100字為限） 　　□實驗失敗 　　□因故實驗中斷 　　□其他原因 說明：
2. 研究成果在學術期刊發表或申請專利等情形（請於其他欄註明專利及技轉之證號、合約、申請及洽談等詳細資訊） 論文：□已發表　■未發表之文稿　□撰寫中　□無 專利：□已獲得　□申請中　■無 技轉：□已技轉　□洽談中　■無 其他：（以200字為限）
3. 請依學術成就、技術創新、社會影響等方面，評估研究成果之學術或應用價值（簡要敘述成果所代表之意義、價值、影響或進一步發展之可能性，以500字為限） 學術貢獻為針對央行獨立性對通膨的影響，做一完整分析，可用於政策分析。
4. 主要發現 本研究具有政策應用參考價值：■否　□是，建議提供機關 （勾選「是」者，請列舉建議可提供施政參考之業務主管機關） 本研究具影響公共利益之重大發現：■否　□是 說明：（以150字為限）