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Did monetary policy fuel the housing bubble in Ireland?

Cindy MOONS & Kevin HELLINCKX

Department of Economics

Faculty of Economics
And Business

Did Monetary Policy Fuel the Housing Bubble in Ireland?

Cindy Moons[1] and Kevin Hellinckx[2]

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Abstract

This paper provides empirical evidence of the role of the euro in the genesis of the recent Irish financial-economic crisis caused by a regional housing boom and its inevitable bust. By using a Taylor-rule, estimated by Generalized Method of Moments, we measure the appropriateness of the ECB's one-size-fits-all policy rate for the Irish economy. A counterfactual analysis suggests that the Irish interest rate should have been on average 6.5% higher. In addition, using a multivariate housing model, we provide econometric evidence for the causal relationship between the low interest rate and the Irish housing boom. Under an alternative sovereign monetary policy, the average house price would have been 25 to 30 percent lower just before the housing bust in the second quarter of 2007. In addition, it shows that a monetary policy tailored to the needs of the member state is enough to prevent housing prices from dramatically increasing and suggests that it is not necessary to include a lean against house price fluctuations in monetary policy strategies.

Jel Code: E5, F6

***Keywords**:* EMU, Monetary policy, Taylor-rule, Housing bubble

[1] Kuleuven, Campus Brussel (HUB), Department of Economics and Business, Warmoesberg 26, 1000 Brussel, Belgium, cindy.moons@kuleuven.be, 00322-608 14 25.
KULeuven and UA, Department of Economics, Belgium.

2 Kuleuven, Campus Brussel (HUB), Department of Economics and Business, Warmoesberg 26, 1000 Brussel, Belgium, 00322-608 14 25.

1. Introduction

Ireland, once hailed as the Celtic Tiger for its roaring economy since the mid-90s, has experienced a severe economic crisis and subsequent debt crisis. A devastating boom-bust cycle in the Irish property market brought Irish banks into financial trouble. This led, in concurrence with the global financial market turmoil since 2007, to a collapse of the Irish banking system and subsequent government intervention. However, taking over private bank liabilities led to double-digit fiscal deficits and a spectacular increase of the national debt from 24.8% of GDP in 2007 to more than 90% by the end of 2010. Over the same period also nominal GDP shrunk by almost 21% (Eurostat, 2013). Due to the fragile design of the EMU and because the Greek example revealed that default of a euro country was no longer unthinkable, bond yields were pushed up to unsustainable levels (De Grauwe, 2012; De Grauwe and Ji, 2013). The government was ultimately no longer able to refinance its debts and in November 2010 it had to fall back on an emergency loan from the EU and IMF. Also other European countries got into trouble and in the course of 2010, the financial-economic crisis led to a full blown European debt crisis. Anno 2013, Irish unemployment peaks at 14,6% and government debt accounts already for more than 120% of GDP (European Commission, 2013). The consequences of the crisis will resonate for years.

The question is how the Irish economic miracle ended up in a deep recession and why Ireland is one of the European countries that was hit the hardest. In this context the advent of the euro does not remain undiscussed. The theory of optimum currency areas (Mundell, 1961) indicates a crucial consequence of euro membership, namely the delegation of a sovereign monetary policy to a common central bank. A member of the euro area is necessarily consigned to the one-size-fits-all policy from the ECB which only takes euro-area-wide conditions into account. But when economic performance among countries within a currency union diverges significantly, the costs of a common currency, e.g. the Walters effect, can be very high. It is very likely that different euro countries prefer different interest rates. Moons and Van Poeck (2008) show on the basis of a normative Taylor-rule that the monetary policy of the ECB clearly does not fit the needs of all individual euro members. Such inappropriate policy rates can potentially lead to domestic macroeconomic imbalances. So heterogeneity in terms of inflation, economic growth and competitiveness in general, represents a major challenge for the ECB since it has to implement a policy for the euro area as a whole (De Grauwe, 2012; Krugman and Obstfelt, 2009; Wyplosz, 2006). Price dispersions are inevitable, even in a currency area like the US, but it becomes critical when those regional differences are significant and persist over time, as in the euro area (Fendel and Frenkel, 2009). In such an environment, a temporary shock can initiate a mechanism that reinforces itself even when the underlying cause has disappeared (Honohan and Lane, 2003). The consequence is that the ECB's interest rate policy, focused on price stability in the aggregate euro zone, can have asymmetric real effects and can even strengthen national macro-economic imbalances through the well-known Walters critique. The danger is that, when a country is in full economic expansion and experiences a higher than average inflation rate, a too low nominal interest rate leads to a decline in real interest rates. This further stimulates domestic demand and induces additional pro cyclical inflationary pressure and can for example trigger a property bubble and related dislocations. Conversely, countries with relative low inflation experience an unintended contra-cyclical impact (Busetti et al., 2007; Fendel and Frenkel, 2009).

The degree of heterogeneity within the euro zone, with focus on Ireland, is discussed in the following figures. A first dimension to look at is the inflation rate. Figure 1, which presents a Philips curve, clearly illustrates that between 1997 and 2008 peripheral countries like Ireland, Spain, Portugal and Greece persistently had an average inflation or HICP rate that was far higher than the euro-average. Germany, Finland, Austria, France and Belgium on the contrary, formed a cluster of low inflation countries (Busetti et al., 2007). Despite the differences, the ECB succeeded relatively well in keeping the aggregate inflation close to 2%[3].

Between 2000 and 2004 Ireland in particular witnessed a strong rise in inflation and was, with about 4%, twice the euro zone average. The most important explanation for this is the low real interest rate. However exposure to exchange rate volatility, namely the depreciation of the euro against the dollar and the British pound, can explain a significant part of the divergent inflation rates. Given the high degree of openness of the Irish economy and the fact that the percentage of intra-EMU trade in its trade volume is the smallest among all euro countries (Honohan and Lane, 2003), the weakness of the euro in the initial years of the EMU, making exports cheaper and imports more expensive, can partly explain the higher Irish inflation rates compared to the euro average.

The correlation between the cumulative growth of real GDP and the cumulative inflation rate is 0.54. This indicates that besides structural factors cyclical or national demand factors played a role

[3] Note that in 2012, 78% of the HICP was determined by the prices in Germany, France, Italy and Spain. Their weights in the HICP between 1997 and 2012 was always around 75 to 80% while Ireland had only a weight of 1.4% (Eurostat, 2013).

in explaining inflation differentials within the euro area (Lane, 2006). The data presented in figure 1 illustrate that economic cycles aren't synchronized, especially when comparing Ireland with the core countries. Between 1997 and 2008, Irish real GDP growth was persistently two to three percent higher than in Germany and the EMU. Furthermore, Ireland was the first to be hit by the crisis, in the first quarter of 2008, while the euro area as a whole was hit one year later. Again, this illustrates the non-synchronization of European business cycles.

A third dimension to look at, crucial for the further analysis in this paper, is the divergent trends in the evolution of the growth of domestic credit and house prices. Figure 2 reveals that the yearly average growth rate of nominal credit, provided by the domestic banking sector, and the yearly change in real house prices have a positive correlation of 0.58. It is striking that Ireland and Spain experienced the biggest growth in credit and house prices and were hit the hardest by the recent crisis (De Grauwe, 2010). Besides that it is also noteworthy that the yearly change in real house prices differs significantly across euro countries. Since welfare is strongly related to property, which is mainly held by local people, it can probably partly explain differences in national consumption levels and private credit accumulation (Lane, 2006).

Furthermore, due to the diverging macro-economic performances and the absence of a national currency, Ireland witnessed between 2003 and 2008 a real appreciation of almost 20%. Domestic prices rose faster than abroad which led to an overshooting effect through the well-known price-wage spiral (Honohan and Lane, 2003) and a loss of competitiveness as a result. So it is no surprise that after 2002 current accounts began to deteriorate. This already indicated that Irish people lived beyond their means, consumed too much and borrowed money from abroad.

Given the above it is unlikely that the ECB's monetary policy was tailored to the needs of the small Irish economy. Many authors like De Grauwe (2010), Honohan (2010), and Taylor (2009) therefore postulate that EMU-membership led to very low, inappropriate Irish interest rates and was the main cause of the Irish credit, consumption and property boom and subsequent bust. However, comprehensive and clear econometric evidence for the precise role and impact of monetary policy rates on Irish property prices and related macro-economic imbalances is still lacking. Taylor (2007 and 2009) shows, using descriptive data, that housing booms in European countries were largest where the deviation of the short term interest rate from the standard Taylor-rule were largest. Similar, De Grauwe (2010) suggests a positive relation between bank credit and rising real house prices. But those analyses do not provide real causal or econometric evidence. Studies that empirically estimate an interest rate reaction function and perform a counterfactual to determine the difference between a benchmark or optimal interest rate and the actual one, like the papers of Hayo (2006) and, Honohan and Leddin (2006), are scarce. Other studies attempt to provide econometric evidence of the impact of monetary policy on property prices. Seyfried (2010), for example, estimates a multivariate regression model of a standard house price model and finds that between 2001 and 2007 the average Irish house price would have increased by only 4.2% per annum while actual prices rose about 9.8%. However, a normative Taylor-rule is used without econometric verification as to whether it can explain the ECB's monetary policy well or not.

The fact is that anno 2013, the euro and the Irish economy are still under pressure. So it is crucial to assess the costs and benefits of euro membership and understand the role of the euro in the genesis of the crisis. Because it is likely that it was not only interest rates that played a role, it is important to come up with clear empirical evidence concerning the causal relationship between euro membership, monetary policy and the property bubble. Therefore this paper seeks to answer the following questions:

1. To what extent was the policy of the ECB in the period leading up to the crisis tailored to the needs of the Irish economy?
2. What relation exists between the monetary policy of the ECB and the genesis of the Irish (property) bubble?
3. How could a sovereign monetary policy, tailored to Irish economic needs, have mitigated the effects of the Irish property bubble and subsequent crisis?

In order to answer these questions and determine whether or not it was a coincidence that Ireland's economic fundamentals began to deteriorate when it joined the euro area, this paper starts with section 2, presenting the analytical framework used to evaluate the appropriateness of the ECB's monetary policy for the Irish economy. Next, we present the empirical GMM estimates of the Taylor-rule for the ECB and Ireland. These results are further used in a counterfactual analysis for the period 1994 - 2012, which allows us to assess the extent to which interest rates were inappropriate given Irish economic conditions. In section 3 we determine the link between the monetary policy and the housing market by deriving a multivariate housing model, discuss the OLS estimates and policy implications. Finally, section 4 summarizes the main findings and concludes.

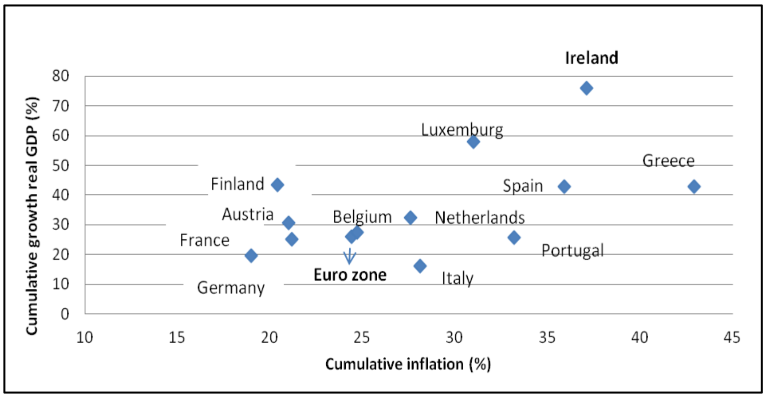

Fig. 1. Philips curve euro countries (1997 – 2008)[4].

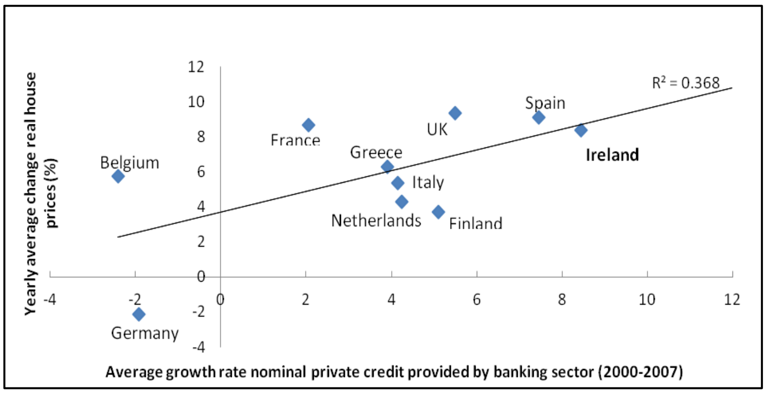

Fig. 2. Change in real house prices and growth nominal credit as percentage of GDP.Specification Taylor-rule, estimation and counterfactual

2. Specification Taylor-rule, estimation and counterfactual

In order to evaluate the extent to which monetary conditions were appropriate for the Irish economy, we need a benchmark that empirically characterizes how the ECB conducts monetary policy. Therefore we present a forward-looking policy reaction function following Clarida, Gali and Gertler (1998). The idea is that a central bank determines its interest rates using information it really possesses, namely predictions of inflation and output gap rather than purely past inflation and output data:

$$i_t = (1-\rho).(\alpha + \beta.E[\pi_{t+n}|I_t] + \gamma.E[y_{t+k}|I_t) + \rho.i_{t-1} + \varepsilon_t \quad (1)$$

Where i_t represents the current interest rate, α is a constant while β and γ represent the long term weights for respectively the inflation rate (π_t) and the output gap (y_t). So the parameters β and γ

[4] Source Eurostat (2013) and OECD (2013).

measure the extent to which the central bank reacts to changing economic conditions. In this way it is possible to quantify the trade-off between the inflation and the output gap in a reaction function. The magnitude of the parameter β is key. If β is bigger than one, the target real interest rate is raised in response to inflationary pressures and is known as the Taylor-principle which is a precondition for a stabilizing monetary policy. *E* represents the rational expectations operator[5], conditional on all relevant information available to policymakers (I_t). The indices *t+n* and *t+k* stand for the number of periods looking forward[6]. Therefore, π_{t+n} represents the annualized inflation rate at point of time *t+n*. Furthermore, ρ represents the smoothing parameter or the weighting factor with a value between zero and one. The size of ρ indicates how fast the central bank reacts to a change in economic conditions. The bigger the parameter, the more gradual the reaction. The econometric reasons in favour of using a lagged interest rate are that the parameter is statistically highly significant, that it improves the model fit and makes it possible to deal with serial correlation (Belke and Polleit, 2007; Gerdesmeier and Roffia, 2005; Sauer and Sturm, 2003). Castelnuovo (2007) shows that it is preferable to model a lagged interest rate instead of using an autoregressive error specification, implicitly providing evidence that the ECB intentionally smoothes its interest rate policy. Also in practice there are several reasons why a central banker probably prefers partial adjustment. Firstly, avoiding significant and discrete interest rate jumps lowers the risk of a shock reaction on financial (bond) markets, bearing in mind that financial stability is crucial for macro-economic stability. Secondly, it implies transparency and consistency. The idea is that private actors form expectations. If policy becomes predictable and actors expect that an initial change will be followed by additional interest rate changes in the same direction, the impact on inflation and output of a modest interest rate change will be bigger. So the effectiveness of monetary policy is strengthened by, and dependent on, the impact of market expectations about the future path of the short term interest rate, hence influencing long term interest rates (Castelnuovo, 2007; Cecchetti, 2000; Woodford, 1999). A third reason is that policy makers do not possess perfect information on the state of the economy when taking policy decisions. There is also uncertainty over the precise impact and timing of a change in interest rates on the real economy. So both arguments justify a gradual, prudent reaction. It buys policy makers some time to assess or redirect policy actions (Cecchetti, 2000; Orphanides, 2003).

[5] The rational expectations theory assumes that predictions are not systematically wrong (Clarida et al., 1998).

[6] If *n* and *k* are zero, the model reduces to the contemporaneous model. If *n* and *k* are negative, it represents a backward looking specification. But this form is both theoretical and empirically rejected. Authors like e.g. Fendel and Frenkel (2009), and Castro (2008) prefer *k* = 12; *n* = 3.

Finally, if the non-observed variables are replaced by the actual, realized variables, we get an estimable forward looking specification[7]:

$$i_t = (1-\rho).(\alpha + \beta.\pi_{t+n} + \gamma.y_{t+k}) + \rho.i_{t-1} + \varepsilon_t \qquad (2)$$

Most empirical literature shows evidence of such forward looking behavior. The idea is that there is a time lag in the monetary transmission mechanism and as a consequence policy makers have to anticipate[8] (ECB, 2011; Gerdesmeier and Roffia, 2003). The central bank cannot wait until higher inflation actually appears since rising price levels can work through the price-wage spiral, making it more difficult to maintain price stability (Orphanides, 2003). And again, it takes time to gather or measure the necessary data. That's why it is reasonable to assume that a central government takes its decisions on the basis of predictions (Fourçans and Vranceanu, 2004; Orphanides, 2003).

In order to estimate the Taylor-rule we use quarterly data on the short term interest rate, the inflation rate and the output gap for the period from 1994 to 2012. Since the EMU officially only started in 1999, prior data relate to a hypothetical euro area. However this approach allows us to enlarge the sample range and improves the reliability of the estimates. In addition, the monetary policy of the euro countries was already aligned since they had to fulfill the Maastricht-criteria. Moreover, national central bankers already coordinated their policy through the European Monetary Institute, set-up in 1994 in order to smooth the transition towards an integrated system of central banks (EMI, 1997). As a result, such an empirical approach is not unusual (Gerdesmeier and Roffia 2003; Gerlach and Schnabel, 2000; Sauer and Sturm, 2003).

For the short term interest rate we use the EONIA retrieved from the ECB database and transformed from monthly to quarterly frequency. To measure inflation, in accordance with the quantitative definition of price stability of the ECB, we use the Harmonised Index of Consumer Prices (HICP) for the aggregate euro zone, and Ireland. Data are ex post, without seasonal correction and retrieved from Eurostat and the ECB database.[9] Monthly data are converted into

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[7] We assume that realized values are a proxy for the expected values. However using ex post data instead of real time data can be criticized (Gerdesmeier and Roffia, 2003; Orphanides, 2003). An alternative is survey data (Sauer and Sturm, 2003).

[8] Output mostly reacts faster than inflation. Clarida et al. (1998) argue that the time between a rate change and impact on output is about six to nine months while for inflation about one year. For EMU, Fourçans and Vranceanu (2004) talk about a lag of six to twelve months.

[9] Data represent the euro 17 group while the data for the output gap only covers the euro 15 group. But because the weight of Slovakia and Estonia in the HICP is only 0.73% and 0.15%, this does not influence the results significantly (Eurostat, 2013).

quarterly data by taking the average. As a proxy for the output gap we use OECD data. The output gap is measured as the percentage difference between real GDP and the estimated potential GDP[10].

A major estimation issue with using Ordinary Least Squares (OLS) techniques to estimate Taylor-rules is that the variables may be correlated with the error term while endogeneity implies biased OLS estimates. Endogenous variables can be detected, using the Hausman specification test. Table 2 reports the p-value of the Hausman test. Since the null hypothesis, which states that OLS results are not biased, is rejected on a 5% significance level, there are clear signs that the interest rate is endogenously determined by the inflation and the output gap. So despite the seemingly good data fit with a high R^2, OLS estimation results are biased and inconsistent which would lead to wrong conclusions. Therefore we must rely on instrumental estimation methods, Two-Stage Least Squares or GMM, which in such circumstances are able to deliver unbiased and consistent estimation results (Carter et al., 2012; Gerdesmeier and Roffia, 2005). However, the tests, reported in the note of table 2, namely the Jarque-Bera test for normality, the White test for homoskedasticity and the Langrange-Multiplier test for serial correlation, all indicate the violation of the necessary assumptions for OLS concerning the structure of the data and the error term. So using OLS causes problems with respect to statistical estimation and inference. Therefore, in line with most empirical literature and taking the advantages into account, we use the Generalized Method of Moments instead of Two-Stage Least Squares. Since 2SLS is based on the OLS technique, it is preferred that the error term is homoscedastic, does not show serial correlation and follows a normal distribution. In large samples the normality assumption may not be a reason for concern. However in this context our 76 observations may not be considered a large sample. GMM, on the contrary, nests many common estimators and has the advantage that it accounts for endogeneity biases as well as non-spherical errors, hence requiring no information about the exact distribution of the error term. All that is required is that the orthogonality condition holds and that the variables are stationary. Furthermore, in contrast to 2SLS, the GMM method does not require that the abundant instruments are reduced in a linear combination. So when over-identification is present GMM will deliver more accurate estimation results, hence producing smaller standard errors and making statistical tests more powerful. However, if normality is fulfilled, GMM reduces to the 2SLS estimates. The only disadvantage is that GMM may perform poorly in small samples when using a lot of instruments (Belke and Polleit, 2007; Florens et al., 2004; Gerdesmeier and Roffia, 2003).

[10] Estimated using a Cobb-Douglas production function approach.

In accordance with the literature the GMM weighting matrix is chosen using the method of Newey and West. This delivers results that are consistent in the presence of heteroscedasticity and auto-correlation of unknown form. The heteroscedastic and auto-correlation covariance approach does not change the point estimates and only results in a correction of the standard errors just like the Newey-West correction with OLS or 2SLS (Baum, 2007; Gerdesmeier and Roffia, 2005). Notwithstanding, we will use 2SLS using the same set of instruments as a robustness check. The reason for this is that GMM estimates can be biased and less efficient when having a small sample and using a relative high number of instruments. Furthermore, cross-checking the results allows us to assess the sensitivity of the results to the chosen estimation technique (Baum, 2003; Florens et al. 2004, Gerdesmeier and Roffia, 2003).

An important estimation challenge in this context is the choice of instruments, since weak or bad instruments lead to weak, unreliable estimates. Lagged values of the explanatory variables are, however, natural candidates. In accordance with the common literature we use lagged values of inflation, the output gap and the interest rate. Specifically for Ireland, we also use the Irish average house prices since we expect it is very likely to be a good predictor of future inflation and output gap (Belke and Polleit, 2007; Carter et al., 2012; Florens et al., 2004).

In the Taylor literature it is often assumed that the variables are stationary without econometric verification. However, in order to get instrumental estimation results with good asymptotic properties and to avoid a spurious regression it is crucial that the variables are stationary (Hall, 2009; Kitamura and Philips, 1997). To be able to exclude non-stationarity we use the ADF or Augmented Dickey Fuller test. The null hypothesis postulates a unit root. If rejected, it gives an indication that the time series follows a stationary process (Carter et al., 2012). An additional test, that in small samples may give a better view of the likelihood of having stationarity around a constant or a trend, is the KPSS test (Gorter et al., 2007). The null hypothesis postulates stationary time series against the alternative of non-stationarity (Kwiatkowski et al., 1992).

For the whole sample covering the period from the first quarter of 1994 till the fourth quarter of 2012, the ADF test clearly rejects a unit root indicating that the Eonia, the HICP and Output gap are stationary. Table 1 represents the test-results together with the appropriate critical values. The results are validated by the KPSS test.

Table 1 ADF test period 1994Q1 - 2012Q4, EONIA and euro average HICP and output gap.

Variable	Number of lags	Test-statistic	Critical values			Order integration	P-value LM test
			1%	5%	10%		
EONIA[a]	1	-3.47	-3.96	-3.41	-3.13	I(0)**	0.71
HICP[b]	4	-2.96	-3.43	-2.86	-2.57	I(0)**	0.46
Output gap[c]	1	-2.70	-3.43	-2.86	-2.57	I(0)*	0.82

Note: ADF null hypothesis (H0) is a unit root. Rejection H0 if test statistic is smaller than critical value. Number of lags needed to avoid serial correlation in the ADF test reported in the second column. Smallest p-value Q-statistic for the test of absence serial correlation: EONIA 0.74; HICP 0.46 and Output gap 0.54. [a]Test with trend. ** significance at 5% *, significance at 10%.

Since the variables are stationary we can safely estimate equation (2). After model selection[11] we decided to continue with the forward-looking specification with n = 2 and k = 1. So for inflation we follow Fourçans and Vranceanu (2004), and Gerdesmeier and Roffia (2003), looking two quarters ahead. For the output gap, we follow Castro (2008) and take one lead since it seems very likely that the ECB also takes predictions of this variable into account.

Table 2 summarizes the estimation results of the interest rate reaction function and the major statistical tests. The parameters alfa, beta and gamma, the coefficients or the long term weights, are deduced from: $\vartheta_0 = (1-\rho)\,.\alpha$, $\vartheta_1 = (1-\rho).\ \beta$ en $\vartheta_2 = (1-\rho).\ \gamma$, with ϑ_x being the parameter regression output. The instruments used to estimate the Taylor-equation (2) are a constant, the EONIA lagged with two and three quarters and the contemporaneous or actual HICP and output gap. This choice of instruments is in accordance with Gorter et al. (2008), and Gerlach and Schnabel (2000). For the detailed regression output we refer to appendix A.

A first finding is that all GMM coefficients are statistically significant at the 99% confidence level. The smoothing parameter (ρ) is relatively high but in line with most empirical literature. Also the coefficients for the inflation (β) and the output gap (γ) have the expected sign and a value in accordance with some earlier findings (Castro, 2008; Fourçans and Vranceanu, 2007; Gerdesmeier and Roffia, 2005; Gorter et al., 2008).

[11] We estimated three models: a contemporaneous one, one with n = 2 and k =0 and with n = 2 and k = 1. But it turned out that the last model resulted in the best model fit. The GMM criterion and the squared residuals were the smallest. Besides that, the strength of instruments was the highest and the number of needed instruments was the smallest, reducing the risk of having biased GMM estimates. Finally, for this specification, the Cusum test for parameter stability was most convincing.

Table 2 Estimation results forward-looking Taylor-rule ECB for the period 1994 – 2012, using GMM and 2SLS.

	Constante (α)	π_{t+2} (β)	y_{t+1} (γ)	i_{t-1} (ρ)	J-test (p-value)	Sargan test	Cragg-Donald	Hausman test
GMM[a]	-1.34	2.09***	1.14***	0.90***	0.91	-	9.84	0.001
2SLS[b]	-1.18	2.02*	1.14 ***	0.90***	-	0.91	9.84	0.001

Note: all residuals follow a normal distribution. [a] GMM criterion = 0.0002; average squared residuals static prediction = 0.27, dynamic prediction = 0.84. No residual bigger than two and a half standard errors. [b] Test for ARCH-effect (of order 1): $Chi^2(1)$ p-value = 0.42. Jarque-Bera test for normality has $Chi^2(2)$ p-value of 0.16. The Pesaron-Taylor test for heteroscedasticity has p-value of 0.02. Godfrey- test, p-value = 0.0007. * significance at 10%, ** significance at 5%, *** significance at 1%.

Parameter β indicates that if inflation increases by 1%, the ECB, ceteris paribus, increases the interest rate by on average 2.09%. Furthermore, the ECB also seems to take economic activity into account since the Eonia is increased by on average 1.14% if the output gap rises with 1%. The smoothing parameter equals 0.90 and implies that the ECB only gradually adapts the interest rate to the desired level. In one quarter the central bank only implements 10% of the targeted interest rate change[12]. Finally, a joint Wald-test[13], which assesses the validity of the joint restriction (β=1.5; γ=0.5) on the coefficients of inflation and the output gap, indicates that the estimated parameters for the ECB significantly differ from the values suggested by Taylor (1993).

As argued earlier, crucial for the validity of the model with IV is the quality of the instruments. In order to test instrument validity table 2 reports the J-statistic which tests the validity of over-identifying restrictions, i.e. when the number of instruments is greater than the number of parameters to be estimated, with the joint hypothesis that the instruments are orthogonal to the error term and that the estimated model is correctly specified. If the hypothesis of non-misspecification can be rejected, the validity of the IV model can be questioned (Gerdesmeier and Roffia, 2003; Hall, 2009). Since the p-value of the test statistic is very high, the null hypothesis is not rejected. The conclusion is that there is a strong indication for the validity of the chosen instruments and proper model specification.

Secondly, instruments should be highly correlated with the instrumented variable. To assess the strength of the chosen instruments we apply the Cragg-Donald F-test proposed by Stock and Yogo (2002). Weak instruments could potentially lead to inaccurate, substantially biased estimates and

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[12] The interpretation can be derived as follows. Remember that $i_t = \rho . i_t - 1 + (1- \rho) . i_t*$ with i_t* the targeted Taylor or policy rate. This expression can be rewritten as: $i_t - i_t -1 = (1-\rho) . (i_t* - i_t - 1)$. The left side of the equation expresses the interest rate change so $1-\rho$ represents the proportion of the difference between the actual and target rate that is being implemented in one quarter.

[13] The Wald-test statistic follows a $\chi^2_{(q)}$ distribution with the number of degrees of freedom equal to the number of hypotheses being tested (Carter et al., 2012).

unreliable test statistics (Carter et al., 2012). Since there are multiple endogeneous variables, the test is carried out for two alternatives. Firstly, there is a test for the maximum relative bias. Using weak instruments can lead to substantially biased estimation results relative to OLS. A high relative bias means that, even in the presence of endogeneity, OLS is preferred to IV estimation methods. Secondly, there are consequences for the rejection rate of hypothesis tests related to the estimated coefficients. When testing the hypothesis with endogenous variables and using weak instruments, the null hypothesis will be rejected more often (the test size) than one would expect using a certain confidence level. The test assesses whether a maximum test size of a certain rejection rate can be rejected. If the Cragg-Donald Wald statistic is smaller than a certain critical level the conclusion is that the instruments are weak (Carter et al., 2012). The test results are presented in table 2 and are based on the first stage regression in the 2SLS procedure. Appropriate critical values are retrieved from Stock and Yogo (2002).

Since the Cragg-Donald test statistic equals 9.84, we conclude that the relative bias is smaller than 5% while also the hypothesis of a test size of 15% can be rejected. The conclusion is that the instruments are relatively strong, making estimates reliable.

In order to cross-check the results we also estimated the same equation, using the same set of instruments, by 2SLS. The results are presented in table 2. Firstly, the Sargan test, which tests the validity of the set of instruments and is equivalent to the Hansen J-test for GMM, clearly confirms the validity of the instruments. Furthermore, normality cannot be reject on a 5% significance level since the p-value is bigger than 0.05. However, the Pesaran-Taylor test for homoscedasticity and the Godfrey-LM test for first order serial correlation reveal the non-spherical character of the residuals. So GMM probably will deliver more accurate estimates which is confirmed by the smaller standard errors. Notwithstanding, we conclude that the results of the 2SLS procedure are in line with those of GMM which makes us think that the GMM estimates are robust and valid. The results further indicate the superiority of IV estimation methods.

Finally, with regard to parameter stability we carried out the Cusum test of the residuals. Plotting recursive residuals against two critical lines helps to assess the presence of parameter instability of the constant term (Baltagi, 2008; Gerdesmeier and Roffia, 2003). Since there are no movements outside the critical lines, the parameters are relatively stable over time (see appendix B).

Using the estimation results it is possible to discuss the economic relevance of the ECB's one-size-fits-all policy and to perform a counterfactual analysis. The idea is that by simulating a counterfactual scenario we can determine what the interest path would have been if the Irish central bank was able to pursue an independent monetary policy and followed a similar reaction function as the ECB.

Figure 3 plots the actual interest rate, namely the quarterly average Eonia, and the Taylor-rate modeled or predicted by the econometric model presented in table 2. So the figure reveals that the estimated Taylor-rule can track the actual monetary policy remarkably well. It confirms that the Taylor-framework is a good methodology for analysing the behavior of the European Central Bank.

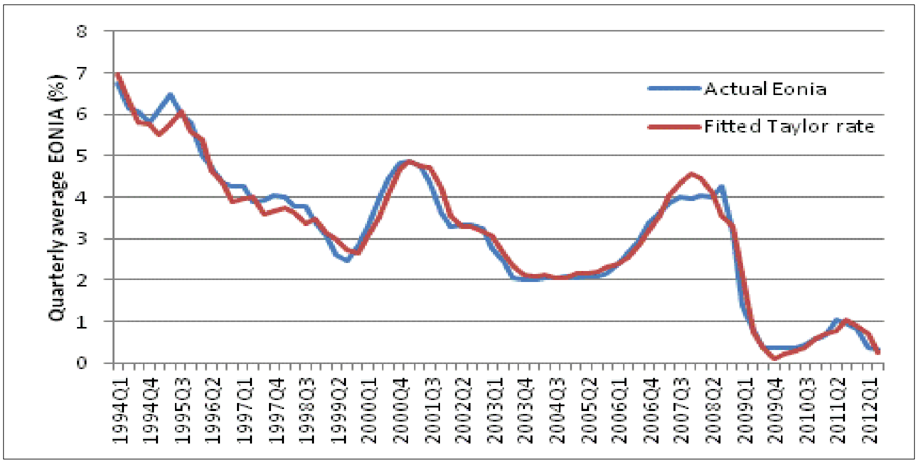

Fig. 3. Actual Eonia and the Eonia modeled by the Taylor-rule for the ECB (data fit).

A rise in expected annual inflation of one percent induces the ECB, ceteris paribus, to increase the interest rate by on average 2.09%. Since β is bigger than one, the estimated reaction function statistically confirms the prediction that the central bank increases real interest rates in response to inflationary pressures. This implies that the past policy of the ECB clearly focused on maintaining price stability in the euro zone as a whole. Also the estimate of the coefficient on the output gap is significant and positive, providing evidence that the ECB also takes economic activity into account. Holding expected inflation constant, a one percent rise in the output gap induces the ECB to increase nominal interest rates by on average 1.14%. Another possible explanation for the output gap being statistically significant is that it only fulfills the role of capturing signals of future price evolutions. However, the fact that this variable is significant in a forward looking fashion makes it very likely that the ECB reacts to the real economy independently of its concern about inflation. This conception is confirmed by the statutes of the ECB since these state that besides safeguarding price stability it also has to support the general economic policy within the euro area (ECB, 2011). So the results provide clear evidence that the ECB conducted a theoretically sound leaning-against-the-wind-policy. It implements an expansionary monetary policy if economic activity is slowing down and conversely, a restrictive policy if economic activity is booming. Furthermore, the coefficients are jointly statistically different from the values proposed by Taylor (1993) meaning that the ECB reacted more aggressively to rising or declining inflation and changes in the output gap than predicted by a normative Taylor-rule.

However, despite the fact that the ECB followed a policy that theoretically fulfills the precondition of a sound policy which is able to stabilize the economy in the aggregate euro zone, since the Taylor-rule was estimated on the basis of weighted euro average data, this does not imply that interest rates were appropriate for every individual euro country. Taking monetary policy decisions based on aggregate developments in the euro area conceals diverse developments at national level.

Using the estimated Taylor-rule for the ECB we can assess the extent to which monetary policy was tailored to the needs of the Irish economy. By taking the estimated model and coefficients as starting point, and replacing the euro average data by the Irish data of inflation and output gap, we can determine an alternative interest rate path under the assumption that Ireland was able to conduct a sovereign monetary policy. The results of those counterfactual simulations for the period between 1994 and 2012 are presented in figure 4. As one can see, there is a huge difference between the actual short term interest rate and what the Taylor-rule prescribes the Irish rate had to be, if the ECB took Irish conditions into consideration for setting its policy. From the first quarter in 1998 till 2007, the average gap or difference between the Irish Taylor-rate and the actual rate, expressed in absolute values or percentage points, was on average 6.67%. More specifically, between 1999 and 2006 the gap was on average 6.72%. It is no coincidence that this period coincided with fast rising Irish house prices, increasingly deviating from their fundamental value.

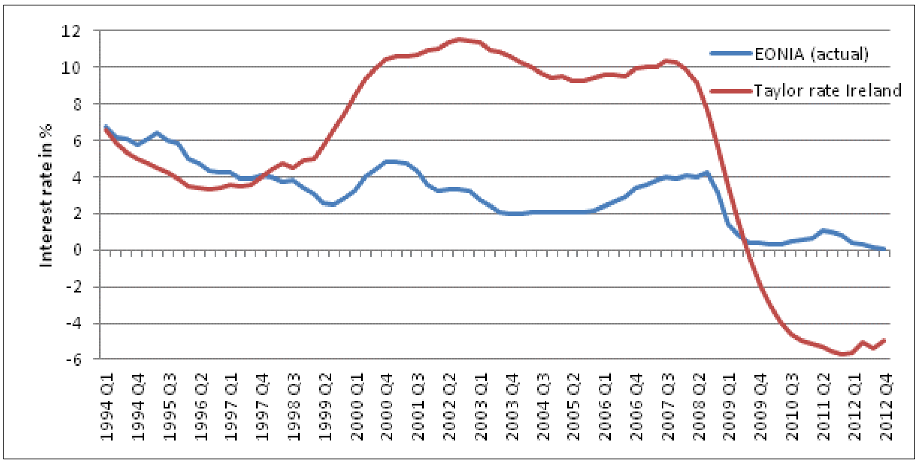

Fig. 4. Actual short term interest rate against the Irish Taylor-rate (1994 – 2012).

These findings are confirmed by an alternative approach, since comparing the estimated coefficients of the Irish Taylor-rate, presented in table 3, with those for the ECB yields the same conclusions. Table 3 reports the GMM estimation results of the Taylor-rule for Ireland and is supplemented by the 2SLS results as a robustness check[14]. The Hausman test reveals endogeneity, favouring IV estimation methods. The J-test and Sargan test both confirm the validity of the instruments and proper model specification. The test of Stock and Yogo (2002) indicates that the relative bias is probably smaller than 5% while the test size is smaller than 20% indicating that the instruments are relatively strong. In this case, in contrast to the ECB estimates, the 2SLS and GMM results differ significantly, a fact which can be explained by the non-spherical and non-normal properties of the residuals, making GMM superior. The coefficients have the expected sign and fulfill the Taylor-principle. However, performing a Wald-test shows that with 95% confidence we can state that the coefficients are statistically smaller than the values proposed by Taylor (1993). So the interest rate was less reactive to Irish inflation and output gap than prescribed by a normative Taylor-rule.

[14] Tests on stationarity can be found in appendix C.

Table 3 The Irish Taylor-rule.

	Constante (α)	π_{t+2} (β)	γ_{t+1} (γ)	i_{t-1} (ρ)	J-test (p-waarde)	Sargan-test	Cragg-Donald	Hausman test
GMM[a]	-1.10	1.34 **	0.31*	0.92***	0.25	-	9.76	-
2SLS[b]	0.91	0.51	0.26*	0.90***	-	0.11	9.76	0.001

Note: results using equation 2 and the following instruments: the Eonia lagged by two, three and four quarters, the actual HICP, the actual output gap and the annual percentage change of the national average house prices [a] GMM criterion: Q = 0.056, average squared residuals static forecast = 0.34, dynamic forecast = 1.42. No error term exceeding two and a half standard deviations. [b] Test for ARCH-effect (1): $Chi^2(1)$ p-value = 0.17. The Jarque-Bera test for normality: p-value = 0.0048. The Pesaran-Taylor test homoskedasticity, p-value = 0.12. Godfrey test (1994) for first order auto-correlation, p-value = 0.0001. * significance at 10%,** significance at 5%, *** significance at 1%.

Since the Irish coefficients for inflation and especially the output gap are noticeably smaller, the view that one size does not fit all is statistically confirmed. So because Irish macro-economic conditions deviate from the average in the euro zone while Ireland is inevitably confined to the monetary policy of the ECB, monetary conditions were far too loose from an Irish perspective. This is the logical result of the fact that Ireland only represents a small weight in aggregate euro zone data, determining 1.4% of aggregate HICP, while its economic performance pertinently diverges from the core countries. Between 1998 and 2004 both the inflation and output gap were far higher than in other euro countries while after 2004 it was especially the output gap that was out of line. This explains why since 1999, the ECB's interest rate was clearly too low for Ireland. Such inappropriate interest rates can be a serious threat to macro-economic stability. Therefore, in the next section, we will assess the link between the interest rate and house prices. The reason is that the unprecedented boom in the housing market was diagnosed to be the main cause of several macro-economic imbalances underlying the Irish bust.

3. The Irish housing bubble

Between 1994 and 2007, the national average Irish house price of residential new and second-hand houses rose from about € 70,000 to respectively € 330,000 and € 386,000. As a result, in the past 15 years Ireland has witnessed the strongest rise in house prices among all OECD-countries. But after house prices reached their peak in the second quarter of 2007, prices fell dramatically by 35% to 45% by 2012 (Department of the Environment, Community and Local Government, 2013). The property crash underlay the Irish crisis while previous analysis proved that the interest rate was far too low. Since the demand for housing is partly driven by mortgage interest rates, which are connected to the Eonia through the monetary transmission mechanism, we develop an econometric model to empirically measure the precise impact of the monetary policy of the ECB on Irish house prices. So the main goal of the model is to measure the influence of monetary policy on house prices while controlling for the impact of other variables. By performing a counterfactual analysis, combining the Irish Taylor-rate with the estimated house price model, we can show the extent to which a monetary policy that was better tailored to the need of the Irish economy could have mitigated the housing boom.

In the literature there is consensus that Irish house prices, the result of supply and demand, were driven by the strong economic growth, low unemployment, increased welfare, the demographic evolution, i.e. the strong population growth, and the too low (real) interest rates (Central Bank of Ireland, 2007; McQuinn and O'Reilly, 2008; Miles and Pillonca, 2008; Rae and Van den Noord, 2006; Stevenson, 2008). Other factors that probably played a role in the demand for credit and hence the demand for housing are the liberalization of financial markets and fiscal encouragements to invest in property (Honjo et al., 2004).

The more prosperous people are, the greater the level of demand since they have more financial resources for financing a home. When the population grows and there is a lot of immigration, it seems logical that demand for housing goes up. Both factors probably explain a big part of the witnessed rise in Irish house prices. However, demand for property is also driven by the affordability and user cost, that, in turn depends on the mortgage interest rate, credit conditions, property taxes and an opportunity cost like the interest rate on a savings account (Miles and Pillonca, 2008; Rae and Van den Noord, 2006).

All these facts provide the building blocks for setting up a model that can explain the evolution of Irish house prices since 1994. In the empirical literature the starting point for such a model is often an inverted demand function, assuming that demand for property is driven by disposable income, interest rate levels, the composition of the population and the price of property (Conefrey and Fitz Gerald, 2010; Duffy et al., 2005; Honjo et al., 2004; Rae and Van den Noord, 2006; Seyfried, 2010; Stevenson, 2008). So a conceptual model usually takes the form of:

$$Q = HSTOCK = f(P, X) \tag{3}$$

where *X* represents a vector with variables having an influence on the demand, inverting yields:

$$P = g(HSTOCK, X) \quad (4)$$

Using the natural logarithm leads to a general log-linear econometric model with ε_t as error term:

$$\ln(P_t) = \alpha + \beta_1.\ln(Y_t) + \beta_2.DEM_t + \beta_3.R_t + \beta_4.\ln(HSTOCK_t) + \varepsilon_t \quad (5)$$

A log-lineair regression is chosen because the variables potentially have a non-linear relationship[15]. Furthermore, using the log rescales the variables and reduces problems with heteroscedasticity, hence improving estimation accuracy while it makes it possible to interpret the coefficients as an elasticity or percentage change[16] (Carter et al., 2012; Verbeek, 2004). P_t represents the house price while Y_t is a measure of disposable income or level of welfare. R_t is an interest rate and $HSTOCK_t$ stands for the stock of housing which can be expressed per capita and is a measure of the supply. The variable DEM_t finally, is a proxy for the demographic evolution and can be the total population or, for example, the proportion of people aged between 25 and 44 which is often considered as the most important group of people buying a house (Roche, 2003). It is important to note that the housing stock variable is often excluded from empirical models, an exclusion that makes the assumption that supply is relatively inelastic, because it is often found to be statistically not significant (Honjo et al., 2004; Rae and van den Noord, 2006; Seyfried, 2010).

In order to model house prices (P_t), in line with most empirical studies, we use the average new house price[17] (Conefrey and Fitz Gerald, 2010; Duffy et al., 2005; Roche, 2003). It is common to use real prices with the underlying thought that real house price inflation is decisive for the investment decision of economic actors. However, in this research we focus on nominal prices since these are important in view of financial stability. Firstly, the nominal price determines the size of the collateral or the value of the underlying property related to the mortgage. If house prices are worth more or less than the underlying mortgage this implies a positive or negative psychological wealth effect. Secondly, since the size of the loss and the probability of default increases with the extent to which the mortgage loan is greater than the nominal value of the collateral, nominal

[15] House prices rather show a log-linear course than a straight line and probably there is a non-linear relationship between the dependent and independent variables. A log-linear transformation allows such relation while maintaining linearity of the model.
[16] A (semi)-elasticity measures the percentage change of the dependent variable if the independent variable changes with 1% (one unit). The log notation can be interpreted in percentages because: $ln(y_0) = b_1 + b_2.x_0$ and $ln(y_1) = b_1 + b_2.x_1$. So $ln(y_1) - ln(y_0) = b_2.(x_1-x_0) = b_2.\Delta x$. Hereby : $100.[ln(y_1) - ln(y_0) \approx \%\Delta y = 100.\ b_2.\ \Delta x$ (Carter et al., 2012).
[17] An alternative was using second-hand house prices. But new and second-hand houses are not perfect substitutes. The reason for not continuing the analysis with second-hand prices is that Roche (2003) argues, using an Engle-Granger causality test for the period between 1979 and 2003, that (Irish) new house prices influence second-hand house prices while causality in the other direction is less present. Though, we decided to perform a robustness check, using the SUR- technique, referring to appendix D.

values are crucial for the potential losses banks have to book in the event of default (Central Bank of Ireland, 2007). The quarterly data come from the Irish Department of the Environment, Community and Local Government (2013) and cover the period 1994 – 2011.

To measure the impact of demographic evolution we used the total population size, redefining DEM_t as $ln(POP_t)$ and it is measured as the natural logarithm of the population at point of time t. The fact is that in the past fifteen years the Irish population has increased enormously. Therefore it is a common thought that it was an important driver of demand for housing which makes it extremely suitable for modeling purposes (Miles and Pillonca, 2008; Roche, 2003). Data are from the IMF (2013) and were interpolated from years to quarters using the Chow-Lin interpolation method (1971).

As supply indicator we use the housing stock per 100 inhabitants which represents a scarcity-effect. If supply goes up, it can be expected that prices go down. Data were retrieved from the Department of the Environment, Community and Local Government (2013) and were interpolated to quarterly data using the method proposed by Chow and Lin.

The income level partly explains how much people can pay and borrow for a house. The chosen proxy is the gross national disposable income per capita, the GDP corrected for outgoing and incoming transfers to or from non-Irish institutional organizations. The time series is only available on a yearly basis till 2011 and was interpolated to quarters (OECD, 2013).

The cost of capital or interest rate is measured as the quarterly average mortgage rate (r_t). The time series is available at the Department of the Environment, Community and Local Government (2013). In order to measure the impact of the monetary policy of the ECB, the mortgage interest rate is explained using an OLS regression with the quarterly average Eonia as explanatory variable. This leads to an estimated interest rate R_t, equation (6), which is recursively completed in the housing equation (7). The idea is that the Irish mortgage rate is strongly connected to the short term money market interest rate, since most loans have a variable rate. More specifically, between 2005 and 2010, 85% of outstanding mortgage debt was linked to a contract stipulating a variable interest rate. In 2010, 30% of all contracts encompassed a variable rate while 55% were a tracker. Such contracts seemed very attractive since policy rates were relatively low. Furthermore, Irish commercial banks relied heavily on short term, wholesale funding. So the Eonia is likely to have a strong impact on the demand for housing (Central Bank of Ireland, 2007; Fitzpatrick and McQuinn, 2007; McQuinn and O'Reilly, 2008; Rae and Van den Noord, 2006).

Finally, equation (5) is complemented by an autoregressive term (*ln(P_{t-1})*), which in contrast to most empirical literature for estimating house prices (Duffy et al., 2005; Stevenson, 2008; Roche, 2003). All mentioned studies intend to estimate the fundamental house price in order to determine the extent to which fundamental factors like the interest rate, increased population and income, can explain the evolution of house prices. Usually it is then assumed that the difference between the actual and estimated price is a good measure of overvaluation. However, the reason that we take the dynamic character of house prices into account is that we want to model prices as accurately as possible. A lagged variable allows us to deal with serial correlation and to capture persistence in the increase in house prices.

If all above mentioned definitions are applied this yields a model resembling the one used in, for example, Conefrey and Fitz Gerald (2010), Duffy et al. (2005), and Stevenson (2008):

$$r_t = \alpha + \beta.EONIA_t \rightarrow \hat{r}_t = R_t = \alpha + \beta.EONIA_t \qquad (6)$$

$$\ln(P_t) = \alpha + \beta_1.\ln(Y_t) + \beta_2.\ln(POP_t) + \beta_3.R_t + \beta_4.HSTOCKPOP_t + \beta_5\ln(P_{t-1}) + \varepsilon_t \qquad (7)$$

The interest rate equation (eq. 6) is estimated for the period between 1994 and 2012. The only purpose is to determine a simple relationship between the mortgage interest rate and the Eonia:

$$R_t = 2.85905 + 0.732557.EONIA_t \qquad (R^2 = 0.7244) \qquad (8)$$

Both the constant and the parameter for the Eonia are statistically significantly different from zero at a 99% confidence level. The equation reflects a base rate, the constant, and an addition varying with the short term interbank interest rate.

Next, before estimating the house price model, the variables are tested for stationarity. If the variables are non-stationary OLS estimates are unreliable and can imply a spurious regression unless there is co-integration (Carter et al., 2012; Verbeek, 2004). The test results of the ADF test are presented in table 4. The results suggest that all variables are non-stationary but integrated of order one (I(1)). Phrased differently, the first order difference of the variables is stationary. Only for the interest rate variable is it difficult to make a distinction between whether or not it is stationary, since the ADF-test indicates stationarity on a 10% significance level but not at 5%. These results are confirmed by the ADF-GLS test while the KPSS test rather indicates stationarity. So the conclusion is that the interest rate variable is nearly integrated of order one.

Since the variables are non-stationary it is necessary to test for co-integration. Co-integration means that the variables are integrated of order one but deliver stationary regression residuals. In that case, time series share a common trend and move together. As a result, OLS estimates are super-consistent, implying that parameter estimates converge much faster to the actual population value when compared to the conventional situation with stationary variables (Verbeek, 2004).

Co-integration is tested using the two-step procedure suggested by Engle and Granger. First, variables are tested for first order integration. Next, it is analyzed whether the regression residuals, the outcome of a linear combination of the variables, follow a stationary process. Since the results in table 4 reveal that the variables are integrated of order one while ε_t is stationary, there is a strong indication of the presence of co-integration. As a result, the OLS estimation results represent a long term relationship between the modeled variables. Hence, the results can be interpreted in an economically meaningful way as a long term equilibrium relationship (Baltagi, 2008; Carter et al.; 2012; McQuinn and O'Reilly, 2008).

Table 4 ADF test variables house price model (1994Q1 – 2011Q4).

Variable	Number of lags	Test statistic levels	Test statistic first order difference	Order of integration
$\ln(P_t)$	1	0.58	-2.91	I(1)***
$\ln(Y_t)$	2 and 1	0.30	-2.82	I(1)***
$\ln(POP_t)$	2 and 1	-2.36	-2.33	I(1)**
R_t^a	1	-3.19	-4.42	I(1)***
$HSTOCKPOP_t^b$	2 and 1	-1.51	-3.32	I(1)***
ε_t	0	-8.81	-	I(0)***

Note: test is performed for the time series covering the period ranging from the first quarter in 1994 till the last quarter in 2011. ADF H0: unit root. Rejection H0 if test statistic is smaller than critical value. The number of lags needed to prevent serial correlation in the ADF test regression is reported in the second column. These were determined by looking, on the one hand at the significance of the lagged terms in de ADF regression, starting from several lags and reducing the number by excluding non-significant lags, and on the other hand by adding lags until the LM-test for serial correlation does no longer indicate significant serial correlation (Carter et al., 2012). Critical values ADF-test with trend on 5% significance equals -3.41; without trend -2.86; and without constant -1.94. For the test on co-integration the ADF-test is applied on the residual term (ε_t) of the model in table 5. The appropriate critical values to assess the null hypotheses at a significance level of 5% with in total five variables is -4.42 (Verbeek, 2004). [a] The interest rate variable is stationary or I(0) at a 10% significance level since the p-value of the ADF-test equals 0.09. ** significance at 5%, * significance at 10%.

The house price model estimation results are presented in table 5. Detailed regression output is reported in appendix D. Note that the interest rate variable is lagged by one quarter which is in line with the approach of Conefrey and Fitz Gerald (2010), and Seyfried (2010). People probably only react with some delay to a change in mortgage interest rates and this way of modeling also delivers proof of causality since the variable represents the impact of a change in interest rates on house prices in the next quarter[18]. The same reasoning is followed for the housing stock variable. Finally, a dummy variable which takes the value of one starting from 2008 was included (Dummy 2008). The reason for this is that, at that time, house prices witnessed a sudden radical collapse and the Dummy allows to capture this unusual market movement.

All variables are statistically significant at a 99% confidence level except the housing stock. However, it looks reasonable to include the variable in the model since it is almost significant at a 90% confidence level. Furthermore, excluding this variable would mean that the model is no longer an inverted demand function since the impact of supply on house prices is no longer taken into account. Referring to appendix E, we also argue that the coefficient estimates, compared with models where the housing stock variable is being treated differently, yield similar results.

Table 5 Estimation results house price model, natural logarithm of average new Irish house prices (1994Q1 – 2011Q4).

Variables	**Final model**[c]
constante (α)	-17.8755 ***
$\ln(Y_t)$	0.5025 ***
$\ln(POP_t)$	1.1160 ***
R_{t-1}	-0.0215 ***
$HSTOCKPOP_{t-1}$	-0.0044 [a]
Dummy 2008	-0.1111 ***
$\ln(P_{t-1})$	0.6762 ***
Adjusted R^2	0.998
Schwarz criterium	-314.485
Jarque-Bera test	0.433
White test	0.103
ARCH(1)-test	0.870
Breush-Godfrey LM test[b]	0.605
Ramsey RESET test	0.074

Note: the lower part of the table reports the p-values of statistical tests. [a] In this specification the p-value is 0.12, thus this variable is almost significant at a 90% confidence level. [b] Test for serial correlation of order one with F-test and null hypothesis of no serial correlation (Carter et al, 2012). [c] See appendix D *** significance at 1%, using HAC-standard errors.

All tests concerning the basic assumptions for reliable OLS estimates indicate no major concerns. The Jarque-Bera test indicates normal residuals, the null hypothesis of the White test postulating homoscedastic residuals is not rejected, the test for ARCH(1) effects is negative and also the

[18] Nonetheless, the model is, as a robustness check, also estimated using the actual interest rate variable. See appendix E.

Breush Godfrey test indicates the absence of serial correlation. Finally, also the Ramsey Reset test seems sufficient to reject non proper model specification.

Since the model contains an autoregressive term, the estimated coefficients represent the immediate or short term effect on the prices in quarter *t*. However, the coefficients can also be interpreted in terms of a long term multiplier or long term impact calculated by: $\frac{\beta_t}{(1-y_{t-1})}$. β_t represents the coefficient of the explanatory variable and y_{t-1} the lagged term $ln(P_{t-1})$. The expression yields the total or cumulative impact over time of a (permanent) change of the explanatory variable with one unit (Carter et al., 2012; Verbeek, 2004).

As expected, a higher level of income causes higher house prices. The coefficient of disposable income is 0.50, so, ceteris paribus, if income increases by 1%, the Irish house prices in the current quarter will on average rise by 0.50%. The long term elasticity equals 1.55. Phrased differently, if the disposable income increases by 1%, in the long run, prices increase by on average 1.55%. But if income decreases, there is a similar but negative impact. These findings are in line with earlier results of Conefrey and Fitz Gerald (2010), Honjo et al. (2004) and Rae and Van den Noord (2006).

Also the population coefficient reveals that the demographic evolution is clearly an important factor in explaining the evolution of Irish house prices. If the population grows by 1%, the average value of a house in quarter *t* will increase by 1.12%. The impact on the long term equals 3.44%. So a rise in the population size by 1% eventually leads to 3.44% higher house prices, keeping in mind that between 1994 and 2012 the Irish population grew by about 27.5% (IMF, 2013). Our findings with respect to the impact of population correspond to those of Stevenson (2008).

The dummy variable (Dummy2008) is significant and negative, capturing the structural change in the model. It may capture the effect of mistrust towards property in the period after house prices started to decrease.

The coefficient of the lagged variable ($ln(P_{t-1})$) equals 0.68 and is relatively high. This means that if houses price in the previous quarter were 1% higher, the price in the current quarter will also be on average 0.68% higher. This may point in the direction of an effect of rising house prices encouraging people to buy houses because they expect prices to increase. So a shock, for example a too low interest rate for a couple of years, can cause a persistent trend of rising house prices even when the underlying trigger of the initial price increases has disappeared or weakened

(Abraham and Hendershott, 1996; Stevenson, 2008). So prospects of future price increases may have created a self-fulfilling wave of optimism that took over rational human behavior. However, it is important to recognize that the effect of expected prices also applies in the other direction. When people expect prices to decrease or recognize that prices are out of line, they can anticipate and quickly sell houses before the market value actually goes down (Shiller, 2007).

Last but not least, there is the impact of the interest rate. If the average mortgage interest rate increases by 1%, average house prices decline by about 2.15%. Conversely, if the interest rate is lowered by 1%, the price of houses will increase. The total impact of one percentage point change in the long run equals 6.65%[19]. This value corresponds to the impact of the interest rate variable estimated in the work of Stevenson (2008). Stevenson estimated a similar model to determine the long term relationship between real house prices and income, demographics and the interest rate. His inverted demand function was estimated for the period between 1978 and 2003 and resulted in a long term interest rate coefficient between 0.05 and 0.07.

So by estimating a model of Irish house prices, controlling for the impact of other variables, we show that the interest rate variable plays an important role. In the previous section, using the Taylor-rule, we determined the stance of monetary policy and how the interest rate path would have looked if the ECB took Irish conditions into account perfectly. Combining both analyses makes it possible to carry out a counterfactual analysis in order to assess the extent to which a monetary policy that was tailored to the needs of the Irish economy could have mitigated the formation of the Irish property bubble. Therefore, we replaced the Eonia in (8) by the Irish Taylor-rate, i.e. the counterfactual short term interest rate, which results in a counterfactual mortgage interest rate under an alternative monetary policy. The results are presented in figure 5, which compares the actual and counterfactual Irish mortgage rate. The simulation results for the period 1999 - 2008 lead to an average counterfactual mortgage rate of 9.9% while the actual rate was on average 4.65%.

[19] The coefficient in table 5 equals -0.022 and the one for the lagged price 0.676. Result: the long-term-multiplier yields -0.021533/(1-0.676229) = - 0.0665. So if the interest rate is increased with one percentage point, the effect is -6.65%.

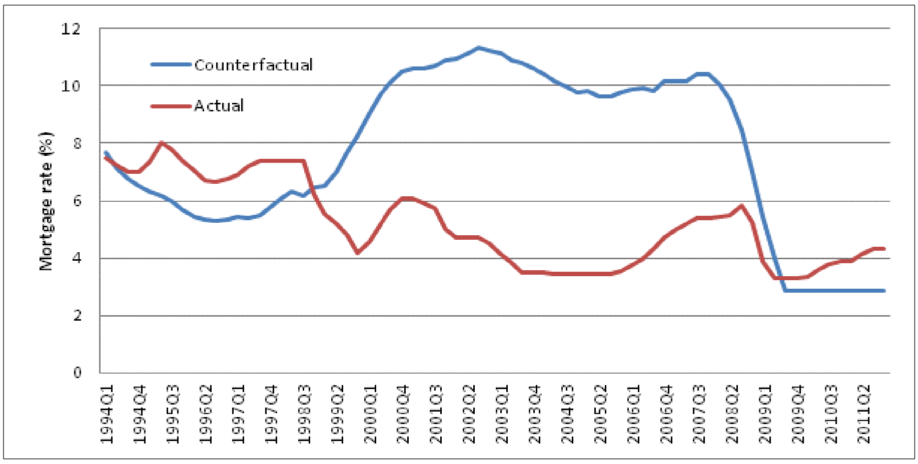

Fig. 5. Actual vs. Counterfactual average Irish mortgage rate (1994 – 2012).

Next, we combined the counterfactual mortgage rate with the housing equation, using the estimated coefficients presented in table 5. This approach shows how the average Irish house price would have evolved over time if there was an alternative monetary policy.

In figure 6 we display the final results of the counterfactual analysis of Irish house prices between 1994 and 2012. It is very striking that since 1999 the actual and counterfactual average Irish house prices start to diverge. So it is no coincidence that this coincided with the introduction of EMU. Especially in the first years after euro accession both time series follow a different path or trend. This provides clear evidence that higher policy interest rates would have led to a much smaller increase in Irish house prices. The actual yearly average growth rate of Irish new house prices between the first quarter of 1999 and the fourth quarter of 2007 was as high as 11.15%. Under the alternative interest rate policy it would have been only 6.84%. As a result, mid 2007[20], just before the bust of the property bubble, Irish house prices would have been about 30%[21] lower. So the consequences of the crisis would be more modest and instead of a sudden-stop, there probably would have been a soft landing as can be seen from figure 6. It also shows that keeping the interest close to the Taylor-rate prevents housing prices from rising dramatically. Therefore it is not necessary to include a lean against house price fluctuations in monetary policy strategies.

[20] The maximum difference between the actual and counterfactual price is biggest in the fourth quarter of 2003, namely 34%.
[21] Calculated as *[(counterfactual price – actual price) / actual price]*100*.

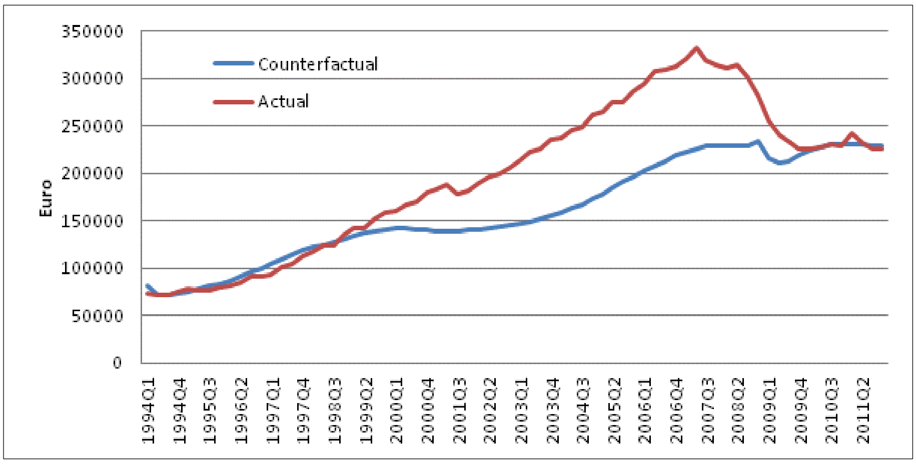

Fig. 6. Counterfactual Irish house prices (1994Q1 – 2011Q4).

It is important to recognize that Irish house prices rose in tandem with the amount of credit provided by the domestic banking sector. So in order to fully explain the severe consequences of the housing boom and inappropriate interest rates, we make one additional step in our reasoning and follow De Grauwe (2010) and Lane (2012), arguing that the property bubble was only able to become so intense because it was supported by an enormous credit expansion. By 2007, at the peak of the real estate boom, the outstanding amount of credit provided by the domestic banking sector was twice as big as GDP against a ratio of only a half in 1994 (The Worldbank, 2013). In addition, in contrast to the past where deposits of private actors were the main source of funding, cross border and interbank wholesale funding became of paramount importance. Euro accession resulted, through a permanent decrease in nominal interest rates, deep short term money market integration and the absence of an exchange liquidity risk premium, in an unprecedented access to cheap interbank wholesale funding (Addison-Smyth et al., 2009; Honohan, 2010; Regling and Watson, 2010). The consequence is that the net foreign debt of the Irish banking system increased from 20% in 1994 to 73% of GDP in 2008. So it is no surprise that between 2002 and 2008 the amount of mortgage credit tripled from 43 to 123 billion euro (Central Bank of Ireland, 2010).

The underlying reasoning for this is the fact that low interest rates makes it attractive for economic actors to borrow money and invest in property while credit providers, mainly banks, will tend to grant more and more credit since the value of houses continuously increases, implying more collateral (ECB, 2011). So the financial accelerator or multiplier mechanism, the mutual interaction between credit and house prices, may explain how the shock in interest rates had huge real effects (White, 2006). With a national currency, increased indebtedness would probably made it more and more difficult to raise money on the financial markets. This would have led to higher funding costs and higher rates for debtors. So sensitivity of domestic interest and exchange rates may have partly curtailed the upward price spiral in Irish property markets (Conefrey and Fitz Gerald, 2010).

This view is confirmed by Addison-Smyth et al. (2009). For the period between 1982 and 2008 they show that a rise of 1% of the amount people can borrow, has a long term impact on Irish house prices of 1%. Furthermore, Fitzpatrick and McQuinn (2007) empirically demonstrate the mutually reinforcing relationship between rising Irish property prices and the average mortgage loan provided by Irish commercial banks. So this picture explains why Irish banks were so vulnerable to a fallback in the property market. When housing became unaffordable, everyone started to smell a rat and recognized that prices could not continue to rise. Public interest witnessed a sudden stop, causing property prices to collapse (Shiller, 2007). As a consequence, in concurrence with high levels of leverage and problems on the interbank market, Irish banks were massively confronted with default and a situation where the outstanding mortgage debt was larger than the collateral.

However, since Ireland is necessarily confined to the policy of the ECB, the question is whether there were policy options, other than the interest rate instrument, that could have mitigated the inflation of the bubble. Our analysis leads us towards fiscal and macro-prudential policy, revealing that there were several alternative policy instruments to drive the economy towards a sustainable equilibrium, ways to overcome the negative implications of giving up a sovereign monetary policy.

Firstly, the absence of a sovereign monetary policy implies that fiscal policy becomes crucial to stabilize the economy and to realize real convergence within the euro zone. But the Irish government conducted a pro-cyclical policy and made the tax structure far too dependent on unsustainable property related revenues. Data of the OECD (2013) show that between 2001 and 2006 there was an average annual structural deficit of about three to four percent of potential GDP. So fiscal policy has probably strengthened excesses and cyclical divergences. Fiscal stimuli for buying property remained till 2007, interest on mortgage lending could be deducted from tax and there was no property tax. Since political economy predicts that democratically elected policy makers probably choose the way of least resistance, this case illustrates the importance of independent supervision by a Euro-wide economic government (Lane, 2012; Regling and Watson, 2010). On the one hand, a contra-cyclical policy could have slowed down building activity, hence mitigating the formation of extreme macro-economic imbalances. The government could have abolished tax deductibility of mortgage interest rates or introduced a tax on ownership, hence influencing the investment behavior of people and subsequent credit accumulation. On the other hand, despite that it significantly reduced national debt levels, the government could have run more significant surpluses in a period of strong economic growth following the example of Finland, Sweden or Denmark. In this way it would be better positioned to face the global financial crisis by having slack resources to stimulate the economy rather than pursuing austerity (Conefrey and Fitz Gerald, 2010; Lane, 2012). However, the smallness and openness of the Irish economy probably implies that the fiscal multiplier, the impact of more government expenditure on GDP, is modest because it is partly diverted towards imports. So euro-wide action is far more effective, although politically difficult due to the absence of a fiscal and political Union.

A second way to intervene is through macro-prudential policy with tighter regulation since credit conditions became far too lax. The idea is that supervisors can affect the fractional reserve system. Loans of 100% loan-to values were introduced without action or regulation of the CBIFSRA (Regling and Watson, 2010). They could have put a legal cap on the loan-to-value ratio or limited the repayment period of loans. By doing so, it could have cooled down the demand for credit and hence the demand for property. This in turn could have mitigated the situation where banks are confronted with default of property related loans while the collateral is less valuable than the outstanding debt. So such measures could have helped to reduce the intensity of the expansion of bank credit and the consequences of the crisis (Conefrey and Fitz Gerald, 2010; White, 2006).

4. Conclusion

Empirical results clearly prove that the monetary policy of the EMI and ECB can be modeled very accurately using a forward-looking Taylor-rule. Furthermore, the policy of the ECB turns out to be attuned to the goal of maintaining price stability and is clearly contra-cyclical, as expected of a sound monetary policy. However, this conceals diverse developments at national level. Ireland entered the EMU at the peak of the Celtic Tiger with economic conditions diverging from the rest of the euro zone. The levels of economic growth, inflation and employment were much higher so that the ECB's monetary policy was not tailored to the needs of the small Irish economy. Our results indicate that between 1999 and 2007 interest rates should have been on average 6.7% higher. Consequently, mortgage rates were persistently about 5% too low. At the same time our estimations showed that under an alternative sovereign monetary policy, the average house price would have been 25 to 30 percent lower just before the housing bust in the second quarter of 2007. Our results thus clearly prove the relationship between too low interest rates, i.e. interest rates that are below the Taylor-rule implied rate, and the housing bubble. In addition, the counterfactual analysis also shows that, to prevent housing bubbles, it is not necessary to include housing prices in monetary policy decisions. A monetary policy tailored to the needs of the member state is enough to prevent housing prices from dramatically increasing. Therefore it is not necessary to include a lean against house price fluctuations in monetary policy strategies.

It thus is shown that EMU-membership led to very low, inappropriate Irish interest rates and that they contributed significantly to the Irish credit, consumption and property boom and subsequent bust. Our findings thus support the view that one size clearly does not fit all and that euro accession contained the seeds of the Irish downfall. Paradoxically, the expectations of lower interest rates and capital costs were an important reason to become a member of the EMU while it created an environment for a non-sustainable credit expansion and other macro-economic imbalances.

However, despite clear evidence that euro membership contributed to a housing boom and subsequent bust, this does not necessarily mean that a sovereign Irish monetary policy would have prevented the crisis although it, given the prolonged time of a too low interest rate, most certainly would have been far less severe. It also does not mean that policy makers could not use alternative counter cyclical policy measures to stabilize the Irish economy during the period leading up to crisis. The absence of an endogenous monetary policy implies that fiscal and macro-prudential policy become more important to guarantee a stable economy while the crisis highlighted the need to realize real convergence within the euro zone. However this remains politically difficult due to the absence of a fiscal and political Union.

Appendix A

Equation: Iterated GMM, using observations 1994:1-2012:2 (T = 74)
Dependent variable: KwgemEONIA
Instrumented: INF2 OutputgapEMU1 INS1
Instruments: const INS2 INS3 HICP OutputgapEMU
HAC standard errors, bandwidth 3 (Bartlett kernel)

	Coefficient	***Std. Error***	***z***	***p-value***	
const	-0.133765	0.192727	-0.6941	0.48764	
INF2[a]	0.208953	0.075910	2.7527	0.00591	***
OutputgapEMU1[b]	0.114431	0.028476	4.0185	0.00006	***
INS1[c]	0.899929	0.023695	37.9797	<0.00001	***

Mean dependent var	3.153715	S.D. dependent var	1.683841

GMM criterion: Q = 0.000168743 (TQ = 0.012487)
J test: Chi-square(1) = 0.012487 [0.9110]

Fig. A. Gretl output forward looking Taylor-rule estimates using GMM. [a] INF2 = inflation with a lead of two quarters. [b] OutputgapEMU1 is the output gap looking one quarter ahead. [c] INS1 is the interest rate smoothing parameter or the Eonia lagged with one quarter. Instruments: const is a constant, INS2 is the second lag of the EONIA; INS3 is the third lag of the EONIA; HICP is the Actual HICP and OutputgapEMU is the actual Output gap. *** Significance at 1%.

Appendix B

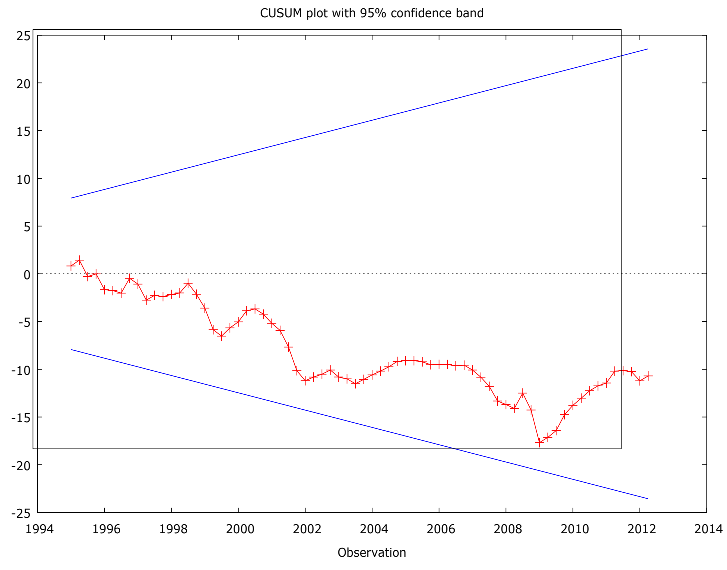

Fig. B. Cusum test Taylor rule ECB

Appendix C

Table C1 ADF-test Irish variables, the HICP and output gap (1994 – 2012).

Variable	Number of lags	Test statistic	Critical values			Order integration	P-value LM test
			1%	5%	10%		
HICP Ireland	1	-2.42	-3.43	-2.86	-2.57	I(1)	0.55
Output gap Ireland	1	-0.86	-3.43	-2.86	-2.57	I(1)	0.29

Note: ADF test with constant, H0: unit root. Rejection H0 if test-statistic is smaller than critical value. Number of lags needed to exclude serial correlation in the ADF test indicated in 2nd column. Smallest p-value Q-statistic for test absence serial correlation: EONIA 0,74, HICP 0,31 and Output gap 0,15. ** significance at 5%.

Table C2 Test stationarity Irish variables, KPSS and ADF-GLS test (1994 – 2012).

KPSS[a]	Test statistic	Critical values			**ADF-GLS** (p-value)
		1%	5%	10%	
HICP Ireland	0.41	0.731	0.47	0.35	0.012
Output gap Ireland	0.67	0.731	0.47	0.35	0.350

[a] Truncation parameters equals three, in line with the standard value calculated as: 4*[T/100] 1/4. Hereby T stands for the number of observations and equals 76. Null Hypothesis of stationarity is rejected if test statistic is larger than critical value.

Tables C1 and C2 summarize the results of the stationarity tests of the Irish HICP rate and output gap. The null hypothesis of a unit root of the ADF test cannot be rejected (see table C1). Therefore the test is complemented by the KPSS and ADF-GLS test which in small samples may yield more powerful results (Elliott et al., 1996; Gorter et al., 2008). The ADF-GLS test indicates that the null hypothesis of non-stationarity can be rejected for the HICP but not for the output gap. The KPSS test turns the null hypotheses around and postulates stationarity against a unit-root (Kwiatkowski et al., 1992). Again, for the HICP the null cannot be rejected while the test statistic of the output gap is smaller than the critical value so indicating a unit root. The conclusion is that the EONIA and HICP are clearly stationary while results for the output gap are less clear. However the results are probably influenced by the limited length of the time series while the recent crisis caused a break. Furthermore, the residuals of the OLS regression of eq. 2 are stationary, indicating that the regression model represents the long term relationship between the modeled variables (Verbeek, 2004). And because the output gap is defined as GDP deviating from a long term trend, we, as is common in the literature, assume that the output gap is stationary in the long run.

Appendix D

OLS, using observations 1994:2-2011:4 (T = 71)
Dependent variable: *ln(P_t)*
HAC standard errors, bandwidth 3 (Bartlett kernel)

	Coefficient	***Std. Error***	***t-ratio***	***p-value***	
const	-17.87550	3.81655	-4.6837	0.00002	***
ln(Y_t)	0.50255	0.08747	5.7456	<0.00001	***
ln(POP_t)	1.11600	0.24532	4.5492	0.00002	***
Dummy2008	-0.11113	0.01828	-6.0791	<0.00001	***
$HSTOCKPOP_{t-1}$	-0.00438	0.00279	-1.5700	0.12135	
R_{t-1}	-0.02153	0.00399	-5.4022	<0.00001	***
ln(P_{t-1})	0.67623	0.04981	13.5748	<0.00001	***

Mean dependent var	12.10386	S.D. dependent var	0.465241
Sum squared resid	0.032555	S.E. of regression	0.022554
R-squared	0.997851	Adjusted R-squared	0.997650
F(6. 64)	6392.723	P-value(F)	6.91e-87
Log-likelihood	172.1619	Akaike criterion	-330.3238
Schwarz criterion	-314.4851	Hannan-Quinn	-324.0253
rho	-0.059004	Durbin's h	-0.543079

Test for normality of residual – H0: error is normaly distributed.
Test statistic: Chi-square(2) = 1.6758 with p-value = 0.432619

White's test for heteroskedasticity – H0: heteroskedasticity not present
Test statistic: LM = 35.4492 with p-value = P (Chi-square(26) > 35.4492) = 0.102273

LM test for autocorrelation up to order 1- H0: no autocorrelation
Test statistic: LMF = 0.270216 with p-value = P (F(1.63) > 0.270216) = 0.605009

Test for ARCH of order 1 – H0: no ARCH effect is present
Test statistic: LM = 0.026817 with p-value = P(Chi-square(1) > 0.026817) = 0.86992

Table D Gretl output OLS estimates final house price model with *ln(P_t)* as dependent variable and the explanatory variables: Y_t, the gross disposable income per capita, POP_t total population, Dummy2008 being one from the first quarter of 2008, R_{t-1} the estimated mortgage interest rate following equation (8) with P_{t-1} the first lag of the average new house price. HAC standard errors (Std. Error). *** Significance at 99%.

Appendix E

In order to prove the robustness of the results, especially to check whether the impact of the interest rate variable remains of the same order, the estimation of the house price model is undertaken using additional model specifications.

In model a (see table E) we estimate a similar model used in the counterfactual but without lagging the interest rate or housing stock variable. But model a suffers from heteroskedasticity since the null hypothesis of the White-test of constant variance is rejected. Also the null hypothesis of proper model specification from the RESET test is rejected. Therefore the model is also estimated without the non-significant housing stock variable (model b). In both models the impact of the mortgage interest rate variable is a bit smaller than the results in table 5. Changing the interest rate by one percentage point has a direct impact on house prices of 1.6% and in the long run 4.8% against 2.15% and 6.65% in the model used for the counterfactual presented in figure 6. Following model b, under an alternative monetary policy for the period 1999 – 2007 house prices would have increased on average by 8%. Hence, the counterfactual house prices of model b are a bit higher. As a result, also the extent of overvaluation would be more modest. More specific, in the second quarter of 2007, counterfactual prices would have been 25% lower.

In model c we use real prices instead of nominal prices, defined as: *100*[(price / (100+ HICP Ireland)]*. As one can see, the coefficient of the interest rate variable is identical to the results presented in table 5. Also the long-term-multiplier remains unchanged. So the choice for nominal prices turns out to be justifiable or robust since the influence of R_t does not change significantly.

In model d nominal prices are estimated using real interest rates as explanatory variable instead of the nominal rates. The real rate is defined as in Conefrey and Gerald (2010): $R_t = [Nominal\ mortage\ rate_{t-1} - \frac{1}{2} * ln(\frac{P_{t-1}}{P_{t-3}})]$. So the mortgage rate is corrected with a three-periodic moving average of the increase in house prices. Using this model, the average counterfactual price for the period 1999 - 2007 would have risen on average by 8.5% per annum while in the secondquarter of 2007 prices would have been 21% lower.

Finally, model e represents the estimation results of average new house prices using the Seemingly Unrelated Regression (SUR) method. The idea is that new and second-hand houses are no perfect substitutes, although their prices isare strongly linked to each other (Rae and van den Noord, 2006). Taking second-hand prices directly into the equation to explain new prices is not common in literature because it is very likely to cause endogeneity. Therefore we use SUR. The SUR-method estimates two equations simultaneously but allows temporal correlation between the error terms. So the relationship between both prices is modeled through the error term (Baltagi, 2008; Rae en van den Noord, 2006). The direct impact of a change of interest rates of one percentage equals 2,1% while the long-run-multiplier is 5.56%. On the basis of this model, the average counterfactual price for the period 1999 – 2007 would have risen by on average 7.6% per annum while in the second quarter of 2007 the counterfactual price would have been 27.8% lower than the actual price. These results remarkably resemble the findings presented in figure 6.

Table E Alternative model specifications house price model: robustness (1994 – 2011).

Variabele	model a	model b	model c real house prices	model d nominal price and real rate	model e (SUR)
Constante	-19.506 ***	-15.494 ***	-13.413 ***	-24.774 ***	-21.385 ***
$\ln(Y_t)$	0.545 ***	0.527 ***	0.472 ***	0.666 ***	0.592 ***
$\ln(POP_t)$	1.205 ***	0.935 ***	0.822 ***	1.537 ***	1.330 ***
R_t	-0.016 ***	-0.016 ***	-	-0.014 ***	
R_{t-1}	-	-	-0.022 ***	-	-0.021 ***
$HSTOCKPOP_t$	-0.005	-	-	-0.008 *	
$HSTOCKPOP_{t-1}$			0.002	-	-0.005
Dummy 2008	-0.106 ***	-0.010 ***	-0.096 ***	-0.010 ***	-0.120 ***
$\ln(P_{t-1})$	0.664 ***	0.669 ***	0.683 ***	0.587 ***	0.624 ***
Adjusted R^2	0.998	0.987	0.998	0.997	0.998
Schwarz criterium	-306.693	-309.933	-318.738	-310.472	-
Jarque-Bera test	0.223	0.223	0.897	0.402	- [b]
White test	0.034	0.058	0.251	0.423	-
ARCH(1)-test	0.934	0.898	0.555	0.560	0.951
Breush-Godfrey [a]	0.387	0.331	0.495	0.433	-
Ramsey RESET test	0.008	0.086	0.145	0.546	-

Note: p-value test statistics in lower part of the table. For model c and d the Engle Granger procedure reveals that the real interest rate and real prices are integrated of the first order but that regression residuals are stationary, hence indicating co-integration [a] Test for absence first order serial correlation. [b] In a system with several equations there is an alternative test, the Doornik-Hansen test. The null hypothesis postulates normal residuals. P-value is 0.44 so the assumption of normality is fulfilled.

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