Did Monetary policy fuel the housing bubble: an application to Ireland.

Cindy Moons¹ and Kevin Hellinckx²

Abstract

This paper provides empirical evidence of the role of the euro in the genesis of the recent Irish financial-economic crisis. By using a Taylor rule 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. Using a BVAR and multivariate housing model, we provide econometric evidence 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 addition, it shows that a monetary policy tailored to the needs of the member state prevents housing prices from dramatically increasing.

Jel Code: E5, F6

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

¹ Ku Leuven campus Brussel, Department of Economics and Business, Warmoesberg 26, 1000 Brussel, Belgium, cindy.moons@kuleuven.be and University of Antwerp, Department of Economics, Belgium.
² Ku Leuven campus Brussel, Department of Economics and Business, Warmoesberg 26, 1000 Brussel, Belgium kevin_hellinckx@hotmail.com.
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. 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 undisputed. The theory of optimum currency areas (Mundel, 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 European Central Bank (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. 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 Obstfert, 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 a consequence, monetary policy can have asymmetric real effects and can even strengthen national macro-economic imbalances through the well-known Walters critique. A too low nominal interest rate leads to a decline in real interest rates which 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 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). Between 1997 and 2008, Irish real GDP growth was persistently two to three percent higher than in Germany and the EMU. The correlation between the cumulative
growth of real GDP and the cumulative inflation rate is 0.54 which indicates that besides structural factors cyclical or national demand factors played a role in explaining inflation differentials within the euro area (Lane, 2006). Economic cycles clearly aren’t synchronized, especially when comparing Ireland with the core countries. 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.

Despite the differences, the ECB succeeded relatively well in keeping the aggregate inflation close to 2%. 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.

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.

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

---

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).
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 2015, the euro and the Irish economy are still under pressure and thus 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, section 2 presents the analytical framework used to evaluate the appropriateness of the ECB’s monetary policy for the Irish economy. We estimate the Taylor-rule for the ECB and Ireland and use these results to perform a counterfactual analysis. It 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 estimating a Bayesian Vector Auto Correction (BVAR) model and discuss the model estimates and policy implications. These results are complemented with an intuitive multivariate housing model. Finally, section 4 summarizes the main findings and concludes.

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

Fig. 2. Change in real house prices and growth nominal credit as percentage of GDP

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 and 2000).

\[ i_t = (1 - \rho). (\alpha + \beta. E[\pi_{t+n} \mid t] + \gamma. E[y_{t+k} \mid t] + \rho. i_{t-1} + \epsilon_t) \]  

(1)


\(^5\) Source OECD (2013a) and The Worldbank (2013).
Where $i_t$ represents the current interest rate, $\alpha$ is a constant while $\beta$ and $\gamma$ represent the long term weights for respectively the inflation rate ($\pi_t$) and the output gap ($y_t$). $E$ represents the rational expectations operator$^6$, 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$^7$. Therefore, $\pi_{t+n}$ represents the annualized inflation rate at point of time $t+n$. The idea is that a central bank determines its interest rates using information it really possesses, namely predictions of inflation and the output gap rather than purely past or current inflation and output data. Next to this, there is a time lag in the monetary transmission mechanism indicating that makers have to anticipate$^8$ (Gerdesmeier and Roffia 2003; Fourçans et al., 2004). Furthermore, $\rho$ represents the smoothing parameter or the weighting factor with a value between zero and one. The size of $\rho$ indicates how fast the central bank reacts to a change in economic conditions. The bigger the parameter, the more gradual the reaction.

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 markets. Secondly, it implies transparency and consistency, strengthening the effectiveness of monetary policy by influencing market expectations about the future path of the short term interest rate, hence influencing long term interest rates (Castelnuovo, 2007; Cecchetti, 2000; Woodford, 1999). Thirdly, 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. All these arguments justify a gradual, prudent reaction which buys policy makers some time to assess or redirect policy actions (Cecchetti, 2000; Orphanides, 2003).

Finally, if the non-observed variables are replaced by the actual, realized variables, we get an estimable forward looking specification$^9$:

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

$^6$ The rational expectations theory assumes that predictions are not systematically wrong Clarida et al. (1998).

$^7$ 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 $k = 12, n = 3$.

$^8$ Output mostly reacts faster than inflation. E.g. Peersman and Smets (2001) find for the euro area that a rise in the short term interest rate in followed by a fall in output after one to two quarters while prices are more sluggish. Similar, 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, Fourcans and Vranceanu (2004) talk about a lag of six to twelve months.

$^9$ We assume that realized values are a proxy for the expected values.
In order to estimate the Taylor-rule\(^{10}\) 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 (see e.g. Gerdesmeier and Roffia 2003; Gerlach and Schnabel, 2000).

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.\(^{11}\) Monthly data are converted into 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.

The reaction function is estimated using the Generalized Method of Moments\(^{12}\) since the Hausman specification test, reported in table 1, indicates that the interest rate is endogenously determined by the inflation and the output gap. GMM 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 (Belke, 2007; Gerdesmeier and Roffia 2003; Floren et al., 2004). In accordance with the literature the GMM weighting matrix is chosen using the method of

\(^{10}\) We are well aware of the identification problems regarding the use of policy rules in New Keynesian models pointed out by Cochrane (2011). However, Sims (2008) demonstrated that non-identification of Taylor rules is not a generic implication of the model, but rather results from a particular (and unrealistic) assumption on the policy rule itself. For a standard specification of the interest rate rule the policy rule parameters are in fact identified and may be estimated using standard techniques. E.g. the non-identification result of Cochrane (2011) requires that the interest rate rule features a stochastic intercept which tracks fluctuations in the Wicksellian natural rate of interest. Furthermore, Mavroeidis (2010) shows that the GMM estimation results of Clarida, Gali and Gertler (2000) are valid and robust. Also, we do not intend to use the estimation results for investigating monetary policy shocks. We merely want to find the actual behavior to determine the (in)appropriateness of monetary policy for the Irish economy.

\(^{11}\) Data of the HICP represent the euro 17 group while the data for the output gap, estimated using a Cobb-Douglas production function approach, only cover the euro 15 group. However, because the weight of Slovakia and Estonia in the HICP is only 0.73% and 0.15%, this does not influence the results significantly.

\(^{12}\) the ADF test in appendix A clearly rejects a unit root indicating that all variables are stationary. 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 (Hail, 2009; Kitamura and Philips, 1997).
Newey and West. This delivers results that are consistent in the presence of heteroscedasticity and auto-correlation of unknown form. 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). Therefore, we will use 2SLS using the same set of instruments as a robustness check. Cross-checking the results allows us to assess the sensitivity of the results to the chosen estimation technique.

After model selection and following Fourçans and Vranceanu (2004), Castro (2008) and Gerdesmeier and Roffia (2003), we continue with the forward-looking specification with $n = 2$ and $k = 1$. Table 1 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: $\theta_0 = (1-\rho)\alpha$, $\theta_1 = (1-\rho)\beta$ and $\theta_2 = (1-\rho)\gamma$; with $\theta_0$ 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 B. The J-test confirms the validity of the chosen instruments and proper model specification. Furthermore, the test of Stock and Yogo (2002) indicates 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 1. Important to note is that 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 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 indicates that the GMM estimates are robust and valid. Finally, with regard to parameter stability we carried out the Cusum test of the residuals. Since there are no movements outside the critical lines, the parameters are relatively stable over time (see appendix C).

Looking at the results, a first finding is that all GMM coefficients are statistically significant at the 99% confidence level. The smoothing parameter ($\rho$) is relatively high but in line with most empirical literature. Also the coefficients for the inflation ($\beta$) and the output gap ($\gamma$) have the expected sign

13 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, also the Cusum test for parameter stability was most convincing.
and a value in accordance with some earlier findings (Castro, 2008; Fourçans and Vranceanu, 2007; Gerdesmeier and Roffia, 2005; Gorter et al., 2008).

### Table 1

<table>
<thead>
<tr>
<th></th>
<th>Constant (α)</th>
<th>π_{t+2} (β)</th>
<th>γ_{t-1} (γ)</th>
<th>ρ_{t-1} (ρ)</th>
<th>J-test (p-value)</th>
<th>Sargan test</th>
<th>Cragg-Donald Hausman test</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>GMM</strong>a</td>
<td>-1.34</td>
<td>2.09***</td>
<td>1.14***</td>
<td>0.90***</td>
<td>0.91</td>
<td>-</td>
<td>9.84</td>
</tr>
<tr>
<td><strong>2SLS</strong>b</td>
<td>-1.18</td>
<td>2.02*</td>
<td>1.14 ***</td>
<td>0.90***</td>
<td>-</td>
<td>0.91</td>
<td>9.84</td>
</tr>
</tbody>
</table>

Notes: all residuals follow a normal distribution. *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. 

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\(^{14}\). Finally, a joint Wald-test\(^ {15}\), 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). So the results provide clear evidence that the ECB conducted a theoretically sound leaning-against-the-wind-policy.

Figure 3 plots the Eonia and the estimated Taylor-rate and 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.

\(^{14}\) The interpretation can be derived as follows: \(h = p \cdot h - 1 + (1- p) \cdot h^*\) with \(h^*\) the targeted Taylor or policy rate can be rewritten as: \(h - h - 1 = (1- p) \cdot h^*\). The left side of the equation expresses the interest rate change so 1-ρ represents the proportion of the difference between the actual and target rate that is being implemented in one quarter.

\(^{15}\) The Wald-test statistic follows a \(X^2_{(q)}\) distribution with the number of degrees of freedom equal to the number of hypotheses being tested (Carter et al., 2012)
Fig. 3. Actual Eonia and the Eonia modeled by the Taylor-rule for the ECB (data fit).

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 are shown 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. The average gap between 1999 and 2006 was 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 2, with those for the ECB yields the same conclusions. Table 2 reports the GMM estimation results of the Taylor-rule for Ireland and is supplemented by the
2SLS results as a robustness check. 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) indicating that the interest rate was less reactive to Irish inflation and output gap than prescribed by a normative Taylor-rule. The test thus confirms the view that one size does not fit all. 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. As explained above, 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 since the unprecedented boom in the housing market was diagnosed to be the main cause of several macro-economic imbalances underlying the Irish bust.

### Table 2 The Irish Taylor-rule.

<table>
<thead>
<tr>
<th></th>
<th>Constant</th>
<th>$\pi_{t+2}$</th>
<th>$y_{t+1}$</th>
<th>$i_{t-1}$</th>
<th>J-test</th>
<th>Sargan-test</th>
<th>Cragg-Donald</th>
<th>Hausman test</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(a)</td>
<td>($\beta$)</td>
<td>(y)</td>
<td>(p)</td>
<td>(p-value)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>GMM$^a$</strong></td>
<td>-1.10</td>
<td>1.34 **</td>
<td>0.31*</td>
<td>0.92***</td>
<td>0.25</td>
<td>-</td>
<td>9.76</td>
<td>-</td>
</tr>
<tr>
<td><strong>2SLS$^b$</strong></td>
<td>0.91</td>
<td>0.51</td>
<td>0.26*</td>
<td>0.90***</td>
<td>-</td>
<td>0.11</td>
<td>9.76</td>
<td>0.001</td>
</tr>
</tbody>
</table>

Notes: 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. Specifically for Ireland, we also use the Irish average house prices since we expect it is likely to be a good predictor of future inflation and output gap [24,36,37]. $^a$ 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(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%.

16 Tests on stationarity can be found in appendix D. 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.
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 for Irish needs. 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. 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. Finally, by performing a counterfactual analysis, combining the Irish Taylor-rate with an 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.

3.1 A Bayesian VAR Model for the Irish property market

In the literature there is consensus that Irish house prices, the result of supply and demand, were driven by the strong economic growth, increased welfare, low unemployment, 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 starting point to model house prices is often an inverted demand equation estimated by OLS. However, in order to examine house price dynamics and causality it is first necessary to take into account the potential endogeneity between house prices and its determinants such as income or the unemployment rate. Therefore we need a model that allows complex lagged interaction effects among the variables. The vector of endogeneous variables comprises the Eonia, average national new house prices \((lnhp)\), gross national disposable income per
capita \((\text{lngndi})\) and the GDP deflator \((\text{indefl})\)\(^{17}\). In contrast to most empirical literature (e.g. Iacoviello, 2005), we expand our benchmark model by adding the unemployment rate \((ur)\) to capture the influence of the favorable Irish demographic evolution and the macroeconomic impact of the booming construction industry. All variables are expressed in natural logarithms except the Eonia and the unemployment rate, which are expressed in percentages.

To avoid identification problems we will estimate the model using a BVAR in which the parameters are treated as random variables and their (posterior) distribution is estimated via the imposition of prior beliefs on their distribution. We will impose a normal inverted Wishart prior (Kadiyala and Karlsson (1997) and Robertson and Tallman (1999)) which is a modification of the Minnesota prior (Litterman, 1986). The Minnesota prior imposes a random walk representation for all variables. The disadvantage of this prior lies in the imposition of a fixed and diagonal covariance matrix of the residuals which rules out possible correlation among residuals of different variables (Auer, 2014). The inverted Whisart prior on the other hand uses the principles of the Minnesota prior but relaxes the assumptions on the covariance matrix structure of the residuals. To estimate the model we follow Canova (2007) and set the overall tightness of the model equal to 0.1 but we perform some robustness exercises on the relevance of the prior tightness to the results\(^{18}\). The lag length is two and chosen according the BIC criterion

For the identification of monetary policy shocks we compute the impulse responses using a Cholesky decomposition (Sims, 1980) of the variance covariance matrix of the reduced-form residuals in model. This method has been widely used to identify monetary policy shocks (see e.g. Christiano and others, 1999, Stock and Watson, 2005). Most of the VAR studies that incorporate house prices identify the model by focusing on recursive contemporaneous restrictions on the interaction between a short term interest rate and house prices. On the one hand we can assume that house prices do not respond immediately to monetary policy as in Goodhart and Hofmann (2008) or Giuliodori (2005). On the other hand one can assume that monetary policy does not react instantaneously to an innovation in house prices as in e.g. Iacoviello (2005)\(^{19}\). In our model, the variables are Cholesky-

\(^{17}\) As such, our model resembles those used in Del Negro and Otrok, 2007; Gattini and Hiebert, 2010; Giuliodori, 2005; Iacoviello, 2005; Jarocinsko and Smets, 2008; McDonald and Stokes, 2013b; Peersmans and Smets (2001), and Tsatzaronis and Zhu, 2004.

\(^{18}\) For the sake of brevity the results are not included and can be provided by the authors upon request. We experimented with a tightness of both 0.05 and 0.15 and found that results did not change dramatically.

\(^{19}\) Jarocinsko and Smets (2008), and Goodhart and Hofmann (2008) recover orthogonalised shocks based on a simple Cholesky decomposition \([\text{output, prices, short term interest rate, house prices, ...}]\). Also e.g. Peersman and Smets (2001) found that output reacts faster than prices while ordering the short term interest rate last. Another approach is used in Del Negro and Otrok (2007) who order house prices first, followed by GDP and inflation while the federal funds rate is ordered last. Iacoviello (2005) on the other hand, applies the recursive identification procedure on a four variable VAR with the federal funds rate ordered first, followed by prices, houses and output.
ordered as follows: unemployment rate, disposable income per capita, house prices, GDP deflator and the Eonia. The main underlying reasoning is that output typically reacts faster than prices, to a change in interest rates, while macroeconomic variables do not respond instantaneously to a policy innovation. As such we follow a standard way to identify a monetary shock by assuming that an interest rate change does not affect output or inflation within a quarter (Christiano, Eichenbaum & Evans, 2005; Goodhart & Hofmann, 2008; Giuliodori, 2005; Jarocinsko & Smets, 2008; Peersman & Smets, 2001; Tstatzaronis & Zhu, 2004). Despite the fact that the chosen ordering is debatable, our results concerning the monetary policy shock and the determination of house prices remain virtually unchanged if the policy interest rate is ranked differently. The same is true for the use of generalized impulse response analysis (see Pesaran & Shin 1998).

The key results of our BVAR analysis are shown in figure 5 and table 3, representing impulse responses and the variance decomposition of some key variables of interest up to twenty quarters. As the primary interest of our analysis is to study the dynamics of house prices, we focus on the responses following a shock in the house price determinants, i.e. the Eonia, house prices, the unemployment rate, gross national disposable income per capita and the GDP deflator.

Results are in line with expectations and clearly show that a one percent deviation in the Eonia has a significant impact on the average Irish house price. From 3 quarters onwards, house prices will decline by more than 3%. Furthermore, the speed of adjustment turns out to be fairly rapid. Our results concerning the long term impact of the short term interest rate on Irish house prices corresponds to most literature on the Irish housing market (see e.g. Stevenson (2008)). The causality also runs from the effect of monetary policy on housing prices rather than in the opposite direction. Furthermore, house prices react modestly positive to a one percentage shock in disposable income per capita. Finally, a positive innovation of one percentage in the unemployment rate causes house prices to decline while an increase in the GDP deflator has a positive effect. Noteworthily is also the large effect of an increase in the interest rate on unemployment. The impulse responses show a disinflationary effect of monetary tightening and thus the results show no evidence of a so-called price puzzle.

---

20 In a price puzzle a tightening of monetary policy leads to a rise in prices, which often can be found in small VARs and is generally attributed to the models’ being too parsimoniously specified. It has been well documented in the literature and VAR models typically included commodity prices to solve the price puzzle (Sims, 1992).
Table 3 reports the variance decomposition of the average Irish house price. It reveals that the unemployment rate and the eonia are the largest components in explaining the variance in house prices. While the unemployment has a large immediate effect, the eonia and gross national disposable income have an increasing effect. The effect of the unemployment rate reaches it peak already after 5 quarters whereafter the effect gradually declines. Monetary policy turns out to be the most important driver of house prices since it explains more than 50% of house price variance after 10 quarters. Furthermore on average 9% of house prices is explained by gross national disposable income. The deflator turns out to have only a modest effect. These results are in line with earlier findings of Miles and Pillonca (2008) and those of McQuinn and O’Reilly (2008) who show that Irish house prices have a strong link with income and interest rates via a standard housing demand and supply equation.

If we combine these results that monetary policy is the largest driver in house prices with our finding that the Irish interest rate was far too low for an extended period of time, i.e. on average 6.7% between 1999 and 2007, it is clear that the ECB’s one size fit all monetary policy significantly contributed to the Irish housing boom. If monetary policy would have been more suited to the Irish domestic needs house prices probably wouldn’t have experienced such intense boom-bust cycle.

---

21 With regard to reverse causality, i.e. the effect of a shock in house price prices on the eonia, we find that the effect is very small. House prices contribute very little to the variance in the monetary policy instrument, more precisely, less than 3%.
3.2 A multivariate housing model

Since we clearly showed the direction of causality, mainly running from the short term interest rate and income towards house prices and only to a small extent vice versa, we may turn to a standard inverted demand equation as in most of the empirical literature examining Irish house prices. (e.g. Stevenson, 2008; Abelson, Joyeux, Milunovich & Chung, 2005). The advantages of this additional exercise are comparability with existing literature, the clear economic interpretation and the ability to perform an intuitive counterfactual analysis. As in most empirical literature we assume that demand for property is driven by disposable income, the composition of the population, interest rate levels and the price of property itself (e.g. Seyfried 2010, Stevenson 2008, Duffy et al. 2005). Using the natural logarithm leads to a general log-linear econometric model with $\varepsilon_t$ as error term:

$$\ln(P_t) = \alpha + \beta_1 \ln(Y_t) + \beta_2 \cdot DEM_t + \beta_3 \cdot R_t + \beta_4 \cdot \ln(HSTOCK_t) + \varepsilon_t \tag{3}$$

Where $Y_t$ is a measure of disposable income, $R_t$ is an interest rate, $HSTOCK_t$ stands for the stock of housing which can be expressed per capita and is a measure of the supply and $DEM_t$ is a proxy for the demographic evolution. In order to model house prices ($P_t$), in line with most empirical studies,
we use the average new house price\textsuperscript{22} (Conefrey and Fitz Gerald, 2010; Duffy et al., 2005; Roche, 2003). In contrast to most existing literature on the Irish property market, we focus on nominal prices instead of real house prices since these are important in view of financial stability. First, 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 values are crucial for the potential losses banks have to book in the event of default (Central Bank of Ireland, Financial stability report 2007 ). To measure the impact of demographic evolution we used the total population size, redefining \( \text{DEM}_t \) as \( \ln(\text{POP}_t) \) 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). 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. 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, as in our BVAR analysis, and equals the GDP corrected for outgoing and incoming transfers to or from non-Irish institutional organizations. The cost of capital or interest rate is measured as the quarterly average mortgage rate \( (r_t) \). 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 (4), which is recursively completed in the housing equation (5). 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 (McQuinn and O'Reilly, 2008; Rae, 2006; Fitzpatrick, 2007).

\textsuperscript{22} 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, but results did not change significantly.
Finally, equation (3) is complemented by an autoregressive term ($\ln(P_{t-1})$). This factor allows us to take the dynamic behaviour of house prices into account. So the 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):

$$r_t = \alpha + \beta_1.EONIA_t \rightarrow \hat{r}_t = R_t = \alpha + \beta_1.EONIA_t$$

(4)

$$\ln(P_t) = \alpha + \beta_1.\ln(Y_t) + \beta_2.\ln(POP_t) + \beta_3.\hat{r}_t + \hat{\beta}_4.HSTOCKPOP_t + \beta_5.\ln(P_{t-1}) + \epsilon_t$$

(5)

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:

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

(6)

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. 

The house price model estimation results are presented in table 4. 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. Note that a similar reasoning was applied in our Cholesky ordering for our BVAR-analysis. In addition, this way of modeling in our multivariate house equation delivers proof of causality since the variable represents the impact of a change in interest rates on house prices in the next quarter. Finally we also include a dummy variable which takes the value of one starting from the third quarter of 2008 (Dummy 2008).

The variables are statistically significant at a 99% confidence level except the housing stock. All tests concerning the basic assumptions for reliable OLS estimates indicate no major concerns. Since the

---

23 Before estimating the variables are tested for stationarity. The test results of the ADF test are presented in appendix F. The results suggest that all variables are non-stationary but integrated of order one (I(1)). However, the standard two-step procedure suggested by Engle and Granger shows that our time series under investigation are cointegrated. 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). Furthermore, it implies that our 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).
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: $\frac{\beta}{1 - \beta_{-1}}$. $\beta$, 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, 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 our BVAR model, hence one percentage lower but corresponds to the 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 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%. 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. Again, as in our BVAR model, this may point in the direction of an effect of rising house prices encouraging people to buy houses because they expect prices to increase. 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 (Stevenson, 2008; Abraham, 1996). 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\%\textsuperscript{24}, 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\%\textsuperscript{25}. This value corresponds to the impact of the interest rate variable estimated in the work of Stevenson (2008). 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.

<table>
<thead>
<tr>
<th>Variables</th>
<th>House price model</th>
</tr>
</thead>
<tbody>
<tr>
<td>constant ($\alpha$)</td>
<td>-17.8755 ***</td>
</tr>
<tr>
<td>$\ln(Y_t)$</td>
<td>0.5025 ***</td>
</tr>
<tr>
<td>$\ln(POP_t)$</td>
<td>1.1160 ***</td>
</tr>
<tr>
<td>$R_{t-1}$</td>
<td>-0.0215 ***</td>
</tr>
<tr>
<td>HSTOCKPOP$_{t-1}$</td>
<td>-0.0044 \textsuperscript{a}</td>
</tr>
<tr>
<td>Dummy 2008</td>
<td>-0.1111 ***</td>
</tr>
<tr>
<td>$\ln(P_{t-1})$</td>
<td>0.6762 ***</td>
</tr>
<tr>
<td>Adjusted $R^2$</td>
<td>0.998</td>
</tr>
<tr>
<td>Schwarz criterium</td>
<td>-314.485</td>
</tr>
<tr>
<td>Jarque-Bera test</td>
<td>0.433</td>
</tr>
<tr>
<td>White test</td>
<td>0.103</td>
</tr>
<tr>
<td>ARCH(1)-test</td>
<td>0.870</td>
</tr>
<tr>
<td>Breush-Godfrey LM test\textsuperscript{b}</td>
<td>0.605</td>
</tr>
<tr>
<td>Ramsey RESET test</td>
<td>0.074</td>
</tr>
</tbody>
</table>

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

### 3.3 A counterfactual analyses

By estimating a model of Irish house prices, controlling for the impact of other variables, we again show that the interest rate variable plays an important role. In the first part of our paper, 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 replace the Eonia in (8) by the Irish Taylor-rate, i.e. the

\[ R_0.022 + 0.676229 = R 0.0665. \]

\textsuperscript{24} To put it in perspective against our BVAR analysis: to get a mortgage interest rate which is one percent higher, the Eonia has to be raised by about 1.4 percent (see eq. 8). Since the effect of a shock of one percent in the Eonia in our BVAR-system has a long term impact of about 4\% it can be calculated that about 5.7\% can be considered as the impact of a change in monetary policy in our BVAR measured on the same scale as in our multivariate approach.

\textsuperscript{25} The coefficient in table 4 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\%.
counterfactual short term interest rate, which results in a counterfactual mortgage interest rate under an alternative monetary policy. The results are presented in figure 6, 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%.

![Counterfactual vs. Actual Mortgage Rates](image.png)

**Fig. 6.** Actual vs. Counterfactual average Irish mortgage rate (1994 – 2012).

Next, we combine the counterfactual mortgage rate with the housing equation, using the estimated coefficients presented in table 4. This approach shows how the average Irish house price would have evolved over time if there was an alternative monetary policy. Figure 6 displays 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. 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 a clear indication 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, just before the bust of the property bubble, Irish house prices would have been about 30% lower. This proves that the consequences of the crisis would have been more modest and instead of a sudden-stop, there probably would have been a soft landing as can be seen

---

26 A remark can be made that these counterfactual experiments ignore the Lucas critique (1976). However, as such we can at least attempt to measure the impact of an alternative monetary policy stance. Furthermore, it is important to emphasize that it is especially in the first years after the introduction of the single currency that actual and counterfactual prices started to diverge, hence dampening the potential importance of the Lucas effect in our counterfactual scenario. In addition, our multivariate house price equation contains an autoregressive house price variable which automatically, recursively, take into account some dynamic effects of a higher short term interest rate.

27 The maximum difference between the actual and counterfactual price is biggest in the fourth quarter of 2003, namely 34%.

28 Calculated as \( \left( \frac{\text{counterfactual price} - \text{actual price}}{\text{actual price}} \right) \times 100 \).
from figure 7. It also shows that keeping the interest close to the Taylor-rate prevents housing prices from rising dramatically. This provides evidence that it is not necessary to include a lean against house price fluctuations in monetary policy strategies.

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.

![Figure 7. Counterfactual Irish house prices (1994Q1 – 2011Q4).](image)

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, applying a BVAR analysis, we showed the causal relationship between the short term interest rate and Irish house prices. The analysis showed that the policy rate and the unemployment rate are the largest components in explaining the variance in house prices. Applying a counterfactual analysis on the basis of a standard inverted demand equation, we also illustrated 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.

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.
References


24


### Appendix A

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

<table>
<thead>
<tr>
<th>Variable</th>
<th>Number of lags</th>
<th>Test-statistic</th>
<th>Critical values</th>
<th>Order integration</th>
<th>P-value LM test</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
<td>1%</td>
<td>5%</td>
<td>10%</td>
</tr>
<tr>
<td>EONIA²</td>
<td>1</td>
<td>-3.47</td>
<td>-3.96</td>
<td>-3.41</td>
<td>I(0)**</td>
</tr>
<tr>
<td>HICP</td>
<td>4</td>
<td>-2.96</td>
<td>-3.43</td>
<td>-2.86</td>
<td>I(0)**</td>
</tr>
<tr>
<td>Output gap</td>
<td>1</td>
<td>-2.70</td>
<td>-3.43</td>
<td>-2.86</td>
<td>I(0)</td>
</tr>
</tbody>
</table>

Notes: 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. * Test with trend. ** significance at 5%, * significance at 10%.
Appendix B

Dependent variable: KwgeomEONIA
Instrumented: INF2 OutputgapEMU1 INS1
Instruments: const INS2 INS3 HICP OutputgapEMU
HAC standard errors, bandwidth 3 (Bartlett kernel)

<table>
<thead>
<tr>
<th>Coefficient</th>
<th>Std. Error</th>
<th>z</th>
<th>p-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>const</td>
<td>-0.133765</td>
<td>0.192727</td>
<td>-0.6941</td>
</tr>
<tr>
<td>INF2</td>
<td>0.208953</td>
<td>0.075910</td>
<td>2.7527</td>
</tr>
<tr>
<td>OutputgapEMU1</td>
<td>0.114431</td>
<td>0.028476</td>
<td>4.0185</td>
</tr>
<tr>
<td>INS1</td>
<td>0.899929</td>
<td>0.023695</td>
<td>37.9797</td>
</tr>
</tbody>
</table>

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. INF2 = inflation with a lead of two quarters. OutputgapEMU1 is the output gap looking one quarter ahead. 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 C

CUSUM plot with 95% confidence band

Fig. Appendix C Cusum test Taylor rule ECB
Appendix D

Table D1 ADF-test Irish variables, the HICP and output gap (1994 – 2012).

<table>
<thead>
<tr>
<th>Variable</th>
<th>Number of lags</th>
<th>Test statistic</th>
<th>Critical values</th>
<th>Order integration</th>
<th>P-value LM test</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
<td>1%</td>
<td>5%</td>
<td>10%</td>
</tr>
<tr>
<td>HICP Ireland</td>
<td>1</td>
<td>-2.42</td>
<td>-3.43</td>
<td>-2.86</td>
<td>-2.57</td>
</tr>
<tr>
<td>Output gap Ireland</td>
<td>1</td>
<td>-0.86</td>
<td>-3.43</td>
<td>-2.86</td>
<td>-2.57</td>
</tr>
</tbody>
</table>

Notes: 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 D2 Test stationarity Irish variables, KPSS and ADF-GLS test (1994 – 2012).

<table>
<thead>
<tr>
<th>KPSS*</th>
<th>Test statistic</th>
<th>Critical values</th>
<th>ADF-GLS (p-value)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>1%</td>
<td>5%</td>
</tr>
<tr>
<td>HICP Ireland</td>
<td>0.41</td>
<td>0.731</td>
<td>0.47</td>
</tr>
<tr>
<td>Output gap Ireland</td>
<td>0.67</td>
<td>0.731</td>
<td>0.47</td>
</tr>
</tbody>
</table>

* 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 D1 and D2 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 E

**Table appendix E: description data and sources Irish house price models (1994Q1 - 2012Q4)**

<table>
<thead>
<tr>
<th>Variable</th>
<th>Source</th>
<th>Frequency/transformation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Eonia</td>
<td>ECB data warehouse</td>
<td>Daily – averaged to quarters</td>
</tr>
<tr>
<td>Irish national new house prices (lnhp)</td>
<td>Department of the Environment, Community and Local Government [48]</td>
<td>Quarterly</td>
</tr>
<tr>
<td>GDP deflator Ireland</td>
<td>OECD [20]</td>
<td>Quarterly</td>
</tr>
<tr>
<td>Irish unemployment rate</td>
<td>OECD [20] - Series harmonised unemployment rate</td>
<td>Quarterly</td>
</tr>
<tr>
<td>Irish population size</td>
<td>IMF [59]</td>
<td>Yearly – interpolated to quarters using Chow-Lin interpolation [60]</td>
</tr>
<tr>
<td>Irish housing stock per 100 inhabitants</td>
<td>Department of the Environment, Community and Local Government [48]</td>
<td>Yearly - interpolated to quarters using Chow-Lin interpolation [60]</td>
</tr>
<tr>
<td>Average Irish mortgage rate</td>
<td>Department of the Environment, Community and Local Government [48]</td>
<td>Quarterly</td>
</tr>
</tbody>
</table>
Appendix F

Table appendix F ADF-test variables house price models

<table>
<thead>
<tr>
<th>Variable</th>
<th>Number of lags</th>
<th>Test statistic levels</th>
<th>Test statistic first order difference</th>
<th>Order of integration</th>
</tr>
</thead>
<tbody>
<tr>
<td>ln(GDPdeflator)</td>
<td>1</td>
<td>-2.77</td>
<td>-3.80</td>
<td>I(1)***</td>
</tr>
<tr>
<td>ln(GNDICAPITA)&lt;sup&gt;a&lt;/sup&gt;</td>
<td>2 and 1</td>
<td>-0.62</td>
<td>-3.17</td>
<td>I(1)***</td>
</tr>
<tr>
<td>UR = (Unemployment rate)</td>
<td>2 and 1</td>
<td>2.04</td>
<td>-2.12</td>
<td>I(1)**</td>
</tr>
<tr>
<td>ln(P&lt;sub&gt;t&lt;/sub&gt;) = ln((hp))</td>
<td>1</td>
<td>0.58</td>
<td>-2.91</td>
<td>I(1)***</td>
</tr>
<tr>
<td>ln(Y&lt;sub&gt;t&lt;/sub&gt;) = ln(GNDICAPITA)</td>
<td>2 and 1</td>
<td>0.30</td>
<td>-2.82</td>
<td>I(1)***</td>
</tr>
<tr>
<td>ln(POP&lt;sub&gt;t&lt;/sub&gt;)</td>
<td>2 and 1</td>
<td>-2.36</td>
<td>-2.33</td>
<td>I(1)**</td>
</tr>
<tr>
<td>R&lt;sub&gt;tb&lt;/sub&gt;</td>
<td>1</td>
<td>-3.19</td>
<td>-4.42</td>
<td>I(1)***</td>
</tr>
<tr>
<td>HSTOCKPOP&lt;sub&gt;tb&lt;/sub&gt;</td>
<td>2 and 1</td>
<td>-1.51</td>
<td>-3.32</td>
<td>I(1)***</td>
</tr>
<tr>
<td>ε&lt;sub&gt;t&lt;/sub&gt;</td>
<td>0</td>
<td>-8.81</td>
<td>-</td>
<td>I(0)***</td>
</tr>
</tbody>
</table>

Notes: First part of the table (Eonia, ln(GDPdeflator) and ln(GNDICAPITA)) relate to the variables used to estimate the BVAR from the first quarter in 1994 till the last quarter of 2012. The second part relates to the variables and Engle-Granger cointegration test for the multivariate housing model where tests are 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 LMR-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<sub>t</sub>) of the model in table 4. The appropriate critical values to assess the null hypotheses at a significance level of 5% with in total five variables is -4.42 [57].<sup>*</sup> Test with trend <sup>b</sup> The interest rate variable, the average Irish mortgage rate, 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%.
### Appendix G

#### Table G1: VAR lag order selection criteria (1994Q1 – 2012Q4)

<table>
<thead>
<tr>
<th>Lag</th>
<th>LogL</th>
<th>LR</th>
<th>FPE</th>
<th>AIC</th>
<th>SC</th>
<th>HQ</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>121.4321</td>
<td>NA</td>
<td>3.39e-08</td>
<td>-3.011677</td>
<td>-2.700317</td>
<td>-2.887472</td>
</tr>
<tr>
<td>1</td>
<td>642.5674</td>
<td>943.6775</td>
<td>5.10e-14</td>
<td>-16.42074</td>
<td>-15.33098</td>
<td>-15.98602</td>
</tr>
<tr>
<td>2</td>
<td>709.9258</td>
<td>112.8709</td>
<td>1.64e-14</td>
<td>-17.56556</td>
<td>-15.69740*</td>
<td>-16.82033*</td>
</tr>
<tr>
<td>3</td>
<td>744.0666</td>
<td>52.59532</td>
<td>1.32e-14</td>
<td>-17.81261</td>
<td>-15.16605</td>
<td>-16.75687</td>
</tr>
<tr>
<td>4</td>
<td>767.5526</td>
<td>33.00725</td>
<td>1.44e-14</td>
<td>-17.77169</td>
<td>-14.34673</td>
<td>-16.40543</td>
</tr>
<tr>
<td>5</td>
<td>798.0161</td>
<td>38.69692</td>
<td>1.35e-14</td>
<td>-17.91935</td>
<td>-13.71599</td>
<td>-16.24258</td>
</tr>
</tbody>
</table>

*Note. Test up to six lags since we work with quarterly data (see Kilian and Ivanov, 2005). *Indicates lag order selected by the criterion. Endogenous variables: EONIA; LNHP; UR; LN(GNDICAPITAII); LNDEFL; Exogenous variables: C Dummy2008. Sample: 1994Q1 – 2012Q4. LR: sequential modified LR test statistic (at 5% level). FPE: final prediction error. AIC: Akaike information criterion. SC: Schwartz information criterion. HQ: Hannan-Quinn information criterion.

#### Table G2: Johansen’s multivariate cointegration tests

Sample: 1994Q1 2012Q4
Included observations: 76
Trend assumption: Linear deterministic trend
Series: EONIA LNHP UR LOG(GNDICAPITAII) LNDEFL
Exogenous series: DUMMY2008
Lags interval (in first differences): 1 to 2

**Unrestricted Cointegration Rank Test (Trace)**

<table>
<thead>
<tr>
<th>Hypothesized No. of CE(s)</th>
<th>Eigenvale</th>
<th>Trace Statistic</th>
<th>Critical Value</th>
<th>Prob.**</th>
</tr>
</thead>
<tbody>
<tr>
<td>None *</td>
<td>0.437235</td>
<td>92.66289</td>
<td>69.81889</td>
<td>0.0003</td>
</tr>
<tr>
<td>At most 1 *</td>
<td>0.277111</td>
<td>48.97096</td>
<td>47.85613</td>
<td>0.0391</td>
</tr>
<tr>
<td>At most 2</td>
<td>0.186906</td>
<td>24.30896</td>
<td>29.79707</td>
<td>0.1877</td>
</tr>
<tr>
<td>At most 3</td>
<td>0.072359</td>
<td>8.583922</td>
<td>15.49471</td>
<td>0.4052</td>
</tr>
<tr>
<td>At most 4</td>
<td>0.037129</td>
<td>2.875496</td>
<td>3.841466</td>
<td>0.0899</td>
</tr>
</tbody>
</table>

Trace test indicates 2 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

**Unrestricted Cointegration Rank Test (Maximum Eigenvalue)**

<table>
<thead>
<tr>
<th>Hypothesized No. of CE(s)</th>
<th>Eigenvale</th>
<th>Max-Eigen Statistic</th>
<th>Critical Value</th>
<th>Prob.**</th>
</tr>
</thead>
<tbody>
<tr>
<td>None *</td>
<td>0.437235</td>
<td>43.69192</td>
<td>33.87687</td>
<td>0.0025</td>
</tr>
<tr>
<td>At most 1</td>
<td>0.277111</td>
<td>24.66201</td>
<td>27.58434</td>
<td>0.1132</td>
</tr>
<tr>
<td>At most 2</td>
<td>0.186906</td>
<td>15.72504</td>
<td>21.13162</td>
<td>0.2413</td>
</tr>
<tr>
<td>At most 3</td>
<td>0.072359</td>
<td>5.708426</td>
<td>14.26460</td>
<td>0.6509</td>
</tr>
<tr>
<td>At most 4</td>
<td>0.037129</td>
<td>2.875496</td>
<td>3.841466</td>
<td>0.0899</td>
</tr>
</tbody>
</table>

Max-eigenvalue test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values